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## The effect of health insurance on workers' compensation filing: Evidence from the affordable care act's age-based threshold for dependent coverage

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# The Effect of Health Insurance on Workers' Compensation Filing: Evidence from the Affordable Care Act's Age-Based Threshold for Dependent Coverage

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# **The Effect of Health Insurance on Workers' Compensation Filing: Evidence from the Affordable Care Act's Age-Based Threshold for Dependent Coverage**

**Upjohn Institute Working Paper 15-232**

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July 2015

## **ABSTRACT**

This paper identifies the effect of health insurance on workers' compensation (WC) filing for young adults by implementing a regression discontinuity design using WC medical claims data from Texas. The results suggest health insurance factors into the decision to have WC pay for discretionary care. The implied instrumental variables estimates suggest a 10 percentage point decrease in health insurance coverage increases WC bills by 15.3 percent. Despite the large impact of health insurance on the number of WC bills, the additional cost to WC at age 26 appears to be small as most of the increase comes from small bills.

JEL Classification Codes: I13, J32, J38

Key Words: Workers' compensation, Moral hazard, Health insurance, Affordable Care Act

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Workers' compensation (WC) is a state-regulated insurance program that provides partial wage replacement and pays medical bills for employees injured on the job. Workers benefit from having WC pay for treatment because they have no copays or deductibles for care, but firms have an incentive to decrease WC claims because treating an injury is typically more expensive under WC than health insurance (Baker and Krueger 1995; Johnson, Baldwin, and Burton 1996; Leigh and Ward 1997) and because WC benefits can encourage workers to delay returning to work. Ex-post moral hazard arises with the filing decision because injured workers choose whether or not to file for WC (Butler and Worrall 1991). By making medical care apart from WC cheaper, health insurance has the potential to influence the worker's filing decision. Studying the effect of health insurance on WC claiming is difficult because health insurance is not randomly assigned to workers or to firms, which means examining the correlation between health insurance and WC filing is not informative about the effect of health insurance on WC filing.

In this paper, I study the health insurance cutoff created by the Affordable Care Act's (ACA) expansion of dependent coverage until the age of 26 to estimate the causal effect of health insurance on WC filing for young adults. Under the ACA, firms that offer dependent coverage have to allow employees' children to stay on their health insurance until the age of 26. Young adults on either side of their twenty-sixth birthdays are very similar except that people just shy of their twenty-sixth birthdays can be on their parents' health insurance while people who just turned 26 cannot. I use Texas WC administrative data and this abrupt decrease in private health insurance coverage to implement a regression discontinuity design by comparing people on either side of their twenty-sixth birthdays. As long as other factors that affect WC filing do not change abruptly at age 26, this strategy provides estimates of the causal effect of health insurance coverage on WC filing for young adults.

In principle, there should be no relationship between having health insurance and filing for WC. If an injury occurs at work, WC should pay for medical treatment. If an injury occurs outside of work, health insurance or the individual should pay for treatment. However, as Card and McCall (1996) explain, workers without health insurance have an incentive to claim their medical issues are work-related even if they are not so that WC will pay for care. If the injury occurs at work, health insurance may deter workers from filing for WC if they feel there is a cost to filing a WC claim. As Lakdawalla, Reville, and Seabury (2007) explain, filing a WC claim is costly if employers dissuade people from filing WC because they fear the claims will increase their premiums. Injured workers also might not want to deal with the paperwork from WC or fear that they will be called on to prove that their injury was caused by work. Some workers may feel there is a stigma associated with filing for WC. Similarly, providers may encourage workers to have health insurance pay for their care since billing WC may have additional administrative burdens (Leigh and Ward 1997). Because manipulating the payment source is more difficult for visible injuries and emergencies, health insurance would be most likely to affect medical care that can easily be delayed.

Biddle and Roberts (2003) study benefit claiming among eligible workers by surveying Michigan workers identified by physicians as likely having work-related injuries. They find that 70 percent of injured workers did not file for WC and that 36 percent of the nonreporting injured workers cited having health insurance as a reason they did not file. Lakdawalla, Reville, and Seabury (2007) use data from the 1979 National Longitudinal Survey of Youth and find that workers at firms that offer health insurance are 16 percentage points more likely to file for WC than workers at firms that do not offer health insurance. Lakdawalla, Reville, and Seabury hypothesize that large firms may be more likely to provide workers with information about WC

and to encourage them to use it. Heaton (2012) studies the impact of the Massachusetts health insurance reform on WC and finds health care reform results in WC paying for fewer emergency room medical bills. Card and McCall (1996) and Campolieti and Hyatt (2006) both study how health insurance factors into the increased number of Monday claims for easy-to-conceal injuries. They both find no difference in Monday claiming by health insurance status, which suggests the Monday effect is not driven by people claiming nonwork injuries from the weekend are work related so that WC would pay the medical costs.

In this paper, I use the National Health Interview Survey (NHIS) and the Survey of Income and Program Participation (SIPP) to study how health insurance and employment change at age 26. The estimates for the fall in insurance coverage for working adults range from 4.4 to 5.3 percentage points. I find no evidence of abrupt employment increases at age 26. To study how health insurance affects WC filing, I use Texas WC administrative data that contain information on all medical bills associated with each WC claim. I find that total WC medical bills increase immediately after young adults turn 26, as do claims for injuries that plausibly do not require immediate medical care. The mean number of bills per claim trends smoothly at age 26, which suggests that claims are not becoming more severe at age 26. Instead, the increase in bills appears to come from claims that began before workers turned 26. These results suggest discretionary care is most responsive to health insurance. The implied instrumental variables (IV) estimates suggest that a 10 percentage point decrease in health insurance coverage results in an increase in WC bills of 15.3 percent. Despite the large impact of health insurance on WC bills, the additional costs to WC at age 26 are small because the increase in total bills comes from small bills. These results imply that the expansion of health insurance from the ACA will

shift many medical bills to being paid for by health insurance instead of WC but that the size of the cost shifting will be small.

The rest of the paper proceeds as follows. The next sections discuss background on WC and dependent coverage and consider health insurance and labor force changes occurring at age 26. The paper then describes the Texas WC data and considers how claims, total care received, the intensity of claims, and the cost of care change at age 26. The final section provides a discussion of the results and concludes.

## **BACKGROUND**

### **Workers' Compensation**

WC insurance pays medical bills and replaces lost wages for employees who are injured during the course of employment. In exchange for being able to receive WC benefits, workers cannot sue their employers for negligence. Thus, WC insurance also protects firms from lawsuits by injured employees. WC insurance is regulated at the state level, and benefits to injured workers are set by the state. Although some states restrict injured workers' choice of physician, a majority of states—including Texas—allow injured workers to choose their own medical providers. While WC pays for medical care immediately after an injury occurs, injured workers become eligible for income replacement benefits after missing three to seven days of work, dependent on the state. In Texas, injured workers become eligible for income replacement benefits after missing at least seven days of work. The injured workers' weekly cash benefits are

a function of their weekly earnings subject to a maximum that varies across states.<sup>1</sup> For thorough overviews of WC, refer to Sengupta, Baldwin, and Reno (2013) and Utterback, Meyers, and Wurzelbacher (2014).

In their summary of WC spending, Sengupta, Baldwin, and Reno (2013) note that \$30 billion was paid to cover injured workers' medical expenses through WC in 2011. They also note that the majority of WC cases—around 76 percent—are medical-only cases and do not involve payments for missed work. Although \$30 billion is small compared to 2011's total medical spending of \$2.7 trillion, the amount makes up a large percentage of WC costs. According to Sengupta et al., the share of medical benefits as a percentage of the total amount paid to workers has risen from around 30 percent in the early 1990s to approximately 50 percent in 2011. As many people with work-related injuries do not file for WC, the medical cost of work injuries is higher than total WC medical costs. Leigh (2011) estimates that the true medical cost of work-related injuries in 2007 was \$60 billion and that the total cost of work-related injuries in 2007 was \$250 billion, which was more than the cost of cancer, coronary heart disease, and stroke.

Texas differs from all states other than Oklahoma in that Texas employers are not required to purchase WC insurance. Despite this, 81 percent of Texas workers work for firms with WC insurance as of 2012 (Texas Department of Insurance [TDI] 2012). Relatively few studies have examined firms' decisions to opt out of WC in Texas (Morantz 2010). Two exceptions are Butler (1996) and Morantz (2010). Butler studies differences in injury rates between subscribing and nonsubscribing firms and finds that both types of firms have similar fatality rates. He finds that nonsubscribing firms have slightly higher nonfatal injury rates, likely

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<sup>1</sup> Previous research has studied how claiming behavior responds to a variety of incentives, such as benefit levels and antifraud measures. For examples, refer to Boden and Ruser (2003); Bronchetti and McInerney (2012); Guo and Burton (2010); Hansen (2014); Meyer, Viscusi, and Durbin (1995); and Neuhauser and Raphael (2004).



because nonsubscribers tend to offer occupational injury plans that provide first-day wage-replacement benefits. Butler concludes that safety levels are likely similar between subscribing and nonsubscribing firms. Morantz surveys large firms who opt out of WC. She finds that most opt out to try to save money, and that nearly all nonsubscribing firms with more than 500 employees offer health insurance to employees, which is true of large firms in general (Kaiser Family Foundation 2013). In the last section of the paper, I discuss how the results might generalize to states with compulsory WC.

Other than not being compulsory, Texas WC is generally similar to other states' WC programs along most dimensions (Morantz 2010). Employers in Texas may self-insure or purchase insurance from an insurance company. The WC insurer pays medical providers a fixed amount for services performed and must report all WC medical spending to TDI, which compiles the information into the data set used in this paper.

WC claims increase costs for employers in several ways. First, a WC claim may include lost wages, while an injury paid for by health insurance will not. Second, medical treatment costs for the same injury are higher under WC than traditional health care (Baker and Krueger 1995; Johnson, Baldwin, and Burton 1996; Leigh and Ward 1997). Third, experience rating is more sensitive to large claims under WC than under group health insurance (Lakdawalla, Reville, and Seabury 2007). Fourth, with health insurance, the injured person is responsible for copays, deductibles, and coinsurance, while WC pays for all of the medical care for a person injured at work. Finally, as is the case with dependent coverage, health insurance may come from outside the employer's plan and thus may not increase the firm's costs at all. For these reasons, employers may dissuade employees from filing for WC.

## **Affordable Care Act and Dependent Coverage for Young Adults**

As of September 23, 2010, the ACA mandated that employers had to allow young adults to stay on their parents' health insurance until the age of 26 beginning with the next renewal date of their plans. Research has found that extending dependent coverage was successful in raising coverage for young adults (Antwi, Moriya, and Simon 2013; O'Hara and Brault 2013). Overall, for people aged 19–25, the likelihood of having employer-sponsored health insurance as a dependent rose by 7 percentage points, while the likelihood of having any health insurance rose by about 3 percentage points. Sommers et al. (2013) find that less-healthy young adults were more likely to sign up for dependent coverage, which has led to increases in the usage of health services for young adults.

Many states, including Texas, extended dependent coverage in the decade before the ACA went into effect. As of 2005, non-self-insured employers in Texas were required to provide coverage to employees' children until the age of 26. Research has found that the effects of states extending dependent coverage on overall health insurance were small compared to the effects of the ACA, that many of the people taking advantage of extended dependent coverage were students, and that it took several years before these laws had effects on coverage (Dillender 2014; Levine, McKnight, and Heep 2011; Monheit et al. 2011). The ACA in contrast had immediate and large effects. The differences are likely because the ACA was well-publicized and people knew they could get back on their parents' health insurance even after they had left. Also, the ACA requires self-insured employers to comply and extends coverage to married young adults, which the state-level laws do not. Antwi, Moriya, and Simon (2013) find no evidence of differential effects of the ACA's dependent coverage mandate between states with and without prior dependent coverage. As most states passed their dependent coverage laws in

the few years before the ACA, finding no differential effect of the ACA based on prior dependent coverage laws is consistent with no early effects of state-level dependent coverage.

The main results use WC data from 2011 to 2013. Even though Texas had a prior dependent coverage law, the estimates are still valid since they are from a regression discontinuity design and do not identify the effect of dependent coverage using variation over time. As a placebo test, I also compute estimates using data from 2005 to 2007. As state-level dependent coverage laws generally rose by 7 percentage points, while the likelihood of having any health insurance rose by about 3 percentage points. Sommers et al. (2013) find that less-healthy young adults were more likely to sign up for dependent coverage, which has led to increases in the usage of health services for young adults.

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## **CHANGES IN HEALTH INSURANCE COVERAGE AND EMPLOYMENT AT AGE 26**

Although research has found that the ACA's dependent coverage provision has increased health insurance coverage for young adults, less is known about the transition from dependent coverage at the age of 26. I begin by examining what happens to health insurance coverage after people are no longer eligible for insurance through their parents' employers. Since the WC administrative data do not have health insurance status, I use data from the 2011 and 2012 NHIS and the 2008 SIPP. The advantage of the NHIS data is that they allow for implementing the same basic strategy as I implement with the WC administrative data. An issue with the NHIS data is that they do not contain state-identifiers. A possible concern is that the discontinuity would be different in Texas because the state had a prior dependent coverage law or because insurance coverage is lower there than the rest of the nation. To examine Texas specifically, I

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<sup>2</sup> Prior to the ACA, there was no set age at which young adults would be removed from their parents' health insurance plans in states that had not formally defined dependent coverage. Many private health insurance contracts covered young adults until the age of 23 if they were in college and until age 19 otherwise. Anderson et al. (2012, 2014) study these discontinuities and find that young adults losing health insurance reduces their usage of health care services.

estimate fixed effects models using the 2008 SIPP to estimate the effect of turning 26 on health insurance coverage and labor force participation.

### **Regression Discontinuity Strategy with NHIS Data**

The NHIS is a survey of a stratified random sample of the U.S. population that asks respondents detailed questions about health, health insurance, and health care access. I focus on insurance through an employer, insurance through the individual market, and overall coverage. I use the individual's birth month and the month of the interview to create the individual's age in months at the time of the interview, which allows for implementing a regression discontinuity design.

In addition to asking about health insurance coverage, the NHIS asks various questions about employment. The need for health insurance has been shown to induce people to enter the labor force.<sup>3</sup> Using a difference-in-differences strategy, Antwi, Moriya, and Simon (2013) find the ACA's dependent provision has decreased the percentage of eligible individuals in the labor force. Young adults beginning work immediately upon their twenty-sixth birthdays would generate an increase in WC claims even if health insurance has no effect on filing. I use the employment variables to test for various labor force discontinuities at age 26 to evaluate concerns about immediate labor force participation changes at 26 hindering the WC identification strategy.

I restrict the sample to include people aged 24–27. Means of key variables are shown in Table 1. About two-thirds of the sample has health insurance. Of those with health insurance,

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<sup>3</sup> For examples, refer to Buchmueller and Carpenter (2012); Dague, DeLeire, and Leninger (2014); Dillender (2015); and Garthwaite, Gross, and Notowidigdo (2014). While research generally finds access to health insurance influences labor force participation, Baicker et al. (2013) find no evidence that the Oregon Medicaid lottery influenced labor market outcomes.

75 percent have it through an employer. Figure 1 shows the age profiles for various outcomes of interest. The age profiles of the health insurance variables are relatively flat until age 26, at which point individuals experience a decrease in employer-sponsored health insurance and an increase in the likelihood of being uninsured. None of the age profiles for the employment variables reveal a trend break at age 26. To quantify the discontinuities and to test their statistical significance, I estimate the following equation:

$$(1) \quad y_i = \gamma_t + X_i\alpha + f(m_i) + Post26_i\beta + \eta_i,$$

where  $i$  indexes the individual;  $t$  indexes the year;  $y$  represents the various outcomes of interest;  $X$  is a set of individual covariates that includes sex, race, education, and marital status;  $\gamma$  is an indicator variable for the year 2011;  $f(m)$  is a smooth function representing the age profile for the dependent variable with respect to age in months  $m$ ;  $Post26$  is an indicator variable equal to 1 if the individual is at least 26; and  $\eta$  is an unobserved error component. I estimate Equation (1) by modeling  $f(m)$  as a quadratic on either side of the threshold. The coefficient on  $Post26$  can be interpreted as the percentage-point change in the outcome variable at 26.

The estimates of the percentage-point discontinuities at age 26 from Equation (1) with health insurance–dependent variables are shown in Panel A of Table 2. Standard errors are clustered by age and shown in parentheses below the estimates. At the age of 26, young adults experience a decrease in having employer-sponsored health insurance of 8.6 percentage points, which represents a nearly 15 percent reduction in employer-sponsored coverage. Privately purchased health insurance may increase, but the coefficient is not statistically distinguishable from zero. Young adults experience an increase in the likelihood of being uninsured of 4.2 percentage points or 14 percent.

Panel A of Table 2 also displays estimates of the above-age-26 coefficients from Equation (1) with employment dependent variables. Although I cannot rule out large employment discontinuities, all of the estimates are statistically indistinguishable from zero for both the intensive and extensive margins. The smoothness at age 26 is likely due in part to the fact that finding a job is not an instant process. Ilg and Theodossiou (2012) find that the average job search for an unemployed worker lasts over two months and that more than a quarter of unemployed workers search for more than six months. Young adults decreasing labor force participation to attend school or to search longer for a better job would not create a sharp discontinuity at age 26. Thus, even though dependent coverage affects labor force participation, the ACA's dependent coverage provision does not appear to create a discontinuous increase in the likelihood of working.<sup>4</sup>

Panel B of Table 2 displays estimates of the above-age-26 coefficients with health insurance variables as the dependent variables with the sample restricted to working young adults. The estimates are generally similar to the estimates for the full sample.

### **Fixed Effects Strategy with SIPP Data**

The SIPP is a longitudinal survey of the civilian noninstitutionalized population. In addition to containing state-level identifiers, the SIPP also follows individuals over time, which allows for estimating fixed effect models. Another advantage of the SIPP is that it asks about the source of employer-sponsored coverage, meaning I know if people have insurance through their own employers or as dependents. I use data from January 2011 to August 2013 and broaden the

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<sup>4</sup> Labor force variables trending smoothly around health insurance cutoffs is consistent with previous research. For example, Card, Dobkin, and Maestas (2008, 2009) find the age profile for employment is smooth at age 65. This is in spite of the fact that research typically finds the availability of health insurance influences retirement decisions (Gruber and Madrian 2002). Anderson, Dobkin, and Gross (2012) provide an example of employment trending smoothly at a health insurance cutoff for young adults.

age window to include young adults aged 23–28. Unfortunately, even with the broadened age ranges, the sample only contains 820 Texans, only 159 of whom turn 26 after 2010. The small sample size makes the estimates for Texas noisy.<sup>5</sup>

Descriptive statistics for young adults in Texas and the rest of the nation are shown in Table 3. Compared to the rest of the U.S., Texas has a higher uninsured rate and a lower percentage of people with college degrees for these age groups. With the SIPP data, I estimate the following equation:

$$(2) \quad y_{it} = \gamma_t + X_{it}\alpha + Post26_{it}\beta + v_i + u_{it},$$

where  $v$  is an individual fixed effect,  $\gamma$  is a full vector of indicator variables for the month and year of the interview,  $X$  now includes an indicator for being married and a linear control for age, and  $u$  is the unobserved error component. The coefficient on  $Post26$  is now identified by how outcomes change for individuals after they turn 26.

Panel A of Table 4 displays estimates of the changes in health insurance coverage at age 26 from Equation (2). The estimate for the loss in employer-sponsored health insurance as a dependent at age 26 in Texas is  $-0.082$ , while the coefficient for the rest of the U.S. is  $-0.046$ . The decrease in insurance as a dependent is not offset by an increase in insurance through one's own employer. The estimates for the decrease in employer-sponsored coverage in Texas of  $-0.075$  is similar to the NHIS estimate of  $-0.086$ , while the estimate for the rest of the nation is smaller at  $-0.043$ . The estimate for the increase in being uninsured at 26 is  $0.031$  in Texas. Although it is insignificant, I cannot reject that it or the national estimate of  $0.050$  is different than the NHIS estimate of  $0.042$ .

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<sup>5</sup> The SIPP interviews people every four months and asks about each month since the previous interview. Chetty (2008) finds that people report the same responses during one interview for all four months associated with the interview period. Therefore, I follow Kolstad and Kowalski (2012) and restrict the data to the interview month.



Panel A of Table 4 also displays estimates of the change in employment at age 26. For Texas, all of the estimates of changes in employment at age 26 are statistically insignificant. For the rest of the United States, the estimates suggest that people may be working less at the intensive margin. I find no evidence that suggests that labor force participation increases dramatically at age 26. Panel B of Table 4 displays estimates that restrict the sample to working people. When I condition the sample on being employed, the point estimate of the change in coverage rises slightly for Texas, while the point estimate for the rest of the United States falls slightly. While the estimates for Texas are noisy, I find no evidence to suggest that the changes in health insurance at 26 are dramatically different in Texas than in the rest of the United States.

Because the SIPP only contains age in years, there are fewer ages in the sample, which makes controlling for the age profile more challenging than with the NHIS. In Appendix A, I consider the robustness of the SIPP results to different controls for age. The health insurance estimates are very similar when I control for a quadratic polynomial in age and when I do not control for age. The labor force participation estimates are slightly more sensitive, but I still find no evidence to suggest that labor force participation increases immediately at age 26 regardless of specification.

## **CHANGES IN WC CLAIMING BEHAVIOR AT AGE 26**

### **Data and Empirical Strategy**

To examine the effect of losing access to insurance through a parent's employer on WC claiming, I use WC administrative data from the Texas Department of Insurance's Division of Workers' Compensation. The data contain the month each claim began, the cost of each bill, the ICD-9 code for each bill, the birth month of each claimant, and a unique identifier for each

claim. Since the data contain information about the underlying claims as well as for all medical bills, they allow for distinguishing between claims and bills. Each injury or illness has one claim associated with it, while each claim generally consists of multiple bills. I create one data set with claims as the unit of observation and another data set with bills as the unit of observation.

Around 60 percent of bills contain the exact date of service. Since knowing age in months is crucial, I only use these bills in the bill-level analysis.<sup>6</sup> I compute the individual's age in months at the time the claim began as well as at the time of each medical treatment when it is known.

For the main analysis, I use data on WC claims that occurred from 2011 to 2013.<sup>7</sup>

One advantage of these data is that they are likely highly accurate because they come directly from WC insurers. Another advantage is that TDI does not truncate or censor any of the bills. Although less than 5 percent of claims are disputed (TDI 2010), the data are updated each month and reflect the resolution of any disputed claims.

The relationship between having health insurance and filing for WC likely differs by injury type. Someone who sustains a cut on his face away from work will likely have a more difficult time filing a WC claim for treatment than someone with a sprained wrist. Someone with carpal tunnel syndrome may not even know whether WC or health insurance should pay for treatment. In addition to showing results for all claims and bills, I also use ICD-9 codes to classify injuries into three broad categories: strains and sprains, injuries that are neither strains nor sprains (nonsprains hereafter), and occupational diseases.<sup>8</sup> As occupational diseases, strains,

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<sup>6</sup> Alternatively, I could use the date the bill was received, which is nonmissing for all bills, to construct an individual's age at the time of the bill. The results are qualitatively very similar whenever I use this date, albeit smoother around the cutoff, which is not surprising since the date the bill was received is a somewhat noisy signal for when the services were received.

<sup>7</sup> The data were obtained from TDI in December 2013, so the information is incomplete for the last few months of the data.

<sup>8</sup> Examples of nonsprain injuries include burns, lacerations, and crushing wounds. The occupational disease category includes all diseases paid for by WC. Examples include carpal tunnel syndrome and herniated disc

and sprains generally require less immediate care, they would likely be more responsive to health insurance than nonsprain injuries.

Public health research finds that health care is more difficult to access on weekends, that elective care comprises a much smaller share of hospital admissions on weekends, and that a larger share of hospital admissions come through the emergency department on weekends (Bell and Redelmeier 2001; Kostis et al. 2007; Ryan, Levit, and Davis 2010). For these reasons, injuries and conditions that are more often treated on weekends are plausibly more likely to be emergencies with treatment that is more difficult to delay than those that are treated during the week. As such, I compute the likelihood that each ICD-9 code is first treated on the weekend and then separate claims into quartiles based on this measure.<sup>9</sup> Injuries least likely to happen on the weekend increasing more at age 26 would suggest discretionary care is more responsive to health insurance.

This strategy is similar in spirit to Dobkin (2003) and Card, Dobkin, and Maestas (2009) who study emergency room admissions in California. Both studies classify ICD-9 codes by the likelihood that admissions occur on weekends and consider ICD-9 codes that have similar admission rates on weekends and weekdays as emergencies. I classify ICD-9 codes into quartiles by weekend likelihood because more work is done during the week, which results in no work-related injury ever being as likely to happen on weekends as during the week. ICD-9 codes in the top quartile of the weekend likelihood distribution are the closest to being evenly distributed on weekends and weekdays. Interpreting weekend likelihood as a measure of urgency has two

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problems. This classification is a more aggregated version of the classifications used in Campolieti and Hyatt (2006) and Card and McCall (1996).

<sup>9</sup> I only use claims where the exact date of the first treatment is known to classify injuries. The likelihood that the exact date of treatment is missing does not vary dramatically for different kinds of injuries. Occupational diseases, sprains and strains, and nonsprain claims are missing the exact start date 47, 49, and 45 percent of the time.

important caveats. First, people in jobs that are more likely to work on weekends may be prone to certain types of injuries. Second, workers injured during the week might delay care until the weekend so they do not have to miss work. If either of these situations is the case, breaking the results down in this way would not yield the expected result that injuries least likely to be reported on weekends are most responsive to health insurance.

Despite these possibilities, the resulting classification is generally consistent with treatment for injuries happening on the weekend being more difficult to delay. Strains and sprains, occupational diseases, and nonsprain injuries are first treated on the weekend 5.5, 5.6, and 9.9 percent of the time, respectively. Over 90 percent of claims in the lower half of the weekend distribution are occupational diseases, sprains, or strains, while over 80 percent of claims in the top half are nonsprain injuries. These numbers provide support for interpreting sprains, strains, and occupational diseases as requiring less immediate care than other injuries. However, within the different injury classifications, there is still variability in the likelihood that the ICD-9 code first receives treatment on the weekend. For instance, 14.1 percent of people with skull fractures first receive treatment on the weekend, while people with blisters or dislocated wrists are first treated on weekends less than 5 percent of the time. All of these claims fall in the nonsprain category. People receiving treatment for epilepsy are first treated on the weekend 15.2 percent of the time, whereas carpal tunnel syndrome is first treated on the weekend around 1 percent of the time. All of these claims are occupational diseases.<sup>10</sup> Table 5 displays the top five injuries for each quartile and how likely they are to happen on weekends.

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<sup>10</sup> Within the sprains and strains category, there is less variation. Ankle injuries are most likely to receive care on weekends at 7.1 percent of the time, while shoulder sprains are the least likely to receive care on weekends at 4.9 percent of the time.

To implement the regression discontinuity design, I restrict the sample to include people within two years of their twenty-sixth birthdays. Descriptive statistics for people on either side of the threshold are shown in Table 6. Males account for a majority of injuries, likely because they work in jobs that require more manual labor. The means for charges are much larger than the medians, as the distribution of charges is right-skewed.

Unlike with the NHIS, with the WC administrative data, I only observe individuals who claim WC. In this way, my sample differs from Biddle and Roberts (2003) and Lakdawalla, Reville, and Seabury (2007), who analyze WC filing decisions conditional on having been injured at work. Because I do not observe individuals if they do not file for WC, I use the logged counts of WC claims and bills as the dependent variables. Assuming the underlying populations at risk for WC claims trend smoothly with age, the estimates of the discontinuities can be interpreted as the percent changes at the age of 26.<sup>11</sup>

Although I found no evidence of abrupt labor force participation changes at age 26 in Section 2, I cannot rule out large effects. As with Anderson et al. (2012) and Card et al. (2008; 2009), I maintain the assumption of no abrupt labor force participation changes at the health insurance cutoff for the main results. Although a common assumption, it is a very important one. An abrupt increase in labor force participation would cause WC claims to increase at age 26 even if health insurance had no effect on WC claiming behavior. To the extent possible, I later relax this assumption by controlling for estimates of the number of people working in Texas from the American Community Survey (ACS), which is the largest data set that allows for estimating the size of the labor force in Texas for the time period studied.

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<sup>11</sup> This follows Card, Dobkin, and Maestas (2008), who study how hospital admissions respond to Medicare eligibility. Unfortunately, using the logged counts precludes controlling for individual covariates in the analysis. Although the WC data are rich in many ways, they do not have much demographic information on claimants that would allow for testing for heterogeneous responses.

To estimate the causal effect of losing access to health insurance on WC behavior, I estimate the following equation:

$$(3) \quad y_m = f(m) + Post26_i\beta + \eta_m,$$

where  $y$  is the log of the number of WC claims or medical bills occurring at age  $m$  and the other variables are defined as in Equation (1). As with Equation (1), I estimate Equation (3) by modeling  $f(m)$  as a quadratic on either side of the cutoff. In Appendix B, I verify that the results are robust to estimating local linear models and to using different bandwidths. In Appendix C, I discuss repeating the claims and bills analysis at different ages as placebo tests.

## **WC Results**

### **WC Claims and Bills**

If losing health insurance affects the extensive margin of claiming behavior, WC claims should increase immediately after people turn 26. The top four graphs of Figure 2 show the overall number of claims and the number of claims for different types of injuries for each age. The overall number of claims appears to trend smoothly at age 26, though there may be a slight increase. Although the age profiles generally become noisier when broken down by injury type, it appears as though there may be an increase in strain and sprain claims at age 26, while claims for other injuries and for occupational diseases trend smoothly. The first column in Panel A of Table 7 reports the above-age-26 coefficients from estimating Equation (3) for all claims and for claims for different injury types using 2011 to 2013 data. All estimates are statistically indistinguishable from zero.

Although total claims do not increase dramatically at age 26, the amount of medical treatment paid for by WC could still respond to health insurance if young adults claim treatment is for a prior work-related injury even when it is not once they lose health insurance through their

parents' employers at age 26. Alternatively, prior to turning 26 young adults could have health insurance pay for some treatment of work-related injuries if they worry about using too many medical services through WC or if providers encourage them to let health insurance pay for care.

The bottom four graphs of Figure 2 display the overall number of bills and the number of bills for different types of injuries at each age. Total bills increase immediately at age 26, as do sprain and strain bills and occupational disease bills. Nonsprain bills trend smoothly at age 26. Column 1 in Panel B of Table 7 displays the above-age-26 coefficients from estimating Equation (3) for bills. The number of bills paid for by WC increases by 8.1 percent immediately after people turn 26. This appears to come from increases in strain and sprain bills and in the number of occupational disease bills, which rise by 11.4 percent and 14.2 percent, respectively. The estimate of the above-age-26 coefficient for bills for nonsprain injuries is statistically indistinguishable from zero.

The second column of Table 7 lists the above-age-26 coefficients from 2005 to 2007 data. As research generally finds no effect of state-level dependent coverage laws during their first few years, these estimates serve as placebo tests and allow for examining the possible identification threat of preexisting discontinuities at age 26.<sup>12</sup> The age profiles of these outcomes are shown in Appendix D. The above-age-26 coefficients for these years are all statistically indistinguishable from zero for bills and claims. The third column of each panel reports the difference between the two estimates.<sup>13</sup> The difference between the two estimates is statistically significant in every instance where the 2011-2013 estimates were significant.

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<sup>12</sup> When I estimate Equation (2) using SIPP data from 2005 to 2007, the standard errors are large but the point estimate for changes in health insurance coverage occurring immediately at age 26 is close to zero and statistically insignificant for Texas, which is consistent with previous research on these laws.

<sup>13</sup> To compute standard errors for the difference between the two estimates, I combine the log counts for 2005–2007 and 2011–2013 and estimate Equation (3) supplemented with interactions of each term with an indicator

In column 4 of Table 7, I supplement Equation (3) with controls for employment. To do this, I set the dependent variable to be the log of counts of claims or bills at each age in each year of the WC data, meaning the regressions have 144 observations instead of 48. I use the 2011–2013 ACS to estimate the number of Texans working at ages 24, 25, 26, and 27 in each year of the data and divide these estimates by 12 to obtain employment estimates for each age in months, which assumes that employment is evenly distributed across all ages in months.<sup>14</sup> I also supplement Equation (3) with year indicator variables. If employment increases at age 26 relative to age 25, the above-age-26 coefficients will fall when I control for employment. The estimates falling dramatically after I control for employment might suggest that changes in employment and not WC behavior drive the results.

The above-age-26 coefficients are reported in column 4 of Table 7 and are slightly larger in size but statistically indistinguishable from the original estimates for all of the estimates that were significant in column 1. The reason that the estimates do not fall is that more 26-year-olds are working in Texas than 25-year-olds in 2012 and 2013. Appendix E displays the total number of employed people in Texas for each age and year. These results provide more evidence that discontinuous increases in employment at age 26 are not responsible for the increased number of WC bills at age 26.

The top four graphs of Figure 3 display counts of claims classified by the likelihood that the ICD-9 code is first reported on weekends. It appears that claims for injuries least likely to first be reported on weekends increase at age 26, while the age profiles for injuries that are more

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equal to one if the logged count is from 2011 to 2013. The coefficient on *Post26* interacted with the 2011–2013 indicator is the difference between the two estimates presented in columns 1 and 2.

<sup>14</sup> Since the ACS does not have age in months, these employment estimates are imperfect. However, if employment increases gradually around age 26, using the employment for each age divided by 12 will bias the estimates towards zero. Employment could increase gradually for the reasons outlined in the previous section and would not be a threat to identification.



likely to occur on the weekend trend smoothly. Panel A of Table 8 displays estimates of the discontinuities at age 26 for claims classified by the likelihood that the ICD-9 code is first reported on the weekend. For injuries in the bottom quartile of the weekend likelihood distribution, WC claims rise by 25.7 percent, which provides evidence that the claiming behavior of people with deferrable injuries and illnesses is influenced by health insurance. For all other kinds of injuries, the estimates are statistically indistinguishable from zero.

The bottom four graphs of Figure 3 display counts of bills organized by the likelihood that the original injury the original injury is first reported on the weekend. As most of the injuries in the bottom half of the weekend distribution are sprains, strains, or occupational disease bills, bills for injuries in the bottom two quartiles of the weekend distribution increase discontinuously at age 26. Panel B of Table 8 displays estimates of bill increases with bills classified by the likelihood that the ICD-9 code is first reported on the weekend. The estimate for the bottom quartile of the weekend distribution is 0.230, while the estimate for the twenty-sixth to fiftieth percentile of the weekend distribution is 0.142. The estimates for the top two quartiles of the weekend distribution are statistically insignificant. These results suggest that follow-up care for open wounds or burns is more easily linked to the original injury than follow-up care for back strains or shoulder sprains.

The second column in Table 8 displays equivalent estimates using data from 2005 to 2007. As expected, I find no evidence of significant discontinuities in WC bills in the early years of the Texas law. As with the estimates in Table 7, the difference between the estimate from columns 1 and 2 is significant in every case, and the estimate in column 1 is statistically significant. Controlling for employment in column 4 does not dramatically alter the estimates.

For a more thorough study of different types of injuries, I follow Campolieti and Hyatt (2006) and classify injuries as back injuries, cuts and lacerations, dislocations, burns, contusions, or fractures based on their ICD-9 codes. Note that this classification largely excludes occupational diseases and sprains and strains.<sup>15</sup> The age profiles generally trend less smoothly when the results are broken down into these narrower categories, so these results should be interpreted with caution.

Table 9 displays estimates of the above-age-26 coefficients for the specific types of injuries using 2011–2013 data. With claims, no category shows a statistically significant discontinuity at age 26. With bills, the discontinuities for back injury bills and dislocation bills are both positive and statistically significant at age 26. As treatment for these two types of injuries is likely discretionary, the results provide more evidence that the most discretionary care responds to health insurance.

### **Type and Intensity of Treatment**

The results suggest that bills are much more sensitive to health insurance than claims. These patterns are consistent with WC paying for more care for workers who were injured prior to turning 26. An alternative story, though, is that more severe claims with more bills are reported immediately after people turn 26, which may cast doubt that health insurance is responsible for the increase since health insurance would likely affect more marginal care. I distinguish between these two explanations empirically by studying how the number of bills per claim changes at age 26. If the increase in bills comes from more severe claims being reported at age 26, the number of bills per claim will increase at age 26.

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<sup>15</sup> Campolieti and Hyatt (2006) include strains and sprains as a separate category, which I did as well in Table 7. Certain injuries, such as crushing injuries or damage to the eye, are excluded from both of our classifications. I do not report the results for strains and sprains and occupational diseases because they are shown in Table 7.

Graph A of Figure 4 displays the mean number of bills per claim by the age of claimants when claims began. The graph provides no evidence to suggest that claims immediately become more severe at age 26. Table 10 displays the above-age-26 coefficient from Equation (3) with the mean number of bills per claim as the dependent variable. The estimated discontinuity is a statistically insignificant  $-0.228$  from a mean of about 6.5. Graph B of Figure 4 displays the mean of the log of number of bills for each claim. Taking the log of the number of bills before calculating the means has the advantage of downweighting outliers and means that the discontinuity at age 26 can be interpreted as the percent change in the number of bills per claim. The graph again provides no evidence that claims are more severe at age 26. When the dependent variable is the mean of the log of the number of bills per claim, the above-age-26 coefficient from Equation (3) is  $-0.035$ . These results are not consistent with claims becoming more severe at age 26 and instead suggest that WC pays for more medical care for already opened claims once young adults lose health insurance at age 26.

One reason the payment source could change is that providers could prefer to be paid through health insurance than through WC. If the increase in bills comes from providers' billing preferences, care that would be covered by both health insurance and WC would likely be more sensitive to health insurance coverage since these bills would allow providers to have their preferred payer pay for the care. I next consider how office visits, chiropractic care, and physical therapy change at age 26 because either WC or health insurance would pay for these treatments.

Graph C considers how counts of WC bills for office visits trend with age, while graph D considers how counts of WC bills for chiropractic care and physical therapy trend with age. Graph E considers how counts of all other bills trend with age. Immediately at age 26, WC is more likely to pay for office visits, chiropractic care, and physical therapy. For other bills, the

increase at age 26 appears to be smaller. Table 10 displays the above-age-26 coefficients for these treatments. The above-age-26 coefficient is 0.081 when the log of the count of bills for office visits is the dependent variable and 0.180 when the log of the count of bills for physical therapy and chiropractic care is the dependent variable. For all other bills, the above age-age-26 coefficient is statistically insignificant. These results provide evidence that care that would be covered by either WC or health insurance is most responsive to health insurance.

### **WC Costs**

Next, I examine the impact of the increased number of bills at age 26 on WC costs. I use the current price index to compute the cost of each WC medical bill in 2013 dollars and classify bills into quartiles by cost. I focus on the 2011–2013 estimates as I did not find evidence of bill increases for 2005–2007. A bill is in the top quartile if it is greater than \$568 and in the bottom quartile if it is less than \$127. Graphs A–D of Figure 5 display counts of bills by cost quartile and reveal discontinuities for the bottom three quartiles of bill costs. The above-age-26 coefficients from Equation (3) are shown in Panel A of Table 11 and range from 8.6 to 12.4 percent for bills in the bottom three quartiles. The number of bills larger than \$568 trends smoothly at age 26.

The most expensive bills trending smoothly at age 26 is significant because expensive bills drive WC medical costs. Though only 25 percent of total bills, bills larger than \$568 accounted for 89 percent of total medical spending in Texas WC from 2011 to 2013. Graphs E–H show the total cost of all bills in each quartile by age in months, and Panel B of Table 11 shows the above-age-26 coefficients when the dependent variable is the log of the total cost of bills in each quartile. Total spending increases at age 26 for all but the most expensive bills. The

estimates of the percent increase in WC costs at age 26 for the bottom three quartiles of the cost distribution range from 9.5 to 12.6.

The estimate of the discontinuity in spending on the most expensive bills is statistically insignificant, but the standard error is large at 0.144. The graph of total spending on the most expensive bills suggests some ages around age 26 have large spending even though the trend does not appear to change. Table 11 reports the equivalent estimate with the 280 bills over \$100,000 dropped from the analysis. The estimate for total spending on the most expensive bills becomes an insignificant  $-0.005$ .<sup>16</sup> The estimate of untrimmed total spending is an insignificant 0.213. Once the outliers are removed, the estimate becomes 0.011. Overall, these results suggest that small medical bills respond to health insurance and that the cost to the overall WC system from young adults losing health insurance at age 26 is small.<sup>17</sup>

### **The Effect of Health Insurance on WC Filing**

The reduced-form estimates presented in the “WC Results” section measure the percent change in WC claims and bills occurring at age 26. I generate IV estimates by dividing the 2011–2013 percent change estimates from Tables 7–11 by the estimate of the decrease in insurance coverage at age 26 for the Texas working subsample of  $-0.053$  and compute standard errors using the delta method. Although it is statistically insignificant, I use the Texas estimate because it is produced using the same population as the WC estimates. As it is close to all of the other estimates of the decrease at age 26, the IV estimates are similar if I use the national estimates of the discontinuity at age 26.

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<sup>16</sup> The transitory increase immediately at age 26 is from two claimants with very large medical bills, which illustrates the sensitivity of estimates of total spending to outliers.

<sup>17</sup> Theoretically, health insurance would likely matter less if an injury causes people to miss work. Unfortunately, the WC data do not contain information on payments for missed work. However, larger medical bills being unaffected by insurance is consistent with little effect on time-loss claims since more severe medical issues usually have higher medical bills and more time missed from work.

The IV estimates are shown in Table 12. The implied IV estimate for all claims is  $-0.19$ , while the implied IV estimate for all bills is  $-1.53$ . These estimates imply that a 10 percentage point decrease in insurance coverage for young adults in Texas would increase WC claims by a statistically insignificant 1.9 percent and WC bills by 15.3 percent. This large IV effect on bills is driven by large effects of health insurance on occupational diseases bills and sprain bills. The IV estimates for these types of bills are  $-2.15$  and  $-2.68$ , respectively. The implied IV estimates for the bottom quartile of claims is especially high. They suggest that a 10 percentage point increase in insurance coverage for young adults in Texas would decrease WC claims by 48.5 percent and WC bills by 43.4 percent. The IV estimates for bills in the lower three cost quartiles range from  $-2.34$  to  $-1.62$  and are statistically indistinguishable from one another. For bills in the top quartile, the IV estimate is a statistically insignificant  $-0.19$ .<sup>18</sup>

## DISCUSSION AND CONCLUSIONS

The results indicate that health insurance has a large impact on WC claims for the most deferrable conditions and on total WC medical bills. It is important to note that the estimates presented in this paper are local in that they apply to young adults in Texas who lose health insurance through their parents' employers at age 26. Young adults differ from older adults in important ways. While younger workers are healthier on average than older workers, they are also more likely to work in jobs with more physical demands and harsher environmental conditions than older workers (Fletcher, Sindelar, and Yamaguchi 2011). Because of this, young

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<sup>18</sup> One important caveat to these implied estimates is that the fall in employer-sponsored coverage for working adults is larger than the fall in coverage. As employer-sponsored coverage is typically better than other health insurance sources, some of the increased care at age 26 could come from people who transition to inferior coverage. If we think of employer-sponsored health insurance as the first stage, the implied IV estimates would be smaller.

adults are more prone to occupational injuries (Salminen 2004) and are thus an important group for WC. During the time period studied, people under 35 accounted for over a third of all Texas WC claims.

Young adults losing health insurance at age 26 in Texas likely differ from other young adults in a number of ways. Texas has a higher uninsured rate than other states, meaning that the fall in health insurance is greater in percentage terms than it would be for the rest of the nation. According to SIPP data, young adults with insurance through their parents' employers at age 25 are less likely to be married and more likely to be working part time. These factors indicate that taking away one source of health insurance matters more for them because they are less likely to have access to other high quality sources. As less-healthy young adults are more likely to sign up for dependent coverage, people losing coverage at age 26 are likely less healthy than other young adults. Since dependent coverage comes from another source besides the individual's own employer, the impact of losing health insurance is likely larger for these people than it would be if people were losing coverage through their own employers.<sup>19</sup> These factors suggest the impact of health insurance on WC may be high for this population, which is notable since I find little effect on overall costs.

The main difference between Texas WC and WC in other states is that all other states other than Oklahoma require all employers to purchase WC insurance. Since the estimates are in percentage terms, requiring all employers to purchase WC insurance would generalize to other states if the relationship between health insurance and WC is constant across firms. Although Morantz (2010) finds that large firms opting out of WC in Texas offer health insurance at rates

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<sup>19</sup> Having health insurance through a source other than one's own employer is common. According to 2013 March CPS data, 30 percent of workers aged 18–64 have health insurance through a source other than their own employers.

similar to other large employers, small firms opting out may be less likely to offer generous health insurance plans. These firms offering no or worse health insurance would mean the discontinuity in WC claims and bills at age 26 would be larger because young adults losing dependent coverage would not be able to access insurance through their own employers.

Texas WC also differs from WC in other states in that injured workers are allowed to see the doctor of their choice, while 17 state WC programs restrict physician choice (Workers Compensation Research Institute 2014). Restrictions on physician choice make both health insurance and self-payment more appealing compared to WC. The increase in WC claims and bills may be larger at age 26 in more restrictive states since more 25-year-olds would be using health insurance instead of WC. On the other hand, some people will prefer to pay for their own medical services out of pocket or forgo care altogether when they lose health insurance rather than receive care through the more restrictive WC, meaning that the increase in WC claims and bills would be smaller at age 26. If provider preferences are the reason health insurance was paying for some treatment, restricting physician choice means that WC would have been paying for all bills prior to 26 and there would be no discontinuity at age 26.

The results from this paper are most comparable to Heaton (2012), who studies the impact of the Massachusetts health insurance reform of 2006 on WC emergency room bills. The implied IV estimates of the effect of health insurance on WC medical bills in his study range from 1.17 to 2.42.<sup>20</sup> He finds similarly large effects of health insurance on large and small bills. My IV estimate for the increase in WC bills of 1.53 is closer to the lower bound of his estimates.

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<sup>20</sup> I calculate the implied IV estimates using Heaton's (2012) estimates of the change in WC emergency room bills and Kolstad and Kowalski's (2012) estimate of the increase in health insurance coverage from the Massachusetts reform of 5.71 percentage points. The increase in health insurance in Massachusetts is from equal-sized increases in employer-sponsored coverage and Medicaid. Though most of Heaton's estimates are for emergency room bills, he also considers inpatient hospital admissions and finds slightly smaller effects.



In contrast to Heaton, I find little impact on emergency care and that large bills do not respond to health insurance. These differences imply that the treatment effect for young adults losing coverage in Texas may differ from uninsured Massachusetts residents gaining coverage, that aspects of the Massachusetts health insurance reform apart from the expansion in coverage decreased WC bills, or that the effect of health insurance on WC was stronger at the height of the Great Recession.

The estimates presented in this paper represent one of the first sets of quasi-experimental estimates of the effect of health insurance on WC claiming behavior. To my knowledge they are the first to use WC administrative data that allow for separately examining the intensive and extensive margins of claiming, and the first to focus on young adults specifically. As people under 35 comprise over half of uninsured nonelderly adults, knowing the effect of health insurance on WC filing for young adults is important in understanding the potential impact of the ACA on WC. The results presented in this paper suggest that the expansion of insurance from the ACA will shift a large amount of discretionary medical care from being paid for by WC to being paid for by health insurance. Despite the large effects of health insurance on claiming behavior, the cost savings to WC systems will likely be small since cheaper services are most responsive to health insurance.

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**Table 1 Means of Key Variables from the NHIS**

|                                   | Aged 24–25 | Aged 26–27 |
|-----------------------------------|------------|------------|
| Male                              | 0.47       | 0.47       |
| Completed high school             | 0.87       | 0.87       |
| Completed college                 | 0.26       | 0.29       |
| Black                             | 0.16       | 0.15       |
| White                             | 0.74       | 0.74       |
| Hispanic                          | 0.26       | 0.26       |
| Married                           | 0.27       | 0.38       |
| Employer-sponsored insurance      | 0.57       | 0.53       |
| Privately purchased insurance     | 0.03       | 0.04       |
| Uninsured                         | 0.31       | 0.34       |
| Working                           | 0.71       | 0.74       |
| Hours worked per week             | 26.71      | 28.57      |
| Works less than 35 hours per week | 0.18       | 0.17       |
| Works less than 30 hours per week | 0.12       | 0.11       |
| Self-employed                     | 0.01       | 0.02       |
| N                                 | 5,055      | 5,322      |

NOTE: The sample includes all individuals aged 24–27 in the 2011 and 2012 NHIS.



Table 2 Regression Discontinuity Models of Health Insurance and Employment by Age from the NHIS

| Dependent variable                     | Mean at age 25 | Discontinuity at age 26 | n      |
|--|----------------|-------------------------|--------|
| <b>Panel A: Full sample</b>            |                |                         |        |
| Employer-sponsored insurance           | 0.58           | −0.086***<br>(0.024)    | 8,356  |
| Privately purchased insurance          | 0.03           | 0.010<br>(0.010)        | 10,013 |
| Uninsured                              | 0.30           | 0.042*<br>(0.023)       | 10,199 |
| Working                                | 0.71           | −0.011<br>(0.017)       | 10,249 |
| Hours worked per week                  | 27.99          | −0.636<br>(1.312)       | 10,290 |
| Log of hours worked per week           | 3.59           | −0.006<br>(0.044)       | 7,291  |
| Works less than 35 hours per week      | 0.17           | 0.002<br>(0.017)        | 10,168 |
| Works less than 30 hours per week      | 0.11           | 0.004<br>(0.019)        | 10,168 |
| Self-employed                          | 0.01           | 0.008<br>(0.008)        | 10,274 |
| <b>Panel B: Conditional on working</b> |                |                         |        |
| Employer-sponsored insurance           | 0.65           | −0.076***<br>(0.028)    | 6,460  |
| Privately purchased insurance          | 0.03           | 0.008<br>(0.012)        | 7,239  |
| Uninsured                              | 0.27           | 0.044*<br>(0.026)       | 7,383  |

NOTE: \* significant at the 0.10 level; \*\* significant at the 0.05 level; \*\*\* significant at the 0.01 level. The sample includes all individuals aged 24–27 in the 2011 and 2012 NHIS. The discontinuities at age 26 are the above-age-26 coefficients from separate regressions of Equation (1) that control for indicators for race, sex, year, and marital status, education, and a quadratic polynomial in age in months fully interacted with an indicator for age greater than or equal to 26. Standard errors are clustered by age in months and are shown in parentheses. The estimates in Panel A are for the full sample, while the estimates in Panel B are restricted to the sample of working individuals.

**Table 3 Means of Key Variables from the SIPP**

|   | Texas      |            | Rest of U.S. |            |
|---|------------|------------|--------------|------------|
|   | Aged 23–25 | Aged 26–28 | Aged 23–25   | Aged 26–28 |
| Male  | 0.51       | 0.48       | 0.50         | 0.49       |
| Completed high school                       | 0.92       | 0.91       | 0.93         | 0.92       |
| Completed college                           | 0.19       | 0.26       | 0.25         | 0.32       |
| Black                                       | 0.17       | 0.13       | 0.13         | 0.11       |
| White                                       | 0.76       | 0.81       | 0.77         | 0.78       |
| Hispanic                                    | 0.39       | 0.36       | 0.14         | 0.14       |
| Married                                     | 0.24       | 0.38       | 0.19         | 0.38       |
| Employer-sponsored insurance as a dependent | 0.22       | 0.09       | 0.27         | 0.12       |
| Insurance through own employer              | 0.24       | 0.36       | 0.26         | 0.39       |
| Employer-sponsored insurance                | 0.46       | 0.45       | 0.52         | 0.50       |
| Uninsured                                   | 0.49       | 0.50       | 0.40         | 0.43       |
| Working                                     | 0.72       | 0.81       | 0.72         | 0.77       |
| Hours worked per week                       | 25.69      | 31.84      | 25.81        | 29.42      |
| Works less than 35 hours per week           | 0.23       | 0.14       | 0.23         | 0.15       |
| Works less than 30 hours per week           | 0.16       | 0.10       | 0.16         | 0.11       |
| Self-employed                               | 0.03       | 0.06       | 0.03         | 0.05       |
| N   | 1,727      | 1,618      | 18,731       | 18,578     |

NOTE: The sample includes all individuals aged 23–28 in the January 2011 to August 2013 SIPP.

Table 4 Fixed Effects Models of Health Insurance and Employment from the SIPP

| Dependent variable                                | Texas          |                      |       | Rest of U.S.   |                      |        |
|---|----------------|----------------------|-------|----------------|----------------------|--------|
|   | Mean at age 25 | Effect of turning 26 | N     | Mean at age 25 | Effect of turning 26 | N      |
| <b>Panel A: Full sample</b>                       |                |                      |       |                |                      |        |
| Employer-sponsored insurance as a dependent       | 0.16           | −0.082**<br>(0.036)  | 3,345 | 0.20           | −0.046***<br>(0.010) | 37,309 |
| Employer-sponsored insurance through own employer | 0.27           | 0.007<br>(0.038)     | 3,345 | 0.33           | 0.005<br>(0.011)     | 37,309 |
| Employer-sponsored insurance                      | 0.44           | −0.075*<br>(0.042)   | 3,345 | 0.53           | −0.043***<br>(0.011) | 40,654 |
| Uninsured   | 0.52           | 0.031<br>(0.040)     | 3,345 | 0.32           | 0.050***<br>(0.010)  | 40,654 |
| Working   | 0.76           | −0.011<br>(0.033)    | 3,345 | 0.75           | −0.012<br>(0.010)    | 37,309 |
| Hours worked per week                             | 28.11          | 0.238<br>(1.407)     | 3,344 | 28.20          | −1.027**<br>(0.406)  | 37,297 |
| Log of hours worked per week                      | 3.54           | −0.008<br>(0.040)    | 2,555 | 3.55           | −0.037***<br>(0.013) | 27,732 |
| Works less than 35 hours per week                 | 0.20           | −0.030<br>(0.032)    | 3,344 | 0.19           | 0.017*<br>(0.010)    | 37,297 |
| Works less than 30 hours per week                 | 0.14           | −0.024<br>(0.029)    | 3,344 | 0.12           | 0.018**<br>(0.009)   | 37,297 |
| Self-employed                                     | 0.04           | −0.013<br>(0.016)    | 3,345 | 0.03           | −0.002<br>(0.004)    | 37,309 |
| <b>Panel B: Conditional on working</b>            |                |                      |       |                |                      |        |
| Employer-sponsored insurance as a dependent       | 0.14           | −0.078**<br>(0.039)  | 2,556 | 0.20           | −0.043***<br>(0.012) | 27,745 |
| Employer-sponsored insurance through own employer | 0.36           | −0.003<br>(0.048)    | 2,556 | 0.43           | 0.011<br>(0.013)     | 27,745 |
| Employer-sponsored insurance                      | 0.49           | −0.084*<br>(0.048)   | 2,556 | 0.62           | −0.036***<br>(0.012) | 30,301 |
| Uninsured   | 0.47           | 0.053<br>(0.044)     | 2,556 | 0.32           | 0.045***<br>(0.012)  | 30,301 |

NOTE: \* significant at the 0.10 level; \*\* significant at the 0.05 level; \*\*\* significant at 0.01 level. The sample includes all individuals aged 23–28 in the January 2011 to August 2013 SIPP. The effects of turning 26 are the above-age-26 coefficients from separate regressions of Equation (2) that control for indicators for the year and month of the observation and for marital status, a linear control in age, and an individual fixed effect. Standard errors are clustered by individual and are shown in parentheses. The estimates in Panel A are for the full sample, while the estimates in Panel B are restricted to the sample of working individuals.

Table 5 Top Claims in Each Quartile of Weekend Likelihood Distribution

|   | Percent of<br>all claims | Percent first receiving<br>care on weekends |
|---|--------------------------|---|
| <u>Top quartile (weekend likelihood <math>\geq 0.096</math>)</u>                              |                          |   |
| Open wound of finger  | 0.055                    | 0.108                                       |
| Unspecified injury  | 0.047                    | 0.128                                       |
| Open wound of hand  | 0.019                    | 0.108                                       |
| Open wound of head  | 0.016                    | 0.097                                       |
| Open wound of elbow, forearm, and wrist   | 0.011                    | 0.096                                       |
| <u>51st to 75th percentile (<math>0.069 &lt; \text{weekend likelihood} &lt; 0.096</math>)</u> |                          |   |
| Contusion of lower limb   | 0.052                    | 0.072                                       |
| Contusion of upper limb   | 0.042                    | 0.080                                       |
| Sprains and strains of ankle and foot   | 0.037                    | 0.071                                       |
| Contusion of trunk  | 0.018                    | 0.069                                       |
| Contusion of face, scalp, and neck  | 0.016                    | 0.089                                       |
| <u>26th to 50th percentile (<math>0.050 &lt; \text{weekend likelihood} &lt; 0.069</math>)</u> |                          |   |
| Sprains and strains of back   | 0.113                    | 0.050                                       |
| Unspecified disorder of joint   | 0.041                    | 0.056                                       |
| Sprains and strains of knee and leg   | 0.037                    | 0.057                                       |
| Unspecified disorder of back  | 0.031                    | 0.059                                       |
| Sprains and strains of wrist and hand   | 0.029                    | 0.065                                       |
| <u>Bottom quartile (weekend likelihood <math>&lt; 0.050</math>)</u>                           |                          |   |
| Sprains and strains of shoulder and upper arm   | 0.041                    | 0.049                                       |
| Disorders of bone attachments   | 0.012                    | 0.021                                       |
| Disorders of tendons, bursa, and synovium   | 0.012                    | 0.020                                       |
| Neck disorders  | 0.006                    | 0.046                                       |
| Intervertebral disc disorder  | 0.005                    | 0.022                                       |

NOTE: The weekend likelihood is constructed using all WC claims from Texas with a nonmissing date of first treatment.

SOURCE: The Texas WC medical database.

**Table 6 Texas WC Claims and Medical Bills Descriptive Statistics**

|                              | Claims         |          |        |                |          |        |
|------------------------------|----------------|----------|--------|----------------|----------|--------|
|                              | Aged 24 and 25 |          |        | Aged 26 and 27 |          |        |
|                              | Mean           | St. dev. | Median | Mean           | St. dev. | Median |
| Age                          | 24.98          | 0.58     | 25.00  | 26.96          | 0.58     | 26.92  |
| Male                         | 0.67           | 0.47     | 1.0    | 0.66           | 0.47     | 1.00   |
| Injury occurs on weekend (%) | 0.08           | 0.04     | 0.07   | 0.08           | 0.04     | 0.07   |
| Total charge (\$)            | 6,557          | 66,910   | 981    | 6,639          | 38,945   | 1,036  |
| Total medical bills          | 6.30           | 11.91    | 3.00   | 6.78           | 12.40    | 3.00   |
| N                            |                | 29,477   |        |                | 30,922   |        |
|                              | Medical bills  |          |        |                |          |        |
|                              | Aged 24 and 25 |          |        | Aged 26 and 27 |          |        |
|                              | Mean           | St. dev. | Median | Mean           | St. dev. | Median |
| Age                          | 24.98          | 0.57     | 25.00  | 26.96          | 0.58     | 27.00  |
| Male                         | 0.76           | 0.43     | 1.00   | 0.76           | 0.43     | 1.00   |
| Injury occurs on weekend (%) | 0.08           | 0.04     | 0.07   | 0.08           | 0.03     | 0.07   |
| Charge (\$)                  | 1,450          | 10,769   | 249    | 1,379          | 9,799    | 247    |
| N                            |                | 104,117  |        |                | 117,455  |        |

NOTE: The data include all WC claims and bills in Texas from 2011 to 2013.

SOURCE: The Texas WC medical database.

**Table 7 Regression Discontinuity Models for All Claims and Medical Bills and for Different Injury Types in Texas, by Age**

|   | 2011–2013           | 2005–2007         | Difference          | 2011–2013<br>w/employment<br>control |
|---|---------------------|-------------------|---------------------|--------------------------------------|
| <b>Panel A: RD models for log (number of initiated WC claims) in Texas by age at start of claim</b> |                     |                   |                     |                                      |
| All claims  | 0.010<br>(0.016)    | 0.001<br>(0.016)  | 0.009<br>(0.023)    | 0.004<br>(0.027)                     |
| <b>by Type</b>  |                     |                   |                     |                                      |
| Sprain/strain claims  | 0.060<br>(0.045)    | 0.073<br>(0.047)  | –0.013<br>(0.065)   | 0.038<br>(0.048)                     |
| Nonsprain/strain claims   | –0.013<br>(0.034)   | –0.019<br>(0.030) | 0.006<br>(0.045)    | –0.004<br>(0.037)                    |
| Occupational disease claims   | 0.010<br>(0.052)    | –0.036<br>(0.030) | 0.046<br>(0.060)    | –0.015<br>(0.069)                    |
| <b>Panel B: RD models for log (number of WC medical bills) in Texas by age at treatment</b>         |                     |                   |                     |                                      |
| All bills   | 0.081***<br>(0.026) | 0.019<br>(0.018)  | 0.062**<br>(0.031)  | 0.092***<br>(0.034)                  |
| <b>by Type</b>  |                     |                   |                     |                                      |
| Sprain/strain bills   | 0.114***<br>(0.040) | 0.030<br>(0.029)  | 0.084*<br>(0.050)   | 0.151***<br>(0.050)                  |
| Nonsprain/strain bills  | 0.048<br>(0.045)    | 0.041<br>(0.028)  | 0.007<br>(0.053)    | 0.046<br>(0.045)                     |
| Occupational disease bills  | 0.142***<br>(0.037) | –0.046<br>(0.042) | 0.188***<br>(0.056) | 0.160***<br>(0.061)                  |

NOTE: \* significant at the 0.10 level; \*\* significant at the 0.05 level; \*\*\* significant at the 0.01 level. The first column displays the above-age-26 coefficients from separate regressions of Equation (3) using 2011–2013 data. The second column displays the above-age-26 coefficients from separate regressions of Equation (3) using 2005–2007 data. The third column is the difference between the 2011–2013 estimates and the 2005–2007 estimates. The fourth column displays the above-age-26 coefficients from separate regressions of Equation (3) using 2011–2013 data that control for the number of people working at each age in each year in Texas from the 2011 to 2013 ACS. The models in columns 1, 2, and 4 include a quadratic polynomial in age in months fully interacted with an indicator for age greater than or equal to 26. In columns 1 and 2, there is a separate observation for each age in months. In column 4, there is a separate observation for each age in months for each year of the data, and the models include a vector of year indicator variables. Robust standard errors are shown in parentheses. In Panel A, the dependent variables are the log of the number of claims. In Panel B, the dependent variables are the log of the number of bills. There are 60,399 WC claims for young adults aged 23–27 from 2011 to 2013 and 84,491 from 2005 to 2007. There are 221,572 WC bills for young adults aged 23–27 from 2011 to 2013 and 351,439 from 2005 to 2007.

SOURCE: The Texas WC medical database.

**Table 8 Regression Discontinuity Models for Claims and Medical Bills by Weekend Likelihood in Texas, by Age**

|   | 2011–2013           | 2005–2007         | Difference          | 2011–2013<br>w/employment<br>control |
|---|---------------------|-------------------|---------------------|--------------------------------------|
| <b>Panel A: RD models for log (number of initiated WC claims) in Texas by age at start of claim</b> |                     |                   |                     |                                      |
| Bottom quartile   | 0.257**<br>(0.098)  | –0.030<br>(0.042) | 0.287***<br>(0.106) | 0.235**<br>(0.101)                   |
| 26th to 50th percentile   | –0.038<br>(0.033)   | 0.047<br>(0.041)  | –0.085<br>(0.053)   | –0.047<br>(0.041)                    |
| 51st to 75th percentile   | –0.008<br>(0.032)   | 0.000<br>(0.048)  | –0.009<br>(0.058)   | –0.005<br>(0.042)                    |
| Top quartile  | –0.001<br>(0.043)   | –0.038<br>(0.035) | 0.037<br>(0.056)    | –0.007<br>(0.042)                    |
| <b>Panel B: RD models for log (number of WC medical bills) in Texas by age at treatment</b>         |                     |                   |                     |                                      |
| Bottom quartile   | 0.230***<br>(0.069) | 0.038<br>(0.045)  | 0.191**<br>(0.082)  | 0.244**<br>(0.121)                   |
| 26th to 50th percentile   | 0.142***<br>(0.052) | 0.004<br>(0.025)  | 0.138**<br>(0.057)  | 0.127**<br>(0.060)                   |
| 51st to 75th percentile   | 0.005<br>(0.053)    | –0.003<br>(0.034) | 0.008<br>(0.063)    | 0.015<br>(0.056)                     |
| Top quartile  | 0.038<br>(0.051)    | 0.052<br>(0.041)  | –0.014<br>(0.065)   | 0.071<br>(0.062)                     |

NOTE: \* significant at the 0.10 level; \*\* significant at the 0.05 level; \*\*\* significant at the 0.01 level. Claims are classified into quartiles by the likelihood that the ICD-9 code was first reported on the weekend as described in the text. The first column displays the above-age-26 coefficients from separate regressions of Equation (3) using 2011–2013 data. The second column displays the above-age-26 coefficients from separate regressions of Equation (3) using 2005–2007 data. The third column is the difference between the 2011–2013 estimates and the 2005–2007 estimates. The fourth column displays the above-age-26 coefficients from separate regressions of Equation (3) using 2011–2013 data that control for the number of people working at each age in each year in Texas from the 2011 to 2013 ACS. The models in columns 1, 2, and 4 include a quadratic polynomial in age in months fully interacted with an indicator for age greater than or equal to 26. In columns 1 and 2, there is a separate observation for each age in months. In column 4, there is a separate observation for each age in months for each year of the data, and the models include a vector of year indicator variables. Robust standard errors are shown in parentheses. In Panel A, the dependent variables are the log of the number of claims. In Panel B, the dependent variables are the log of the number of bills. There are 60,399 WC claims for young adults aged 23–27 from 2011 to 2013 and 84,491 from 2005 to 2007. There are 221,572 WC bills for young adults aged 23–27 from 2011 to 2013 and 351,439 from 2005 to 2007.

SOURCE: The Texas WC medical database.

**Table 9 Regression Discontinuity Models for Claims and Medical Bills by Narrow Injury Categories in Texas, by Age**

|   | Discontinuity at age 26 | N      |
|---|-------------------------|--------|
| <b>Panel A: RD models for log (number of initiated WC claims) in Texas by age at start of claim</b>   |                         |        |
| Back claims   | −0.066<br>(0.089)       | 6,683  |
| Contusion claims  | 0.037<br>(0.065)        | 7,268  |
| Laceration claims   | −0.042<br>(0.040)       | 10,159 |
| Fracture claims   | −0.013<br>(0.129)       | 2,749  |
| Burn claims   | −0.042<br>(0.144)       | 1,212  |
| Dislocation claims  | 0.221<br>(0.411)        | 406    |
| <b>Panel B: RD models for log (number of initiated WC medical bills) in Texas by age at treatment</b> |                         |        |
| Back bills  | 0.252***<br>(0.077)     | 28,247 |
| Contusion bills   | 0.073<br>(0.051)        | 15,076 |
| Laceration bills  | −0.093*<br>(0.053)      | 21,977 |
| Fracture bills  | 0.037<br>(0.114)        | 22,486 |
| Burn bills  | 0.223<br>(0.189)        | 4,243  |
| Dislocation bills   | 0.319***<br>(0.106)     | 5,127  |

NOTE: \* significant at the 0.10 level; \*\* significant at the 0.05 level; \*\*\* significant at the 0.01 level. The estimates are the above-age-26 coefficients from separate regressions of Equation (3) using 2011–2013 data. The models include a quadratic polynomial in age in months fully interacted with an indicator for age greater than or equal to 26. Robust standard errors are shown in parentheses. In Panel A, the dependent variables are the log of the number of claims. In Panel B, the dependent variables are the log of the number of bills. The numbers of each type of claims and bills in the sample are shown in column 3. SOURCE: The Texas WC medical database.



**Table 10 Regression Discontinuity Models for Intensity and Type of Treatment in Texas, by Age**

|   |                     |
|---|---------------------|
| Mean bills per claim                    | –0.228<br>(0.229)   |
| Mean of log (number of bills per claim) | –0.035*<br>(0.020)  |
| Office visit bills                      | 0.081***<br>(0.027) |
| Physical therapy and chiropractic bills | 0.180**<br>(0.075)  |
| Other bills                             | 0.039<br>(0.038)    |

NOTE: \* significant at the 0.10 level; \*\* significant at the 0.05 level; \*\*\* significant at the 0.01 level. The estimates are the above-age-26 coefficients from separate regressions of Equation (3) using 2011–2013 data. The models include a quadratic polynomial in age in months fully interacted with an indicator for age greater than or equal to 26. Robust standard errors are shown in parentheses. For the first estimate, the dependent variable is the mean number of bills per claim at each age. For the second estimate, the dependent variable is the mean of the log of the number of bills per claim at each age. For the third, fourth, and fifth estimates, the dependent variables are logs of counts of bills at each age.

SOURCE: The Texas WC medical database.

**Table 11 Regression Discontinuity Models for Cost of WC Medical Services in Texas, by Age****Panel A: RD models for log (number of WC medical bills) in Texas by age at treatment**


---

|                         |          |
|-------------------------|----------|
| Bottom quartile         | 0.086*   |
|                         | (0.045)  |
| 26th to 50th percentile | 0.124*** |
|                         | (0.045)  |
| 51st to 75th percentile | 0.101*** |
|                         | (0.033)  |
| Top quartile            | 0.010    |
|                         | (0.041)  |

**Panel B: RD models for log (total costs of WC medical bills) in Texas by age at treatment**

|                         |          |
|-------------------------|----------|
| Bottom quartile         | 0.095*   |
|                         | (0.049)  |
| 26th to 50th percentile | 0.126*** |
|                         | (0.046)  |
| 51st to 75th percentile | 0.093**  |
|                         | (0.036)  |
| Top quartile            | 0.227    |
|                         | (0.144)  |
| Top quartile-trimmed    | -0.005   |
|                         | (0.076)  |
| All bills               | 0.213    |
|                         | (0.128)  |
| All bills-trimmed       | 0.011    |
|                         | (0.066)  |

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NOTE: \* significant at the 0.10 level; \*\* significant at the 0.05 level; \*\*\* significant at the 0.01 level. The cost of bills is calculated in 2013 dollars. Bills are classified into quartiles by cost as described in the text. The estimates are the above-age-26 coefficients from separate regressions of Equation (3) using 2011–2013 data that include a quadratic polynomial in age in months fully interacted with an indicator for age greater than or equal to 26. Robust standard errors are shown in parentheses. In Panel A, the dependent variables are the log of the number of bills. In Panel B, the dependent variables are the log of total costs. For the trimmed bills, the 280 bills greater than \$100,000 are removed from the sample. There are 221,572 WC bills for young adults aged 23–27 from 2011 to 2013.

SOURCE: The Texas WC medical database.

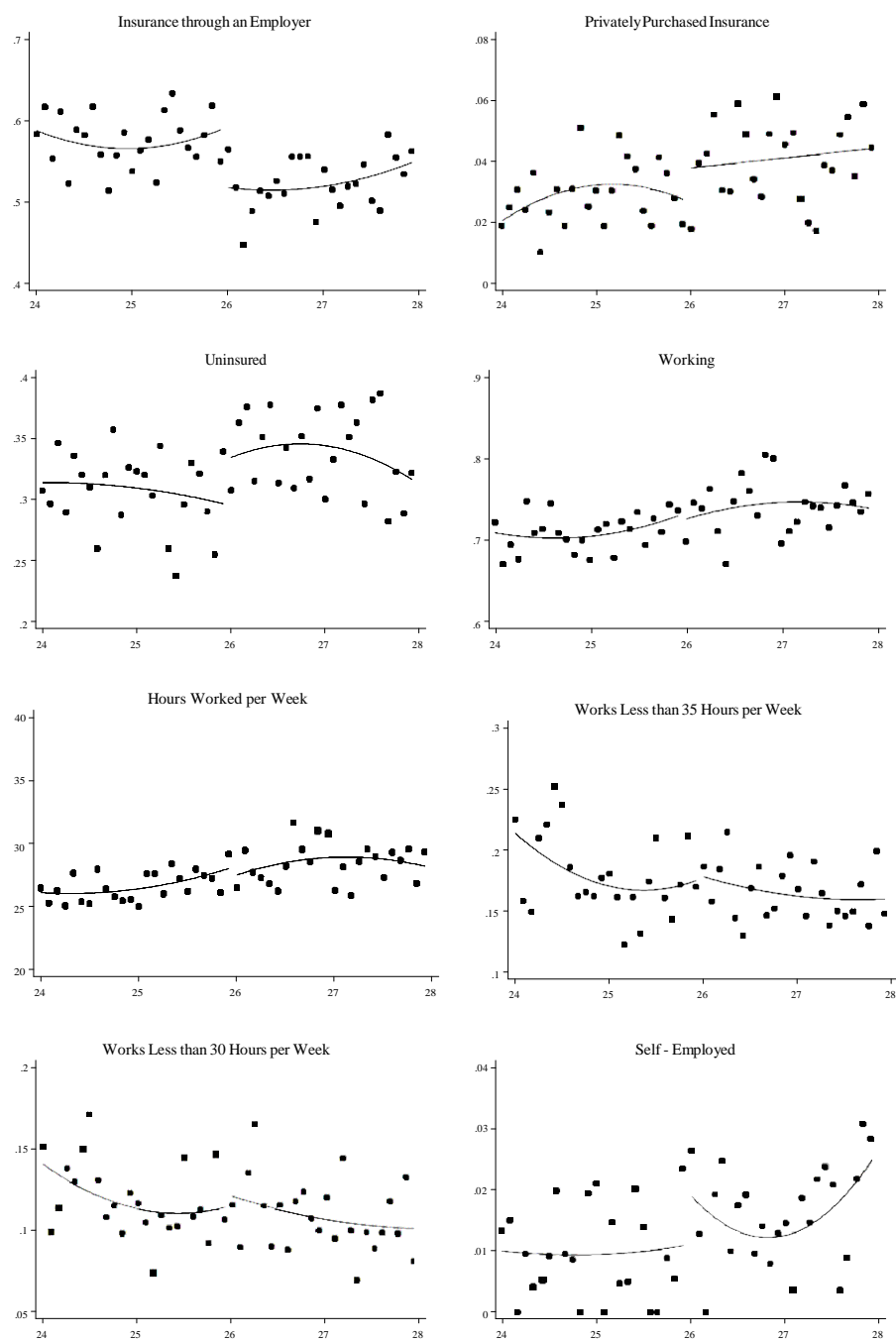
**Table 12 The Impact of Insurance on WC Claims, Bills, and Costs**

| <b>Panel A: Claims</b>       |                    | <b>Panel B: Bills</b>        |                    | <b>Panel C: Costs by cost quartile</b> |                    |
|------------------------------|--------------------|------------------------------|--------------------|--|--------------------|
| All claims                   | -0.19<br>(0.30)    | All bills                    | -1.53***<br>(0.49) | Bottom quartile                        | -1.62*<br>(0.85)   |
| <b>by Type</b>               |                    | <b>by Type</b>               |                    | 26th to 50th percentile                | -2.34***<br>(0.85) |
| Sprain/strain claims         | -1.13<br>(0.85)    | Sprain/strain bills          | -2.15***<br>(0.75) | 51st to 75th percentile                | -1.91***<br>(0.62) |
| Nonsprain/strain claims      | 0.25<br>(0.64)     | Nonsprain/strain bills       | -0.91<br>(0.85)    | Top quartile                           | -0.19<br>(0.77)    |
| Occupational disease claims  | -0.19<br>(0.98)    | Occupational disease claims  | -2.68***<br>(0.70) |  |                    |
| <b>by Weekend likelihood</b> |                    | <b>by Weekend likelihood</b> |                    | <b>Total costs by cost quartile</b>    |                    |
| Bottom quartile              | -4.85***<br>(1.85) | Bottom quartile              | -4.34***<br>(1.30) | Bottom quartile                        | -1.79*<br>(0.92)   |
| 26th to 50th percentile      | 0.72<br>(0.62)     | 26th to 50th percentile      | -2.68***<br>(0.98) | 26th to 50th percentile                | -2.38***<br>(0.87) |
| 51st to 75th percentile      | 0.15<br>(0.6)      | 51st to 75th percentile      | -0.09<br>(1.00)    | 51st to 75th percentile                | -1.75**<br>(0.68)  |
| Top quartile                 | 0.02<br>(0.81)     | Top quartile                 | -0.72<br>(0.96)    | Top quartile-trimmed                   | 0.09<br>(1.43)     |

NOTE: \* significant at the 0.10 level; \*\* significant at the 0.05 level; \*\*\* significant at the 0.01 level. Each coefficient is the ratio of the reduced-form coefficient from Tables 7–11 divided by the estimate of the change in insurance coverage at age 26 for working Texans of -0.053. Standard errors are computed using the delta method and are shown in parentheses.

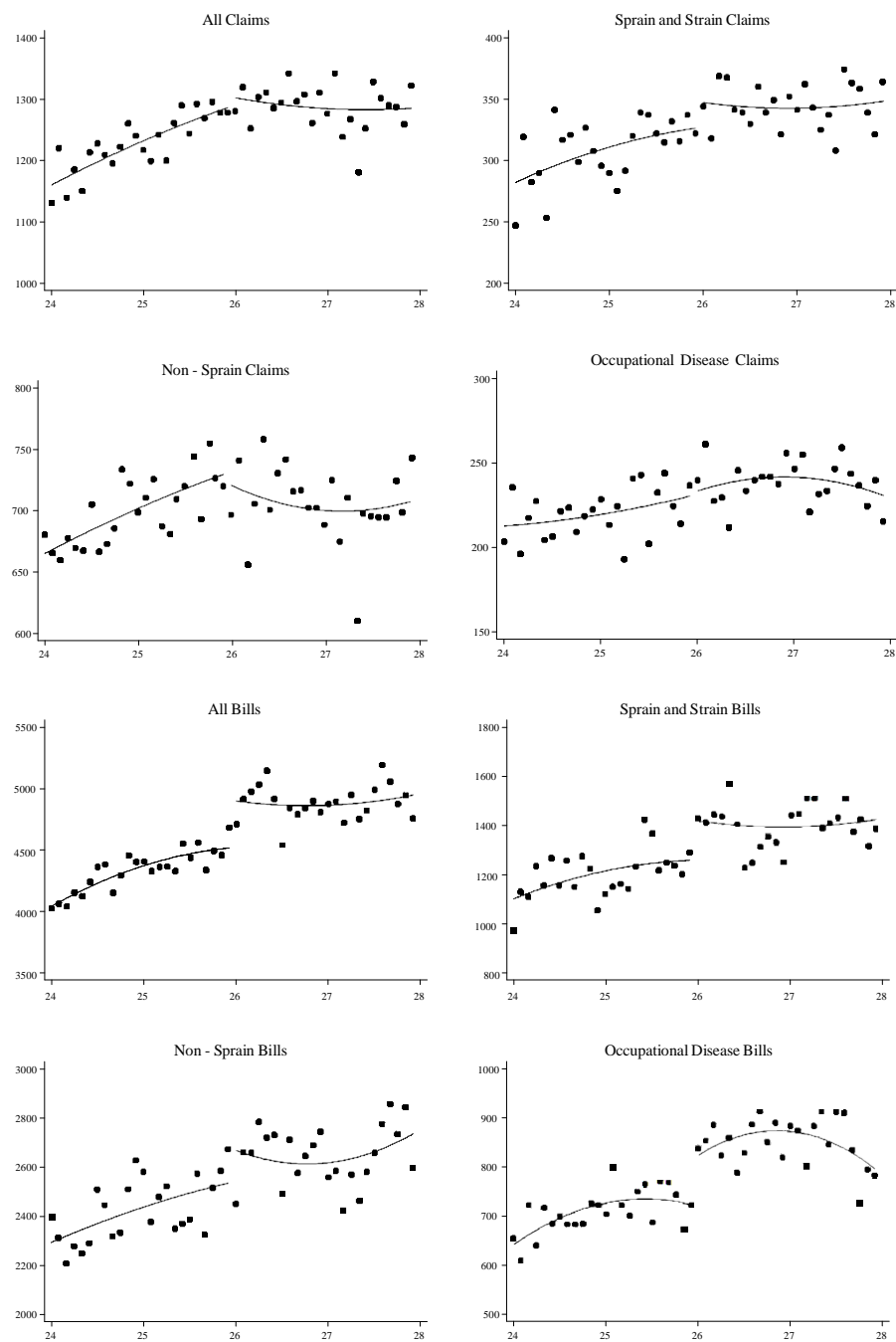
SOURCE: The Texas WC medical database.

**Figure 1 Means of Insurance and Employment Variables by Age from 2011 and 2012 NHIS**



NOTE: The points represent means of NHIS variables at each age in months.  
SOURCE: The Texas WC medical database.

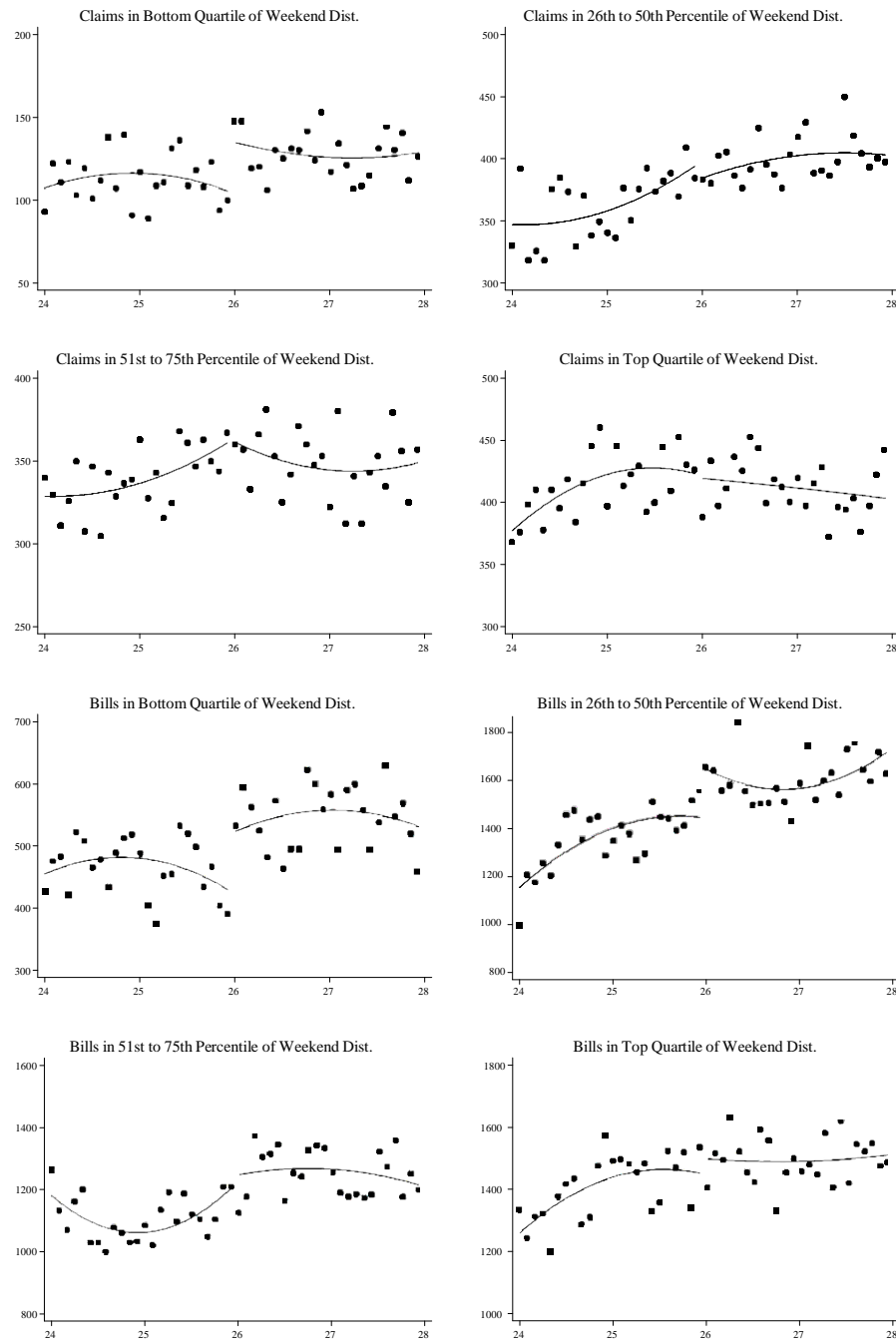
**Figure 2 Counts of WC Claims and Medical Bills in Texas by Age from 2011 to 2013**



NOTE: The points represent counts of WC claims or medical bills at each age in months.

SOURCE: The Texas WC medical database.

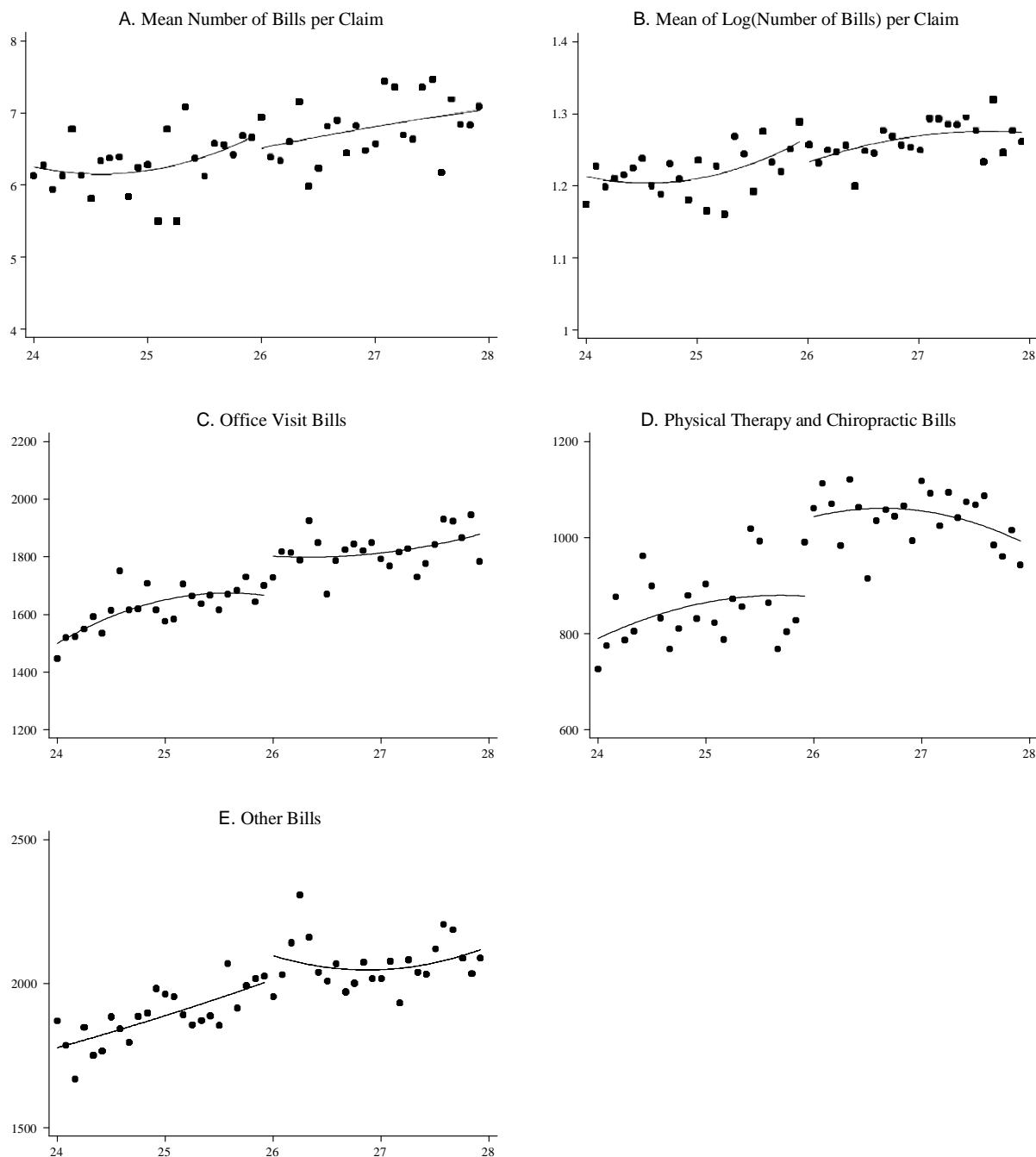
**Figure 3 Counts of WC Claims and Medical Bills in Texas by Age from 2011 to 2013 by Weekend Likelihood**



NOTE: Claims are classified into quartiles by the likelihood that the ICD-9 code was first reported on the weekend as described in the text. The points represent counts of WC claims or medical bills in each quartile of the weekend likelihood distribution at each age in months.

SOURCE: The Texas WC medical database.

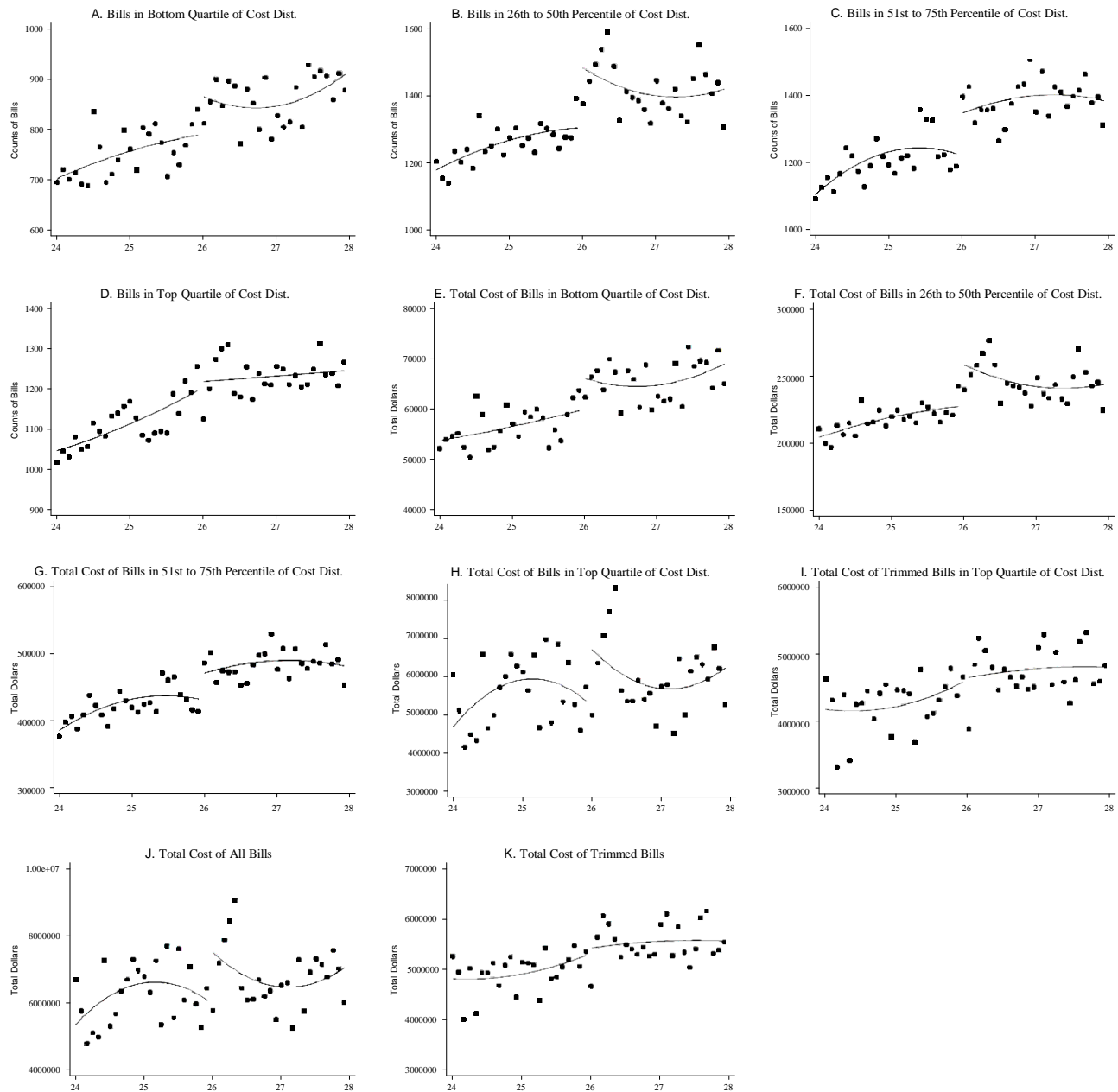
**Figure 4 Intensity and Type of WC Treatment in Texas by Age from 2011 to 2013**



NOTE: In graph A, each point represents the mean number of bills per claim by age in months at the start of the claim. In graph B, each point represents the mean of the log of the number of bills per claim by age in months at the start of the claim. In graphs C-E, each point represents counts of bills by the type of treatment paid for by age in months at the time of treatment.

SOURCE: The Texas WC medical database.

**Figure 5 WC Bill Cost in Texas by Age from 2011 to 2013**



NOTE: Bills are classified into quartiles by cost as described in the text. The points in graphs A-D represent counts of WC medical bills in each quartile of the cost distribution at each age in months. The points in graphs E-I represent total costs of all WC bills in each quartile of the cost distribution at each age in months. The points in graphs J and K represent total costs of all WC bills at each age in months. For graphs I and K, the 280 bills greater than \$100,000 are removed.

SOURCE: The Texas WC medical database.



## **Appendix A**

### **Robustness to Specification—SIPP**

In the second section of the paper, I include a linear control for age with the SIPP analysis. Table A.1 considers the sensitivity of the estimates in Table 4 to not controlling for age and to controlling for a quadratic polynomial in age.

**Table A.1 Sensitivity of Fixed Effects Models of Health Insurance and Employment from the SIPP**

| Dependent variable                                | Mean at age 25 | No age control      | Quadratic age control | N     | Mean at age 25 | No age control       | Quadratic age control | N      |
|---|----------------|---------------------|-----------------------|-------|----------------|----------------------|-----------------------|--------|
| <b>Panel A: Full sample</b>                       |                |                     |                       |       |                |                      |                       |        |
| Employer-sponsored insurance as a dependent       | 0.16           | −0.074**<br>(0.034) | −0.082**<br>(0.036)   | 3,345 | 0.20           | −0.048***<br>(0.009) | −0.045***<br>(0.010)  | 37,309 |
| Employer-sponsored insurance through own employer | 0.27           | 0.010<br>(0.035)    | 0.007<br>(0.038)      | 3,345 | 0.33           | 0.008<br>(0.010)     | 0.005<br>(0.011)      | 37,309 |
| Employer-sponsored insurance                      | 0.44           | −0.065*<br>(0.039)  | −0.075*<br>(0.042)    | 3,345 | 0.53           | −0.039***<br>(0.010) | −0.039***<br>(0.011)  | 40,654 |
| Uninsured   | 0.52           | 0.033<br>(0.036)    | 0.031<br>(0.040)      | 3,345 | 0.32           | 0.051***<br>(0.010)  | 0.051***<br>(0.010)   | 40,654 |
| Working   | 0.76           | 0.000<br>(0.031)    | −0.010<br>(0.033)     | 3,345 | 0.75           | −0.001<br>(0.009)    | −0.012<br>(0.010)     | 37,309 |
| Hours worked per week                             | 28.11          | 0.743<br>(1.340)    | 0.261<br>(1.405)      | 3,344 | 28.20          | −0.339<br>(0.369)    | −1.033**<br>(0.406)   | 37,297 |
| Log of hours worked per week                      | 3.54           | 0.005<br>(0.038)    | −0.010<br>(0.040)     | 2,555 | 3.55           | −0.021*<br>(0.012)   | −0.039***<br>(0.013)  | 27,732 |
| Works less than 35 hours per week                 | 0.20           | −0.026<br>(0.029)   | −0.030<br>(0.032)     | 3,344 | 0.19           | 0.006<br>(0.009)     | 0.017*<br>(0.010)     | 37,297 |
| Works less than 30 hours per week                 | 0.14           | −0.016<br>(0.027)   | −0.025<br>(0.029)     | 3,344 | 0.12           | 0.012<br>(0.008)     | 0.018**<br>(0.009)    | 37,297 |
| Self-employed                                     | 0.04           | −0.006<br>(0.011)   | −0.013<br>(0.016)     | 3,345 | 0.03           | −0.000<br>(0.004)    | −0.002<br>(0.004)     | 37,309 |
| <b>Panel B: Conditional on working</b>            |                |                     |                       |       |                |                      |                       |        |
| Employer-sponsored insurance as a dependent       | 0.14           | −0.072*<br>(0.037)  | −0.078**<br>(0.039)   | 2,556 | 0.20           | −0.045***<br>(0.011) | −0.043***<br>(0.012)  | 27,745 |
| Employer-sponsored insurance through own employer | 0.36           | 0.006<br>(0.043)    | −0.003<br>(0.048)     | 2,556 | 0.43           | 0.014<br>(0.012)     | 0.010<br>(0.013)      | 27,745 |
| Employer-sponsored insurance                      | 0.49           | −0.068<br>(0.043)   | −0.084*<br>(0.048)    | 2,556 | 0.62           | −0.029**<br>(0.011)  | −0.030**<br>(0.013)   | 30,301 |
| Uninsured   | 0.47           | 0.051<br>(0.039)    | 0.053<br>(0.044)      | 2,556 | 0.32           | 0.044***<br>(0.011)  | 0.044***<br>(0.012)   | 30,301 |

NOTE: \* significant at the 0.10 level; \*\* significant at the 0.05 level; \*\*\* significant at the 0.01 level. The sample includes all individuals aged 23–28 in the January 2011 to August 2013 SIPP. The effects of turning 26 are the above-age-26 coefficients from separate regressions of Equation (2) that control for indicators for the year and month of the observation and for marital status and an individual fixed effect. Standard errors are clustered by individual and are shown in parentheses. The estimates in Panel A are for the full sample, while the estimates in Panel B are restricted to the sample of working individuals.

## **Appendix B**

### **Robustness to Specification–WC**

The main analysis estimated discontinuities at age 26 parametrically with two years of data on either side of the cutoff. I now consider the robustness of these results to different bandwidths. I also use a triangular kernel to fit local linear regressions on each side of age 26 and estimate the limit of the expectation function from the left and the right of age 26 to obtain nonparametric estimates of the discontinuity, as in Hahn, Todd, and Van der Klaauw (2001).

Panel A of Table B.1 considers the robustness of the claims results for all injuries and for different types of injuries. The first column is the original estimate. The second column uses 18 months of data on either side of the threshold. The third column uses three years of data on either side of the threshold. Columns 4–6 display nonparametric estimates of the discontinuity in claims at age 26 for each bandwidth. Panel B of Table B.1 displays the corresponding estimates for bills. Table B.2 displays the equivalent results with ICD-9 codes classified by the likelihood they are first reported on weekends.

The results are qualitatively very similar across specifications. When the estimate is statistically significant in the original specification, the estimates are statistically significant in all other specifications as well. In a few cases, estimates that are not statistically significant become statistically significant in the new specifications. They are usually significant at the 10 percent level and are only significant in one of the six specifications. The only exceptions are the estimates for bills in the third quartile of the weekend distribution, which are significant in half of the specifications.

**Table B.1 Sensitivity of Regression Discontinuity Models for All Claims and Medical Bills and for Different Injury Types in Texas, by Age**

|   | Parametric specification |                     |                     | Local linear models |                     |                     |
|---|--------------------------|---------------------|---------------------|---------------------|---------------------|---------------------|
|   | 24–27                    | 24.5–26.5           | 23–28               | 24–27               | 24.5–26.5           | 23–28               |
| <b>Panel A: RD models for log (number of initiated WC claims) in Texas by age at start of claim</b> |                          |                     |                     |                     |                     |                     |
| All claims  | 0.010<br>(0.016)         | –0.019<br>(0.017)   | 0.018<br>(0.015)    | 0.004<br>(0.010)    | 0.002<br>(0.011)    | 0.003<br>(0.009)    |
| <b>by Type</b>  |                          |                     |                     |                     |                     |                     |
| Sprain/strain claims  | 0.060<br>(0.045)         | –0.008<br>(0.049)   | 0.065*<br>(0.036)   | 0.042<br>(0.031)    | 0.036<br>(0.034)    | 0.030<br>(0.025)    |
| Nonsprain/strain claims   | –0.013<br>(0.034)        | –0.043<br>(0.035)   | 0.028<br>(0.057)    | –0.024<br>(0.022)   | –0.024<br>(0.025)   | –0.019<br>(0.017)   |
| Occupational disease claims   | 0.010<br>(0.052)         | –0.026<br>(0.027)   | 0.086*<br>(0.044)   | 0.030<br>(0.040)    | 0.029<br>(0.045)    | 0.032<br>(0.033)    |
| <b>Panel B: RD models for log (number of WC medical bills) in Texas by age at treatment</b>         |                          |                     |                     |                     |                     |                     |
| All bills   | 0.081***<br>(0.026)      | 0.064**<br>(0.027)  | 0.071***<br>(0.020) | 0.066***<br>(0.020) | 0.072***<br>(0.024) | 0.052***<br>(0.016) |
| <b>by Type</b>  |                          |                     |                     |                     |                     |                     |
| Sprain/strain bills   | 0.114***<br>(0.040)      | 0.118***<br>(0.041) | 0.110***<br>(0.037) | 0.091***<br>(0.029) | 0.095***<br>(0.029) | 0.052*<br>(0.027)   |
| Nonsprain/strain bills  | 0.048<br>(0.045)         | –0.000<br>(0.042)   | 0.023<br>(0.032)    | 0.032<br>(0.031)    | 0.036<br>(0.038)    | 0.027<br>(0.023)    |
| Occupational disease bills  | 0.142***<br>(0.037)      | 0.180***<br>(0.039) | 0.169***<br>(0.035) | 0.141***<br>(0.029) | 0.150***<br>(0.032) | 0.133***<br>(0.024) |

NOTE: \* significant at the 0.10 level; \*\* significant at the 0.05 level; \*\*\* significant at the 0.01 level. Columns 1–3 display the above-age-26 coefficients from separate regressions of Equation (3) using 2011–2013 data that include a quadratic polynomial in age in months fully interacted with an indicator for age greater than or equal to 26. Columns 4–6 display nonparametric estimates of the discontinuities at age 26. Columns 1 and 4 use two years of data on either side of the threshold. Columns 2 and 5 use 18 months of data on either side of the threshold. Columns 3 and 6 use three years of data on either side of the threshold. Robust standard errors are shown in parentheses. In Panel A, the dependent variables are the log of the number of claims. In Panel B, the dependent variables are the log of the number of bills. There are 60,399 WC claims for young adults aged 23–27, 45,563 for young adults aged 24.5–26.5, and 90,256 for young adults aged 23–28. There are 221,572 WC bills for young adults aged 23–27, 167,040 for young adults aged 24.5–26.5, and 333,811 for young adults aged 23–28. SOURCE: The 2011–2013 Texas WC medical database.

**Table B.2 Sensitivity of Regression Discontinuity Models for Claims and Medical Bills by Weekend Likelihood in Texas, by Age**

|   | Parametric specification |                     |                     | Local linear models |                     |                     |
|---|--------------------------|---------------------|---------------------|---------------------|---------------------|---------------------|
|   | 24–27                    | 24.5–26.5           | 23–28               | 24–27               | 24.5–26.5           | 23–28               |
| <b>Panel A: RD models for log (number of initiated WC claims) in Texas by age at start of claim</b> |                          |                     |                     |                     |                     |                     |
| Bottom quartile   | 0.257**<br>(0.098)       | 0.248**<br>(0.119)  | 0.228***<br>(0.076) | 0.198***<br>(0.070) | 0.225***<br>(0.084) | 0.139**<br>(0.057)  |
| 26th to 50th percentile   | –0.038<br>(0.033)        | –0.071*<br>(0.040)  | 0.027<br>(0.031)    | –0.013<br>(0.023)   | –0.026<br>(0.026)   | 0.008<br>(0.019)    |
| 51st to 75th percentile   | –0.008<br>(0.032)        | –0.024<br>(0.036)   | –0.003<br>(0.028)   | –0.007<br>(0.025)   | –0.006<br>(0.027)   | –0.007<br>(0.022)   |
| Top quartile  | –0.001<br>(0.043)        | –0.047<br>(0.041)   | –0.032<br>(0.035)   | –0.027<br>(0.026)   | –0.030<br>(0.029)   | –0.030<br>(0.021)   |
| <b>Panel B: RD models for log (number of WC medical bills) in Texas by age at treatment</b>         |                          |                     |                     |                     |                     |                     |
| Bottom quartile   | 0.230***<br>(0.069)      | 0.263***<br>(0.080) | 0.235***<br>(0.065) | 0.208***<br>(0.048) | 0.224***<br>(0.055) | 0.117***<br>(0.042) |
| 26th to 50th percentile   | 0.142***<br>(0.052)      | 0.053*<br>(0.031)   | 0.078**<br>(0.036)  | 0.075**<br>(0.033)  | 0.084**<br>(0.037)  | 0.049*<br>(0.028)   |
| 51st to 75th percentile   | 0.005<br>(0.053)         | 0.017<br>(0.054)    | 0.101*<br>(0.051)   | 0.073*<br>(0.043)   | 0.053<br>(0.052)    | 0.103***<br>(0.033) |
| Top quartile  | 0.038<br>(0.051)         | 0.045<br>(0.062)    | –0.016<br>(0.041)   | 0.005<br>(0.039)    | 0.023<br>(0.047)    | –0.010<br>(0.031)   |

NOTE: \* significant at the 0.10 level; \*\* significant at the 0.05 level; \*\*\* significant at the 0.01 level. Claims are classified into quartiles by the likelihood that the ICD-9 code was first reported on the weekend as described in the text. Columns 1 through 3 display the above-age-26 coefficients from separate regressions of Equation (3) using 2011–2013 data that include a quadratic polynomial in age in months fully interacted with an indicator for age greater than or equal to 26. Columns 4–6 display nonparametric estimates of the discontinuities at age 26. Columns 1 and 4 use two years of data on either side of the threshold. Columns 2 and 5 use 18 months of data on either side of the threshold. Columns 3 and 6 use three years of data on either side of the threshold. Robust standard errors are shown in parentheses. In Panel A, the dependent variables are the log of the number of claims. In Panel B, the dependent variables are the log of the number of bills. There are 60,399 WC claims for young adults aged 23–27, 45,563 for young adults aged 24.5–26.5, and 90,256 for young adults aged 23–28. There are 221,572 WC bills for young adults aged 23–27, 167,040 for young adults aged 24.5–26.5, and 333,811 for young adults aged 23–28. SOURCE: The 2011–2013 Texas WC medical database.

## **Appendix C**

### **Placebo Analysis at Other Ages**

Tables 7 and 8 displayed the above-age-26 coefficients from Equation (3) using 2005–2007 data. None of the placebo coefficients were statistically different from zero. As another placebo test, I now show estimates of discontinuities at age 30. Figures C.1 and C.2 show the age profiles around 30. The results for bills and claims are shown in Table C.1. One of the coefficients is significant at the 10 percent level, while another is significant at the 5 percent level. This amounts to 12.5 percent of the estimates for discontinuities at 30 being significant at the 10 percent level or higher, which is not far from the 10 percent we would expect.

As these placebo tests are arbitrary, I also repeat the analysis at 10 consecutive ages starting with age 28 using 2011–2013 data. The coefficients are significant in 20.6 percent of these placebo regressions, which suggests that the RD strategy overrejects the null hypothesis of no effect. I put the placebo estimates through the same sensitivity tests that the main estimates were all robust to. Only 5.6 percent of the placebo estimates are robust to these minor alterations of the empirical strategy. The placebo estimates being much more sensitive to the specification than the original estimates suggests that it is unlikely that the results presented in the paper and in Appendix B are driven by chance.

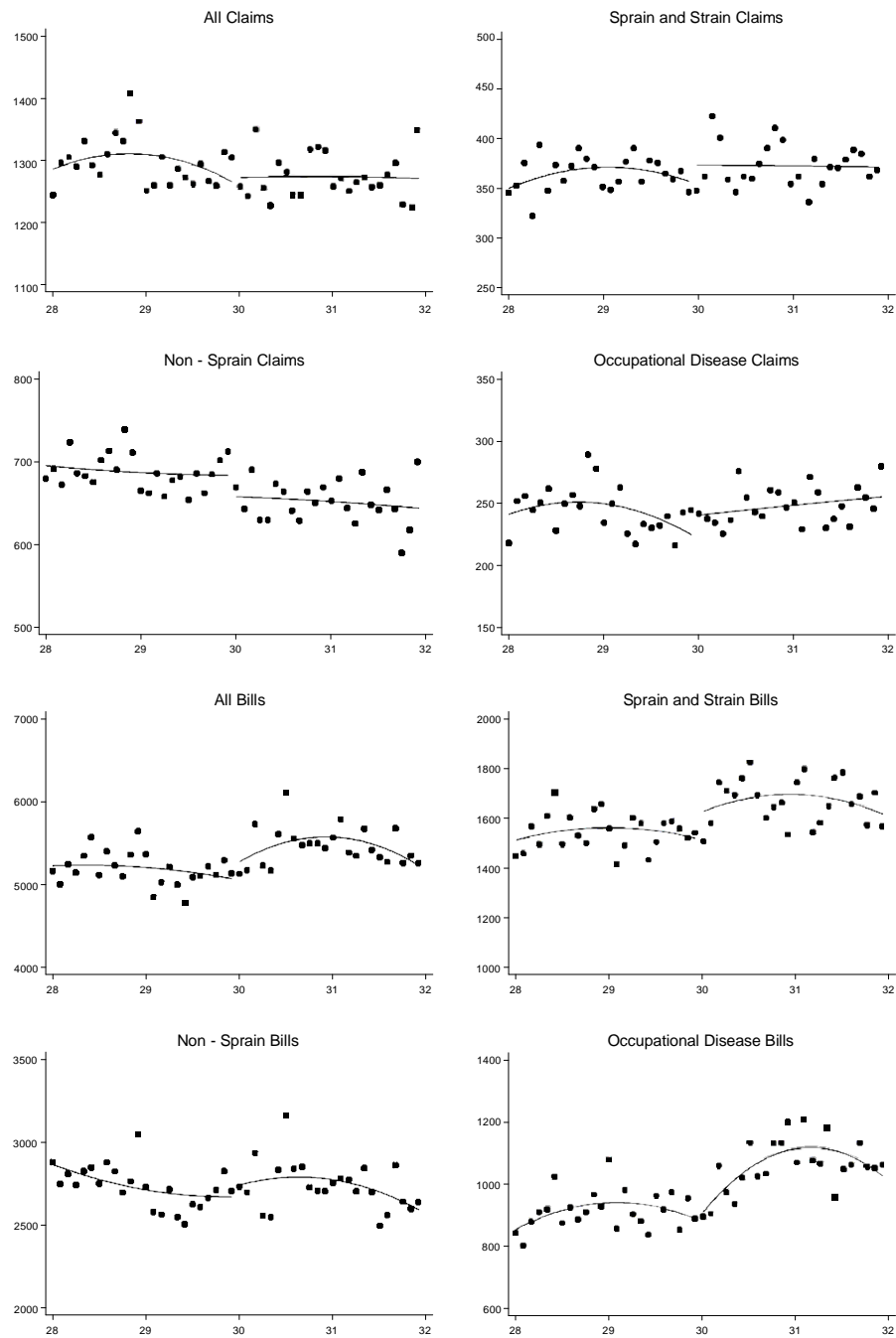
**Table C.1 Placebo Analysis at Age 30 for Regression Discontinuity Models Claims and Medical Bills in Texas, by Age**

| <b>Panel A: RD models for log (number of initiated WC claims) in Texas by age at start of claim</b> |                   | <b>Panel C: RD models for log (number of initiated WC claims by weekend likelihood) in Texas by age at start of claim</b> |                    |
|---|-------------------|---|--------------------|
| All claims  | 0.008<br>(0.026)  | Bottom quartile   | 0.185**<br>(0.078) |
| <b>by Type</b>  |                   |   |                    |
| Sprain/strain claims  | 0.050<br>(0.047)  | 26th to 50th percentile   | 0.025<br>(0.036)   |
| Nonsprain/strain claims   | -0.041<br>(0.030) | 51st to 75th percentile   | 0.051<br>(0.042)   |
| Occupational disease claims   | 0.079<br>(0.054)  | Top quartile  | -0.009<br>(0.051)  |
| <b>Panel B: RD models for log (number of WC medical bills) in Texas by age at treatment</b>         |                   | <b>Panel D: RD models for log (number of WC medical bills by weekend likelihood) in Texas by age at treatment</b>         |                    |
| All Bills   | 0.040<br>(0.030)  | Bottom quartile   | 0.070<br>(0.048)   |
| <b>by Type</b>  |                   |   |                    |
| Sprain/strain bills   | 0.074*<br>(0.039) | 26th to 50th percentile   | 0.065<br>(0.041)   |
| Nonsprain/strain bills  | 0.022<br>(0.035)  | 51st to 75th percentile   | -0.043<br>(0.036)  |
| Occupational disease bills  | 0.041<br>(0.045)  | Top quartile  | -0.031<br>(0.043)  |

NOTE: \* significant at the 0.10 level; \*\* significant at the 0.05 level; \*\*\* significant at the 0.01 level. For Panels C and D, claims are classified into quartiles by the likelihood that the ICD-9 code was first reported on the weekend as described in the text. The estimates are the above-age-26 coefficients from separate regressions of Equation (3) using 2011–2013 data that includes a quadratic polynomial in age in months fully interacted with an indicator for age greater than or equal to 26. Robust standard errors are shown in parentheses. In panels A and C, the dependent variables are the log of the number of claims. In Panel B and D, the dependent variables are the log of the number of bills. There are 79,650 WC claims and 324,408 WC bills for people aged 28–31 from 2011 to 2013.

SOURCE: The Texas WC medical database.

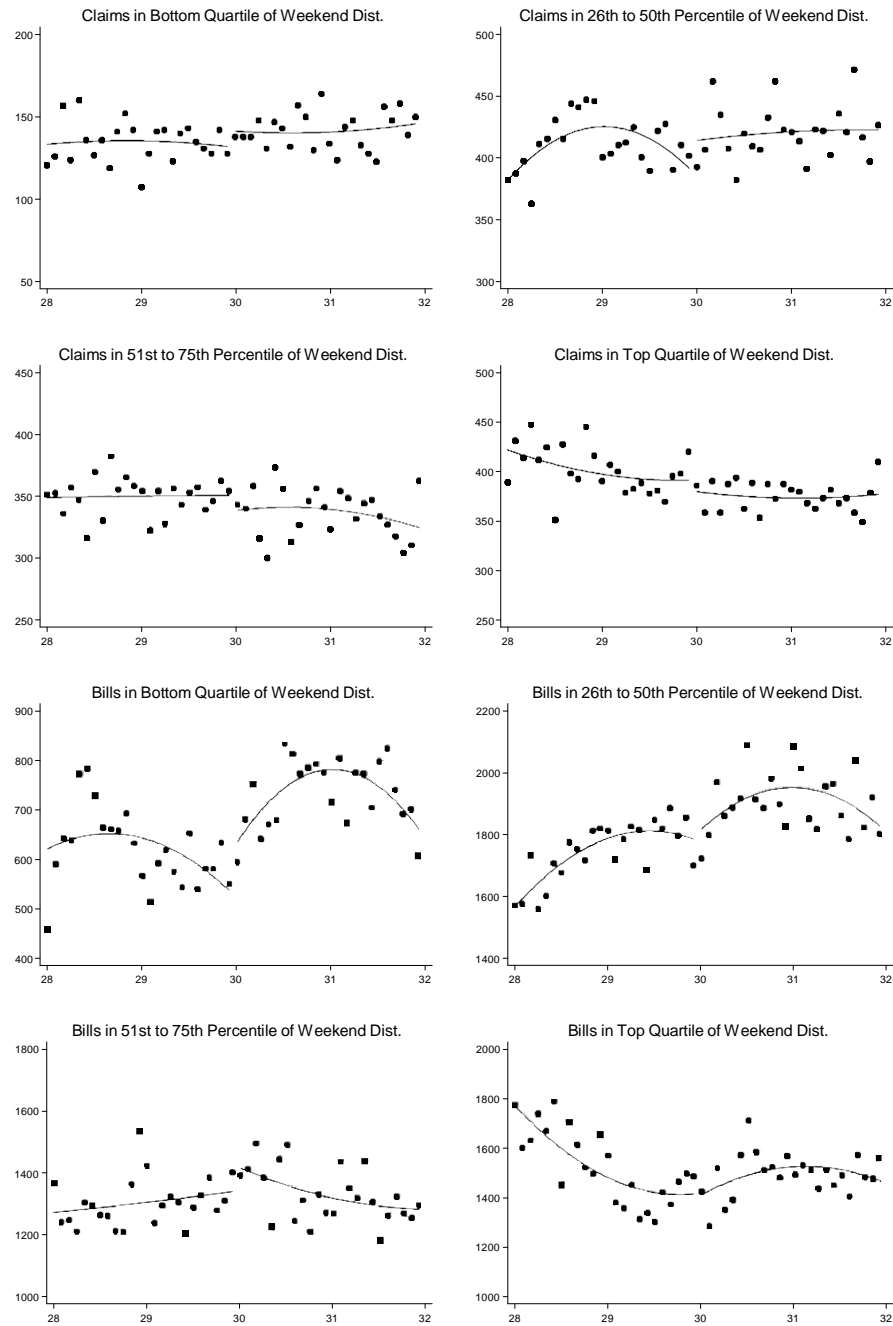
**Figure C.1 Counts of WC Claims and Medical Bills in Texas by Age from 2011 to 2013**



NOTE: The points represent counts of WC claims or medical bills at each age in months.  
SOURCE: The Texas WC medical database.



**Figure C.2 Counts of WC Claims and Medical Bills in Texas by Age from 2011 to 2013 by Weekend Likelihood**



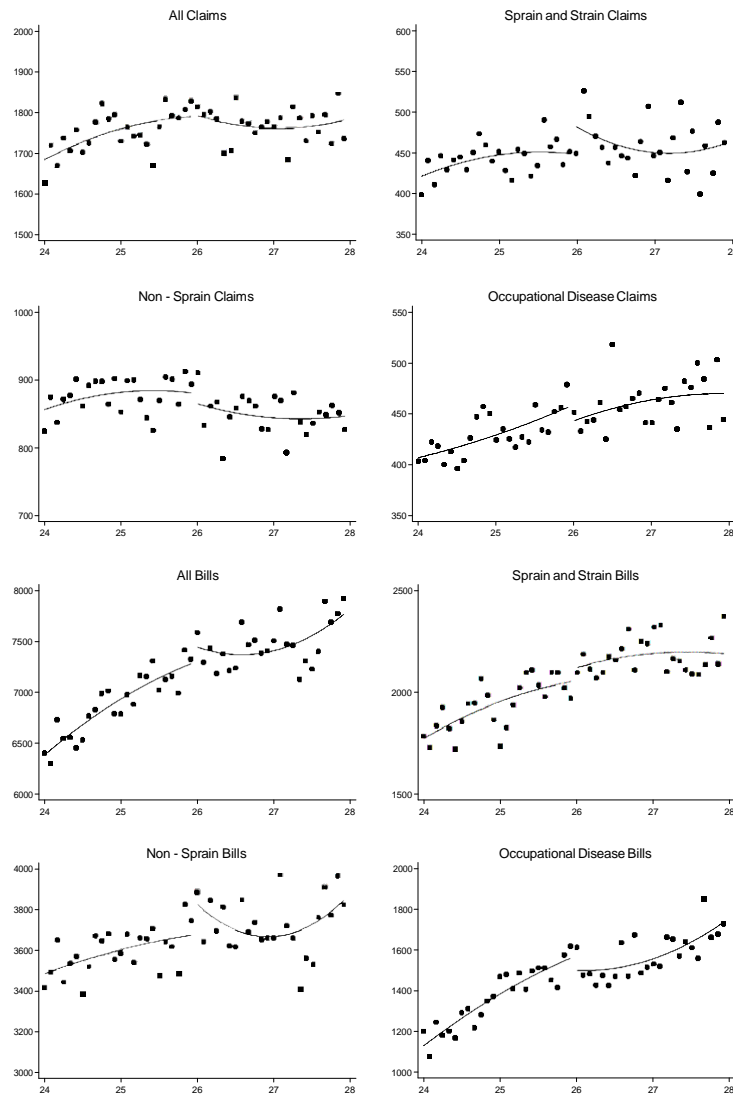
NOTE: Claims are classified into quartiles by the likelihood that the ICD-9 code was first reported on the weekend as described in the text. The points represent counts of WC claims or medical bills in each quartile of the weekend likelihood distribution at each age in months.

SOURCE: The Texas WC medical database.

## Appendix D

### Age Profiles for 2005 to 2007

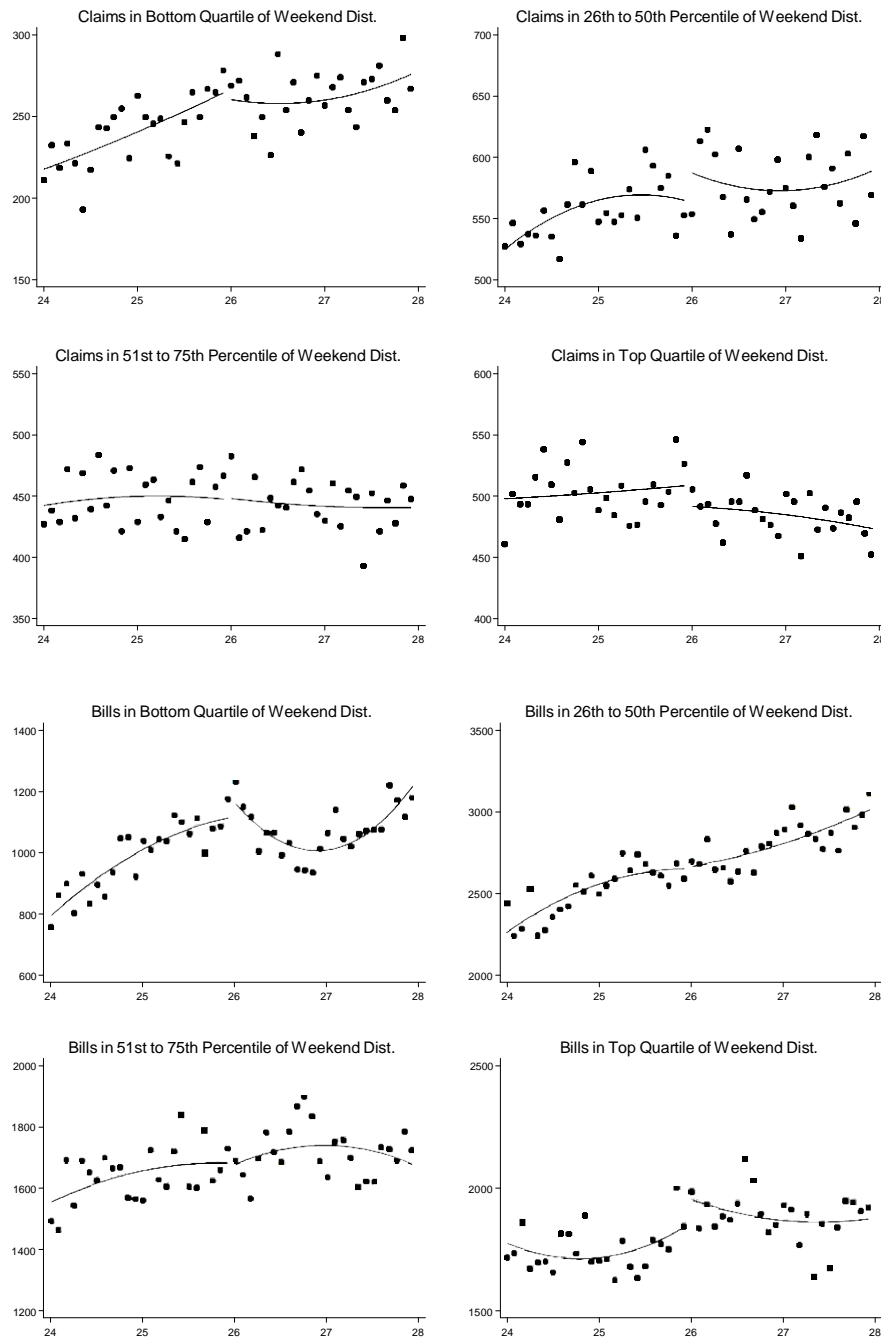
Figure D.1 Counts of WC Claims and Medical Bills in Texas by Age from 2005 to 2007



NOTE: The points represent counts of WC claims or medical bills at each age in months.

SOURCE: The Texas WC medical database.

**Figure D.2 Counts of WC Claims and Medical Bills in Texas by Age from 2005 to 2007 by Weekend Likelihood**



NOTE: Claims are classified into quartiles by the likelihood that the ICD-9 code was first reported on the weekend as described in the text. The points represent counts of WC claims or medical bills in each quartile of the weekend likelihood distribution at each age in months.

SOURCE: The Texas WC medical database.

## Appendix E

### Texas Employment by Age and Year from 2011 to 2013 ACS

Figure E.1 Texas Employment by Age and Year from 2011 to 2013 ACS

