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

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

Stephanie Rennane*

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Abstract

I examine the roles of moral hazard and increased liquidity in explaining the relationship between Workers' Compensation (WC) benefit levels and injury duration. Using a discrete proportional hazard model and exploiting variation in the timing and size of a retroactive lump-sum WC payment, I decompose the benefit-duration elasticity into the response to increased liquidity and to a decreased opportunity cost of missing work. I estimate that the liquidity effect accounts for 50 - 60 percent of the increase in claim duration, suggesting that WC provides important insurance value for injured workers.

Keywords: liquidity, moral hazard, duration, disability, workers' compensation

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1 Introduction

Workers' Compensation (WC) insures individuals against the health and income shock of an illness or injury that occurs on the job. In 2014, annual expenditures on WC were $62 billion (Baldwin and McLaren 2016), nearly two times annual expenditures on Unemployment Insurance (UI) in the same year ($36 billion, U.S. Department of Labor 2015). Injured workers can experience substantial consumption losses after a disabling event on the job (Galizzi and Zagorsky 2009; Bronchetti 2012), meaning that WC could play an important role in smoothing consumption for injured workers. On the other hand, WC benefits lead workers to extend their absences from work, and could increase the incidence of on-the-job injuries (e.g., Butler and Worrall 1985; Krueger 1990*a*,*b*; Meyer, Viscusi and Durbin 1995; Neuhauser and Raphael 2004; Bronchetti and McInerney 2011). This literature has generally been viewed to provide evidence of moral hazard, but the overall benefit-duration relationship could reflect both moral hazard via a substitution effect, and a potential response to the income effect (Baily 1978; Nyman 2003). The former response reduces social welfare, while the latter is welfare-enhancing if it helps a risk-averse worker smooth consumption.

In recent years, 33 states have reduced WC benefits or introduced changes making it more difficult to qualify for the program (Qiu and Grabell 2015). These changes have been followed by stories in popular press showing that injured workers are struggling to make ends meet (Grabell and Berkes 2015). However, there is little empirical evidence about the relative magnitude of the insurance value and distortionary costs of WC with which to determine the welfare consequences of these reforms (Meyer 2002). I fill this gap by providing new evidence on the liquidity-moral hazard tradeoff for WC.

This paper adds evidence to a broader literature estimating the insurance value of social insurance programs (e.g., Chetty 2006, 2008; Gruber 1997; Bronchetti 2012; Shimer and Werning 2007; LaLumia 2013; Meyer and Mok 2013; East and Kuka 2016). As outlined in the framework by Baily (1978), the optimal level of benefits in a social insurance program equates the marginal benefit of the program, measured by the associated gain in consumption smoothing, with the marginal cost of the program, measured by the associated behavioral response in work effort. Estimates of the marginal benefit of social insurance can be obtained by directly analyzing changes in consumption (e.g., Gruber 1997; Bronchetti 2012; East and Kuka 2016), or by using revealed preference methods (e.g., Chetty 2008; Shimer and Werning 2007). Chetty (2006, 2008) shows that examining the duration response to an unconditional lump sum payment can identify the marginal benefit of a social insurance. If the worker is liquidity constrained, such a payment increases consumption smoothing, alleviates the need to rush back to work prematurely, and allows the worker to choose a more optimal recovery length (Chetty and Finkelstein 2013). If the worker is *not* liquidity constrained, the lump sum should not affect the length of recovery, since the worker can already attain his or her optimal claim length.

In addition to this potential liquidity effect, all workers have an incentive to lengthen their claims in response to the substitution effect in WC benefits, which are paid conditional on remaining out of work. Thus comparing the duration response to an unconditional payment and the overall WC benefit-duration elasticity informs the relative magnitude of the liquidity and moral hazard effects. Chetty (2008) shows that, under certain assumptions, the liquidity and moral hazard effects are sufficient statistics to determine the social welfare consequences of a marginal change in social insurance benefits.

I identify liquidity effects in WC by analyzing a retroactive lump-sum payment (RP) to WC claimants in Oregon. In all states, injured workers face a waiting period before they receive any WC cash benefits. The waiting period in Oregon is three *consecutive* days from the beginning of the claim, including holidays, weekends, and unscheduled workdays. 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, increasing their second bi-weekly WC check by 10 percent on average. Eligible claimants will receive the RP in their second WC check regardless of when they return to work. As a result, if claimants with larger retroactive payments 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. The retroactive payment reimburses benefits for *scheduled* workdays during the waiting period rather than *calendar* days. This means that identical claimants with identical work schedules who are injured on different days of the week will have different sized retroactive payments. I use the date of injury as a source of exogenous variation in the size of the retroactive payment.

Using state administrative claims data, I estimate a discrete proportional hazard model and decompose the elasticity of claim duration into moral hazard and liquidity responses by estimating changes in the rate of exit from WC before and after eligibility for the retroactive payment. To use this retroactive payment to identify the liquidity effect, I assume 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 assess the validity of these assumptions and find that the frequency and distribution of observable characteristics of claims are balanced across the date of injury. I also analyze a subgroup of claimants who are most likely to have similar levels of cash on hand regardless of their date of injury, and find similar results to my baseline estimates.

I obtain moral hazard and liquidity elasticities in the range of 0.13-0.22 and 0.24-0.30, respectively. These estimates imply an overall duration-benefit elasticity of approximately 0.4-0.5, which is similar to estimates in the extant literature (Meyer, Viscusi and Durbin 1995; Neuhauser and Raphael 2004). Based on these estimates, the liquidity effect accounts for 50-60 percent of the overall duration response to a change in benefits. Applying my liquidity and moral hazard elasticities to the optimal benefit formula from Chetty (2008), I conclude that increasing benefits would increase overall social welfare, implying that the current benefit level is lower than optimal for this population of injured workers.

I identify these liquidity effects based on variation in a one-time payment during a WC claim that typically lasts a few weeks. The marginal WC claimant who responds to this payment is sensitive to small, short-term fluctuations in their income. While other work has found evidence of liquidity effects with large lump-sum payments (e.g., Chetty 2008; Schmieder, von Wachter and Bender 2012; LaLumia 2013), and some work has analyzed longer term variation in consumption for injured workers (Bronchetti 2012), this research demonstrates that small and temporary income shocks can affect disabled workers' behavior. This is consistent with findings that individuals are sensitive to variation in the timing of paychecks or regular benefit checks (Stephens 2003, 2006), and this sensitivity could occur due to low liquidity or consumption commitments (Soueles, Parker and Johnson 2006; Chetty and Szeidl 2007; Pollak 2017). Regardless of the source of constraints, however, my results demonstrate that the timing, as well as the level, of support is an important consideration in the design of public benefits.

2 Waiting Periods and Retroactive Payments in Workers' Compensation

I identify liquidity and moral hazard effects by taking advantage of a common feature of WC programs: waiting periods and retroactive payments. All state WC programs have a waiting period at the beginning of the claim, and in 46 states, claimants can eventually receive a retroactive payment (RP) for this waiting period.[1] The waiting period in Oregon is three *consecutive* days from the beginning of the claim, including holidays, weekends, and unscheduled workdays. If the injury lasts longer than two weeks, workers become eligible for an RP equal to the benefits they *would have received* during the waiting period.[2] WC checks

[1] The length of the waiting period and claim duration before claimants receive the RP both vary across states (Tambe 2012). See Information Technology and Research Section (2012) for details on the general structure of WC payments in Oregon.

[2] Workers also are eligible for the RP if they are admitted to the hospital, regardless of how long their claim lasts. Unfortunately, the Oregon Workers' 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 Workers' 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).

are paid every two weeks relative to the injury date. Since claimants are not eligible for the RP during the first two weeks, any response to a 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.[3] If workers cannot borrow against the future benefit, the response during the first two weeks represents a moral hazard effect (Shavell and Weiss 1979; Chetty and Finkelstein 2013).[4]

As an example of how the RP creates variation in the size of this one-time unconditional payment, 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. The effect of cash on hand will be most important for lower

[3]Based on conversations with insurers in Oregon, it is common policy for claims adjusters to explain the waiting period during their first conversation with workers at the beginning of the claim. The RP is also explained on the Oregon Workers' Compensation Division website and explained in brochures and forms provided to injured workers at the beginning of the claim.

[4]Relaxing this assumption would imply that workers can consume some of the RP during the first two weeks, prior to receiving it. In this case, the response during the first two weeks reflects a combination of liquidity and moral hazard effects, and my estimate of the liquidity effect based on the second two weeks would be a lower bound on the true liquidity effect.

SUN	MON	TUE	WED	THU	FRI	SAT
			X	X	X	
	$	$	$	$	$	
	$	$XXX	$	$	$	
				X	X	X
	$	$	$	$	$	
	$	$	$XX	$	$	
					X	X
X	$	$	$	$	$	
	$	$	$	$X	$	

Figure 1: Variation in Retroactive Payment by Day of the Week

Notes: This figure describes the interaction between the three day waiting period and a claimant's regular work schedule. The first panel demonstrates the scenario where a worker is injured on Wednesday, typically works Monday - Friday. This worker's the three-day waiting period will occur on Wednesday- Friday (indicated by X). The worker will be eligible for regular WC payments beginning the following Monday (indicated by $). If the claimant is still out of work two weeks later, they will be eligible for an RP equal to 3 days of benefits (indicated by $XXX). The subsequent panels repeat the exercise for a worker injured on Thursday or Friday. In these cases one or more of the waiting period days occurs over the weekend, when the worker was not scheduled to work, and thus the retroactive payment is only provided days when the claimant would have been at work. The smaller RP is indicated by $XX or $X indicating RPs equal to 2 and 1 days of benefits for Thursday and Friday injuries, respectively.

income claimants 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.

3 Data and Summary Statistics

I analyze a rich administrative dataset from the Oregon Department of Consumer and Business Services, Workers' Compensation Division (ORWC) which contains complete information on closed claims for which cash benefits were paid between roughly 1987 and 2012 (Oregon Department of Consumer and Business Services 2015). The dataset includes detailed information about the amount of benefits and duration of the claim. 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.

I make several restrictions to derive the sample used for this analysis. I exclude workers receiving permanent disability benefits. 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. Individuals excluded from the sample are older, have slightly higher wages, and are more likely to have suffered severe injuries. These restrictions predominantly exclude claimants who are unlikely to be responsive to the RP.

Table 1 shows the observable characteristics of claimants in the sample across days of the week. Over 70 percent of the sample is male, and the average age of claimants is 36. Approximately 60 percent of all injuries are muscle strains or sprains, 10 percent are bone breaks or fractures, and 20-24 percent of injuries are wounds (cuts or burns). Nearly 65 percent of claimants worked agriculture, construction, trade, transportation, or manufacturing prior to their injury. The mean weekly wage ranges between \$720-\$740; the median weekly wage ranges from \$630-\$650 in 2012 dollars.[5]

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 claims are more frequent on Monday and Tuesday, particularly those lasting less than two weeks. Additionally, table 1 shows that injuries occurring on Monday and Tuesday are significantly more likely to occur in the morning (P-value < 0.01), have a shorter average duration (P-value < 0.01), and are more likely to be hard-to-verify injuries such as muscle strains (P-value < 0.01). 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 2016). 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

[5] I inflate all monetary variables to 2012 dollars using the nominal growth rate in Oregon's state average weekly wage, approximately \$700 (Peniston 2014).

Table 1: Claimant Characteristics by Day of the Week of Injury

	Mon	Tue	Wed	Thu	Fri
Male	0.74	0.73	0.71	0.72	0.71
Age	36	36	36	36	36
Wage and benefit information					
Weekly wage	743	732	725	728	717
WC days paid	13.5	13.6	14.8	14.2	13.9
Retroactive payment	292	287	282	190	94
Daily benefit	97	96	95	95	94
Medical cost	2,104	2,122	2,211	2,204	2,177
Afternoon	0.46	0.48	0.50	0.50	0.50
Injury type					
Trauma	0.04	0.04	0.04	0.04	0.04
Fracture	0.09	0.09	0.10	0.10	0.10
Strain	0.63	0.60	0.61	0.59	0.59
Wound	0.21	0.23	0.22	0.23	0.24
Other	0.03	0.04	0.04	0.04	0.04
Industry					
Agriculture	0.06	0.06	0.06	0.06	0.05
Construction	0.13	0.12	0.11	0.12	0.11
Manufacturing	0.21	0.20	0.20	0.20	0.20
Trade	0.16	0.17	0.17	0.17	0.18
Transportation	0.10	0.10	0.10	0.10	0.09
Other	0.34	0.36	0.37	0.36	0.37
Observations	38,517	35,446	31,898	32,704	32,092

Notes: 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.

Friday, where the frequency of injuries is stable. Because the weekend creates variation in the size of the RP, this restriction still allows me to identify claimant responses to the RP. While an F-test formally rejects the hypothesis that the observable characteristics are jointly balanced across the latter days of the week, the magnitude of the differences in these characteristics is small. Still, I include these characteristics as controls in the analysis. To account for more notable differences in Friday injuries, such as a lower average weekly wage, I restrict the analysis to Wednesday and Thursday injuries in a robustness check. As shown in Table 6, the estimates with this restriction are similar.

Figure 3 shows a histogram of claim length, measured by the number of workdays for which benefits were paid. There is a long and thin right tail to the distribution of claims: approximately 92 percent of claim durations in my sample are less than 40 workdays, and 96

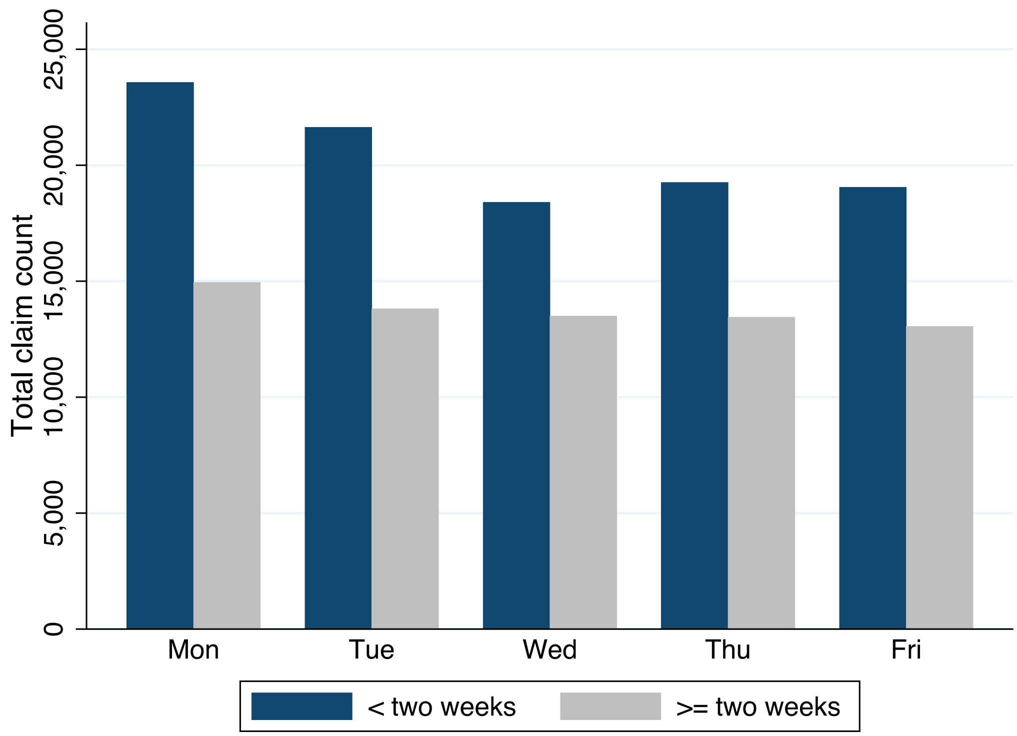

Figure 2: Frequency of WC Claims by Day of the Week of Injury
Notes: Oregon Department of Consumer and Business Services, WC claims 1987-2012. Bars show the frequency of claims by date of injury. The dark bars on the left show the frequency of claims lasting less than two weeks; the light bars on the right show the frequency of claims lasting at least two weeks.

percent of durations are less than 60 workdays. Additionally, figure 3 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]

4 Hazard Analysis

I estimate the discrete proportional hazard model with the complementary log-log function shown below, which allows me to observe how observable characteristics affect the probability of exit during grouped time intervals, in this case, two-week intervals (Allison 1982; Meyer

[6]Figures available upon request.

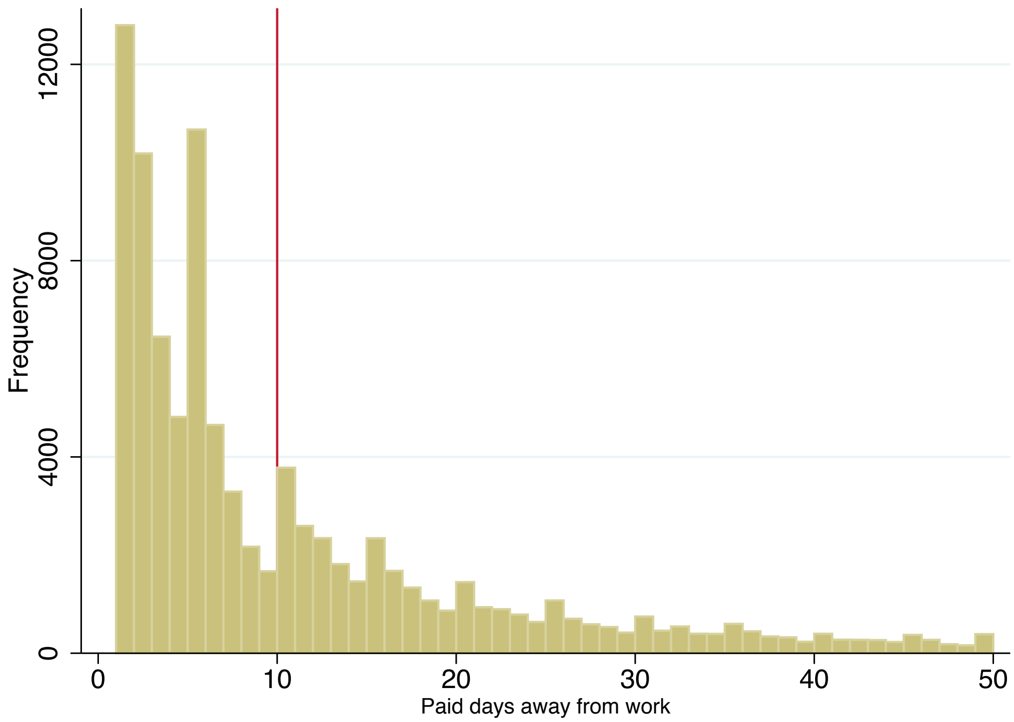

Figure 3: Actual distribution of WC Claim Duration

Notes: Oregon Department of Consumer and Business Services, WC claims 1987-2012. Sample includes claims occurring on a Wednesday, Thursday or Friday that lasted at most one year. The x-axis represents 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.

1990; Butler, Baldwin and Johnson 2001; Jenkins 2005). The equation is as follows:

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

where h_{it} represents the probability of individual i leaving WC during period t, conditional on *not* leaving WC prior to period t. I control for duration dependence over time with indicators representing every two week period in the claim, represented by the γ_k terms, where each $k \in (1, t)$ indicates a two-week period.[7] Then, I interact these indicators with $ln(RP_i)$, allowing the elasticity of the hazard rate with respect to the RP to vary over the duration of the claim. I adjust the hazard rate for time-invariant individual observable char-

[7] Results are robust to including a flexible polynomial for each two-week period of the claim instead of an indicator.

acteristics in $X_i'\beta$ including the claimants' pre-injury wage, weekly WC benefit, gender, age, and total WC-paid medical costs. I include a parsimonious set of indicators for broad injury categories and key occupation groups. Because of the spikes in the frequency of claim exits shown in figure 3, I also include indicators for durations in multiples of five. 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 coefficients appreciably.[8]

While an ideal experiment would use changes in benefits for the entire population of beneficiaries to estimate liquidity and moral hazard effects, the liquidity estimate in this analysis is driven by the subsample of claims that last longer than two weeks. Because their claims are longer, this select sample of claimants could have a lower elasticity with respect to liquidity than the average claimant in the overall population, perhaps due to the severity of their injury or a better ability to smooth income over a longer absence. As a result, I estimate a propensity score of the probability of remaining out of work at least two weeks on a set of pre-determined observable covariates using the stepwise regression procedure outlined in Imbens (2015). Then, I reweight claims shorter and longer than two weeks to reflect the overall distribution of claims using the estimated propensity score (Nichols 2008).[9] Furthermore, I show in Table 3 that the liquidity results are driven by the types of injuries with claims that are expected to last longer than two weeks.

In addition to estimating the hazard rates on the overall sample, I interact the model with an indicator for different wage groups to allow flexibility in the baseline hazard rate for claimants with different pre-injury earnings. Indeed, a test of the proportionality rejects the hypothesis that the hazard rates for different wage groups are proportional. While WC claimants tend to be a lower income population overall, liquidity effects are likely to be most

[8]Censoring claims above 40 and 100 days yield similar results.

[9]Results without propensity score re-weight are available in the appendix

pronounced for those with the lowest wages and presumably, lowest wealth.[10]

Figure 4 shows the unadjusted survival curves for different claims by day of the week. The survival curve for Wednesday claims is higher than the curve for Thursday or Friday, and the difference widens after the 10-day mark. Wilcoxon and Logrank tests both confirm that this difference is statistically significant. To formalize the descriptive pattern shown in the figure, table 2 shows the coefficients for $ln(RP_i) * \gamma_t$ for the overall sample and interacted with an indicator for claimants earning above or below Oregon's median wage. 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 0.025 - 0.046 percent. Workers also significantly lengthen their claims in response to receiving the lump-sum payment. A 1 percent increase in the RP reduces the hazard of leaving WC during the *second* two weeks by an additional 0.04 - 0.05 percent. The marginally significant coefficient for lower earning individuals in weeks 5-6 likely reflect some spillover of the liquidity effect into period 3.

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 1 - 2 percent decrease in the probability that a worker's WC claim will end during the first two weeks. For workers whose claim exceeds two weeks, a 50 percent increase in the RP decreases the probability of leaving during the second two week period by 2-2.5 percent. Furthermore, the coefficient on the first two week period is larger for higher earners. While I cannot reject the equality of these coefficients, the results suggest that the moral hazard effect could be stronger among higher earners.

Figure 5 and 6 display the coefficients from equation 1 interacted with wage quartiles. Despite fairly wide confidence intervals, the figures generalize the trend of a larger moral hazard as incomes increase, and a smaller liquidity response as incomes increase. This trend is again consistent with the RP playing a larger role in consumption smoothing for lower earners. While all individuals receive information about the RP at the time they file a

[10] The survival curves for claims above and below the median wage in Oregon are available in the appendix.

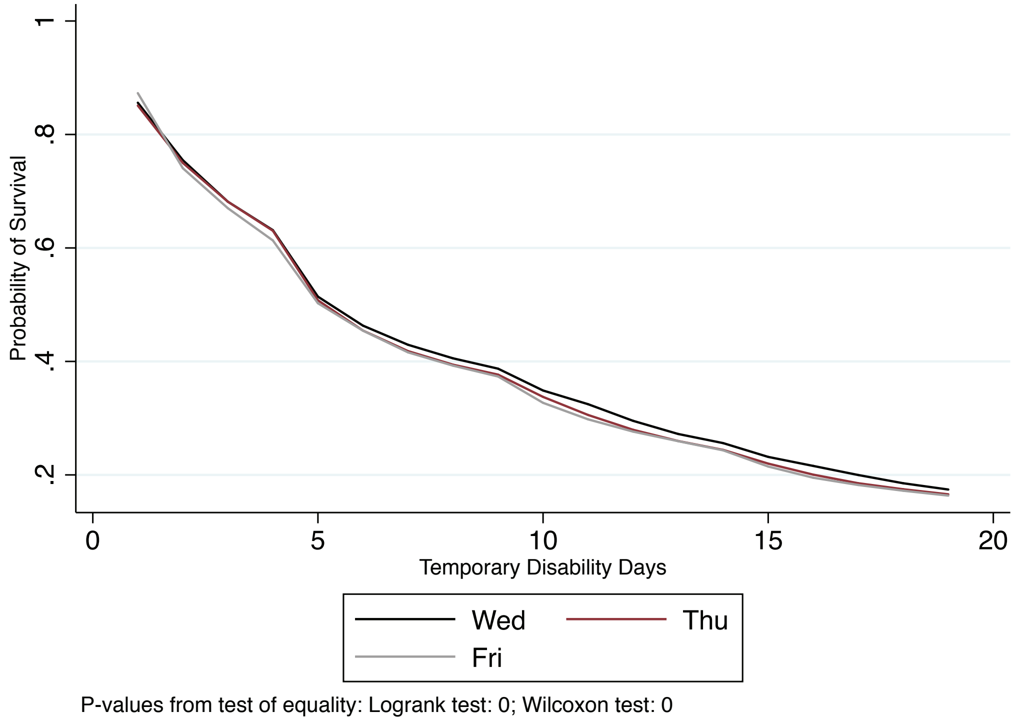

Figure 4: Survival Curves of WC Claim Duration by Day of the Week
Notes: Oregon Department of Consumer and Business Services, WC claims 1987-2012. Sample includes claims occurring on a Wednesday, Thursday or Friday that lasted at most one year. The x-axis represents the duration of claims, measured by 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.

claim, the larger moral hazard response for higher earners could reflect the effect of increased information about the RP, perhaps through an attorney or third-party representative.

To put these results in context, a 50 percent increase in the RP amounts to an 8 percent increase in the bi-weekly WC benefit, on average. Similarly, 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 if he returns to work during period 1. Scaling the coefficients in table 2 by these amounts, I obtain a liquidity elasticity of 0.28-0.30 and a moral hazard elasticity of approximately 0.14-0.22. This implies an overall elasticity in the range of 0.4-0.5, which is consistent with other elasticities in the literature (Meyer, Viscusi and Durbin 1995; Neuhauser and Raphael 2004). Based on my estimates, the liquidity effect amounts to approximately 50-60 percent of the total response to a change in benefits, with the liquidity

Table 2: 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	-0.046	-0.035	-0.025
	(0.010)	(0.010)	(0.011)
Weeks 3-4	-0.041	-0.048	-0.049
	(0.013)	(0.014)	(0.015)
Weeks 5-6	-0.008	-0.032	-0.042
	(0.018)	(0.020)	(0.023)
Weeks 7-8	0.007	-0.022	-0.037
	(0.024)	(0.027)	(0.031)
Observations		92,652	

Notes: Oregon Department of Consumer and Business Services, WC claims 1987-2012. Standard errors, clustered at the claimant level, in parenthesis. Coefficients show interaction of ln(RP) with an indicator for each two-week time interval during the claim. Columns (2) and (3) show the coefficients from one regression where the RP term is additionally interacted with indicators for earnings above/below the median wage. Sample includes claims for injuries occurring on a Wednesday, Thursday or Friday that lasted at most one year. Regression also controls for the claimant's weekly benefit, wage, total medical costs, age, gender, indicators for claims post-2002, injury type, occupation, five-day multiples in duration, and a spline in total duration. Claims are reweighted using the inverse propensity of the claim lasting more than two weeks. $p<0.1$ +, $p<0.05$ *, $p<0.01$ **

effect comprising a higher share of the total response for lower earners.

Injured workers, however, are not completely in control of the length of their 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.[11] Workers facing fairly minor injuries should be less likely to have a doctor certify that their injury warrants two weeks away from work, while workers facing severe injuries will likely remain out of work longer than two weeks, regardless of how large their RP might be. Furthermore, workers with variable recovery times and less observable injuries may be able to adjust their claim length in response to the RP (Dionne and St. Michel 1991). In order

[11] Under Oregon WC law, injured workers have control over choosing their doctor.

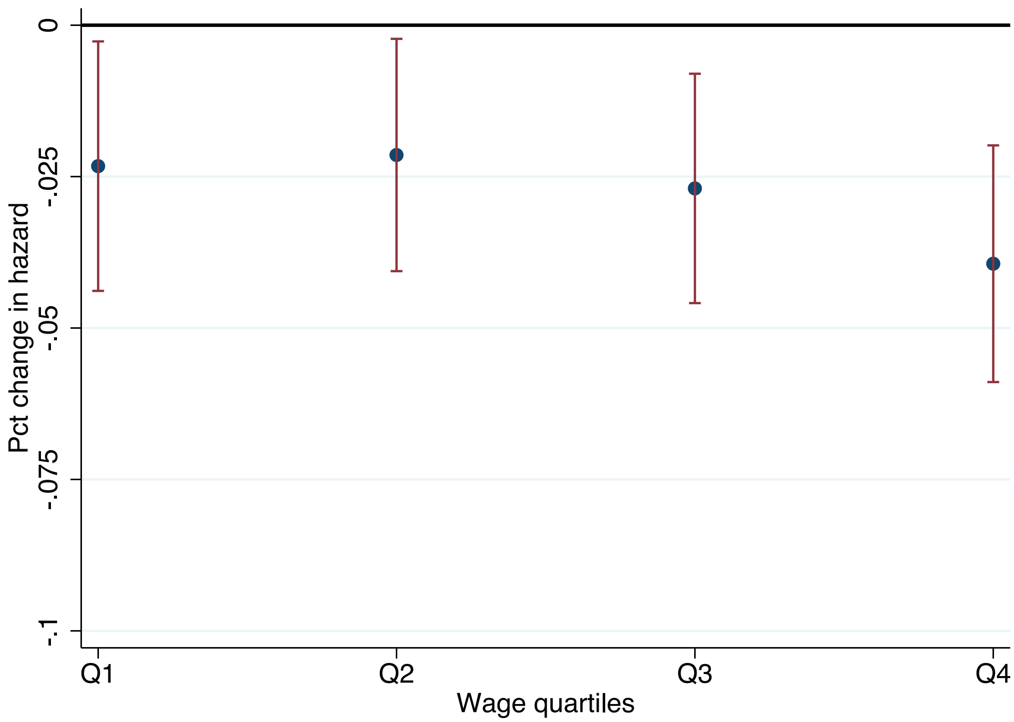

Figure 5: Coefficients on the Moral Hazard Effect, by Wage Quartile

Notes: Oregon Department of Consumer and Business Services, WC claims 1987-2012. Graph plots coefficients on $ln(RP_i) * \gamma_t$ interacted with wage quartiles. Coefficients shown for Weeks 1-2 for each quartile. Bars indicate 95% confidence intervals. Sample includes claims occurring on a Wednesday, Thursday or Friday that lasted at most one year.

to test this hypothesis, table 3 presents estimates from equation 1 where the coefficients are additionally interacted with broad injury categories.

The coefficients interacted with injury categories are shown in table 3. First of all, the results show that the estimated liquidity response is driven by claims that tend to last longer than two weeks on average. Column 4 shows that there is a statistically significant moral hazard effect, but not a significant liquidity effect, for claimants with cuts and burns. The average duration for these claims is 9.7 days, slightly less than 2 workweeks, meaning that these workers are on the margin of qualifying for the RP and thus have an incentive to try to lengthen their claim during the first two weeks. However, there is not a liquidity response during the second period, perhaps because these claims tend not to be severe enough to persist after two weeks.

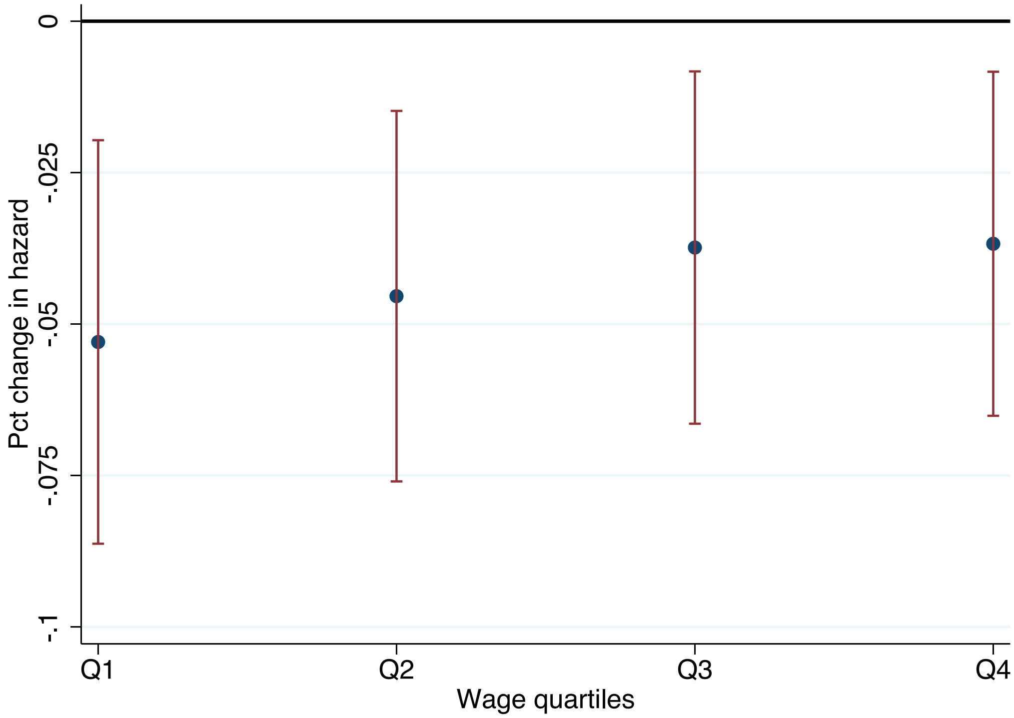

Figure 6: Coefficients on the Liquidity Effect, by Wage Quartile

Notes: Oregon Department of Consumer and Business Services, WC claims 1987-2012. Graph plots coefficients on $ln(RP_i) * \gamma_t$ interacted with wage quartiles. Coefficients shown for Weeks 3-4 for each quartile. Bars indicate 95% confidence intervals. Sample includes claims occurring on a Wednesday, Thursday or Friday that lasted at most one year.

By contrast, I estimate a statistically significant liquidity effect for all other injury types (including sprains, fractures, traumatic injuries). Average claim duration for these injury types ranges between 15 and 20 days, or 3-4 weeks. Unlike workers with cuts or burns, workers with the other types of injuries receive the RP at the time that they are on the margin of returning to work, and having extra cash on hand could influence their decision about when to return. Moral hazard effects are largest for sprains, particularly among lower wage workers. This is consistent with earlier work showing that the moral hazard response is important for workers with injuries that are difficult to diagnose and verify (Dionne and St. Michel 1991). Traumatic injuries and fractures, which are easier to verify, have smaller moral hazard responses. Workers with traumatic injuries or fractures also represent the most severe injuries in my sample and are the least likely to return to work in the first two weeks of their injury, another explanation for the smaller moral hazard response for these injury groups.

Table 3: Proportional Hazard Estimates of the Effect of the Retroactive Payment on the Probability of Exit from WC, by Injury Type

	(1) Trauma	(2) Fracture	(3) Sprain	(4) Wound	(5) Other
Panel A: All Claimants					
Weeks 1-2	-0.027	-0.040	-0.052	-0.048	-0.044
	(0.011)	(0.010)	(0.010)	(0.011)	(0.011)
Weeks 3-4	-0.035	-0.121	-0.045	0.010	-0.051
	(0.015)	(0.014)	(0.013)	(0.013)	(0.015)
Panel B: Claimants earning below median wage					
Weeks 1-2	-0.005	-0.015	-0.032	-0.018	-0.017
	(0.013)	(0.011)	(0.011)	(0.011)	(0.013)
Weeks 3-4	-0.043	-0.140	-0.057	0.009	-0.055
	(0.018)	(0.017)	(0.015)	(0.016)	(0.020)
Panel C: Claimants earning above median wage					
Weeks 1-2	-0.014	-0.029	-0.036	-0.043	-0.035
	(0.012)	(0.011)	(0.010)	(0.011)	(0.012)
Weeks 3-4	-0.044	-0.121	-0.050	-0.006	-0.063
	(0.017)	(0.015)	(0.014)	(0.016)	(0.018)
Average claim duration (workdays)	15.4	21.6	14.7	9.7	14.9
Observations			92,652		

Notes Oregon Department of Consumer and Business Services, WC claims 1987-2012. Clustered standard errors in parenthesis. Regression repeats baseline specification, now additionally interacting ln(RP)*γ_t with injury type. Panels (B) and (C) show the coefficients from one regression where the RP term is additionally interacted with indicators for earnings above/below the median wage.

4.1 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, this would lead to a large share of claims ending exactly at the point where workers become eligible for the RP. Indeed, figure 3 exhibits a spike in claim exits at exactly two weeks. If 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, indicating the behavioral response is driven by moral hazard and not liquidity.

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 the distribution of claim length would be smooth without the discrete change in the RP eligibility after two weeks. However, figure 3 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:

$$n_d = \sum_{t=1}^{5} f(d) * \mathbb{I}[d \in \{10(t-1), 10 * t\}] + \beta S_d + \epsilon_d \tag{2}$$

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.[12] Additionally, S_d is an indicator for exits occurring at any interval of 5 days. 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 ending on each day. Then, I calculate the number of claims ending at exactly 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 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 7 compares the actual density of claim exit with the estimated counterfactual density of claim exit. Comparing the two densities suggests a small amount of excess bunching around the two week mark, when claimants would become eligible for the RP. Additionally, figures in the appendix show that the excess bunching is 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 claimant's pre-injury bi-weekly wage. Under the assumption that the effective "tax" of 13 percent represents a small change, any excess bunching is a function of a pure substitution effect, and can thus be interpreted as the moral hazard effect in a static model framework (Chetty 2006; Saez 2001, 2010). I scale the estimate of excess bunching to obtain an alternative estimate of the moral hazard elasticity in the following equation (Saez 2010):

$$e = \frac{dn \big/ n}{dr \big/ (1-r)}, \tag{3}$$

where dn is the estimate of excess mass at day 10, n is the time interval for claim exit. I set this to the second week of the claim to focus on cases most likely to be on the margin

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

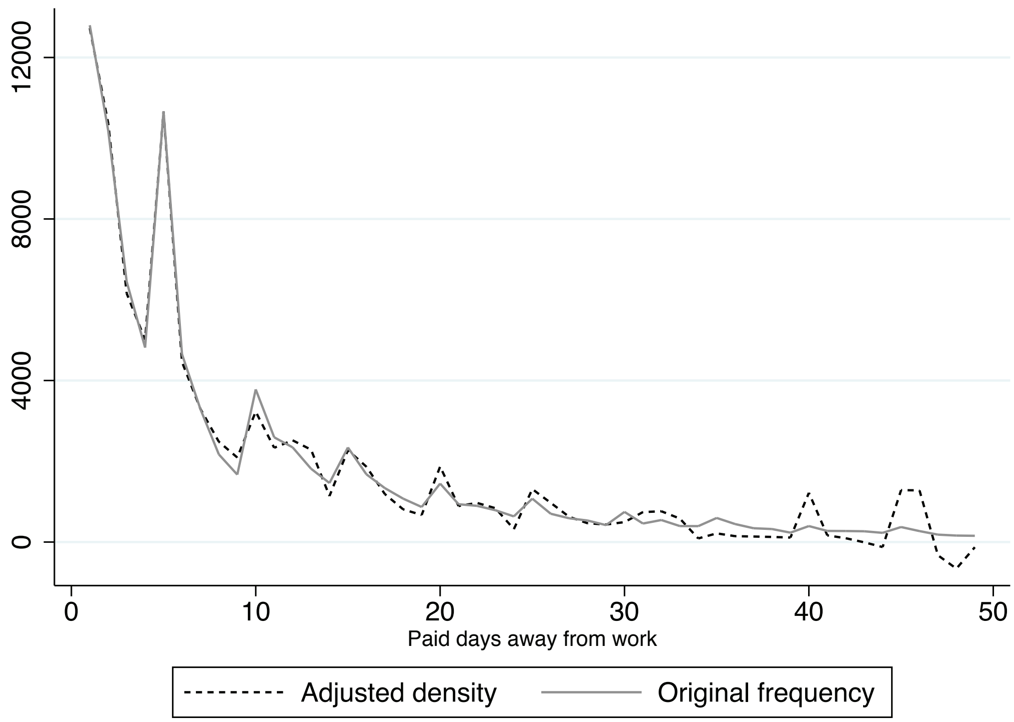

Figure 7: Actual vs counterfactual distribution of WC Claim Duration

Notes: Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Sample includes claims occurring on a Wednesday, Thursday or Friday that lasted at most one year. The x-axis represents 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. 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.

of extending to two weeks, but obtain similar results when setting the time period to be the first two weeks. dr represents the 13 percent of "tax" that claimants incur 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 4 shows the estimates of excess bunching and elasticities. 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 percent more claims ending on day

10, rather than some other day during the second weeks. The estimate of excess bunching is larger for workers above the median wage: I estimate excess bunching of approximately 4.7 percent for workers earning above the median wage, compared to 2.6 percent for workers below the median wage.

Table 4: Excess bunching and the elasticity of claim exit at the threshold for retroactive payment eligibility

	(1) Bunching	(2) Elasticity
All	0.035	0.101
se	(0.020)	(0.057)
Below Median	0.026	0.073
se	(0.020)	(0.055)
Above Median	0.047	0.136
se	(0.031)	(0.089)

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). 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.

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 elasticities derived from the proportional hazard model, but have overlapping confidence intervals.[13]

[13]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 larger than the elasticity calculated with excess bunching.

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. However, the day of the week also creates variation in the size of the worker's paycheck during the week of the injury (hereafter, the period-of-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 period-of-injury paycheck during the first two weeks of their claim, meaning workers with larger RPs have less cash on hand during period 1.[14]

The depleted cash on hand during the first two weeks could create a bias going in either direction. On one hand, workers may not be able to extend their claims during the first two weeks if they are cash constrained, meaning that the most sensitive workers could be more likely to exit prior to RP eligibility. On average, this effect could attenuate the moral hazard response during period 1. If 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, leading to an upward bias on the liquidity effect.[15]

Without data on savings or other assets, it is impossible to test which effect dominates directly. However, I address this concern in several ways. First of all, the estimates by injury type in table 3 allow comparison of the effects across workers with different levels of ability to return to work during the first two weeks. Workers with fractures and traumatic injuries are unlikely to return during the first two weeks, even if this means they would be more

[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.

[15]See the appendix for an extended explanation of these potential biases.

cash constrained due to a smaller paycheck. The response during period 1 is insignificant for these injuries, while the liquidity effect is on the larger end of the range of estimated liquidity effects, suggesting that the variation in period-of-injury paychecks could indeed introduce an upper bound on the liquidity effect.

Secondly, I use access to sick leave as a proxy to test how sensitive workers are to a change in the size of their period-of-injury paycheck. Because the employer manages sick leave and insurers separately manage WC payments, a worker may use sick leave during the waiting period without affecting his or her eligibility for the RP. Using sick leave during the waiting period thus equalizes the size of the period-of-injury paycheck for workers who are injured on different days of the week without affecting the size of, or eligibility for, the RP.

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).[16] As shown in the appendix, 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, approximately 48 percent of the total sample has access to sick leave. I divide the sample into high and low sick day 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. I control for the observable differences shown in the appendix, including differences in age, gender, wage, and industry. However, the overall severity of claims is similar: there is no significant difference in total medical costs between claims with a high and low propensity of sick leave. Additionally, the proportion of claims with high and

[16] Unfortunately neither state-specific estimates nor industry-specific estimates of the prevalence of sick leave were available prior to 1999.

low propensity of sick leave that exceed two weeks is very similar, at 41 and 42 percent, respectively. Thus, claims most sensitive to the period-of-injury paycheck (i.e., those with a low probability of having sick leave) do not appear more likely to exit prior to RP eligibility.

Table 5 shows that while moral hazard effects are similar for claimants high and low probability of sick leave, the liquidity effects are larger for claimants with low probability of sick leave, suggesting that variation in period-of-injury paycheck could exacerbate sensitivity to the RP. Estimates for the sample of claimants with sick leave are similar to the overall sample. Increasing the RP by 1 percent decreases the probability of exit during the first two weeks by 0.024-0.042 percent, and a 1 percent increase in the RP decreases the probability of exit by approximately 0.040-0.047 percent for workers who are more likely to have access to sick leave.

Table 5: Proportional Hazard Model Estimates of the Effect of the Retroactive Payment on the Probability of Exit from WC, by Prevalence of Sick Days in Worker Industry

	(1) All	(2) Above median wage	(3) Below median wage
	Panel A: High Probability of Sick Leave		
Weeks 1-2	-0.042	-0.033	-0.024
	(0.010)	(0.010)	(0.011)
Weeks 3-4	-0.040	-0.041	-0.047
	(0.013)	(0.014)	(0.015)
Weeks 5-6	-0.007	-0.025	-0.040
	(0.018)	(0.020)	(0.023)
Weeks 7-8	0.061+	-0.014	-0.039
	(0.032)	(0.027)	(0.031)
	Panel B: Low Probability of Sick Leave		
Weeks 1-2	-0.043	-0.034	-0.025
	(0.013)	(0.010)	(0.011)
Weeks 3-4	-0.059	-0.068	-0.059
	(0.019)	(0.014)	(0.016)
Weeks 5-6	-0.032	-0.056	-0.057
	(0.019)	(0.014)	(0.024)
Weeks 7-8	0.007	-0.027	-0.031
	(0.025)	(0.028)	(0.032)
Observations		92,652	

Notes Oregon Department of Consumer and Business Services, WC claims 1987-2012, 1999 Employee Benefits Survey, and 2010 National Compensation Survey. Clustered standard errors in parenthesis. Regression repeats baseline specification, now interacting ln(RP)*γ_t with an indicator for whether less than 50 percent of workers have access to paid sick leave in the claimant's industry. p<0.1 +, p<0.05 *, p<0.01 **

These findings imply that variation in the period-of-injury paycheck could introduce a slight upward bias in the main results. Estimates based on the sample of those who have sick leave are the least susceptible to this bias due to these workers' ability to eliminate the differences in the period-of-injury paycheck. The results based on the sick leave sample yield a similar estimate of the overall elasticity in the range of 0.4-0.5, and a similar liquidity to moral hazard ratio as in the overall sample, suggesting the upward bias is small.

5.2 Comparison of Tuesday and Wednesday Claims

Next, I perform falsification tests to confirm that the response to the RP does not result from a systematic correlation with the day of the week. For example, if claimants tend to "round up" the length of their claim until the end of the week, this would also result in Wednesday claims having a longer duration. First, I show in an appendix table that the probability of returning to work on any given day of the week (e.g., Monday or Friday) is not significantly related to the day of injury for claims exceeding two weeks. The one exception to this finding is that Wednesday claims are more likely to return to work on a Wednesday, and Thursday claims are slightly more likely to return to work on a Thursday. This pattern of returning on the same day of the week as injury, rather than extending claims through the end of the week, works against any concerns of Wednesday claims being differentially longer because they "round up" their duration until the end of the week.

Conditional on all other observable characteristics, claimants injured on Tuesdays and Wednesdays will receive the same RP and there should not be a significant relationship between the RP and claim duration. As a result, I re-estimate the model with Tuesday and Wednesday claims. Table 6 shows that the coefficients for this specification are wrong-signed, and with the exception of period 1, insignificant. These findings also do not provide any evidence of a systematic pattern of claimants "rounding up" to the end of the week effect.[17]

Limiting the analysis to two days also reduces the overall sample size by one-third meaning that the lack of a significant response between Tuesday and Wednesday could result from imprecision. I compare the results in Panel A with Panel B, where I perform the analysis on Wednesday and Thursday injuries only. This regression preserves the exogenous variation in

[17] The positive and significant effect during period 1 could result from a Monday effect lingering for Tuesday injuries.

the RP, but also reduces the sample size by one-third. However, the results in Panel B are negative, and of a similar magnitude and significance to the baseline results.

Table 6: Proportional Hazard Model Estimates of the Effect of the Retroactive Payment on the Probability of Exit from WC, Tuesday and Wednesday vs Wednesday and Thursday

	(1)	(2)	(3)
	Tuesday and Wednesday		
	All	Above median	Below median
Weeks 1-2	0.074	0.119	0.125
	(0.054)	(0.057)	(0.057)
Weeks 3-4	0.086	0.056	0.045
	(0.056)	(0.061)	(0.062)
Weeks 5-6	0.106	0.047	0.028
	(0.060)	(0.069)	(0.073)
Weeks 7-8	0.170	0.074	0.046
	(0.066)	(0.081)	(0.087)
Observations		64,523	
	Wednesday and Thursday		
Weeks 1-2	-0.073	-0.057	-0.048
	(0.022)	(0.023)	(0.024)
Weeks 3-4	-0.059	-0.072	-0.073
	(0.025)	(0.029)	(0.031)
Weeks 5-6	0.043	-0.079	-0.090
	(0.041)	(0.039)	(0.044)
Weeks 7-8	0.052	0.047	0.052
	(0.051)	(0.052)	(0.058)
Observations		64,523	

Notes: Oregon Department of Consumer and Business Services, WC claims 1987-2012. Clustered standard errors in parenthesis. Regression repeats baseline specification, but only includes injuries occurring on a Tuesday or Wednesday in Panel (A), and Thursday or Friday in Panel (B). p<0.1 +, p<0.05 *, p<0.01 **

6 Implications for Optimal Benefits

The presence of liquidity effects establishes the fact that the injured worker population is sensitive to short-term income shocks. Additionally, it provides evidence that WC enhances injured workers' ability to smooth consumption. To understand the implications for social welfare, I turn to the optimal benefit formula derived in Baily (1978) and Chetty (2008). 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. In each period, the claimant exerts effort e_t into trying to return to work. Under the assumption that agents have maximized their private welfare, the optimal benefit level is determined by the first order condition on the social planner's problem:

$$\frac{dW_t}{db_t} = \frac{(1-\sigma_t)}{\sigma_t}\left(\underbrace{\frac{\partial e_t/\partial A_t}{\partial e_t/\partial w_t}}_{(a)} - \underbrace{\frac{\epsilon_{1-d_t,b}}{\sigma_t}}_{(b)}\right). \tag{4}$$

The optimal benefit level depends on (a) the ratio of liquidity ($\partial e_t/\partial A_t$) to moral hazard ($\partial e_t/\partial w_t$) effects; (b) the elasticity of the probability of not working with respect to benefits ($\epsilon_{1-d_t,b}$); and the frequency of the risky event (σ_t). The benefit level maximizes social welfare when equation (4) equals zero. If equation (4) yields a positive number when applying the estimated liquidity to moral hazard ratio, this indicates that increasing the benefit level will increase overall social welfare. If the equation yields a negative number, this indicates instead that *decreasing* the benefit level would increase social welfare.[18] Between the baseline estimates and the subsample of claimants with a high probability of sick leave, I estimate a liquidity elasticity of approximately 0.24-0.30, and moral hazard elasticities of 0.13 - 0.23. The sum of the two effects is the overall elasticity of the probability of not working with

[18]See the appendix for an extended discussion and derivation of the optimal benefits formula in Chetty (2008).

respect to benefits during the first four weeks.[19] By using 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 4 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.[20] I apply these estimates to the Chetty (2008) formula to determine the effects of WC on social welfare.

Table 7 shows the estimated welfare change based on a 1.6 percent incidence rate of workplace injury over the time frame of the analysis.[21] The metric of the welfare change 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 (3) of panel A indicates that increasing WC benefits by $1 would increase lower earner's utility by approximately 1 cent per week, or $0.50 per year. These approximations indicate small welfare gains to increasing benefits; however, they do imply that the optimal benefit level is higher than the current level for all workers. Furthermore, the welfare gains associated with increasing benefits are largest for the lower wage workers, suggesting that a more progressive benefit structure could be welfare-enhancing. For workers earning less than the median wage in Oregon, increasing benefits by $1 would increase utility by nearly 3 cents per week, or $1.50 per year.

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. There are several possible explanations for the larger welfare gains under UI. First, UI benefits have a lower replacement rate than WC benefits on average. The lower incidence of workplace injury relative to unemployment spells, and longer average duration of UI spells could also play a

[19] 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.

[20] 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.

[21] This represents the average incidence rate over the years in the data, as approximated by the Survey on Occupational Illness and Injury.

role. Furthermore, the marginal utility of income may be affected by an individual's health, potentially leading leading to a smaller liquidity effect for injured workers than unemployed workers.[22]

Table 7: Welfare Effects of WC Benefits

	(1) Liquidity Effect	(2) Moral Hazard Effect	(3) Welfare Change
	Panel A: Overall		
Baseline	-0.28	-0.22	0.012
	(0.04)	(0.06)	(0.01)
Sickleave	-0.24	-0.23	0.009
	(0.07)	(0.04)	(0.004)
	Panel B: Below median wage		
Baseline	-0.30	-0.14	0.028
	(0.09)	(0.04)	(0.02)
Sickleave	-0.29	-0.13	0.028
	(0.09)	(0.04)	(0.03)
	Panel C: Above median wage		
Baseline	-0.29	-0.19	0.017
	(0.08)	(0.04)	(0.01)
Sickleave Only	-0.25	-0.18	0.015
	(0.08)	(0.04)	(0.01)

Notes Bootstrapped standard errors reported in parenthesis. Oregon Department of Consumer and Business Services, 1987-2012, and Survey of Occupational Injuries and Illnesses. Column (1) contains the liquidity elasticities and column (2) contains the moral hazard elasticities, scaling the coefficients in table 2 by the equivalent percentage change in the bi-weekly wage induced by the RP. Column (3) applies these estimates to the formula in Chetty (2008) to obtain the money metric of the estimated welfare gains associated with a $1 change in benefits.

My empirical strategy allows me to identify liquidity effects for workers with short term injuries. Presumably, liquidity effects could be larger for workers whose injuries last months, or even years. Bronchetti (2012) analyzes consumption of injured workers in the Health and Retirement Study, where consumption is measured on a bi-annual basis. Under plausible

[22]Importantly, this approach holds even in the case of state-dependent utility; see the appendix and Chetty and Finkelstein (2013) for further discussion.

levels of risk aversion and assumptions about the utility function, Bronchetti applies these estimates to the Bailey-Chetty framework and obtains a range of possible optimal replacement rates for WC between 0.1 to 0.6. Her estimated range of optimal replacement rates suggests that the current benefit level is higher than the optimal replacement rate. However, these estimates are based on a sample of workers over age 50 who are at a lower risk of incurring a workplace injury than the average WC claimant in his or her mid-30s. Bronchetti also notes that older workers could be more likely to have accrued savings by the time of their injury, suggesting that the consumption smoothing benefit of WC could be lower for older individuals than for the overall population at risk for on-the-job injuries. Given that the current replacement rate in Oregon is 0.66, my estimates suggest that the optimal benefit level for younger workers is likely above the upper end of the range of replacement rates in Bronchetti (2012). Still, future research is needed to understand the effect of WC on younger injured workers with long term injures.

7 Conclusion

Relatively little is known about the optimal design of WC despite the program's large size, and the policy discussion about potential reforms to WC benefits. I examine how claimants adjust the duration of their WC claims in response to variation in a retroactive payment, allowing me to isolate the liquidity and moral hazard effects for short-term recipients of WC. I find that the liquidity effect accounts for 50 - 60 percent of the increase in claim duration. These results are especially pronounced for injuries that last longer than to weeks, meaning that claimants' decision to return to work could be influenced by small fluctuations in WC payments introduced by the RP.

Variation in period-of-injury paychecks may lead some workers to be more sensitive to the RP, potentially introducing a 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 period-of-injury paycheck, the estimates are very similar to the overall sample. Furthermore, a falsification exercise comparing claimants with similar RPs but different days of the week does not yield significant evidence of a liquidity effect. 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). The results suggest that increasing the benefit level would increase social welfare, particularly for lower wage workers.

These results add to a broader literature finding that individuals are sensitive even to small changes in income, evidence that workers could face liquidity constraints (Soueles, Parker and Johnson 2006), and that the timing of income receipt affects individuals' behavior (Stephens 2003, 2006). Risk-averse workers could reduce their consumption right after an on the job injury due to immediately binding liquidity constraints, consumption commitments that prevent them from optimally re-allocating their household expenditures, or to increase their precautionary savings (Chetty 2005; Chetty and Szeidl 2007; Carroll and Kimball 2008). The presence of significant liquidity effects implies that these WC claimants reduce their consumption even after short spells away from work. My estimates are based on the responses of claimants with short term injuries, but future research should examine the effects on longer term injuries, a group who could face even greater liquidity constraints. These findings provide evidence that liquidity constraints are an important consideration for the population of workers at risk of an on-the-job injury, and that timely changes in income can significantly affect injured workers' welfare.

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A FOR ONLINE PUBLICATION: APPENDIX

A.1 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 the 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. The function $v_t(c_t)$ represents utility in the working state, and $u_t(c_t)$ represents utility in the non-working state, thus allowing for state-dependent utility functions (Viscusi and Evans, 1990). As shown below, $\psi_t(e_t)$ is an additively-separable function reflecting the extent of effort the worker spends on controlling when they return to work.

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

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

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.[1] 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 $u(A_1 - s_1 + b_1) = u(c_1^n)$ is also concave, and $J_2(A_2, RP)$ represents the expected value of the claimant's decision in the next period:

$$J_2(A_2, RP) = \max_{e_2} e_2 V_2(A_2, RP) + (1 - e_2) U_2(A_2, RP) - \psi(e_2).$$

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:

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

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

The first order condition for each period t reflects the marginal effort level at which the claimant is indifferent between returning to work or receiving WC for another period:

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

[1] Future versions of this framework could 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.

Note that $E[V_{t+1}(A_{t+1}) - J_{t+1}(A_{t+1}, RP_{t+1})] = E[OV_t]$ represents the claimant's expected option value associated with deciding whether or not to work. The RP increases the expected option value of staying out of work during period 1.

With this framework, changes in the level of effort will affect the rate at which individuals return to work during period t. Empirically, I estimate changes in the hazard rate, or probability of return to work. By examining how changes in each of the parameters in $\Omega_t = \{b_t, w_t, A_t, RP\}$ affect the duration of claims, I examine how these parameters influence the claimants' effort 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:

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

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, Viscusi and Durbin 1995; Neuhauser and Raphael 2004). On the other hand, increasing the wage during any one period yields:

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

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:

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

In this case, the change in cash on hand affects utility in *both* the working and non-working state. The sign of $\frac{\partial e_t}{\partial A_t}$ depends on how A_t affects utility when individuals are working, relative to when they are not. In an ideal world where workers are able to attain their desired consumption level when out of work, then they will equate marginal utility of consumption between the two states of the world, such that $v'(c_t^e) = u'(c_t^n)$ and $\frac{\partial e_t}{\partial A_t} = 0$ (Chetty 2008). 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 $\frac{\partial e_t}{\partial A_t} < 0$, 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. Importantly, this approach will reflect the effect of an injury on an individual's utility, regardless of whether the optimal level of consumption is different between the two states of the world. Analyzing changes in return to work rates in response to a change in cash on hand thus identifies the gap (or lack of a gap) in marginal utilities in the two states of the world, even if utility is state-dependent (Chetty and Finkelstein, 2013).

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

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

Hence, an increase in benefits could decrease effort of return to work and lengthen claims through two distinct channels: by relaxing liquidity constraints and by reducing the opportunity cost of missing work. Importantly, while the first term captures the extent to which claimants value the additional income while out of work, the second term reflects the extent to which claimants respond to the change in incentive to work. The ratio of hazard rates with respect to a change in cash on hand and a change in the implied change in price of missing work, which are estimated in data, yields the ratio of $\frac{\partial e_t}{\partial A_t}$ and $\frac{\partial e_t}{\partial w_t}$, 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:

$$\frac{\partial e_1}{\partial RP} = \beta E\left[\frac{\partial OV_1}{\partial RP}\right] = \begin{cases} -\beta\frac{\partial V_2}{\partial RP} : V_2 > U_2 \\ -\beta\frac{\partial U_2}{\partial RP} : V_2 \leq U_2 \end{cases} < 0 \tag{5}$$

Since both $-\beta\frac{\partial V_2}{\partial RP} < 0$ and $-\beta\frac{\partial U_2}{\partial RP} < 0$, increasing the RP always lowers effort, and return to work rates, 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 e_2 as a change in A_2:[2]

$$\frac{\partial e_2}{\partial RP} = v'(c_2^e) - u'(c_2^e) \leq 0.$$

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 implies that workers primarily lengthen claims in response to income that offsets the gap in their consumption: the liquidity

[2]In practice, 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.

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.[3]

A.1.1 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. Under the assumption that agents have maximized their private welfare, the optimal benefit level is determined by the first order condition on the social planner's problem (Chetty 2008; Baily 1978):

$$\frac{dW_t}{db_t} = \frac{(1-\sigma_t)}{\sigma_t}\left(\underbrace{\frac{u'(c_t^n) - v'(c_t^e)}{v'(c_t^e)}}_{(1)} - \underbrace{\frac{\epsilon_{1-d_t,b}}{\sigma_t}}_{(2)}\right). \tag{6}$$

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. The benefit level maximizes social welfare when equation 6 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 6 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.

Importantly, the local change in benefits does not have a first-order effect on other inputs that are endogenous to policy changes if agents have already maximized their expected private utility. Under this key assumption, the ratio of liquidity to moral hazard elasticities estimated from changes in the current benefit level informs whether a local change in benefits would increase or decrease social welfare. If equation 6 yields a positive number when applying (1) the liquidity to moral hazard ratio estimated around current benefit levels, this indicates that increasing the benefit level will increase overall social welfare. 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. While the estimated elasticities are informative about marginal welfare effects, the results cannot be extrapolated beyond local policy changes because the optimal benefits formula relies on the assumption that private utility is at the optimum (Chetty 2008; Baily 1978).

Furthermore, 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

[3]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.

points in time during a claimant's absence from work. Hence, applying my estimated liquidity to moral hazard ratio to equation 6 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. 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 for the data used in estimating the liquidity and moral hazard effects.

Additionally, equation 6 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). As a result, the conclusions about optimal benefits in this case hold under the assumption that workers bear the full cost of WC premiums. Research on the incidence of WC premiums finds that the majority of costs are indeed fully passed through to the employee, suggesting that this is a reasonable assumption (Dorsey and Walzer 1983; Krueger and Gruber 1990; Fortin and Lanoie 2000).

A.1.2 Variation in Period-of-Injury Paycheck

The day of the week also creates variation in the size of the worker's period-of-injury paycheck: workers who would receive larger RPs also earn fewer days of wages during the week of their injury. The depleted cash on hand during the first two weeks could create a bias going in either direction. Consider a revised version of equation 3 to understand the implications of this fact:

$$\psi'(e_1) = v(A_1 - s_1 + w_1) - u(A_1(d) - s_1 + b_1) + \beta[V_2(A_2) - J_2(A_2, RP(d))].$$

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:

$$\frac{\partial e_1}{\partial d} = -u'(c_1^n)\frac{\partial A_1}{\partial d} + \beta\frac{\partial OV_1}{\partial d} = -u'(c_1^n)\frac{\partial A_1}{\partial d} - \begin{cases} \beta\frac{\partial V_2}{\partial d} : V_2 > U_2 \\ \beta\frac{\partial U_2}{\partial d} : V_2 \leq U_2 \end{cases} \tag{7}$$

The second term in equation 7 is the same as in equation 5, implying that increasing the RP decreases $\frac{\partial e_1}{\partial d}$. But the first term is positive and potentially increases $\frac{\partial e_1}{\partial d}$, since $\frac{\partial A_1}{\partial d}$ is negative. Ultimately, whether $\frac{\partial e_1}{\partial d}$ 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, 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 $\frac{\partial e_1}{\partial d}$, 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.

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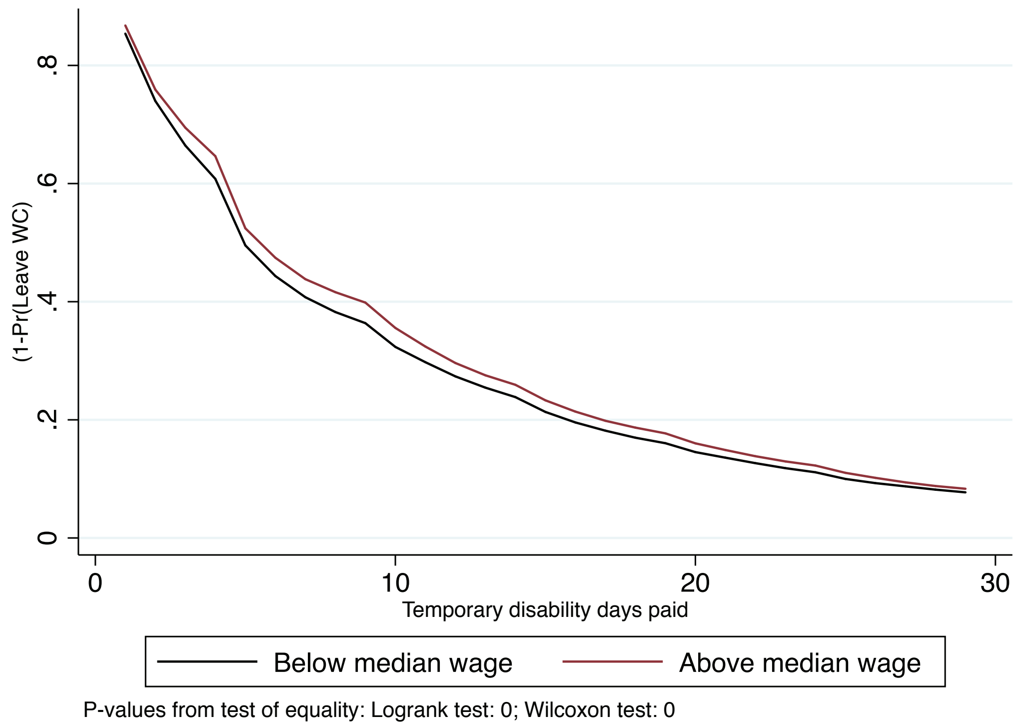

Figure A1: Survival curves for WC claims, by Pre-injury Wages

Notes Oregon Department of Consumer and Business Services, WC claims 1987-2012. Sample includes claims occurring on a Wednesday, Thursday or Friday that lasted at most one year.

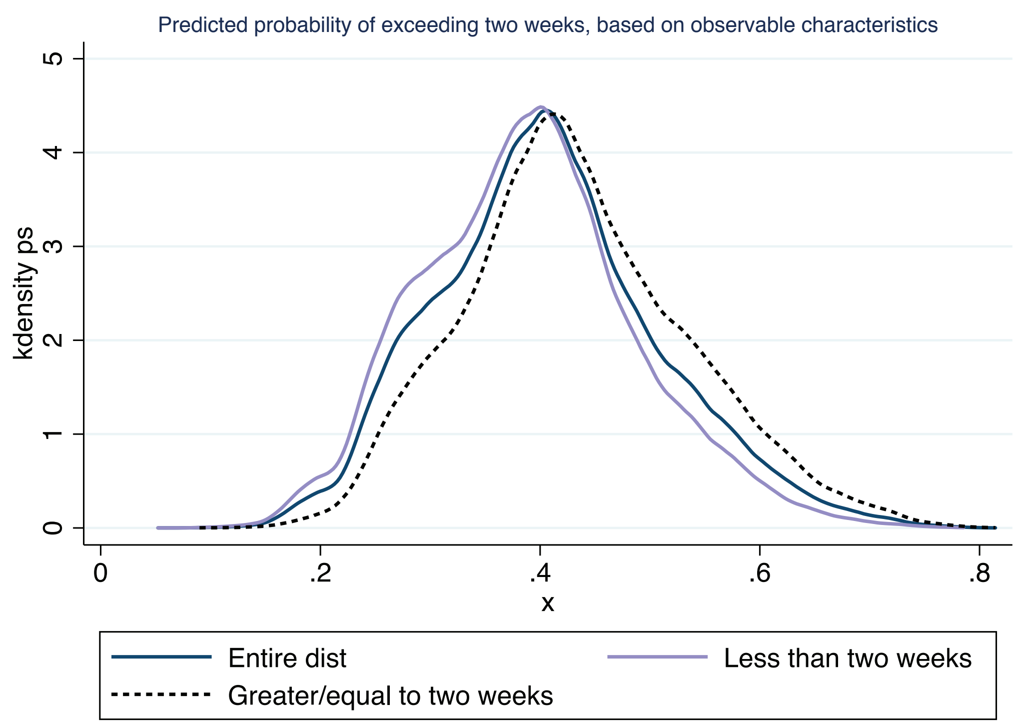

Figure A2: Distribution of Propensity Score of Claim Lasting Longer than Two Weeks

Notes Oregon Department of Consumer and Business Services, WC claims 1987-2012. Sample includes claims occurring on a Wednesday, Thursday or Friday that lasted at most one year.

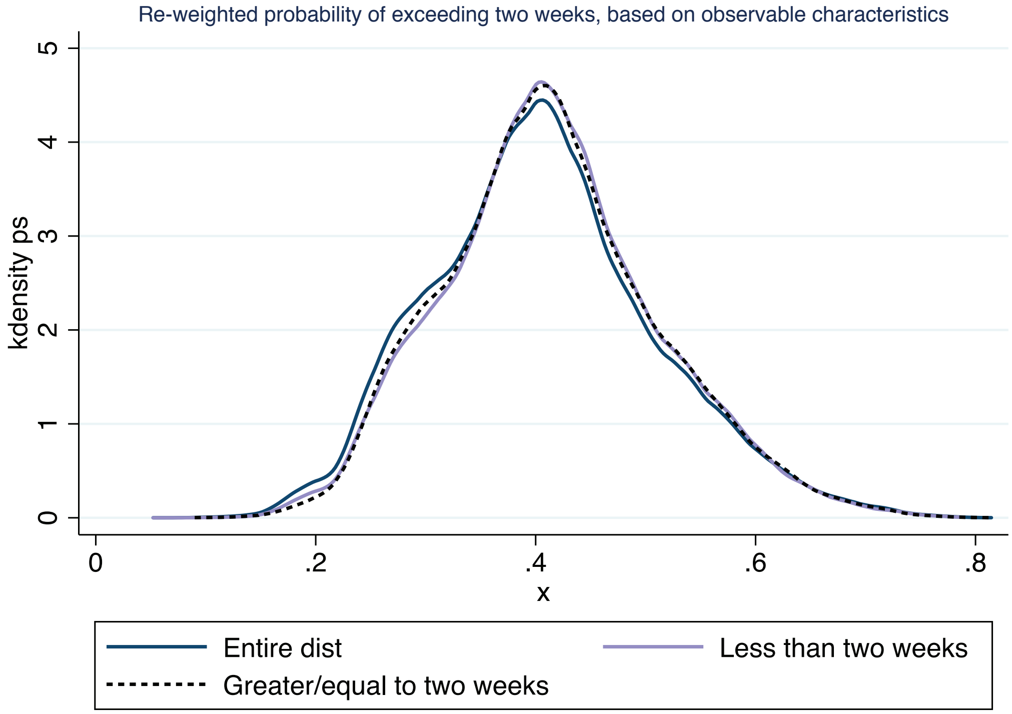

Figure A3: Distribution of Propensity Score of Claim Lasting Longer than Two Weeks, Reweighted

Notes Oregon Department of Consumer and Business Services, WC claims 1987-2012. Sample includes claims occurring on a Wednesday, Thursday or Friday that lasted at most one year.

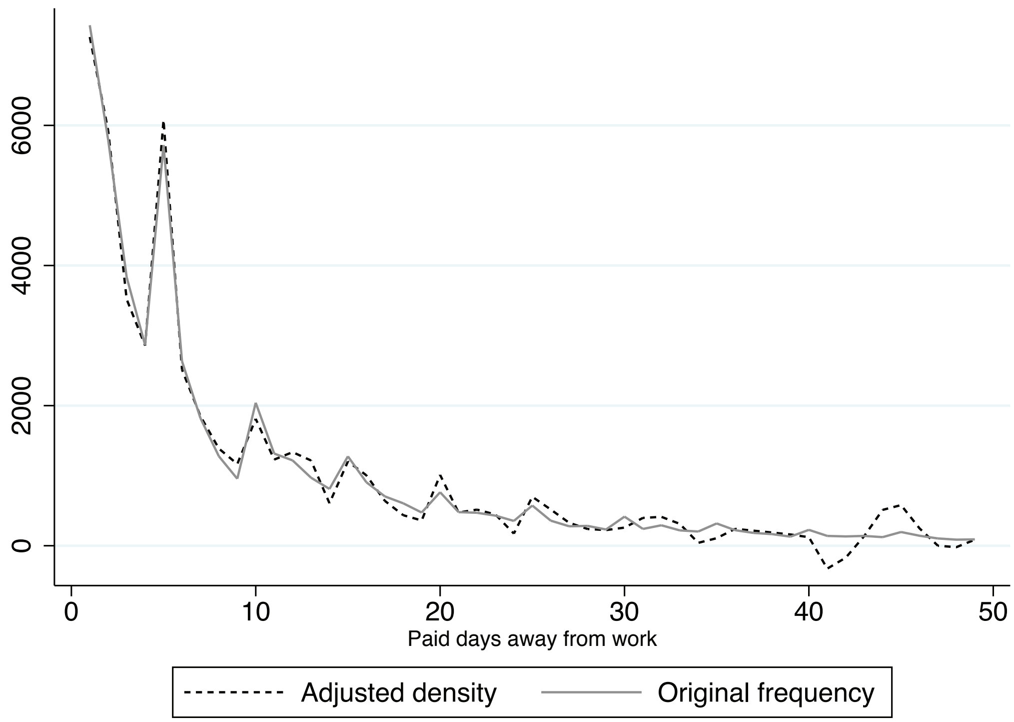

Figure A4: Actual vs counterfactual distribution of WC Claim Duration: Below Median Wage

Source Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Sample includes claims occurring on a Wednesday, Thursday or Friday that lasted at most one year. The x-axis represents 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. 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 represents 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.

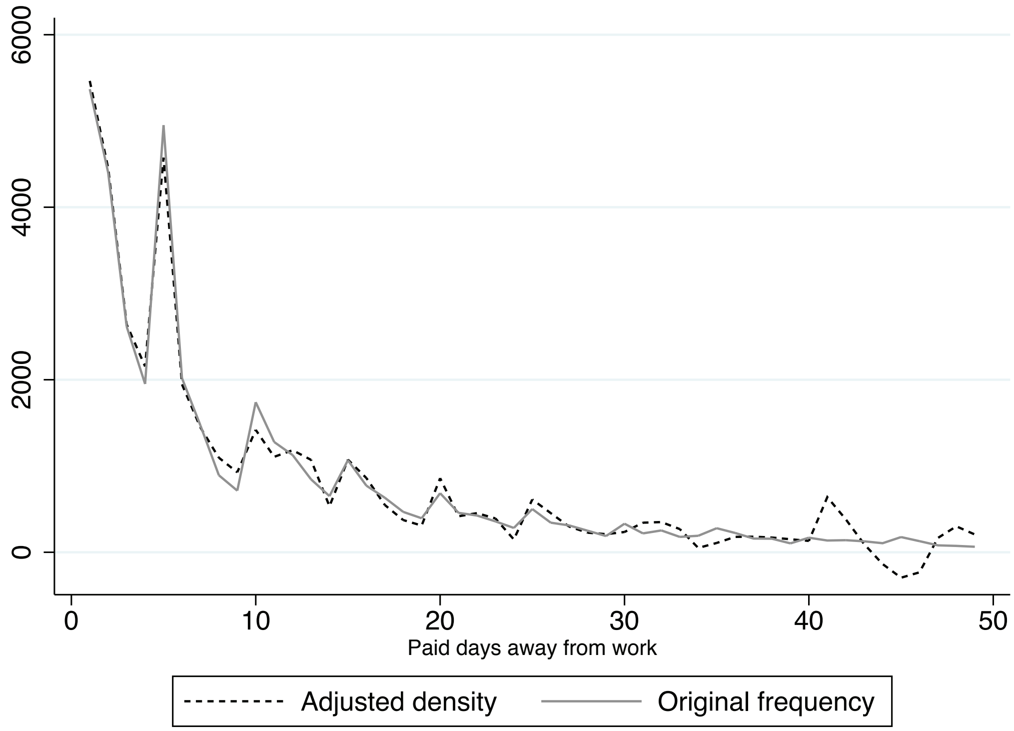

Figure A5: Actual vs counterfactual distribution of WC Claim Duration: Above Median Wage

Source Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Sample includes claims occurring on a Wednesday, Thursday or Friday that lasted at most one year. The x-axis represents 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. 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 represents 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.

Table A1: Probability of Claim Exit by Day of the Week, Claims Longer than Two Weeks

	(1) Mon	(2) Tue	(3) Wed	(4) Thu	(5) Fri
Wed Injury	-0.005	0.008	0.019	0.004	-0.007
	(0.005)	(0.005)	(0.004)	(0.004)	(0.004)
Thu Injury	0.001	0.002	0.006	0.008	-0.008
	(0.005)	(0.005)	(0.004)	(0.004)	(0.004)
Observations	38,674	38,674	38,674	38,674	38,674

Notes Notes: Data from Oregon Department of Consumer and Business Services, WC claims from 1987-2012. Sample includes claims occurring on a Wednesday, Thursday or Friday that lasted at most one year. Each column shows an indicator for the claim ending on the day of the week listed in the column header,regressed on indicators for the claim ending on a Wednesday or Thursday (Friday is the omitted category), as well as controls for worker age, gender, wage, injury type, industry, and claim year. Robust standard errors in parentheses. ** $p<0.01$, * $p<0.05$, + $p<0.1$

Table A2: Baseline Results without Propensity Score Reweight

	(1) All	(2) Above median wage	(3) Below median wage
Weeks 1-2	-0.058	-0.041	-0.031
	(0.009)	(0.009)	(0.010)
Weeks 3-4	-0.013	-0.030	-0.042
	(0.012)	(0.014)	(0.015)
Weeks 5-6	0.015	-0.021	-0.043
	(0.018)	(0.020)	(0.023)
Weeks 7-8	0.033	-0.011	-0.038
	(0.023)	(0.026)	(0.030)
Observations		92,735	

Notes Notes: Data from Oregon Department of Consumer and Business Services, WC claims 1987-2012. Clustered standard errors in parenthesis. Regression repeats baseline specification, now additionally interacting ln(RP)*γ_t with injury type. Panels (B) and (C) show the coefficients from one regression where the RP term is additionally interacted with indicators for earnings above/below the median wage. p<0.1 +, p<0.05 *, p<0.01 **

Table A3: Frequency of Sick Days 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:		0.484

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 A4: Observable Characteristics by Sick Day Prevalence

	Low sickday	High sickday	Pvalue
Male	0.74	0.69	0.00
Age	34.00	38.00	0.00
Weekly wage	658.00	777.00	0.00
Median wage	0.37	0.54	0.00
WC days paid	15.00	13.70	0.00
Claim > 10 Days	0.42	0.41	0.00
Retroactive payment	170.00	203.00	0.00
Daily benefit	86.00	102.00	0.00
Medical cost	2,212.00	2,185.00	0.40
Wed	0.33	0.33	0.64
Thu	0.33	0.34	0.00
Fri	0.34	0.33	0.00
Afternoon	0.50	0.50	0.29
Trauma	0.04	0.04	0.57
Fracture	0.11	0.09	0.00
Strain	0.56	0.62	0.00
Wound	0.26	0.20	0.00
Other	0.04	0.04	0.01
Agriculture	0.13	0.00	0.00
Construction	0.24	0.02	0.00
Manufacturing	0.00	0.36	0.00
Trade	0.27	0.09	0.00
Transportation	0.00	0.17	0.00
Other	0.37	0.36	0.04
Observations	43,030	53,664	

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