Buying Time: The Insurance Value and Distortionary Effects of Worker's Compensation

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Abstract

Optimal design of social insurance must balance the welfare gains of additional liquidity against the welfare costs of unintended distortions in claimant behavior. While a growing literature estimates the insurance value of Unemployment Insurance, little is known about the value of other social insurance programs. I examine the extent to which the liquidity-enhancing benefits of social insurance outweigh the moral hazard costs in the context of Worker's Compensation (WC). Analyzing administrative data from Oregon, I estimate a discrete proportional hazard model exploiting variation in the timing and size of a retroactive lump-sum WC payment to decompose the elasticity of claim duration with respect to benefits into two components: the elasticity with respect to an increase in cash on hand (a liquidity effect) and a decrease in the opportunity cost of missing work (a moral hazard effect). I find the liquidity effect is 1.5 times as large as the moral hazard effect for claimants with pre-injury earnings below the median wage in Oregon. By contrast, I find little evidence of a liquidity effect for claimants with pre-injury earnings above the median wage. The results suggest that WC relaxes liquidity constraints even for claimants with fairly short absences from work. Using the framework from Chetty 2008, I conclude that the insurance value of WC exceeds the distortionary cost for lower-wage workers, and that increasing the benefit level for these workers could increase overall social welfare.

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1 Introduction

Social insurance programs are designed to provide protection for individuals against losses in consumption owing to some unanticipated negative shock, such as unemployment, disability onset, or injury on the job. If individuals cannot fully insure against an unexpected health or income shock through private insurance or other alternatives, public social insurance programs provide claimants with needed cash (liquidity) during a time when they cannot earn a wage. However, the payments from such a program also lower the opportunity cost of missing work, and thereby have a distortionary "moral hazard" effect. As is well-recognized in the public finance literature, the optimal design of social insurance depends critically on balancing the welfare gains of providing additional liquidity against the welfare costs of unintended distortions in claimant behavior.

There is a growing body of research estimating the benefits of social insurance programs, in particular for the unemployment insurance (UI) program. These studies consistently find evidence that UI provides considerable insurance value to unemployed workers (e.g., Gruber 1997; Card *et al.* 2007; Chetty 2008; Schmieder *et al.* 2012; LaLumia 2013).[1] Bronchetti (2012) investigates the consumption smoothing benefits of the Workers Compensation program for older workers. Taking advantage of within-state variation in benefit levels, Bronchetti estimates that a 10 percent increase in WC benefits would offset approximately 3-5 percent of the consumption loss following an on-the-job injury.

My study builds on these literatures with an examination of the liquidity-enhancing benefits and moral hazard costs in the context of Worker's Compensation (WC). Analyzing administrative data from Oregon, I estimate a discrete proportional hazard model exploiting variation in the timing and size of a retroactive lump-sum WC payment to decompose the elasticity of claim duration with respect to benefits into two components: the elasticity with respect to an increase in cash on hand (a liquidity effect) and a decrease in the opportunity cost of missing work (a moral hazard effect). Typically, UI or WC benefits provide claimants with cash on hand that allows them to stay out of work while maintaining a particular level of consumption. At the same time, they effectively lower the claimant's net wage, distorting the decision to return to work. However, a payment that is made regardless of when claimants return to work, such as with the retroactive payment in WC, separately identifies the liquidity effect. Chetty (2008) outlines this approach

[1] This complements a set of studies investigating the distortionary labor supply effects of the unemployment insurance program (see Krueger and Meyer (2002) for a review of that literature.) Those studies tend to find that higher levels of UI benefits lead to longer unemployment duration, but it is debated as to whether that increased duration is socially costly or beneficial.

in the context of UI. If WC claimants extend their claims after receiving the retroactive payment, this implies that the additional income affords them more time to recover, moving closer to the claim duration they would choose in a world without liquidity constraints that force them to return to work prematurely. This is the approach I take to separately identify these two effects in the context of WC.

The WC program provides approximately \$60 billion annually to insure workers against the health and income shock of an illness or injury on the job (National Academy of Social Insurance 2014). Since the majority of WC claims occur in physical jobs, WC benefits could be essential in affording claimants sufficient recovery time to return to work successfully. On the other hand, injuries are often difficult to observe, and claimants typically return to the same job they had prior to their injury, so there is little uncertainty about future employment prospects. These factors could increase moral hazard costs relative to UI, or imply less need for the liquidity that WC provides. Many states have recently started reducing benefits and making it more difficult to qualify for WC, in order to lower costs (Grabell and Berkes 2015). However, there is little empirical evidence about the relative magnitude of the insurance value and distortionary costs to determine the welfare consequences of these reforms (Meyer 2002).

Using administrative data from Oregon, I take advantage of a small retroactive lump-sum payment to WC claimants that separates the liquidity and moral hazard effects. As I explain in detail below, WC claimants are paid a small lump sum (equal to 25 percent of their weekly wage, on average) if their claim lasts longer than two weeks. This means that claimants first have an incentive to extend their claim, and later receive additional cash regardless of when they return to work. I estimate a discrete proportional hazard model and examine changes in the rate of exit from WC before and after eligibility for the retroactive payment to decompose the elasticity of claim duration with respect benefits into the elasticity with respect to a change in moral hazard and liquidity. Among claimants with pre-injury wages below the median wage in Oregon (i.e., claimants earning less than \$700 per week), I obtain moral hazard and liquidity elasticities of .16 and .24, indicating that the liquidity effect is 1.5 times as large as the moral hazard effect. In contrast, I estimate a moral hazard elasticity of approximately .23 for high wage workers, but find no significant liquidity effect. These estimates suggest that WC plays an important role in providing cash on hand for lower-wage workers, but that higher-wage workers may have alternative forms of insurance (e.g., savings) to help them smooth their consumption during temporary spells away from work.

By observing how the retroactive payment affects behavior during the first few weeks of the WC claim, I demonstrate that claimants are sensitive to changes in their income even after short spells away from work.

This sensitivity is additional evidence that WC relaxes claimant liquidity constraints, affording claimants more time to recover from an injury or illness. Longer recoveries could additionally improve workers' long-term health, reduce the probability of re-injury on the job, or may increase adjustment costs when a worker returns. I carry out an additional analysis to explore this possibility using linked claims and wage data that I obtained from the state of Oregon. In general, the results do not provide strong evidence that claim length significantly affects post-injury outcomes for those claimants whose return to work decisions are influenced by the retroactive payment.

In the setting I examine, WC claimants face a three-consecutive day waiting period after their injury before they receive any cash benefits. If the injury lasts longer than two weeks, claimants are retroactively paid a lump sum equal to the benefits they would have received during the waiting period, effectively increasing their second bi-weekly WC check by 10 percent, on average. The retroactive payment only reimburses benefits for scheduled work days during the waiting period, meaning that identical claimants injured on different days of the week will have different sized retroactive payments. Under the assumption that injuries occur randomly across different days of the week and that existing levels of cash on hand are uncorrelated with the date of injury, this variation in the size of the claimant's retroactive payment identifies the liquidity and moral hazard effects. I assess the validity of these assumptions and find that the frequency and distribution of observable characteristics is balanced across the date of injury among the claims in my sample. Additionally, I find that my baseline results are comparable to results for a subgroup of claimants who are most likely to have similar levels of cash on hand, regardless of the date of their injury.

I obtained access to an original administrative dataset of WC claims from the Oregon Department of Business and Consumer Services for this study. The database contains rich information on cash benefit claims over more than twenty years and also includes detailed worker and injury characteristics that provide valuable information about other factors that would affect claim length. Additionally, I worked with the Department of Business and Consumer Services and the Oregon Employment Department to obtain a file of matched claims data to employment data. I use these records to examine the effects of longer claims on post-injury outcomes. I supplement this administrative data with survey data from the National Compensation Survey, the Survey on Occupational Illness and Injury and the Current Employment Statistics Survey. I use additional statistics from these surveys in combination with my estimates of liquidity and moral hazard to analyze the welfare effects of a change in WC benefits, and to test my identifying assumptions.

I use the setting of the retroactive payment in Oregon to analyze how WC affects claimant behavior and well-being. The findings in this paper offer additional evidence that social insurance provides lower-income claimants with insurance value, relaxing their liquidity constraints. Under the assumption that claimants maximize their private welfare, the elasticity of claim duration with respect to liquidity and moral hazard are sufficient statistics to determine the effect of a local change in social insurance benefits on social welfare (Baily 1978; Chetty 2008, 2009). Applying my liquidity and moral hazard elasticities to the optimal benefit formula from Chetty (2008), I conclude that the current WC benefit level in Oregon could be too high for higher-wage workers, but is lower than optimal for lower-wage workers. Thus, increasing benefits for lower-wage workers could increase overall social welfare.

2 Identification and data

2.1 Identification strategy

In order to separate the liquidity and moral hazard channels, I take advantage of a common feature of WC payments that separates these effects. In nearly all states, workers face a waiting period at the beginning of their WC claim. Benefits are withheld for the first few days of the claim, and if the claim's duration exceeds a certain length, claimants are reimbursed for the withheld benefits in a lump sum.[2] In Oregon, the setting for my analysis, workers have a three-consecutive day waiting period before they receive cash benefits. If the injury lasts longer than two weeks, they become eligible for a retroactive payment equal to the benefits they would have received during the waiting period.[3] WC checks are paid every two weeks relative to the injury date, and eligible claimants will receive the retroactive payment (RP) in their second WC check regardless of when they return to work. As a result, if claimants with larger RPs differentially lengthen their WC spells after they are eligible for the RP, this can be attributed solely to the effect of receiving additional income after a negative shock: the liquidity effect. Since claimants are not eligible for the RP during the first two weeks, any response to change in the RP during the first two weeks of the claim can be attributed to the increased *incentive* to lengthen claims in order to satisfy the eligibility condition for the RP. If workers cannot borrow against this future benefit, the response to this incentive during the first two weeks represents

[2] See Information Technology and Research Section (2012) for details on the general structure of WC payments.

[3] Workers also are eligible for the retroactive payment if they are admitted to the hospital, regardless of how long their claim lasts. Unfortunately, the Oregon Worker's Compensation Division does not maintain data on hospitalizations; however, as long as hospitalizations are orthogonal to the date of injury, potential hospitalizations should not bias my analysis. Conversations with staff in the Oregon Worker's Compensation Division confirm that hospitalizations during the first two weeks of WC claims are infrequent. While statistics on the share of claimants admitted to the hospital are not available, inpatient hospital services only account for approximately 13 percent of total medical costs (Information Technology and Research Section 2012).

a moral hazard effect (Shavell and Weiss, 1979; Chetty and Finkelstein, 2013).

I take advantage of a unique source of variation in the RP to identify these two effects. As noted above, the waiting period in Oregon is three *consecutive* days from the beginning of the claim, including holidays, weekends, and unscheduled work days. Since the RP only reimburses benefits for scheduled work days during the waiting period, the date of the injury creates variation in the size of this one-time unconditional payment. As an example, consider a typical worker with a Monday to Friday work schedule. Figure 1 shows that for workers injured on a Friday, only one of the waiting period days occurs on a day he was scheduled to work, and the other two waiting period days fall on the weekend. As a result, the worker only has one day of benefits withheld and reimbursed as a lump sum in the RP. However, an identical worker injured on Wednesday or earlier would receive an RP equal to three times his daily benefit, since the entire waiting period falls during the workweek. Under the assumption that injuries occur randomly across different days of the week and that existing levels of cash on hand are uncorrelated with the date of injury, I use this variation in the size of the retroactive payment to estimate liquidity and moral hazard effects.

On average, eligible claimants receive \$100 to \$300 in a lump sum due to the RP. For comparison, the average WC claimant in my sample earns approximately \$650 per week, meaning the RP ranges between 15 and 45 percent of gross weekly earnings. While the absolute value of this payment is small, it provides claimants with a lump sum that is large relative to their typical income stream, precisely at a point in time when they face reduced income due to their injury. In other words, the cash on hand effects could be substantial. The effect of the RP is likely most salient for claimants with some degree of liquidity constraint who are on the margin of staying out of work, rather than claimants with extremely severe or minor injuries. I examine heterogeneity in the effect of the RP across injury type and income level to test these hypotheses.

Since the date of the injury is the main source of variation in the size of the retroactive payment, I address several concerns that the results could be driven by other unobservable characteristics that are correlated with the day of the week. First of all, research has documented a higher frequency of WC claims are filed on Mondays, suggesting the date of injury is not entirely random (Card and McCall 1996). I conduct my main analysis on claims occurring in the second half of the week, where the frequency and distribution of observable characteristics of claims is balanced. Secondly, variation in the day of the week of the injury could affect the size of the worker's final paycheck, which could also affect consumption and claim duration decisions. I estimate liquidity and moral hazard effects on a subsample of workers whose paycheck is less likely to be affected by the date of the injury and find a similar pattern of results as in my main estimates. I also reweight claims in my sample to address the fact that I estimate the liquidity effect on the select sample of claimants who remain out of work at least two weeks, and my results are broadly robust to this correction. Finally, I find that the results are also robust to employers' use of return to work interventions.

2.2 Data and summary statistics

I analyze a rich administrative dataset from the Oregon Department of Consumer and Business Services, Worker's Compensation Division (ORWC) which contains information on closed claims for which cash benefits were paid between roughly 1974 and 2013 (Oregon Department of Consumer and Business Services 2015). The dataset includes detailed information needed to determine the length of the claim, including the date of injury, date of first and last timeloss payments, total workdays for which timeloss benefits were paid, and the number of days typically worked per week. It also contains information about the worker's pre-injury wage, total amount of timeloss payments, total amount of medical payments, age, gender, occupation and industry. Injury information is categorized with ICD-9 codes and includes the nature of the injury, the event causing the injury, and the body part(s) affected. I impute a worker's potential RP using the date of injury, the number of days worked per week, and the worker's pre-injury wage.

Additionally, the database contains several measures of post-injury outcomes for claims occurring after 1999. ORWC matched these more recent claims to closure reports containing information about the worker's employment immediately following their claim, including whether the worker was released to return to work, whether the worker returned to the same employer and/or the same job, and whether the worker required modifications to his work activities. The data also includes a count of the number of times the claim was re-opened due to an aggravation of the injury. Finally, together with the Oregon Employment Department, ORWC matched claims to quarterly earnings records from 1999-2013, allowing me to observe changes in hours and wages before and after the injuries occurring within this time frame (Oregon Employment Department 2015). For all injuries occurring after 1999, I observe wages at least 2 quarters before, and 4 quarters after the event.

I make several restrictions to derive the sample used for this analysis. Because the RP likely will not affect claim decisions for workers with extremely severe injuries, I exclude workers receiving permanent benefits. I restrict my sample to years where the database contains the complete record of claims: between

1987 and 2012. I also restrict the sample to claims lasting at most one year and to cases where the claimant stopped working immediately after the injury. In order to impute the RP, I restrict the sample to injuries occurring on weekdays and to claimants reporting a five-day workweek. Table 1 provides a complete list of all sample restrictions, and the appendix provides more information about the criteria used in making these restrictions. As shown in table A.1, individuals excluded from the sample are older and have slightly higher wages. Additionally, the excluded observations also are more likely to have suffered a fracture and less likely to have suffered a cut or burn. These restrictions predominantly exclude claimants who are unlikely to be responsive to the RP.

Table 2 shows the observable characteristics of my sample across days of the week. Over 70 percent of the sample is male, and the average age of claimants is 36. Table 2 also shows that 60 percent of all injuries are muscle strains or sprains, approximately 10 percent are bone breaks or fractures, and an additional 20-24 percent of injuries are wounds (cuts or burns). The remaining share of injuries are traumatic injuries or other occupational illnesses and diseases (approximately 5 percent for each category). Nearly 65 percent of claimants worked in one of five industries prior to their injury: agriculture, construction, trade, transportation, or manufacturing. The mean weekly wage ranges between \$720-\$740; the median weekly wage ranges from \$630-\$650 in 2012 dollars.[4] On average, WC claimants earn a lower wage than the typical worker in Oregon: the median weekly wage in Oregon is approximately \$700 (Penitson 2014).

As a first test of my identifying assumption, I examine whether WC claimants are similar across different days of the week. First of all, figure 2 confirms that injuries, particularly among claims lasting less than two weeks, are more frequent on Monday and Tuesday. Additionally, table 2 shows that injuries occurring on Monday and Tuesday are slightly more likely to occur in the morning, and have a shorter average duration than claims on other days of the week. Relative to the second half of the week, a higher frequency of Monday and Tuesday injuries are muscle strains. Indeed, the p-values in column 6 confirm that although the differences are small in magnitude, observable characteristics of Monday claims are significantly different from Wednesday claims. These differences in the observable characteristics at the beginning of the week are consistent with the "Monday effect" documented in the literature (Smith 1989; Card and McCall 1996; Ruser 1998; Hansen 2014). The Monday effect could occur if workers try to receive benefits for non-work

[4] I inflate all monetary variables to 2012 dollars using the nominal growth rate in Oregon's state average weekly wage.

related injuries occurring over the weekend or if workers are more careless on Mondays, perhaps due to fatigue. If workers are more likely to report false claims on Monday, these claims are likely less severe and would result in the shorter claims observed at the beginning of the week. The higher frequency of claims and shorter duration of injuries on Tuesday could result from this effect spilling over to Tuesdays after several Monday holidays throughout the year, or due to workers taking long weekends (Smith 1989).

As a result, I restrict the analysis to injuries occurring on Wednesday, Thursday and Friday, where the frequency of injuries is relatively stable. Because the weekend creates variation in the size of the RP, this restriction still allows me to identify claimant responses to the RP. The p-values in column 7 demonstrate that overall, the observable characteristics of workers are balanced between Wednesday and Thursday. Additionally, the composition of injuries and industries is similar between Wednesday, Thursday, and Friday. However, the p-values in column 8 reveal that there are significant differences in the weekly wage and medical costs for Friday injuries. I control for these observable differences in the analysis and restrict the analysis to Wednesday and Thursday injuries to test whether observable or unobservable differences in characteristics of Friday injuries affect the results. The estimates are robust to this restriction, suggesting that differences in Friday injuries do not appear to affect the results significantly.[5]

Figure 3a shows the distribution of claim length in my sample of claims. The measure of duration is the number of workdays for which benefits were paid, so five days represents one work week. These figures reveal two important facts about the distribution of claims: first, there is a long and thin right tail to the distribution of claims: approximately 92 percent of claims in my sample are less than 40 work days, and 96 percent of claims are less than 60 work days. Additionally, figure 3a demonstrates a spike in the frequency of exits at five-day intervals (corresponding to work weeks). This pattern is consistent across injuries on each day of the week, suggesting that the pattern is due to the weeks since the claim began, rather than the day of the week.[6] I return to a more thorough discussion of the implications of these spikes in section 4.

[5] Results available upon request.
[6] Figures available upon request.

3 Distinguishing liquidity from moral hazard

To show how liquidity and moral hazard can be separated conceptually, I draw upon frameworks for the optimal design of benefits from Chetty (2006, 2008) and Diamond and Sheshinski (1995) as well as a dynamic decision-making model from Manoli and Weber (2011), which describes how workers respond to the option value of receiving a future payment. Consider a WC claimant injured at the beginning of period $t = 1$ who must decide whether or not to return to work during periods $t \in \{1, 2, ...T\}$, where each period represents a two-week interval since the injury. For each period in which the claimant remains out of work, he will receive a WC benefit b_t. If he returns to work in period t, he will earn a net wage w_t, but will experience disutility from working, measured by α_t. This disutility of work α_t represents a combination of the claimant's preference for leisure over work, as well as any additional disutility associated with working after an injury. Because workers are uncertain about how long their recovery will take, disutility of work is determined by $\alpha_t = \delta\alpha_{t-1} + \varepsilon_t$, where $\varepsilon_t \sim F_t(\sigma_t)$ represents unexpected variation in the recovery process. Additionally, the worker has cash on hand A_t, and must decide how much to save for the next period, $s_t \geq L$, where L is the lower-bound on borrowing.

At the beginning of period 1, the worker must decide whether to stay out of work or return to work in the current period, and must also consider the fact that remaining out of work during period 1 maintains the option to receive the RP during period 2. The claimant's value function of returning to work in period 1 can be written as

$$V_1 = \max_{s_1 \geq L} v(A_1 - s_1 + w_1) - \alpha_1 + \beta V_2(A_2),$$

where $v(A_1 - s_1 + w_t) = v(c_1^e)$, with $v'(c_1^e) > 0, v''(c_1^e) < 0$. If the claimant decides to return to work in period 1, he does not receive the RP, and I assume he remains at work in all subsequent periods.[7] The claimant's value function of choosing WC during period 1 can be written as:

$$U_1 = \max_{s_1 \geq L} u(A_1 - s_1 + b_1) + \beta J_2(A_2, RP),$$

where $u(A_1 - s_2 + b_1) = u(c_2^n)$ is also concave, and $J_2(A_2, RP)$ represents the expected value of the claimant's

[7]Future versions of this framework will relax this assumption. Realistically, the claimant could face a risk of being injured again in the future, and this risk could be correlated with the length of his recovery time.

decision in the next period:

$$J_2(A_2, RP) = E[max\{U_2(A_2, RP), V_2(A_2, RP)\}].$$

If the worker chooses WC during period 1, he receives the RP during period 2, regardless of his work decision. This leads to the following value functions in period 2:

$$V_2 = \max_{s_2 \geq L} v(A_2 - s_2 + w_2 + RP) - \alpha_2 + \beta V_2(A_3) \tag{1}$$

$$U_2 = \max_{s_2 \geq L} u(A_2 - s_2 + b_2 + RP) + \beta J_2(A_3) \tag{2}$$

In each period $t \in \{1, 2\}$, the claimant has a reservation disutility level α_t^* at which he is indifferent between returning to work or receiving WC for another period:

$$\alpha_t^* = v_t(c_t^e) - u_t(c_t^n) + \beta E[OV_t]. \tag{3}$$

The claimant will choose to work if his realized disutility of work is lower than his reservation disutility level, $\alpha_t < \alpha_t^*$. Note that $E[V_{t+1}(A_{t+1}) - J_{t+1}(A_{t+1}, RP)] = E[OV_t]$ represents the claimant's expected option value associated with deciding whether or not to work. Since he loses the option to receive the RP if he returns to work during period 1, the RP increases the expected option value of staying out of work during period 1.

With this framework, the hazard rate of returning to work during period t can be represented as the probability that a worker's disutility is below his reservation disutility during period t (Manoli and Weber 2011):

$$h_t = Pr(\alpha_t < \alpha_t^*).$$

Factors that increase α_t^* indicate an increase in the worker's reservation disutility, shortening claims. Similarly, factors that decrease α_t^* lower the threshold disutility level and lengthen claims. Changes in the hazard rate translate to changes in α scaled by the probability density function of α. Empirically, I estimate changes in the hazard rate, or probability of return to work. Examining how changes in each of the parameters in $\Omega_t = \{b_t, w_t, A_t, RP\}$ affect the duration of claims demonstrates how these parameters influence the claimants' decision to return to work and, as a result, how changes in the parameters affect claimants' utility in different states of the world.

First, consider the effect of a one-time change in the WC benefit level in any period:

$$\frac{\partial \alpha_t^*}{\partial b_t} = \frac{\partial U_t}{\partial b_t} = -u'(c_t^n) < 0.$$

Increasing b_t increases utility while on WC, but does not affect utility while working. Given this result, an increase in b_t decreases the hazard of leaving WC and lengthens claims. This prediction has been confirmed in previous work finding that more generous WC benefits lead to longer claims (e.g., Krueger 1990; Butler and Worrall 1985; Meyer *et al.* 1995; Neuhauser and Raphael 2004). On the other hand, increasing the wage during any one period yields:

$$\frac{\partial \alpha_t^*}{\partial w_t} = \frac{\partial V_t}{\partial w_t} = v'(c_t^e) > 0.$$

Here, a change in w_t only increases utility if the claimant returns to work. Since the opportunity cost of missing work is increasing in the wage, this implies that increasing the wage will increase the rate at which claimants return to work.

Now, consider the effect of a change in the level of cash on hand during any one period:

$$\frac{\partial \alpha_t^*}{\partial A_t} = v'(c_t^e) - u'(c_t^n) \leq 0.$$

In this case, the change in cash on hand affects utility in *both* the working and non-working state. The sign of $\frac{\partial \alpha_t^*}{\partial A_t}$ depends on how A_t affects utility when individuals are working, relative to when they are not. If workers are able to maintain their desired consumption level when out of work, then their marginal utility of consumption will be the same in each state of the world, such that $v'(c_t^e) = u'(c_t^n)$ and $\frac{\partial \alpha_t^*}{\partial A_t} = 0$. However, since $b_t < w_t$, claimants may lower their consumption while on WC if they cannot completely offset the gap in income with savings, or if they have precautionary savings motives. If workers reduce their consumption such that $v'(c_t^e) < u'(c_t^n)$, then $\frac{\partial \alpha_t^*}{\partial A_t} < 0$, indicating that additional cash on hand is *more valuable* to individuals when they are not working. In this case, an increase in A_t allows workers to move closer to their desired consumption level while out of work. Since they are now able to consume more while out of work, their reservation disutility falls.

As shown in Chetty (2008), $\frac{\partial \alpha_t^*}{\partial b_t}$ can be decomposed into the response to change in the level of cash on hand and a change in the wage:

$$\begin{aligned}\frac{\partial \alpha_t^*}{\partial b_t} &= [v'(c_t^e) - u'(c_t^n)] - v'(c_t^e) \\ &= \frac{\partial \alpha_t^*}{\partial A_t} - \frac{\partial \alpha_t^*}{\partial w_t}\end{aligned} \tag{4}$$

Hence, an increase in benefits could increase the reservation disutility level and lengthen claims through two distinct channels: by relaxing liquidity constraints and reducing the opportunity cost of missing work. The ratio of $\frac{\partial h_t}{\partial A_t}$ and $\frac{\partial h_t}{\partial w_t}$, which are estimated in data, yields the ratio of $\frac{\partial \alpha_t^*}{\partial A_t}$ and $\frac{\partial \alpha_t^*}{\partial w_t}$, informing the relative size of these two channels. To see how the RP helps to identify these effects, consider comparative statics on RP during period 1 and period 2. Because workers who stay out of work during period 1 maintain the option of receiving the RP, the payment effectively lowers the opportunity cost of missing work during period 1. For these workers, the RP changes the expected value of utility during period 2, J_2:

$$\frac{\partial \alpha_1 *}{\partial RP} = \beta E\left[\frac{\partial OV_1}{\partial RP}\right] = \begin{cases} -\beta \frac{\partial V_2}{\partial RP} : V_2 > U_2 \\ -\beta \frac{\partial U_2}{\partial RP} : V_2 \leq U_2 \end{cases} < 0 \tag{5}$$

Since both $-\beta \frac{\partial V_2}{\partial RP} < 0$ and $-\beta \frac{\partial U_2}{\partial RP} < 0$, increasing the RP always lowers the reservation disutility level during period 1. Since workers do not receive the income from the RP until period 2, the response to the RP during period 1 is solely due to the increased option value of receiving the RP during period 2.

Once the worker is eligible for the RP, equations 1 and 2 show that the RP increases his utility during period 2 regardless of the decision to work, and has an identical effect on α_2* as a change in A_2:[8]

$$\frac{\partial \alpha_2 *}{\partial RP} = v'(c_2^e) - u'(c_2^e) \leq 0.$$

The separation between the time when claimants face the change in their opportunity cost and the time when claimants actually receive the *payment* allow me to distinguish the response to receiving additional cash from the response to a change in the incentive to return to work. If the response to the option value in period 1 is small relative to the response of receiving the non-distortionary payment during period 2, this

[8]In reality, workers will not receive the RP at the beginning of period 2. However, since the value of the RP is guaranteed upon reaching period 2, the effect of the RP could also be interpreted as decreasing the borrowing constraint L during period 2. Conceptually, this one-time decrease in L has the same effect on utility during period 2 as an increase in A_2. If workers instead wait until they receive the payment at the end of period 2, the RP will relax liquidity constraints during period 3.

implies that workers primarily lengthen claims in response to income that offsets the gap in their consumption: the liquidity effect. On the other hand, if the response during period 1 is larger than the response during period 2, this suggests that claimants primarily respond to the change in incentives: the moral hazard effect.[9]

4 Empirical analyses

As a first assessment of how the RP affects claim length, table 3 shows the estimated coefficients from a linear regression of the log of the total number of workdays in the claim on the log retroactive payment, controlling for the claimant's pre-injury wage, injury, occupation, gender, age, and total medical costs in the claim. The results show that claim length does respond to the RP: a 1 percent increase in the retroactive payment lengthens claim durations by approximately .02 percent overall, and .03 percent among claims lasting longer than two weeks. Relative to the average claim duration of approximately 14-15 days, a 50 percent increase in the retroactive payment translates to an increase in duration of approximately half a day. Additionally, the retroactive payment has a small negative effect on claims lasting less than two weeks, which could result from the offsetting effect of the smaller paycheck during period 1, or due to compositional changes from workers who extend their claims to two weeks or longer to claim the RP. Column 3 shows that the RP does not significantly affect claims lasting longer than eight weeks. This finding is reasonable, as a claimants with injuries severe enough to warrant a long claim are is unlikely to be influenced by a small change in the structure of their payment during the first four weeks of their claim. While these estimates demonstrate that the RP has an effect on claim length, the effect of the RP on total claim length does not inform whether workers respond to the distortionary effects in the RP, or if the lump-sum payment relaxes their liquidity constraints after two weeks. In the sections below, I analyze claimants' responses before and after the payment of the RP to decompose these two channels for longer claims.

I decompose the elasticity of claim duration with respect to benefits into the liquidity and moral hazard effects in two ways. First, I estimate a discrete proportional hazard model to determine the extent to which the RP affects claim duration at different points in time. Then, I obtain an alternative estimate of the moral hazard effect by estimating excess bunching around the eligibility threshold for the RP. Finally, to examine whether the liquidity and moral hazard effects have consequences after claimants return to work, I investigate

[9]If workers have some ability to borrow and have a strong expectation that their claim will last long enough to receive the RP, they could choose to "spend" the RP prior to the two week mark. If this occurs, the response during period 1 could be an over-estimate of moral hazard, and an under-estimate of the overall liquidity-moral hazard ratio. See section 4 for more details.

the effects of longer claims on post-injury outcomes using the retroactive payment as an instrument for claim length.

4.1 Hazard analysis

First, I estimate the following discrete proportional hazard model:

$$h_{it} = 1 - \exp[-\exp(\sum_{t=1}^{k} \theta_k ln(RP_i)\mathbb{I}[k=t] + X_i'\beta + \gamma_k)] \tag{6}$$

where h_{it} represents the hazard rate: the probability of individual i leaving WC during period t, conditional on *not* leaving WC prior to period t. I estimate the discrete proportional hazard model with the complementary log-log function shown above, which allows me to observe how observable characteristics affect the probability of exit during grouped time intervals (Allison 1982; Meyer 1990; Jenkins 2005), in this case, two-week intervals. I obtain the coefficients in equation 6 using maximum likelihood estimation in Stata. While this specification does not identify the underlying baseline hazard rate, it relies on fewer assumptions than a fully parametric specification with little loss in efficiency (Meyer 1986, 1990).

The main assumption in proportional hazard estimation is that observable characteristics and the baseline hazard are multiplicative: the effect of an observable characteristic X_i scales the baseline hazard rate by $X_i'\beta$. Similarly, time-varying coefficients scale the hazard rate during each period t (Kalbfleish and Prentice 2002; Jenkins 2005). As a result, this specification allows me to understand how a relative change in observable characteristics affects the hazard rate. For example, if $X_i = \{0,1\}$ is a dummy variable indicating whether or not a worker is female, the corresponding coefficient identifies the proportional difference in the hazard rate for women, relative to the hazard rate for men. Additionally, I interact the coefficient on the RP with two week intervals to allow the effect of the RP to vary over time. I control for the claimants' pre-injury wage and weekly WC benefit, meaning the remaining variation in the RP comes from the date of the injury as explained in section 2. Because I include the log of the RP in this specification, the θ_t coefficients represent the elasticity of the hazard rate with respect to the RP during period t. For example, a coefficient of -0.04 would indicate that claimants with a 100 percent higher RP have 4 percent lower hazard rate, indicating that claimants with larger RPs have longer claims.

Additionally, $X_i'\beta$ adjusts the hazard rate for time-invariant individual observable characteristics. I control for a four-piece spline in pre-injury log wage, gender, age, the worker's log weekly benefit, and total

WC-paid medical costs in X_i. I also include a parsimonious set of indicators for broad injury categories and key occupation groups and an indicator for claims occurring after 2002, a year when the maximum benefit increased and several other WC policy changes occurred in Oregon. I control for duration dependence over time using a spline in duration with knots every two weeks, represented by the γ_t terms. Because of the spikes in the frequency of claim exits shown in figure 3a, I include an indicator for durations in multiples of five. Finally, I censor claims exceeding 60 workdays (12 weeks), since accurate estimation of the long right tail of the distribution would require parametric assumptions about the baseline hazard rate (Meyer 1990). Less than 5 percent of claims in my sample exceed 60 workdays and, in practice, this restriction does not affect the $ln(RP)$ coefficients appreciably.[10]

Liquidity effects are likely to be most important for workers who have low levels of wealth prior to their injury. As a proxy for wealth, I estimate the hazard model separately for workers earning above and below the median wage in Oregon, which is approximately \$700 per week (Penitson 2014). If claimants with lower wages are more sensitive to small changes in their payment, i.e., if they are more liquidity constrained, then there should be a larger effect of the RP during period 2 for claimants earning less than the median wage.

Table 4 and figure 5 show the coefficients from equation 6. In the overall sample, column 1 shows that the RP significantly reduces the probability of exit during period 1: a 1 percent increase in the RP reduces the probability of exit by .058 percent. Columns 2 and 3 show that conditional on the other covariates, a 1 percent increase in the RP reduces the hazard of leaving WC during the first two weeks of the claim by .042 percent for workers above the median wage, and .029 percent for workers below the median wage. For reference, a 50 percent increase in the RP increases the payment from 2 days of benefits to 3 days of benefits, on average. A change of this size leads to a 2.1 and 1.5 percent decrease in the probability that a worker's WC claim will end during the first two weeks for workers above and below the median wage, respectively.

While all workers respond to the *option value* to recoup the benefits withheld during the first two weeks, only low wage workers significantly lengthen their claims in response to receiving the unconditional payment. For low wage workers, a 1 percent increase in the RP additionally reduces the hazard of leaving WC during the *second* two weeks by .039 percent. Conditional on having a claim lasting at least two weeks, these estimates imply that a 50 percent increase in the RP leads to an approximate 2 percent decrease in the probability of exit for a low-wage worker. The fact that the coefficient on period 2 is not significantly different from zero for high wage workers implies that high wage worker's response to the RP is dominated

[10] Censoring claims above 40 and 100 days yield similar results, which are available upon request.

by the change in incentives during period 1, the moral hazard effect.

To put these results in context, a 50 percent increase in the RP amounts to an 8 percent increase in the bi-weekly benefit, on average. Similarly, if the claimant returns to work during period 1 and gives up the option to receive the payment, a 50 percent increase in the RP implies that the claimant would give up a payment equal to an additional 6 percent of the average bi-weekly wage. Scaling the coefficients in table 4 by these amounts, I obtain a liquidity elasticity of approximately .24 and a moral hazard elasticity of approximately .16 for low-wage workers. Based on equation 4, this implies an overall elasticity of approximately .4, and shows the liquidity effect amounts to approximately two-thirds of the total response to a change in benefits for low wage workers. The overall elasticity is approximately .23 for claimants above the median wage, and based on the insignificant coefficient after the first two weeks for claimants above the median wage, I conclude that this elasticity for higher-wage workers is driven by moral hazard.

This finding is important for several reasons. First of all, these results add to a broader literature finding that individuals are sensitive even to small lump-sum payments, and this sensitivity suggests that workers could face liquidity constraints (Soueles *et al.* 2006). While previous research finds that workers who have experienced injuries on the job reduce their consumption relative to when they are employed (Bronchetti 2012), this research does not provide information about *when* the decline in consumption occurs following the injury, or whether consumption only falls for severe injuries with long durations. The significant liquidity effect in table 4 suggests that low-wage claimants reduce their consumption even after fairly short spells away from work. Workers could reduce their consumption right after an on the job injury due to immediately binding liquidity constraints, or to increase their precautionary savings to hedge against the risk of facing a binding liquidity constraint later in their claim (Carroll and Kimball 2008). The fact that this payment affects the duration of claims suggests that timely changes in income can significantly affect workers' welfare during their recovery from an injury, in particular for those with low incomes and presumably, low assets.

In general, the decision to leave WC is not completely determined by the claimant; doctors also play an important role in determining the length of a claim. A claimant's doctor must initially certify that a claimant cannot work for a certain period of time, and workers must revisit the doctor in order to be granted additional time away from work. Workers facing fairly minor injuries should be less likely to have doctors certify that their injury warrants two weeks away from work, and workers facing severe injuries will likely remain out of work longer than two weeks, regardless of how large their RP might be. However, workers with less obvious recovery times may be able to adjust their claim length in response to the RP. In order to test this hypothesis, table 5 presents estimates from equation 6 separately for broad injury categories. While the decreasing sample size leads to fairly wide confidence intervals on some of these estimates, these results suggest that the response is driven by injury types with variable recovery times. Indeed, column 1 shows that claimants with severe injuries like traumas do not respond significantly to the RP. Additionally, column 4 shows that claimants with fairly minor injuries, like cuts and burns, only respond to the option value of the RP during the first two weeks.

4.2 Excess Bunching

The estimates from the proportional hazard model are based on the assumption of a semi-parametric functional form for the hazard rate, and allow me to identify the relative effect of the RP on the rate at which people end their WC spells. I provide further evidence about the magnitude of the moral hazard effect using a different estimation procedure that does not rely on these parametric assumptions. If claimants respond to the incentive stay out of work until they are eligible for the RP but do not use the additional income to further extend their claim past two weeks, this would lead to a large share of claims ending exactly at the point where workers become eligible for the RP. Indeed, figure 3a exhibits a spike in claim exits at exactly two weeks. Additionally, because these claimants do not extend their claim beyond two weeks, it indicates that claimants are able to reach their optimal claim length without the non-distortionary payment from the RP.

I estimate the amount of excess mass in the distribution of claim exits at the two week threshold as an alternative estimate of moral hazard. The main assumption in estimating excess bunching is that absent the change in incentives, the distribution of claim length would be smooth. However, figure 3a shows spikes in the frequency of claim exit every 5 workdays, indicating a seasonal pattern in exits of WC after each week of the claim. As a result, I estimate a counterfactual distribution of claims that allows for a pattern of seasonality, but smooths the spike at two weeks, similar to what might exist in a world where workers do not have incentives to lengthen claims due to the option value of receiving the RP. I draw upon methodologies in Saez (2010) and Manoli and Weber (2011) to estimate excess bunching. In particular, I estimate the following regression:

$$n_d = \sum_{t=1}^{5} f(d) * \mathbb{I}[d \in \{10(t-1), 10*t\}] + \beta S_d + \varepsilon_d \tag{7}$$

where n_d is the number of claims ending after d days of benefits, $f(d)$ is a fourth-degree polynomial, interacted with an indicator for each 10-day duration interval.[11] Additionally, S_d is an indicator for exits occurring at any interval of 5. Finally, I interact this equation with indicators for each day of the week included in the main analysis. Using this regression, I predict a counterfactual count of claims on each day. Then, I calculate the share of claims ending exactly after 10 workdays under the original and counterfactual distribution, and attribute the difference between these two shares as excess bunching due to the option value incentive of the RP. I estimate the excess mass as a fraction of two claims intervals: a fraction of total claims ending during the second week, and as a fraction of all claims ending during the first two weeks.

Figure 3b compares the actual density of claim exit with the estimated counterfactual density of claim exit. Comparing the two densities suggests the spikes in the distribution are driven in part by seasonality in claim length. Still, there is a small amount of excess bunching around the two week mark, when claimants would become eligible for the RP. Additionally, figures 4a and 4b show that the excess bunching appears to be larger for claimants above the median wage.

Claimants who leave WC prior to the two week mark "give up" the option of receiving the RP, which is equal to approximately 13 percent of the claimants pre-injury bi-weekly wage, on average. Chetty (2006) shows that liquidity and moral hazard effects can also be represented as a Slutsky decomposition of income and substitution effects in a static model, where the income effect corresponds to the liquidity effect and the substitution effect corresponds to moral hazard. Furthermore, in estimating excess bunching in earnings with respect to a change in taxes, Saez (2001, 2010) show that when the change in the tax rate is small, any excess bunching is a function of the compensated elasticity, a pure substitution effect. Under the assumption that the effective "tax" of 13 percent represents a small change, the estimate of excess bunching can thus be translated into a substitution elasticity of duration in a static model framework. I obtain this elasticity for an alternative estimate of moral hazard by scaling the estimate of excess bunching in the following equation

[11] The results are robust to interacting the polynomial with 9 or 11 day intervals instead of 10 day intervals.

(Saez 2010):

$$e = \frac{dn \Big/ n}{dr \Big/ (1-r)}, \tag{8}$$

where dn is the estimate of excess mass at day 10, n is the time interval for claim exit: either the second week or the first two weeks; and r represents the 13 percent of the wage that claimants give up by leaving WC prior to two weeks. I estimate the excess mass, and the elasticities separately for workers earning above and below the median wage in addition to estimating these statistics for the overall sample. For each estimation, I bootstrap the estimation of excess bunching and the elasticity to obtain standard errors

Table 6 shows the estimates of excess bunching and elasticities. Panel A reports the estimates calculated over the second week; panel B reports the estimates calculated over the first two weeks. Column 1 reports the estimated excess bunching, and column 2 scales this calculation by the average change in the share of wages "given up" by returning to work prior to eligibility for the RP. I estimate that the option value of the RP leads to approximately 3.5 (1) percent more claims ending on day 10, rather than some other day during the second (first two) weeks. The estimate of excess bunching is larger for workers above the median wage: I estimate excess bunching of approximately 4.7 (1.3) percent for workers earning above the median wage, compared to 2.6 (0.6) percent for workers below the median wage. Overall, these estimates are broadly consistent with the hazard estimates in section 4.1. For example, recall that a 50 percent increase in the RP leads to a 1.2 percent decline in the probability that a low-wage worker's claim will end during the first two weeks of the claim. Approximately 56 percent of low-wage workers have claims ending during the first two weeks (excluding the 10th workday), and 1.2 percent of this share is 0.7 percent - similar to the estimate of excess bunching for low wage workers reported in Panel B, column 1. Approximately 54 percent of high-wage workers end their claims before day 10, and 1.7 percent of this total is .92 - slightly lower than the estimate of excess bunching for high wage workers in Panel B of 1.2.

Once scaling these excess bunching estimates by the change in the "tax" $r \big/ (1-r)$, I obtain alternative estimates of the substitution elasticity, the moral hazard effect (Chetty 2006; Saez 2010). As a result, column 2 in panel A shows that the elasticity of claim duration with respect to a change in option value RP is approximately .14 for claimants above the median wage, and .07 for claimants below the median wage. These elasticities are slightly smaller than the estimates derived from the proportional hazard model, but have overlapping confidence intervals, with the elasticities calculated in section 4.1.[12] In general, the evidence of excess bunching provides visual and non-parametric evidence of the moral hazard effect, again suggesting that the moral hazard effect is fairly small, in particular for low-wage claimants.

4.3 Effects on return to work outcomes

Longer claims could also affect outcomes once claimants return to work. On one hand, if the liquidity effect affords workers to more time to recover, this could lead a better match with the employer upon return, potentially increasing earnings relative to what the claimant would have earned if he had returned to work earlier (Boden *et al.* 2001). On the other hand, employers may have a harder time re-integrating employees into the workforce, or may penalize their workers for their longer absence. Higher adjustment costs could lead to lower wages or fewer working hours once a claimant returns to his job (Butler *et al.* 1995). Since the duration of a claim is endogenous to injury severity, it is difficult to determine the effect of claim duration on these outcomes. After demonstrating that the RP lengthens claims, I use the retroactive payment as an instrument for the duration of a claim and estimate the following instrumental variables (IV) regression:

$$\begin{aligned} y_{it} &= \alpha + \gamma d_i + X_{it}'\beta + \varepsilon_{it} \\ d_i &= \theta + \phi RP_i + X_{it}'\delta + \nu_{it} \end{aligned} \tag{9}$$

I examine the effect of longer claims on return to work outcomes by estimating equation 9 with two-stage least squares (2SLS). d_i measures the duration of the claim. I examine several outcome variables in y_{it} including the change in average hours worked per quarter and the change in the average hourly wage, where I take the average over the quarter before and after the injury, omitting the quarter(s) in between the date of injury and the last day for which benefits were paid. Additionally, I estimate the effect of a longer duration on the probability that the claimant returns to the same work as before and the probability that the claimant returns to modified work after the injury.

Panel C in table 7 gives the first stage coefficients: a 100 percent increase in the RP lengthens claims by

[12] If the moral hazard estimates from the proportional hazard model include an income effect due to claimants "spending" the RP in advance of qualifying for it, this could explain why the moral hazard elasticity from the proportional hazard model is slightly larger than the elasticity calculated with excess bunching.

approximately 0.7 days in the overall sample, and 1 day among claims lasting longer than two weeks. Panels A and B in table 7 show the IV and reduced form coefficients on the change in hours and wages between the quarters before and after the injury. Columns 1 and 2 do not suggest that longer claims significantly affect the probability of a claimant returning to the same job, or requiring modifications on their work activities after their injury. Similarly, the coefficients in column 3 do not provide evidence that an increase in claim length significantly affects the hourly wage earned one quarter after the injury, relative to the hourly wage earned one quarter prior to the injury. If liquidity and moral hazard have conflicting effects on post-injury outcomes, this could explain the lack of a result. On the other hand, these negligible effects are consistent with research finding that liquidity effects in unemployment insurance do not significantly improve subsequent job matches (Card *et al.* 2007).

However, column 4 shows that longer claims lead to a small reduction in the number of hours worked after the injury. The instrumental variables estimate implies that increasing claim length by one day leads to a 10 hour decrease in hours worked during the first quarter after an injury. These results are similar for claimants above and below the median wage.[13] Because the first stage represents the total change in duration in response to a change in the RP, it does not distinguish between an increase in duration due to liquidity or moral hazard. As a result, the reduction in hours could reflect a positive or negative outcome, and should be examined further in future work. However, these results suggest that in general, changes in claim length may not have a significant effect on post-injury outcomes for the population on the margin of increasing their claim in response to the RP. Longer claim length may have a larger impact for claimants with significantly longer durations and more severe injuries, a population whose return to work decision would not be influenced by the RP.

5 Robustness checks

5.1 Variation in cash on hand

My empirical strategy exploits variation in the RP generated by the day of the week of the injury to identify the liquidity and moral hazard effects. However, the day of the week also creates variation in the size of

[13] Results available upon request.

the worker's last pre-injury paycheck: workers who would receive larger RPs also earn fewer days of wages during the week of their injury. Approximately 85 percent of workers in Oregon receive their final paycheck during the first two weeks of their claim, meaning workers with larger RPs have less cash on hand during period 1.[14] Consider a revised version of equation 3 to understand the implications of this fact:

$$\alpha_1^* = v(A_1 - s_1 + w_1) - u(A_1(d) - s_1 + b_1) + \beta[V_2(A_2) - J_2(A_2, RP(d))].$$

Assume that d is increasing in the number of waiting period days on which benefits are withheld, increasing the *RP* in period 2 and decreasing A_1. Then, the effect of variation in the date of the injury is as follows:

$$\frac{\partial h_1}{\partial d} = -u'(c_1^n)\frac{\partial A_1}{\partial d} + \beta\frac{\partial OV_1}{\partial d} = -u'(c_1^n)\frac{\partial A_1}{\partial d} - \begin{cases} \beta\frac{\partial V_2}{\partial d} : V_2 > U_2 \\ \beta\frac{\partial U_2}{\partial d} : V_2 \leq U_2 \end{cases} \tag{10}$$

The second term in equation 10 is the same as in equation 5, implying that increasing the *RP* decreases h_1. But the first term is positive and potentially increases h_1, since $\frac{\partial A_1}{\partial d}$ is negative. Ultimately, whether h_1 rises or falls during period 1 will depend on which one of these two effects dominates. If workers have a large amount of cash on hand, then they are better able to smooth their consumption and d likely only has a small effect on A_1, making the first term small. As a result, the incentive in option value will dominate for workers with a high ability to smooth.

However, if workers have limited cash on hand or have a precautionary savings motive (Chetty 2005), d could have a relatively large effect on A_1 and they will reduce consumption while on WC. If $u'(c_1^u)\frac{\partial A_1}{\partial d} < \beta\frac{\partial OV}{\partial d}$, then the option value will dominate, and workers will lengthen their claims. On the other hand, if marginal utility is sufficiently large, any small change in c_1^n will result in $u'(c_1^u)\frac{\partial A_1}{\partial d} > \beta\frac{\partial OV}{\partial d}$, and the reduction in the benefit will increase α_1^*, shortening claims. On average, this effect could attenuate the moral hazard response to the option value during period 1. Additionally, since the workers who would be *most* sensitive to receiving the RP are more likely to leave the sample *prior* to their RP eligibility, using variation in the day of the week to identify the response to the RP could lead to a lower-bound estimate of the liquidity effect. On the other hand, if claimants deplete their cash on hand to smooth through the smaller paycheck during period 1, they could be more sensitive to receiving the RP during period 2.

[14] Based on special calculations from Burgess (2014), approximately 71 percent of workers are paid at least twice a month, and one-half of the remaining 28 percent of workers paid monthly would receive their monthly check during any given two-week period.

I use access to sick leave as a rough proxy to test how sensitive workers are to a change in the size of their final pre-injury paycheck. Because sick leave is managed by the employer and WC payments are managed separately by the insurer, a worker may use sick leave during the waiting period without affecting their eligibility for the RP. However, using sick leave during the waiting period equalizes the size of the final paycheck for workers who are injured on different days of the week. If the smaller paycheck differentially "selects out" workers who are sensitive to small variations in income, workers without sickdays should be less sensitive to the RP than workers who can use sickdays to make up the difference in wages in their final paycheck. On the other hand, the lack of sick leave may lead claimants to deplete their assets during period 1, making them more sensitive to the RP during period 2. I examine whether the results vary with access to sick leave to test for these potential biases.

I obtain national estimates of the share of workers in each industry who have sick leave from the 2010 National Compensation Survey. I adjust the industry-specific estimates by the total share of workers in the West region who have sick leave based on data from the 1999 Employee Benefits Survey (U.S. Bureau of Labor Statistics, 2010, 1999).[15] Table A.2 shows there is considerable variance in the share of workers per industry who have sick leave. While only 24 percent of workers in the food and accommodation industry have sick leave, over 77 percent of workers in utilities have sick leave. Based on the composition of industries in my sample, I approximate that 48 percent of the sample has access to sick leave. I divide the sample into high and low sickday prevalence categories depending on whether at least 50 percent of workers in the industry have access to sick leave, the median industry share in my sample.

Admittedly, workers with and without sick leave could be different along many other characteristics. Tables A.3a and A.3b show that workers without sick leave are younger, more likely to be male, earn a slightly lower wage, and have longer WC spells. While I control for these characteristics, workers without sickdays may claim WC for more severe injuries, if they are willing to "tough out" fairly minor injuries to avoid missing wages that would not be replaced by sick leave. On the other hand, workers without sick leave could claim WC for minor injuries since they don't have an alternative way to cover their wages during missed work time. While claimants with a low likelihood of sick leave are more likely to have cuts or burns, which are typically less severe injuries, they are also more likely to have fractures, which are typically more severe injuries. On average, the difference in total medical costs across groups, another measure of injury

[15] Unfortunately state-specific estimates of the prevalence of sick leave were not available, nor were industry-specific estimates prior to 1999.

severity, is small. For low-wage workers, there is no significant difference in medical costs; for workers above the median wage, the difference less than $200, but is statistically significant.

Table 8 reports the RP coefficients for workers above and below the median wage in industries with a high and low prevalence of sick days, respectively. The results are consistent with the hypothesis that the variation in the final paycheck could attenuate the response to the RP during the first two weeks, in particular for low-wage workers. As shown in column 1 of Panel B, the coefficient during the first two weeks for claimants unlikely to have sickleave is smaller than in the overall sample of claimants earning less than the median wage, and is insignificant. By contrast, the coefficient during period 1 is slightly larger for low-wage claimants who do have access to sick leave: increasing the RP by 1 percent decreases the probability of exit during the first two weeks by approximately .043 percent. This pattern is consistent with the hypothesis that the smaller final paycheck could offset claimants' ability to lengthen their claims in order to receive the RP. Furthermore, the effect after the first two weeks is similar in magnitude and significance across the two samples below the median wage, although it is more persistent for claimants who are likely to have access to sickleave.

If workers above the median wage are less liquidity constrained, they should also be less sensitive to variation in their final pre-injury paycheck, regardless of the likelihood of receiving sick leave. Indeed, Panel A shows that the RP has a significant effect on the hazard rate for both groups: increasing the RP by 1 percent reduces the hazard by approximately 0.047 percent for claimants likely to have sickleave, and 0.037 percent for claimants unlikely to have sick leave. In general, this effect is comparable to the effect in the overall sample of claimants earning above the median wage. Additionally, the coefficient on the RP during period 2 indicates that a 1 percent increase in the RP significantly reduces the rate of leaving WC by 0.04 percent for claimants above the median wage who do not have access to sick leave. While these higher-wage claimants are able to "hold on" during period 1 to become eligible for the RP, this significant result after the first two weeks could suggest that the lower paycheck depletes these claimants' cash on hand, increasing their sensitivity to later receiving the RP.

In general, these findings are broadly consistent with the hypothesis that the smaller paycheck could induce some claimants to return to work prior to receiving the RP, attenuating the moral hazard effect during the first two weeks. Claimants without sick leave who remain in the sample after the first two weeks appear to have depleted their assets, making them more sensitive to the RP. However, the confidence intervals between the coefficients in table 8 and table 4 overlap. As a result, I cannot reject the hypothesis that the trends in these sub-samples and the overall sample of low-wage workers are the same.

5.2 Changes in the composition of claimants

An ideal experiment would use changes in benefits for the entire population of beneficiaries to estimate liquidity and moral hazard effects. In this analysis, however, the liquidity estimate is based on claims that last longer than two weeks. If these claimants are less sensitive to small fluctuations in their benefits, either due to the severity of their injury or a better ability to smooth income, this select sample of claimants could have a lower elasticity with respect to liquidity than the average claimant in the overall population. To examine the extent to which this affects my estimates, I reweight the sample of claimants who have claims less than and greater than two weeks to reflect the overall distribution of claims in the sample.

First, I estimate a propensity score of the probability of remaining out of work at least two weeks on a set of observable covariates including age, gender, pre-injury wage, industry and occupation. I determine which linear and quadratic covariates should be included in the propensity score using the stepwise regression procedure outlined in Imbens (2014). Then, I reweight the sample using the estimated propensity scores to minimize the difference between the sample of claims longer and shorter than two weeks so that the distribution of each group is similar to the overall distribution of claims (Nichols 2008). Figure 6a shows the distribution of propensity scores for the overall sample, the sample of claims less than 10 days and the sample of claims greater than or equal to 10 days. After reweighting the claims, the distribution of propensity scores is better matched across the three groups, as shown in figure 6b.

Table 9 provides coefficients from equation 6 with the reweighted sample. As expected, the coefficients on the liquidity effect during period 2 are slightly larger for claimants below the median wage on the reweighted sample compared to the baseline estimates, suggesting that selection in the sample of claimants could attenuate the baseline estimates of the liquidity effect. Additionally, the coefficient on the first period is smaller in the reweighted sample. Additionally, in the overall sample and the sample of claimants above the median wage, the moral hazard coefficient is smaller in magnitude and the liquidity coefficient is larger and significant. While the confidence intervals on these results do overlap with the baseline results, these estimates nevertheless suggest that the moral hazard effect could be slightly larger without accounting for selection, and the liquidity effect is slightly smaller without accounting for selection, meaning that the baseline estimates could yield a lower-bound estimate of the liquidity to moral hazard ratio. I additionally estimate an alternative hazard model which corrects the hazard model for unobservable differences among claimants that could lead some claimants to systematically leave more quickly than others (Kalbfleish and Prentice 2002). The estimates have the same pattern as correcting the sample for observable differences with the reweighting technique outlined above.[16]

5.3 Effects of employer incentives

WC is unique from other forms of social insurance because the claimant maintains a relationship with his employer. Employers face several costs associated with injuries on the job: the cost of WC insurance premiums, the costs of improving the safety of the workplace, and direct and indirect costs associated with an accident, including productivity losses and repair costs. Employers seeking to minimize these costs may encourage workers to return to work more quickly (Krueger and Burton 1990; McInerney 2010; Bronchetti and McInerney 2015).[17] This could mitigate the overall elasticity of duration with respect to benefits in equation (4). However, if this incentive is correlated with the day of the week, employer incentives could introduce bias in my estimation of the liquidity and moral hazard effects. Since a larger RP would have a larger impact on premium costs, employers may have a greater incentive to encourage workers with the largest potential RPs to return to work before their eligibility for the RP. If true, employer incentives could bias my estimate of moral hazard during period 1 towards zero. Employers may also seek to mitigate the increase in claim duration after the RP is received. However, one additional day of benefits has the same effect on premium costs regardless of the date of injury. As a result, after workers have earned the RP, employer incentives could lead to a smaller elasticity, but likely do not introduce bias. Moreover, while an increase in claim duration increases premium costs, the change in total costs is likely small relative to the costs incurred from an additional injury on the job, and employers likely devote more time to reducing accident costs along other margins.[18]

[16] Available upon request.

[17] Employers may encourage workers to take a longer absence to ensure a complete recovery. This employer response would depend on the severity of the injury, which table 2 shows is broadly balanced across different days of the week. As a result, this incentive does not bias my estimates. See the appendix for more information about the how the role of the employer affects this analysis.

[18] Employers could also reduce the frequency of injury by enhancing safety features. Since most safety features are designed to reduce the frequency of injury, rather than the severity (Occupational Safety and Health Administration 2012), employer's control over safety features likely has a larger effect on the extensive margin of injuries, rather than the intensive margin that would

I use information on Oregon's Employer at Injury Program (EAIP) to empirically examine how employer incentives might affect the response to the RP. The EAIP subsidizes wages for injured workers who return to the same employer, but require modifications to their work activities. If the employer finds transitional work for the injured worker, it receives a subsidy of 45 percent of the injured worker's wages for the first two months after their return to work, and receives additional subsidies for accommodation equipment.[19] Since the EAIP makes it more affordable to accommodate injured workers, it may facilitate an employer's ability to reduce claim length, offsetting the response to the RP. As a result, I split the sample for claimants whose employers have and have not used the EAIP to examine the extent to which employer activity could offset the incentive to lengthen claims during the first two weeks.

Table 10 shows the results for high and low wage workers separately based on whether the employer has ever taken advantage of the Employer at Injury Program (EAIP). Column 3 in table 10 shows that claimants whose employers have not used the EAIP have the same pattern of results above and below the median wage as in the overall sample. On the other hand, claimants whose employers have used the EAIP are slightly less sensitive to the RP during the first two weeks of the claim. While this could be due to employer behavior that makes it easier for workers to return to the job, it also could result if workers who use the EAIP have more severe injuries. However, table A.4 does not show a clear pattern indicating differences in severity depending on employer's use of the EAIP. Hence, these results provide suggestive evidence that employer incentives could offset the response to the option value during the first two weeks of the claim, but broadly employer incentives do not appear to systematically interact with the effect of the RP.

6 Implications for optimal benefits

The ultimate goal of estimating liquidity and moral hazard effects is to determine how a local change in benefits could affect social welfare. Chetty (2008) shows that the optimal benefit level is determined by the first order condition on the social planner's problem:

$$\frac{dW_t}{db_t} = \frac{(1-\sigma_t)}{\sigma_t}\left(\underbrace{\frac{u'(c_t^n)-v'(c_t^e)}{v'(c_t^e)}}_{(1)} - \underbrace{\frac{\varepsilon_{1-d_t,b}}{\sigma_t}}_{(2)}\right). \tag{11}$$

determine eligibility for the RP.

[19] For more details on EAIP, see http://www.cbs.state.or.us/wcd/rdrs/rau/eaip/eaip.html.

The optimal benefit level depends on (1) the relative difference in marginal utilities of consumption in the working and non-working state; and (2) the elasticity of the probability of not working with respect to benefits.[20] The benefit level maximizes social welfare when equation 11 equals zero. While extensive research in WC has yielded estimates of (2), only one paper has attempted to estimate (1) for WC. Bronchetti (2012) uses within-state variation in WC benefits over time to estimate how WC affects consumption following a workplace injury. Under plausible levels of risk aversion and assumptions about the utility function, she combines her estimates on the effects of WC on consumption to a variant of equation 11 and obtains a range of possible optimal replacement rates for WC between 0.1 to 0.6. However, Chetty (2008) shows that the ratio of liquidity to moral hazard effects is a sufficient statistic for (1), without requiring additional assumptions about consumption or utility.

The liquidity to moral hazard elasticities estimated from changes in the *current* benefit level inform whether a local change in benefits would increase or decrease social welfare. If equation 11 yields a positive number when setting (1) equal to the current liquidity to moral hazard ratio, this indicates that increasing the benefit level will increase overall social welfare, and similarly, if the equation yields a negative number, this indicates that the current benefit level is too high: *decreasing* the benefit level would increase social welfare. By taking advantage of the separation of the liquidity and moral hazard responses to the RP, I estimate the liquidity and moral hazard effect that occur at specific points in time during a claimant's absence from work. Hence, applying my estimated liquidity to moral hazard ratio to equation 11 requires assuming that workers' elasticity with respect to lump sum payments is the same across all points in time, and that the elasticity is constant for payments of all sizes. Approximately 80 percent of claimants in my sample exit WC during the first four weeks. As a result, these estimates are based on responses during a time frame when most claimants make a decision about when to return to work.

Scaling the baseline estimates from table 4 by the percentage change in income due to the RP, I approximate that a liquidity elasticity of approximately .24 for low-wage claimants, and moral hazard elasticities of .16 and .23 for low and high wage claimants, respectively. Under the interpretation of the response during period 2 as the liquidity effect and the response during period 1 as the moral hazard effect, and the sum of the two effects as the overall elasticity of the probability of not working with respect to benefits during the

[20]This equation is based on the social planner's problem where that benefits are financed via a lump-sum tax on individuals. In the case of WC, firms pay for insurance directly. See the appendix for details on when this equation will hold when considering the firm's role in the social planner's problem.

first four weeks, I apply these estimates to equation 11 to determine the effects of WC on social welfare.[21] I use the Survey of Occupational Illness and Injury to obtain estimates of $(1-\sigma_t)$, the incidence rate of workplace injury (U.S. Bureau of Labor Statistics 2013). As the rate of workplace injury has declined over time, I present results based on two incidence rates: the incidence rate in 2013, and the average incidence rate between 1994 and 2013, to approximate the incidence over the same time frame used in estimating the liquidity and moral hazard effects.

Table 11 shows the application of equation 11 for individuals above and below the median wage based on both incidence rates. While the magnitude of $\frac{dW}{db}$ is small, the signs of the equation inform whether a marginal increase in benefits would increase or decrease overall social welfare. $\frac{dW}{db}$ is scaled such that the magnitude of the equation can be interpreted as the monetary value of a change in benefits. In other words, column 4 indicates that increasing WC benefits by \$1 would increase an individual's utility by approximately 2 cents per week, or \$1 per year. On the other hand, increasing weekly WC benefits by \$1 would decrease individuals' utility above the median wage by approximately 20 cents per year.[22]

These approximations indicate only small welfare gains to increasing and decreasing benefits; however, they do imply that the optimal benefit level is higher than the current level for lower-wage workers, and vice-versa for higher-wage workers. Additionally, given the potential that the liquidity effect could be under-estimated due to the selection effects, this estimate likely represents a lower-bound on the welfare gains from increasing the WC benefit level. Panel B shows the welfare estimates based on the reweighted coefficients in table 9, which suggest slightly larger welfare gains, in particular for claimants below the median wage.

To understand how estimates from WC claims in Oregon might compare with estimates based on claims in other states, table A.5 compares current WC benefit parameters in Oregon with the average parameter across all other states. While a few states have slightly larger (75-80 percent) or smaller (60 percent) replacement rate, the two-thirds replacement rate in Oregon is quite standard. The minimum benefit level in Oregon, \$50 or 90 percent of the workers' average weekly wage, is low compared to an average of approxi-

[21] As noted in Bronchetti (2012), the elasticity of the probability of not working with respect to benefits is the same as the elasticity of claim duration with respect to benefits if benefits do not influence the frequency of claims. Bronchetti and McInerney (2011) find very small elasticities of the frequency of claims with respect to benefit levels once they flexibly control for pre-injury wages, suggesting that the elasticity of duration with respect to benefits is a reasonable approximation for the overall elasticity in equation 11.

[22] For comparison, Chetty (2008) finds that increasing UI benefits by \$1 per week would increase an individual's utility by approximately 4 cents per week, or \$2 per year.

mately \$150 across all other states.[23] On the other hand, Oregon's maximum benefit is much more generous than the average across other states - approximately \$1120, compared to approximately \$830 across other states (Tambe 2012). In practice, very few people in the claims data reach the maximum benefit level. Finally, the median weekly wage in Oregon is slightly larger, but fairly close to the median wage across all other states.

Additionally, table A.6 shows that workers in Oregon are similar across demographic characteristics and savings habits, using data from the Survey on Income and Program Participation. Oregonians are slightly more likely to have a checking or interest accruing savings account, suggesting that liquidity constraints could be a smaller concern in Oregon than in other states. They are also more likely to owe debt, meaning they could be less constrained in their borrowing as well. Both of these facts mean that the liquidity effect could be smaller in Oregon than in other states, but this hypothesis should be verified with additional research.

While these characteristics suggest that Oregon's WC system and the characteristics of the Oregonian population is broadly similar to other states, the differing minimum and maximum benefits in other states could lead to different conclusions about the welfare impacts of a change in the current benefit level in other states. In particular, if liquidity and moral hazard effects are similar in other states, the benefits paid to lower-wage workers could be closer to optimal in places that have more generous minimum benefits.

7 Conclusion

Despite the large expenditures on social insurance in the United States, relatively little is known about the social welfare effects of many social insurance programs. In particular, little is known about the magnitude of the positive and negative welfare consequences of WC despite the growing policy discussion about reforming WC benefits. I observe how claimants adjust the duration of their WC claims in response to unique variation in a retroactive payment, allowing me to isolate the liquidity and moral hazard effects for WC. I find that the liquidity effect accounts for 60- 65 percent of the increase in claim duration among lower-wage workers, but do not find evidence of a significant liquidity effect for high wage workers. These results are primarily driven by injuries that have variable recovery times, where claimants' decision to return to work could be influenced by small fluctuations in WC payments.

[23] All dollar values in 2012 dollars.

Under the assumption that the elasticity of duration is constant over the size and frequency of the payment, I apply these estimates to the optimal benefit formula outlined in Chetty (2008). While the results suggest that the increasing the benefit level for low-wage workers would increase social welfare, it suggests the opposite result for high wage workers. Variation in pre-injury paychecks may lead some workers to return to work before becoming eligible for the RP, potentially introducing a downward bias in my estimates. However, when I restrict the sample to workers who likely can use sick days to make up the difference in their final paycheck, this does not change the results dramatically, suggesting that this potential bias is small. While changes in the composition of claimants in the sample during the first and second two weeks of the claim could also bias the liquidity to moral hazard ratio, reweighted estimates suggest that bias due to this selection is again likely not very large and, if anything, imply my baseline estimates could present a lower-bound on the potential welfare gains associated with increasing the benefit level.

This analysis also provides evidence that WC claimants respond to small payments (Soueles *et al.* 2006). Additionally, my results demonstrate that low-wage workers are sensitive to fluctuations in income even at the beginning of their WC spell, either due to an immediately binding liquidity constraint, or precautionary savings to prevent a constraint from binding in the future. Both of these findings provide evidence that liquidity constraints are an important consideration for the population of workers at risk of an on-the-job injury.

My analysis of post-injury outcomes does not suggest that small increases in claim duration have a significant effect on the probability of returning to the same work, or on post-injury wages. The analysis does find that longer claims lead to fewer hours worked after an injury; however, the reduction in hours is quite small. The reduction in hours could reflect a positive or negative consequence of longer claims on the post-injury job match. However, even without a substantial effect on post-injury outcomes, an increase in duration due to a liquidity effect itself implies that WC benefits provide insurance value to injured workers, and as a result, this provides evidence that WC benefits relax claimants' liquidity constraints during recovery from an on-the-job injury. Future work could look more in depth at return to work outcomes. A better understanding of whether an increase in claim duration is beneficial or costly to workers could provide information about the welfare effects of WC once a worker returns to the labor force.

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Table 1: Sample Selection Criteria

Restriction	N
Adults >18	614,217
No PPD/TPD	367,249
Five-day workweeks	309,363
Weekday injuries	274,222
Continuous WC spell	170,657
Wed-Fri injuries	96,694

Notes: Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Main sample restrictions exclude: claimants under age 18; claimants who received permanent disability payments (who are unlikely to respond to the retroactive payment) or temporary partial disability payments (likely ineligible or the retroactive payment); claimants who worked more than five days per week or were injured on the weekend, Monday or Tuesday (to improve accuracy of retroactive payment calculation and avoid the Monday effect). Claimants are also excluded if they did not leave work right after the injury, or if they returned to work intermittently during their WC claim.

Table 2: Summary Statistics by Day of the Week of Injury

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Mon	Tue	Wed	Thu	Fri	Pval - MW	Pval - WTh	Pval - WThF
Male	0.74	0.73	0.71	0.72	0.71	0.00	0.00	0.00
Age	36.0	35.87	35.91	36.10	36.08	0.25	0.04	0.07
Wage and benefit information								
Weekly wage	743.37	732.39	725.41	728.50	717.14	0.00	0.26	0.00
WC days paid	13.53	13.59	14.79	14.16	13.89	0.00	0.00	0.00
Retroactive payment	291.59	287.22	281.80	189.54	93.85	0.00	0.00	0.00
Daily benefit	97.20	95.74	94.82	95.24	93.85	0.00	0.21	0.00
Medical cost	2,103.86	2,122.38	2,211.46	2,203.74	2,176.54	0.00	0.83	0.61
Afternoon injury	0.46	0.48	0.50	0.50	0.51	0.00	0.74	0.14
Injury type								
Trauma	0.04	0.04	0.04	0.04	0.04	0.87	0.18	0.21
Fracture	0.09	0.09	0.10	0.10	0.10	0.00	0.68	0.90
Strain	0.63	0.60	0.61	0.59	0.59	0.00	0.00	0.00
Wound	0.21	0.23	0.22	0.23	0.24	0.00	0.00	0.00
Other	0.03	0.04	0.04	0.04	0.04	0.00	0.99	0.96
Industry								
Agriculture	0.06	0.06	0.06	0.06	0.06	0.08	0.74	0.42
Construction	0.13	0.12	0.11	0.12	0.11	0.00	0.20	0.23
Manufacturing	0.21	0.20	0.20	0.20	0.20	0.00	0.23	0.30
Trade	0.16	0.17	0.17	0.17	0.18	0.01	0.18	0.02
Transportation	0.10	0.10	0.10	0.10	0.09	0.97	0.75	0.06
Other	0.34	0.36	0.37	0.36	0.37	0.00	0.64	0.22
Observations	38,517	35,446	31,898	32,704	32,092			

Notes: Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Sample includes claims that lasted at most one year. All dollar values in 2012 dollars. Medical costs reflect total medical expenditures during the WC spell. Injuries are recorded in the data with ICD-9 codes and are grouped here into five broad categories. Industry is recorded with six-digit NAICS codes in the data and grouped here into six broad categories. P-values test the equality of means across different days of the week: column (6) shows the p-values on a test of equality between Monday and Wednesday, column (7) shows the p-values on a test of equality between Wednesday and Thursday, and column (8) shows the p-values on a test of equality between Wednesday, Thursday and Friday.

Table 3: OLS estimates of the effect of the retroactive payment on claim duration

	(1)	(2)	(3)	(4)
Dependent variable: ln(duration)	All	$<$ two weeks	$\geq$ two weeks	$\geq$ eight weeks
Log (RP)	0.020	-0.015	0.029	-0.001
	(0.008)*	(0.007)*	(0.007)**	(0.009)
Observations	96,605	56,656	39,949	8,171

Notes: Standard errors, clustered at the claimant level, in parenthesis. + $p<0.1$, * $p<0.05$, ** $p<0.01$. Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Duration is measured by the number of workdays for which benefits were paid. Column (1) includes all claims in the sample; column (2) limits the sample to claims lasting less than two weeks (i.e., claims ineligible for the retroactive payment); column (3) limits the sample to claims lasting at least two weeks (i.e., claims eligible for the retroactive payment); and column (4) limits the sample to claims lasting at least eight weeks (i.e., claims with durations unlikely to be responsive to the retroactive payment). Sample includes claims for injuries occurring on a Wednesday, Thursday or Friday, lasting at most one year. Regression includes controls for gender, age, pre-injury wage, total medical costs, and year fixed effects.

Table 4: Proportional hazard estimates of the effect of the retroactive payment on the probability of exit from WC at varying points during the claim

	(1) All	(2) Above median wage	(3) Below median wage
Weeks 1-2	-0.058*	-0.042*	-0.029*
	(0.009)	(0.014)	(0.012)
Weeks 3-4	-0.013	-0.033	-0.039*
	(0.012)	(0.022)	(0.019)
Weeks 5-6	0.015	-0.018	-0.043
	(0.018)	(0.033)	(0.028)
Weeks 7-8	0.032	-0.046	-0.016
	(0.023)	(0.044)	(0.037)
Observations	92,735	41,175	51,560

Notes: Standard errors, clustered at the claimant level, in parenthesis. + $p<0.1$, * $p<0.05$, ** $p<0.01$. Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Each column contains the coefficients from a separate regression. Column (1) includes all claims in the sample; column (2) limits the sample to claimants who earned less than the Oregon median wage ($700) prior to their injury; column (3) limits the sample to claimants who earned at least the Oregon median wage prior to their injury. Sample includes claims for injuries occurring on a Wednesday, Thursday or Friday that lasted at most one year. Duration is censored at 60 workdays. All dollar values in 2012 dollars and represented in logs.Regression also includes controls for the claimant's weekly benefit, wage, total medical costs, age, gender, an indicator for claims occurring after 2002, an indicator for experiencing a trauma, fracture, muscle sprain, or cut/burn, indicators for participating in a health support or transportation occupation, a spline in total duration with knots every two weeks, and an indicator for five-day multiples in duration to control for weekly spikes.

Table 5: Proportional hazard estimates of the effect of the retroactive payment on the probability of exit from WC at varying points during the claim, by injury type

	(1) Trauma	(2) Fracture	(3) Sprain	(4) Wound	(5) Other
Panel A: Claimants earning above median wage					
Weeks 1-2	-0.061	-0.055+	-0.030*	-0.041*	-0.090*
	(0.046)	(0.032)	(0.012)	(0.018)	(0.046)
Weeks 3-4	0.030	0.046	-0.051*	-0.000	-0.163*
	(0.069)	(0.043)	(0.017)	(0.029)	(0.071)
Weeks 5-6	-0.083	-0.087+	-0.018	0.022	-0.017
	(0.096)	(0.052)	(0.025)	(0.050)	(0.096)
Weeks 7-8	0.009	-0.037	-0.023	0.077	-0.177
	(0.139)	(0.061)	(0.034)	(0.069)	(0.130)
Panel B: Claimants earning below median wage					
Weeks 1-2	-0.068	-0.052	-0.019	-0.036+	-0.065
	(0.049)	(0.035)	(0.012)	(0.018)	(0.048)
Weeks 3-4	0.028	0.035	-0.062*	-0.017	-0.163*
	(0.077)	(0.048)	(0.019)	(0.033)	(0.081)
Weesk 5-6	-0.100	-0.147*	-0.030	-0.013	-0.011
	(0.109)	(0.060)	(0.029)	(0.058)	(0.112)
Weeks 7-8	-0.058	-0.065	-0.044	0.046	-0.190
	(0.165)	(0.071)	(0.039)	(0.078)	(0.156)
Observations	3,482	8,857	55,024	21,803	3,569

Notes: Standard errors clustered at the claimant level in parenthesis. + $p<0.1$, * $p<0.05$, ** $p<0.01$. Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Each column contains the coefficients from a separate regression restricting the sample to the injury type listed in the column header, where the ln(RP) was interacted with indicators for whether the claimant's pre-injury wage was above or below the median wage in Oregon ($700). Includes claims for injuries occurring on a Wednesday, Thursday or Friday that lasted at most one year. Duration is censored at 60 workdays. All dollar values in 2012 dollars and represented in logs. Regression also includes controls for the claimant's weekly benefit, wage, total medical costs, age, gender, an indicator for claims occurring after 2002, an indicator for experiencing a trauma, fracture, muscle sprain, or cut/burn, indicators for participating in a health support or transportation occupation, a spline in total duration with knots every two weeks, and an indicator for five-day multiples in duration to control for weekly spikes.

Table 6: Excess bunching and the elasticity of claim exit at the threshold for retroactive payment eligibility (two weeks)

	(1) Excess mass of claims at threshold	(2) Elasticity of claim exit at the threshold
	Panel A: Claims ending during the second week	
All	0.035	0.101
	(0.020)	(0.057)
Below Median	0.026	0.073
	(0.020)	(0.055)
Above Median	0.047	0.136
	(0.031)	(0.089)
	Panel B: Claims ending during the first two weeks	
All	0.009	0.061
	(0.009)	(0.061)
Below Median	0.006	0.083
	(0.006)	(0.033)
Above Median	0.013	0.044
	(0.013)	(0.046)

Notes: Bootstrapped standard errors reported in parenthesis. Data from Oregon Department of Consumer and Business Services, includes WC claims from 1987-2012. Column (1) shows the excess mass in the distribution of claim durations exactly at the two week threshold (i.e., the point of eligibility for the RP). To estimate the excess mass I predict a counterfactual count of claims on each day. Then, I calculate the share of claims ending exactly after two weeks under the original and counterfactual distribution, and attribute the difference between these two shares as excess bunching due to the incentive of the RP. Column (2) scales the estimate of excess mass by the relative gain in benefits due to the RP to obtain the elasticity of claim exit at the two week threshold. Panel A shows the estimate of excess mass as a fraction of all claims ending during week 2, and Panel B shows the estimates of excess mass and elastcity as a fraction of all claims ending during the first two weeks.

Table 7: Instrumental variables regressions of claim duration on post-injury labor force outcomes

	(1)	(2)	(3)	(4)
Dependent var:	Same work = 1	Modified work = 1	Chg-log wage	Chg- hours
Panel A: IV Coefficients				
WC days	0.001	0.005	-0.011	-10.703 +
	(-0.006 - 0.017)	(-0.008 - 0.001)	(-0.033 - 0.011)	(-23.500 - 2.095)
Panel B: RF Coefficients				
Log RP	0.001	0.005	-0.008	-7.070 +
	(-0.005 - 0.017)	(-0.007 - 0.001)	(-0.022 - 0.007)	(-13.572 - -0.568)
Mean of dependent variable	0.869	0.016	0.018	-13.81
Panel C: First stage				
	(1)	(2)	(3)	(4)
Dependent variable: claim duration	All	$<$ two weeks	$\geq$ two weeks	$\geq$ eight weeks
Log RP	0.726	0.008	1.050	0.716
	(0.249)**	(0.037)	(0.521)*	(1.498)
Mean of dependent variable	14.3	3.67	31.3	69.9
Obs	38,069	40,901	38,538	41,121

Notes: Standard errors, clustered at the claimant level, in parenthesis. + $p<0.1$, * $p<0.05$, **$p<0.01$. Data from Oregon Department of Consumer and Business Services and the Oregon Employment Department, WC claims from 1999-2012. Sample includes claims for injuries occurring on a Wednesday, Thursday or Friday that lasted at most one year. All dollar values in 2012 dollars and represented in logs. Regression includes controls for gender, age, pre-injury wage, total medical costs, and year fixed effects. In panel C, column (1) includes all claims in the sample; column (2) limits the sample to claims lasting less than two weeks (i.e., claims ineligible for the retroactive payment); column (3) limits the sample to claims lasting at least two weeks (i.e., claims eligible for the retroactive payment); and column (4) limits the sample to claims lasting at least eight weeks (i.e., claims with durations unlikely to be responsive to the retroactive payment). Duration in panel C is measured by the number of workdays for which benefits were paid. F-statistic from the first stage is 8.76.

Table 8: Proportional hazard model estimates of the effect of the retroactive payment on the probability of exit from WC at varying points during the claim, by prevalence of sick days in worker industry

	(1) Baseline	(2) No sickdays	(3) Sickdays
Panel A: Claimants earning above median wage			
Weeks 1-2	-0.042*	-0.037*	-0.047*
	(0.014)	(0.014)	(0.012)
Weeks 3-4	-0.033	-0.042*	-0.029
	(0.022)	(0.020)	(0.018)
Weeks 5-6	-0.018	-0.016	-0.038
	(0.033)	(0.029)	(0.028)
Weeks 7-8	-0.046	0.030	-0.048
	(0.044)	(0.037)	(0.038)
Panel B: Claimants earning below median wage			
Weeks 1-2	-0.029*	-0.017	-0.043*
	(0.012)	(0.014)	(0.013)
Weeks 3-4	-0.039*	-0.039+	-0.048*
	(0.019)	(0.022)	(0.021)
Weeks 5-6	-0.043	-0.021	-0.071*
	(0.028)	(0.033)	(0.031)
Weeks 7-8	-0.016	0.025	-0.093*
	(0.037)	(0.043)	(0.043)
Observations	92,735	41,083	51,652

Notes: Standard errors, clustered at the claimant level, in parenthesis. + p<0.1, * p<0.05, **p<0.01. Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012, the 1999 Employee Benefits Survey, and the 2010 National Compensation Survey. Each column contains the coefficients from a separate regression, where the ln(RP) was interacted with indicators for whether the claimant's pre-injury wage was above or below the median wage in Oregon. Column (1) includes all claims in the sample; column (2) includes claimants who worked in industries where less than 50% of workers have access to paid sick leave; column (3) includes claimants who work in industries where at least 50% of have access to paid sick leave. Sample includes claims for injuries occurring on a Wednesday, Thursday or Friday that lasted at most one year. Duration is censored at 60 workdays. All dollar values in 2012 dollars and represented in logs. Regression also includes controls for the claimant's weekly benefit, wage, total medical costs, age, gender, an indicator for claims occurring after 2002, an indicator for experiencing a trauma, fracture, muscle sprain, or cut/burn, indicators for participating in a health support or transportation occupation, a spline in total duration with knots every two weeks, and an indicator for five-day multiples in duration to control for weekly spikes.

Table 9: Proportional hazard model estimates of the effect of the retroactive payment on the probability of exit from WC at varying points during the claim, reweighted

	(1) All	(2) Above median	(3) Below median
Weeks 1-2	-0.046*	-0.033*	-0.026+
	(0.010)	(0.016)	(0.014)
Weeks 3-4	-0.041*	-0.044*	-0.051*
	(0.013)	(0.022)	(0.019)
Weeks 5-6	-0.009	-0.021	-0.048+
	(0.018)	(0.034)	(0.028)
Weeks 7-8	0.007	-0.051	-0.021
	(0.024)	(0.045)	(0.038)
Observations	92,735	41,175	51,560

Notes: Standard errors, clustered at the claimant level, in parenthesis. + p<0.1, * p<0.05, **p<0.01. Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Each column contains the coefficients from a separate regression. Sample includes claims for injuries occurring on a Wednesday, Thursday or Friday that lasted at most one year. Column (1) includes all claims; column (2) includes claimants who earned more than the median wage in Oregon prior to their injury ($700); column (3) includes claimants who earned less than the median wage in Oregon prior to their injury. In each regression, the sample of claims is reweighted using the predicted probability of the claim lasting longer than two weeks to minimize the distance between the distribution of claims less than two weeks and greater than two weeks, in order to mirror the overall distribution of claims. All dollar values in 2012 dollars and represented in logs. Duration is censored at 60 workdays. Regression also includes controls for the claimant's weekly benefit, wage, total medical costs, age, gender, an indicator for claims occurring after 2002, an indicator for experiencing a trauma, fracture, muscle sprain, or cut/burn, indicators for participating in a health support or transportation occupation, a spline in total duration with knots every two weeks, and an indicator for five-day multiples in duration to control for weekly spikes.

Table 10: Proportional hazard model estimates of the effect of the retroactive payment on the probability of exit from WC at varying points during the claim, by employer use of return to work incentives

	(1) Baseline	(2) EAIP	(3) No EAIP
Panel A: Claimants earning above median wage			
Weeks 1-2	-0.042*	-0.031*	-0.048*
	(0.014)	(0.015)	(0.012)
Weeks 3-4	-0.033	-0.038+	-0.028
	(0.022)	(0.022)	(0.017)
Weeks 5-6	-0.018	-0.015	-0.025
	(0.033)	(0.032)	(0.025)
Weeks 7-8	-0.046	0.014	-0.022
	(0.044)	(0.044)	(0.033)
Panel B: Claimants earning below median wage			
Weeks 1-2	-0.029*	-0.012	-0.043*
	(0.012)	(0.015)	(0.012)
Weeks 3-4	-0.039*	-0.049*	-0.038*
	(0.019)	(0.024)	(0.019)
Weeks 5-6	-0.043	-0.040	-0.044
	(0.028)	(0.037)	(0.029)
Weeks 7-8	-0.016	-0.020	-0.044
	(0.037)	(0.051)	(0.038)
Observations	92,735	36,602	56,133

Notes: Standard errors, clustered at the claimant level, in parenthesis. + p<0.1, * p<0.05, **p<0.01. Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Sample includes claims for injuries occurring on a Wednesday, Thursday or Friday, lasting at most one year. Each column contains the coefficients from a separate regression, where the RP was interacted with indicators for whether the claimant's pre-injury wage was above or below the median wage in Oregon. Column (1) includes all claims in the sample; column (2) includes claimants whose employer ever used the Employer At Injury Program; column (3) includes claimants whose employer never used the Employer at Injury Program (EAIP). All dollar values in 2012 dollars and represented in logs. Duration is censored at 60 workdays. Regression also includes controls for the claimant's weekly benefit, wage, total medical costs, age, gender, an indicator for claims occurring after 2002, an indicator for experiencing a trauma, fracture, muscle sprain, or cut/burn, indicators for participating in a health support or transportation occupation, a spline in total duration with knots every two weeks, and an indicator for five-day multiples in duration to control for weekly spikes.

Table 11: Welfare effects of WC benefits

	(1) Injury incidence rate $(1-\sigma)$	(2) Liquidity elasticity $\frac{\partial h_t}{\partial A_t}$	(3) Moral Hazard Elasticity $\frac{\partial h_t}{\partial w_t}$	(4) Change in welfare $\frac{dW}{db}$
Panel A - Baseline estimates				
Below median wage				
2013	1%	-.24	-.16	.012
	–	(0.12)	(0.07)	(0.06)
Avg 1994 - 2013	1.6%	-.24	-.16	.018
	–	(0.12)	(0.07)	(0.09)
Above median wage				
2013	1%	0	-.23	-.0003
	–	(0.14)	(0.06)	(0.01)
Avg 1994 - 2013	1.6%	0	-.23	-.0004
	–	(0.14)	(0.06)	(0.01)
Panel B - reweighted estimates				
Below median wage				
2013	1%	-.31	-.14	.019
	–	(0.12)	(0.07)	(0.18)
Avg 1994 - 2013	1.6%	-.32	-.12	.028
	–	(0.12)	(0.07)	(0.26)
Above median wage				
2013	1%	-.27	-.18	.011
	–	(0.14)	(0.07)	(0.04)
Avg 1994 - 2013	1.6%	-.27	-.18	.017
	–	(0.14)	(0.07)	(0.06)

Notes: Bootstrapped standard errors reported in parenthesis. Data from Oregon Department of Consumer and Business Services, 1987-2012, and the Survey of Occupational Injuries and Illnesses. Column (1) contains the incidence rate of workplace injury for the relevant time frame as documented by the Survey of Occupational Injuries and Illnesses. Column (2) contains the liquidity elasticities - scaling the coefficients on the ln(RP) during Weeks 3-4 from the proportional hazard model by the percentage change in benefits due to the RP. Similiarly, column (3) contains the moral hazard elasticities, scaling the coefficients on the ln(RP) during weeks 1-2 of the proportional hazard model by the equivalent percentage change in the bi-weekly wage induced by the RP. Under the assumption that the elasticities with respect to liquidity and moral hazard are constant over time, column (4) applies these benefits to equation 11. The value in column (4) represents the monetary value of a change in welfare in response to a $1 change in benefits. Panel A represents the welfare calculations using the baseline estimates, and panel B represents the welfare calculations using the reweighted estimates.

Figure 1: Variation in retroactive payment by day of the week

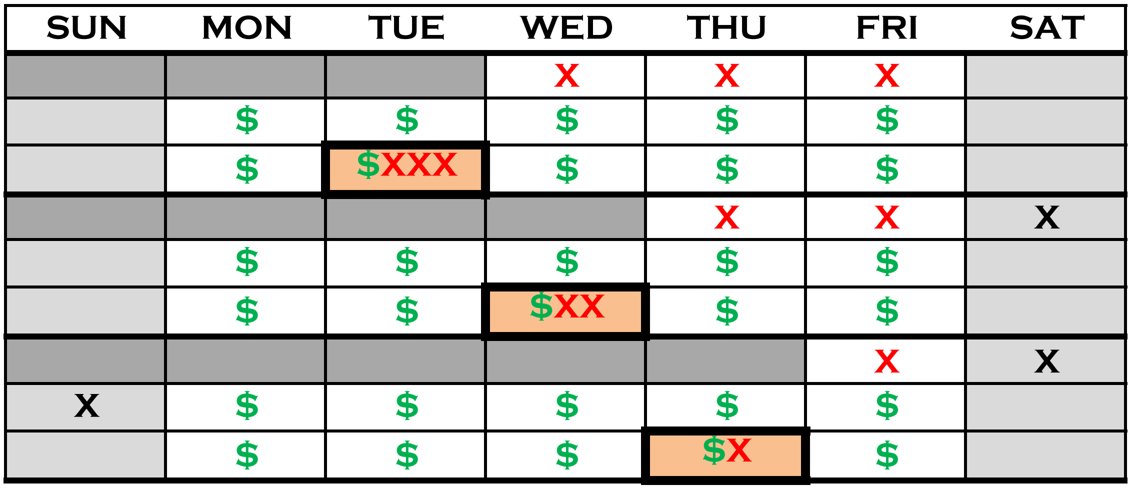
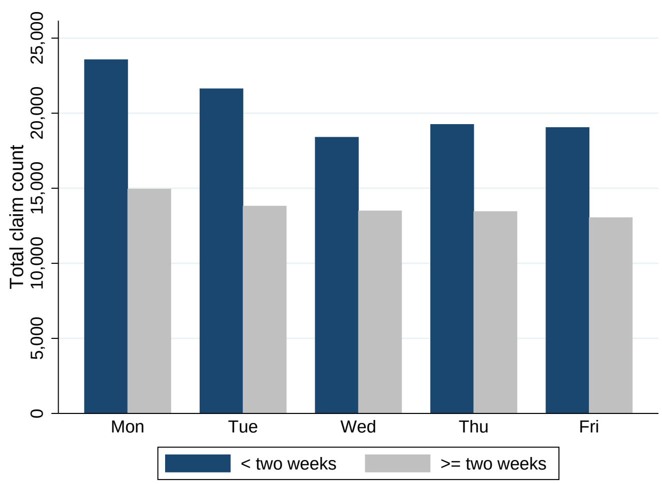


Figure 2: Frequency of WC claims by day of the week of injury

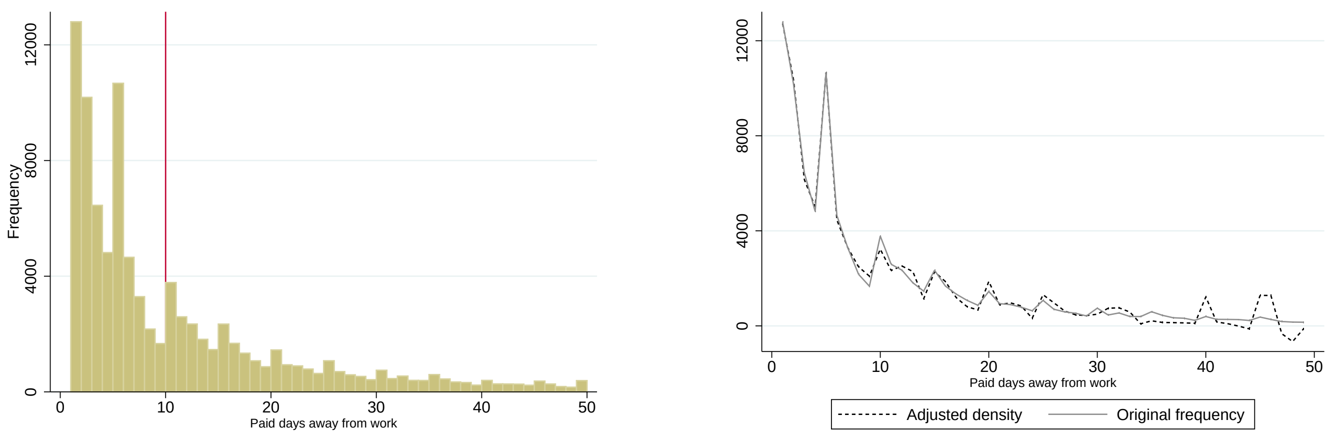


Notes: Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Sample includes claims occuring on a Wednesday, Thursday or Friday that lasted at most one year. The dark bars on the left show the frequency of claims lasting less than two weeks by weekdate of injury; the light bars on the right show the frequency of claims lasting at least two weeks by weekdate of injury.

Figure 3: Distribution of of WC claim duration

(a) Actual distribution of WC claim duration

(b) Actual vs counterfactual distribution

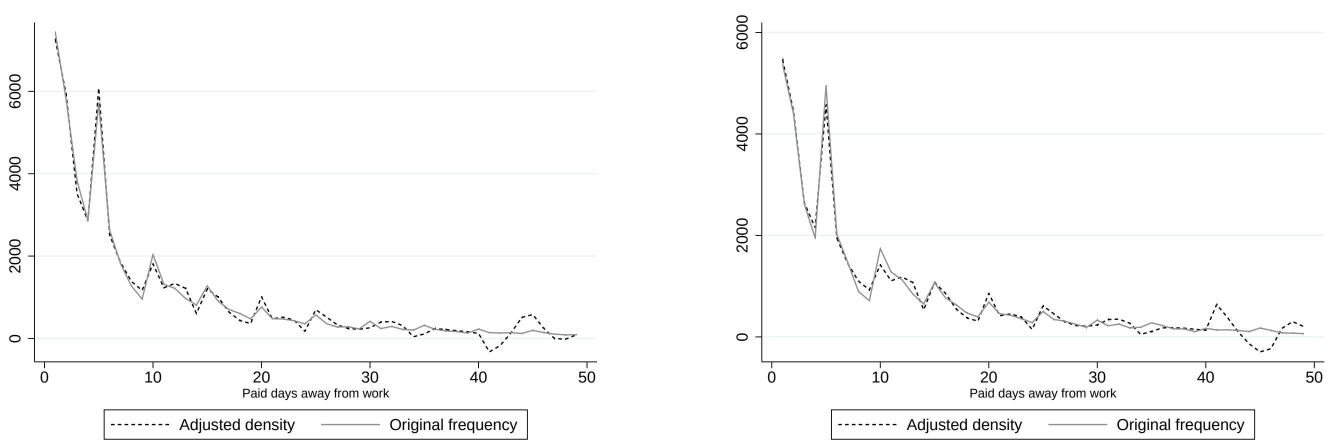


Notes: Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Sample includes claims occuring on a Wednesday, Thursday or Friday that lasted at most one year. Counterfactual distribution is predicted from a regression of the total count of claims ending per each duration on a flexible polynomial interacted for each ten day interval of claim length. In each panel, the x-axis respresnts the duration of claims, measured by the number of workdays for which benefits were paid. Because the sample is limited to claimants working five days per week, 10 days corresponds to two weeks.

Figure 4: Actual vs counterfactual distribution of WC claim duration, by median wage

(a) Claimants earning below Oregon median wage

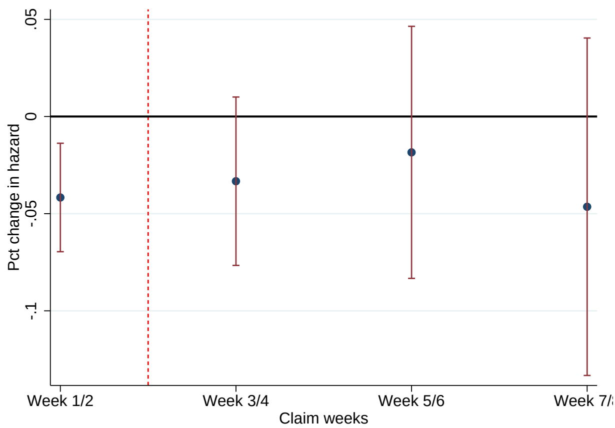
(b) Claimants earning above Oregon median wage



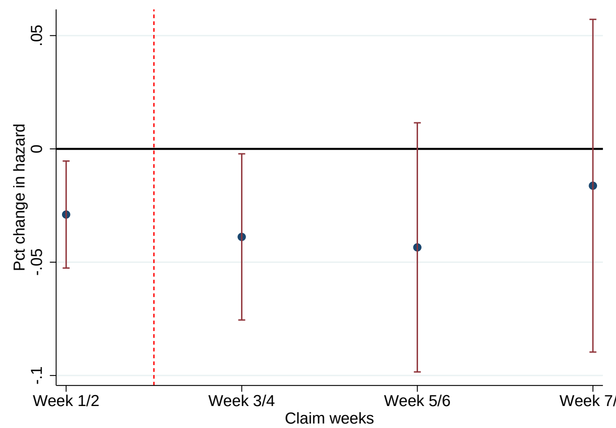
Notes: Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Sample includes claims occuring on a Wednesday, Thursday or Friday that lasted at most one year. Counterfactual distribution is predicted from a regression of the total count of claims ending per each duration on a flexible polynomial interacted for each ten day interval of claim length. In each panel, the x-axis respresnts the duration of claims, measured by the number of workdays for which benefits were paid. Because the sample is limited to claimants working five days per week, 10 days corresponds to two weeks.

Figure 5: Proportional hazard estimates of the effect of the retroactive payment on claim duration, by median wage

(a) Workers earning more than Oregon median wage ($700/wk)



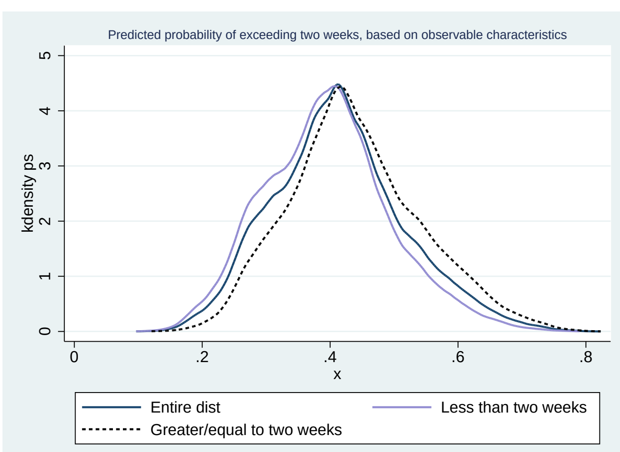
(b) Workers earning less than Oregon median wage ($700/wk)



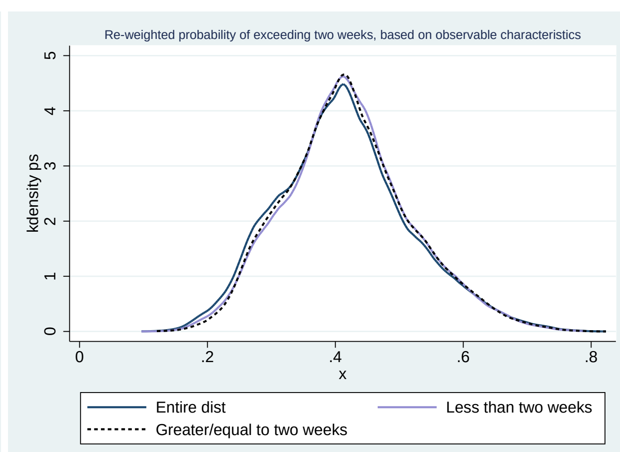
Notes: Bars indicate 95% confidence intervals. Dashed vertical line indicates the two week threshold for RP eligibility. Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Sample includes claims occuring on a Wednesday, Thursday or Friday that lasted at most one year. Duration censored at 60 workdays. Regression also includes controls for the claimant's weekly benefit, wage, total medical costs, age, gender, an indicator for claims occurring after 2002, an indicator for experiencing a trauma, fracture, muscle sprain, or cut/burn, indicators for participating in a health support or transportation occupation, a spline in total duration with knots every two weeks, and an indicator for five-day multiples in duration to control for weekly spikes.

Figure 6: Distribution of propensity scores for the probability of a claim lasting at least two weeks

(a) Propensity score, baseline



(b) Propensity score, reweighted



Notes: Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. ISample includes claims occuring on a Wednesday, Thursday or Friday that lasted at most one year. Panel (a) shows the distribution of predicted probability that the claim lasts longer than two weeks for the entire distribution, claims lasting less than and longer than two weeks, respectively. Panel (b) shows the distribution of the predicted probability or claims lasting less than and longer than two weeks, respectively, reweighted to minimize the distance between the overall distribution and the distribution of the subsample, using estimated propensity score weights.

Appendix

Data sample selection

In order to construct the sample of claims used in analysis, I make the following restrictions to the dataset:

1. Exclude individuals under age 18: these workers comprise less than one percent of the total sample, and are likely to have unusual work schedules and possibly other sources of income from their parents, and so do not represent the typical WC claimant.

2. Exclude the bottom and top 0.5% of the wage distribution: most of these cases represent extreme outliers.

3. Exclude claims prior to 1987: the dataset does not contain the full set of claims in years prior to 1987, so they are dropped from the analysis. These early claims represent approximately 5% of the original sample.

4. Exclude claims lasting more than one year: The distribution of claims has a long right tail. Claims lasting more than one year are likely so severe that they would not be influenced by the retroactive payment: the retroactive payment comprises less than 1 percent of the total WC payments these beneficiaries receive during their claim. Power calculations suggest that I am unable to detect a change in duration in response to a payment comprising less than 1 percent of the total WC payments and so these claims are dropped.

5. Exclude claims where claimants did not stop working immediately: I exclude claims where the date of first payment occurs more than one week after the date of injury or claims where the date the employer was informed of the injury occurs more than one week after the date of the injury. These claims could represent injuries that occur more gradually over time, and workers could have time to adjust to the injury and plan their exit from work.

6. Exclude weekends, Monday and Tuesday injuries: Due to the Monday effect explained in the draft and the fact that claimants who work on the weekend likely do not have typical work schedules, I exclude all of these claims from the analysis.

Optimal benefits formula with the firm

Equation 11 represents the first order condition from the social planner's problem, assuming that individuals pay for the benefit through a lump-sum tax. In the case of WC, the government mandates that firms provide benefits, rather than providing them directly. Under the assumption that employees value the benefit at its full cost, the costs of providing WC will be fully passed through to employees, lowering wages by the full cost (Summers 1989). In this case, the optimal benefits formula in equation 11 is identical when firms pay directly for benefits, rather than workers. Research on the incidence of WC finds that the majority of costs are indeed passed through to the employee, suggesting that this is a reasonable assumption (Dorsey and Walzer 1983; Krueger and Gruber 1990; Fortin and Lanoie 2000). For more detail, consider the static social welfare model from Chetty (2008), where the social planner chooses the benefit level b to maximize the worker's expected utility with following equation:

$$\begin{aligned}\max_b W(b) &= [1-s(b)]u(A+b)+s(b)v(A+w-\tau)-\phi(s(b)) \\ \text{s.t. } b[1-s(b)] &= s(b)\tau\end{aligned} \tag{12}$$

In this equation, s percent of individuals work, receive a wage w, and pay a lump-sum tax τ to finance benefits. $1-s$ percent of individuals do not work and receive the WC benefit b. Note that

$$\frac{d\tau}{db} = \frac{1-s}{s} - \frac{b}{s^2}\frac{ds}{db}.$$

In this equation there are two effects of an increase in benefits on taxes: the first term shows that the tax increases in proportion to the share of individuals that receive the benefit, relative to the share that pays for it. However, the second term represents the fact that since the share of people who are working declines when benefits increase. Since there are now fewer workers to pay the taxes needed to finance benefits, taxes must increase more than they would if there were no adjustment in the duration or incidence of claims. Using this equation, the first order condition for 12 is equal to:

$$\frac{dW}{db} = (1-s)[u'(c^n)-v'(c^e)]+v'(c^e)\frac{b}{s}\frac{ds}{db} = 0. \tag{13}$$

Now, consider the social planner's problem with a constraint of maintaining firm profits above a given level c, rather than a balanced budget constraint:

$$\begin{aligned}\max_b W(b) &= [1-s(b)]u(A+b)+s(b)v(A+w-\tau)-\phi(s(b)) \\ \text{s.t. } F(s)-w\cdot s-b(1-s) &\geq C\end{aligned} \tag{14}$$

In this case, s percent of individuals work, produce $F(s)$ for the firm and earn w. Firms also pay a premium cost based on the share of claimants who do not work. Here, the premium is assumed to be equal to the benefit workers receive, reflecting a case of perfect experience rating. (See Ruser 1985; National Council on Compensation Insurance 2014 for detailed explanations of experience rating in WC). Assume the firm's profit condition in 14 holds with equality, and consider the effect of a change in benefits on wages:

$$\frac{dw}{db} = -\frac{(1-s)}{s} + \frac{F'(s)s-F(s)+b+C}{s^2}\frac{ds}{db}.$$

Here, a change in benefits affects the net wage via a mechanical effect equal to the share of employees who now receive the larger benefit, as well as by an additional amount due to fact that higher benefits reduce the share of employees who work, and this affects productivity and firm costs above and beyond the mechanical effect. Plugging this back equation 14 yields the following equation:

$$\frac{dW}{db} = (1-s)[u'(c_n) - v'(c_e)] + v'(c_e)\left[\frac{F'(s)s - F(s) + b + C}{s^2}\frac{ds}{db}\right]. \tag{15}$$

Under the additional assumption that workers are paid their marginal product (i.e., $F'(s) = w$), this reduces to:

$$\frac{dW}{db} = (1-s)[u'(c_n) - v'(c_e)] + v'(c_e)\left[\frac{b}{s}\frac{ds}{db}\right].$$

Just as in the standard problem, the worker decides whether to stay out or return to work by choosing between his benefit level and his net wage. If employers are able to shift the cost of higher benefits onto the employee, they essentially lower the worker's net wage in the same way as an increase in τ, driving a larger wedge between the market wage and the net return to work.

If firms do not pass the full amount of benefits through to employee wages, higher firm costs will lead to a lower equilibrium level of employment (Summers 1989). Since WC provides insurance for workers who are injured on the job, the welfare consequences of non-employment that is not a direct result of a disability would not be incorporated in the benefit formula. However, as mentioned above, the best estimates of the incidence of WC find that the majority of WC costs are passed through to workers, suggesting that full pass-through is a reasonable assumption.

If $F'(s) > w$, then equation 15 will not reduce to equation 13. As a result, the optimal benefit level will also depend on the effects on firm profits. Reductions in firm profits that are not passed through to wages would increase the cost of social insurance, and an optimal benefit formula that does not incorporate this effect will likely overstate the optimal benefit level. Alternatively, if $F'(s) < w$, then the optimal benefit level could be understated.

Table A.1: Summary stats comparing sample to excluded observations

	Mon Sample	Mon Other	Tue Sample	Tue Other	Wed Sample	Wed Other	Thu Sample	Thu Other	Fri Sample	Fri Other
Male	0.74	0.67	0.73	0.66	0.72	0.66	0.72	0.66	0.71	0.66
Age	36.73	39.78	36.69	39.91	36.81	39.86	36.91	39.80	36.94	39.90
Wage and benefit information										
Weekly wage	740.47	779.96	730.14	778.08	725.89	778.38	728.01	774.94	719.12	770.70
WC days paid	20.18	95.20	20.46	96.80	22.01	96.14	21.03	94.70	20.89	95.81
Retroactive payment	290.91	305.17	286.84	304.40	282.45	301.78	189.66	201.36	94.23	100.49
Daily benefit	96.97	101.72	95.61	101.47	95.05	101.51	95.31	101.11	94.23	100.49
Medical cost	3,216.33	12,123.76	3,270.93	12,541.87	3,447.86	12,319.75	3,352.20	12,060.50	3,372.65	12,307.94
Afternoon Injury	0.46	0.46	0.48	0.48	0.50	0.50	0.50	0.51	0.50	0.53
Injury type										
Trauma	0.05	0.05	0.05	0.06	0.05	0.05	0.05	0.05	0.05	0.05
Fracture	0.12	0.13	0.13	0.13	0.13	0.13	0.13	0.13	0.13	0.13
Strain	0.59	0.65	0.57	0.64	0.57	0.64	0.56	0.64	0.56	0.64
Wound	0.20	0.09	0.22	0.09	0.21	0.09	0.22	0.10	0.22	0.10
Other	0.03	0.08	0.04	0.08	0.04	0.08	0.04	0.08	0.04	0.08
Industry										
Agriculture	0.06	0.05	0.06	0.05	0.06	0.06	0.06	0.05	0.06	0.06
Construction	0.13	0.12	0.13	0.12	0.12	0.12	0.12	0.12	0.12	0.12
Manufacturing	0.21	0.22	0.20	0.22	0.20	0.22	0.20	0.22	0.19	0.22
Trade	0.16	0.16	0.17	0.16	0.17	0.16	0.16	0.16	0.17	0.17
Transportation	0.10	0.09	0.10	0.09	0.10	0.08	0.09	0.08	0.09	0.08
Other	0.34	0.35	0.35	0.36	0.36	0.36	0.36	0.36	0.37	0.36
Observations	52,560	47,996	49,185	44,542	45,180	44,290	45,717	44,520	44,748	42,705

Notes: Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Main sample restrictions exclude weekend claims, claims opened more thanonce and lasted more than one year. All dollar values in 2012 dollars.

Table A.2: Frequency of sickdays by industry

Industry	Share of workers in industry	Share of industry with sick leave
Agriculture	0.06	0.30
Mining	0.006	0.51
Utilities	0.01	0.77
Construction	0.11	0.30
Manufacturing	0.19	0.51
Wholesale trade	0.05	0.66
Retail trade	0.12	0.43
Transportation/warehousing	0.10	0.60
Information	0.01	0.74
Finance and insurance	0.01	0.76
Real estate	0.01	0.67
Professional/technical	0.01	0.61
Management	0.002	0.74
Waste management	0.07	0.33
Educational services	0.04	0.65
Health care/social assistance	0.09	0.65
Leisure/hospitality	0.01	0.26
Accommodation/food services	0.06	0.24
Other services	0.03	0.44
Public administration	0.03	0.74
Weighted average:		48.4%

Notes: Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012, National Compensation Survey, 2010 and Employee Benefits Survey, 1999.

Table A.3: Observable characteristics by sick day prevalence

(a) Below median wage

	Low sickday	High sickday	Pvalue
Male	0.65	0.56	0.00
Age	32.72	35.25	0.00
Weekly wage	438.45	481.38	0.00
Median wage	0.00	0.00	
WC days paid	16.70	15.81	0.00
Retroactive payment	115.46	128.33	0.00
Daily benefit	58.47	64.20	0.00
Medical cost	2,278.79	2,297.04	0.67
Wed	0.33	0.33	0.18
Thu	0.33	0.34	0.00
Fri	0.34	0.32	0.00
Afternoon	0.54	0.51	0.00
Trauma	0.04	0.04	0.72
Fracture	0.10	0.09	0.00
Strain	0.58	0.62	0.00
Wound	0.26	0.22	0.00
Other	0.04	0.04	0.04
Agriculture	0.09	0.00	0.00
Construction	0.14	0.01	0.00
Manufacturing	0.00	0.35	0.00
Trade	0.30	0.09	0.00
Transportation	0.00	0.12	0.00
Other	0.47	0.44	0.00
Observations	32,999	29,168	

(b) Above median wage

	Low sickday	High sickday	Pvalue
Male	0.91	0.78	0.00
Age	36.98	40.56	0.00
Weekly wage	1,004.64	1,013.52	0.00
Median wage	1.00	1.00	
WC days paid	17.14	15.40	0.00
Retroactive payment	258.04	263.82	0.00
Daily benefit	129.58	132.16	0.00
Medical cost	2,553.61	2,388.61	0.00
Wed	0.33	0.33	0.90
Thu	0.34	0.34	0.19
Fri	0.33	0.33	0.23
Afternoon	0.44	0.49	0.00
Trauma	0.04	0.04	0.77
Fracture	0.13	0.10	0.00
Strain	0.55	0.64	0.00
Wound	0.24	0.18	0.00
Other	0.04	0.04	0.62
Agriculture	0.19	0.00	0.00
Construction	0.41	0.02	0.00
Manufacturing	0.00	0.34	0.00
Trade	0.21	0.10	0.00
Transportation	0.00	0.22	0.00
Other	0.19	0.32	0.00
Observations	19,133	34,527	

Notes: Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Claims are included in the high sickday prevalence category if at least 50 percent of workers in the the claimant's industry has access to paid sick leave. Sickday prevalence estimates obtained using the National Compensation Survey, 2010 and Employee Benefits Survey, 1999.

Table A.4: Observable characteristics by EAIP participation

(a) Below median wage

	Never used EAIP	Used EAIP	Pvalue
Male	0.63	0.56	0.00
Age	32.99	34.28	0.00
Weekly wage	455.24	469.50	0.00
Median wage	0.00	0.00	
WC days paid	14.37	14.16	0.30
Retroactive payment	120.30	124.43	0.00
Daily benefit	60.72	62.61	0.00
Medical cost	2,077.60	2,168.59	0.08
Wed	0.33	0.33	0.97
Thu	0.33	0.34	0.20
Fri	0.34	0.33	0.21
Afternoon	0.53	0.52	0.30
Trauma	0.04	0.04	0.61
Fracture	0.09	0.08	0.00
Strain	0.57	0.63	0.00
Wound	0.26	0.22	0.00
Other	0.04	0.04	0.21
Agriculture	0.05	0.03	0.00
Construction	0.10	0.03	0.00
Manufacturing	0.16	0.18	0.00
Trade	0.19	0.23	0.00
Transportation	0.05	0.07	0.00
Other	0.45	0.46	0.04
Observations	33,463	17,317	

(b) Above median wage

	Never used EAIP	Used EAIP	Pvalue
Male	0.88	0.77	0.00
Age	38.01	40.00	0.00
Weekly wage	1,001.56	1,031.90	0.00
Median wage	1.00	1.00	
WC days paid	14.78	13.63	0.00
Retroactive payment	256.77	267.84	0.00
Daily benefit	129.38	133.95	0.00
Medical cost	2,356.56	2,220.84	0.00
Wed	0.33	0.33	0.28
Thu	0.34	0.35	0.12
Fri	0.33	0.32	0.01
Afternoon	0.46	0.49	0.00
Trauma	0.04	0.04	0.26
Fracture	0.12	0.10	0.00
Strain	0.57	0.64	0.00
Wound	0.23	0.18	0.00
Other	0.04	0.04	0.41
Agriculture	0.10	0.03	0.00
Construction	0.21	0.09	0.00
Manufacturing	0.22	0.23	0.13
Trade	0.14	0.14	0.11
Transportation	0.12	0.17	0.00
Other	0.20	0.35	0.00
Observations	25,186	20,728	

Notes: Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Claims are separated depending on whether or not the employer ever used the Employer at Injury Program (EAIP), a program that subsidizes employer efforts to induce employees to return to work after a WC claim.

Table A.5: Workers compensation benefit parameters in Oregon vs all other states, 2012

	(1) Oregon	(2) All other states
Replacement rate	0.67	0.68
Minimum weekly benefit	50	151
Maximum weekly benefit	1120.55	832.08
Median hourly wage	17.14	16.64
Median weekly wage	685.60	665.61

Notes: Data from the Worker's Compensation Research Institute, 2012, and the Occupational Employment Statistics, 2012. The all other states column represents an average of the values from all states excluding Oregon. All dollar values in 2012 dollars.

Table A.6: Demographic and savings habits in Oregon vs all other states, 2009

	Other states	Oregon	P-value
Demographics			
Female	0.51	0.52	0.33
Nonwhite	0.20	0.14	0.00
Married	0.40	0.40	0.97
Ed $<$ HS	0.34	0.30	0.01
Ed - HS	0.20	0.21	0.41
Ed- some college	0.26	0.28	0.04
Ed - BA+	0.20	0.20	0.89
Work-limiting disability	0.04	0.04	0.15
Monthly earnings	1,516.27	1,410.79	0.22
Benefit receipt			
On WC	0.00	0.00	0.37
Rec noncash ben	0.33	0.33	0.81
Rec cash ben	0.08	0.08	0.53
Debt and Savings			
Total debt owed	4,306.98	3,659.21	0.13
Non-interest checking acct value	242.23	302.07	0.05
Interest acct value	5,558.53	6,412.43	0.12
Have any debt	0.34	0.39	0.00
Have non-int check acct	0.18	0.25	0.00
Have an interest acct	0.43	0.50	0.00
Observations	90,177	1,042	

Notes: Data from the 2008 Survey on Income and Program Participation, wave 4. The all other states column represents an average of the values from all states excluding Oregon. All dollar values in 2012 dollars. Statistics calculated with SIPP respondent weights.