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Abstract

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Disciplines

Benefits and Compensation | Labor and Employment Law | Workers' Compensation Law

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This article was published in a journal that has since ceased to be produced.

Rate Regulation, Safety Incentives, and Loss Growth in Workers' Compensation Insurance

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Rate Regulation, Safety Incentives, and Loss Growth in Workers' Compensation Insurance*

I. Introduction

Well-designed tort liability and workers' compensation systems and well-functioning insurance markets encourage efficient safety by internalizing expected injury costs. Insurance markets with prices that vary in relation to expected claim costs allow potential injurers to reduce risk while still providing incentives for loss control. Because insurance prices cannot perfectly reflect the expected cost of risky behavior, however, incentives for loss control are weakened by moral hazard (Shavell 1982). Limited wealth and limited liability also dilute incentives for loss control. This problem can be reduced but not eliminated by compulsory insurance rules (Shavell 1986).

An issue that has received comparatively less attention is that insurance price regulation also can distort incentives for safety. Both the adoption and administration of price regulation are influenced by political pressure. This pressure can lead to temporary or even chronic differences between approved rates and the expected costs of providing coverage for many buyers. Natural in-

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surer supply responses to price constraints can be substantially constrained, at least in the short run, by mandates that insurers provide coverage to all buyers at regulated prices, often on a pooled basis through residual market mechanisms. Material deviations between regulated rates and costs are especially likely during periods of rapid loss growth due to increased consumer pressure for regulatory constraints (e.g., Baumol 1991), especially when compulsory coverage requirements reduce demand elasticity.

If severe regulatory lag in the presence of rapid loss growth drives a large wedge between prices and expected costs, consequences usually include a wealth transfer from producers to consumers (e.g., due to loss of insurer quasi rents; see Harrington [1992]) and some degree of ex ante cross-subsidy from low-risk buyers to high-risk consumers (see Kwon and Grace 1996). The efficiency consequences of attendant distortions in incentives for loss control may be less obvious. But if rate regulation suppresses rates in relation to expected losses for some consumers, the resulting distortions in incentives can be expected to increase loss growth. The implication is that price regulation that attempts to shield consumers from the full effects of exogenous cost surges will increase loss growth and thus ultimately be self-defeating.

This study investigates the relationship between insurance rate regulation, inflationary cost surges, and incentives for loss control using state-level data on workers' compensation insurance loss growth for 24 states during 1984–90. Many state workers' compensation insurance markets experienced large cost surges during this period (see fig. 1).[1] Regulators often responded by denying rate increases or by approving increases materially lower than initially requested by insurers. The market share of the workers' compensation insurance residual market grew rapidly in many states.[2] We exploit cross-state and time-series variation in loss growth and regulatory responses to test the prediction that rate suppression increased loss growth. Our regressions indicate a positive and statistically reliable relationship between loss growth and lagged measures of regulatory price constraints, suggesting that rate regulation led to some increase in the frequency and/or severity of employee injuries.[3]

Our loss growth models incorporate fixed time and state effects and condition on several variables that could affect loss growth at the state level. The use of fixed effects and lagged measures of regulation should

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1. For example, benefit costs per $100 of payroll increased from $0.95 in 1978 to $1.56 in 1989, an average annual rate of increase of 4.2% for 1978–84 and 6.2% for 1984–89 (Klein, Nordman, and Fritz 1993, pp. 79 ff).

2. Workers' compensation insurance residual markets are briefly described in Section II below.

3. The finding that rate suppression aggravated loss growth in part may reflect increased claims reporting or delays in returning to work (e.g., Butler 1994).
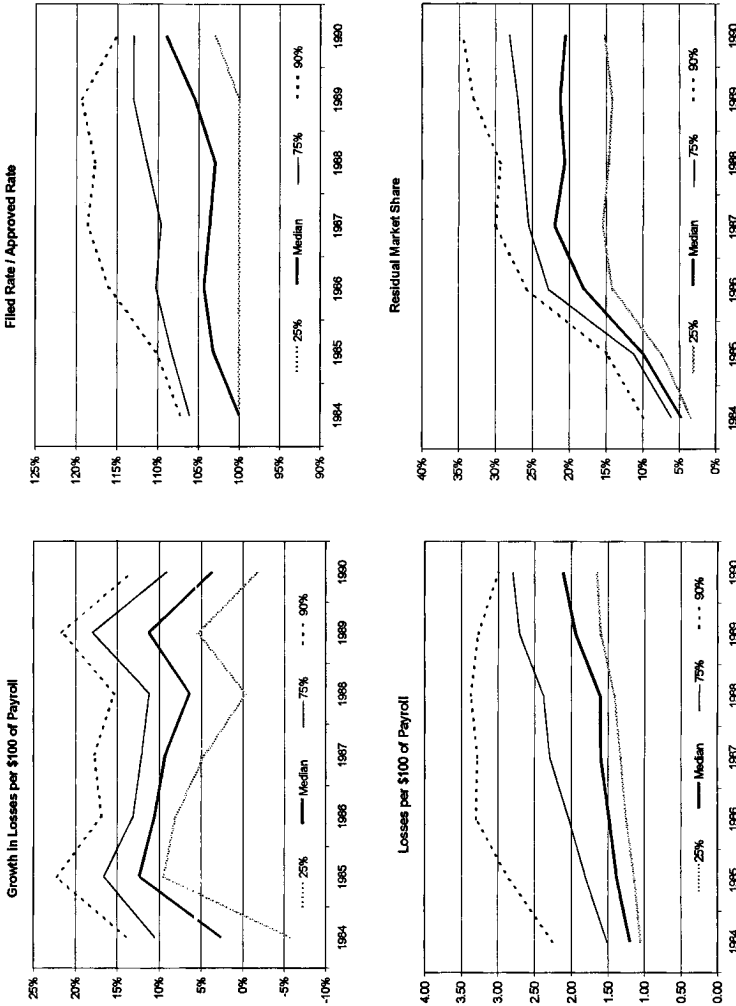
FIG. 1—Losses, rate regulation, and residual markets in workers' compensation insurance: 1984–90, selected percentile values for 24 states.

substantially allay concern that the results could reflect correlation between omitted cost drivers and our measures of regulation. In addition, the estimated positive relationship between loss growth and regulation persists when we condition on estimated growth in self-insurance to control for adverse selection in response to rate suppression.

The possibility that rate regulation may lead to higher loss growth is discussed by Danzon (1992), Harrington (1992), Kramer (1992), and Klein, Nordman, and Fritz (1993). Kaestner and Carroll (1997) provide evidence of larger (one-digit standard industrial classification) injury rates in states with some form of statutory rate ''deregulation'' during 1983–88, arguing that regulation pushes up prices (see Carroll and Kaestner 1995) and therefore encourages loss control. This explanation is inconsistent with evidence of rate suppression during the 1980s (e.g., see fig. 1). Their estimates instead might reflect lower injury rates at the time of the statutory change for the few states in their sample that relaxed rate regulatory statutes (in each case at the beginning of their sample period), and their injury rate models do not include fixed state effects.[4] Barkume and Ruser (1998) test for a relationship between the statutory type of rate regulation and levels of workers' compensation costs and injury rates over longer time periods, suggesting that statutory deregulation was associated with a reduction in injury costs and in the margin between premiums and injury costs. These combined effects are puzzling and would not be expected in an environment characterized by strong cost surges and rate suppression. More generally, market-wide rate suppression cannot be a long-run equilibrium, and the effects of rate regulation will not be time invariant.

Our analysis focuses on rate regulation's possible effects on loss growth during a period characterized by a strong cost surge and attendant failure of regulators to permit premium growth commensurate with expected loss growth. In contrast to previous studies, we employ lagged measures of rate suppression that reflect differences among states regardless of the type of statute: (1) the difference between rates filed by insurers and approved by regulators and (2) residual market share (the market share of the assigned risk plan). This approach should provide a more powerful test of the null hypothesis that rate regulation did not affect loss growth during a period when most states had some type of prior-approval rate regulation and few states changed their laws. Our models include a variety of control variables, including actuarial predictions of the effects of changes in benefit laws on loss growth. As noted, fixed time and state effects also are included to help control for omitted cost drivers.[5] Moreover, the analysis of growth rates should

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4. Kaestner and Carroll (1997) note that ''including state dummy variables in the model makes it difficult to identify significant effects'' (p. 643).

5. Kaestner and Carroll (1997) report a significantly positive relation between injury rates and contemporaneous residual market share. They suggest that this could indicate

be much less vulnerable to omitted variable bias than analysis of cost levels.

Beginning around 1990, numerous states adopted workers' compensation reform legislation to reduce losses, and many states changed their systems of voluntary and residual market price regulation. These changes were associated with slower loss growth, improved insurer financial results, and declining residual markets and residual-market deficits. By the mid-1990s, requests for rate increases largely evaporated and rates began to fall in some states. By year-end 1998, the countrywide residual market share was comparable to the low level of the early 1980s. Recent data, however, suggest renewed growth in claim costs and deterioration in insurer underwriting margins. Whether regulatory responses to a strong cost surge would mimic the 1980s' experience is an open question.

Our main results imply that insurance rate regulatory systems that suppress rates in relation to costs have the inefficient and potentially self-defeating side effect of increasing loss growth. A related implication is that price regulation can weaken specific insurance pricing mechanisms, such as experience rating, that align rates with expected claim costs and thus provide significant incentives for loss control. These cost-increasing effects of rate regulation raise serious policy concerns, further undermining the weak case for price regulation in competitively structured insurance markets. Our evidence is consistent with the prediction that deregulation of rates will improve incentives for efficient loss control.

Section II provides a brief overview of workers' compensation insurance markets, rate regulation, and related research. Section III develops our hypotheses concerning the effects of price regulation on loss growth. Section IV describes the data and methodology, and results follow in Section V. Section VI concludes.

II. Background

The 1980s Cost Surge and Regulatory Responses

Workers' compensation pays for workers' medical expenses and wage loss arising out of work-related injuries and diseases. Employers must purchase insurance or provide proof of financial responsibility in order to self-insure. Regulation of workers' compensation insurance rates in most states historically required all insurers to use the same rates, rating

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reduced safety or that states with more risky jobs could have higher residual market shares. We use lagged residual market share, and our analysis of growth rates and inclusion of fixed state effects should largely control for state differences in job risk that are unrelated to regulation.

classes, and experience rating plans, with rate filings developed on behalf of the industry by rate advisory organizations. These organizations collect loss and expense data from the industry and use these data to develop ''advisory rates'' (or, more recently, ''prospective loss costs'') for hundreds of rate classes. The National Council on Compensation Insurance (NCCI) currently serves in this capacity in a majority of states.

There typically are several hundred rate classes in a state to reflect differences in industry and type of business. Most states require experience rating with the exception of small employers. Class rates for an occupational class are modified upward or downward based on the employer's experience in a prior period compared to the class average. Beginning in the 1970s, some states permitted insurers to file for deviations from rates filed by advisory organizations. In the 1970s and 1980s, some states began to permit schedule rating (except for small employers), which allows an insurer to modify the class rate based on the underwriter's evaluation of the employer's risk of loss.

Worker's compensation insurance residual markets generally require insurers to provide coverage at a regulated rate to applicants who presumably cannot obtain voluntary coverage. The usual form of residual market assigns policyholders to designated servicing carriers, which issue policies and pay claims in exchange for fees without directly bearing underwriting risk on assigned business. Operating deficits for the residual market are apportioned among all workers' compensation insurers in proportion to their share of workers' compensation insurance voluntary market premiums in the state.

Before the 1980s cost surge, regulated rates generally were set high enough to encourage insurers to cover most employers voluntarily. Competitive pressures led insurers to pay dividends to inframarginal employers; service quality was another dimension of competitive strategy. Insurers also competed with rate deviations and schedule rating where permitted. The 1980s cost surge produced considerable turmoil. The increase in workers' compensation losses relative to payroll is partly explained by increases in medical care costs in excess of general wage inflation, benefit growth, and changing workplace demographics (e.g., Butler 1994). Large differences among states in the level and growth of losses suggest that other state-specific factors also were important.

Growth in workers' compensation insurance losses during the 1980s was accompanied by deteriorating financial results for insurers, who argued that many state regulators refused to allow rate increases commensurate with expected loss growth. Consistent with binding regulatory constraints on rate increases in the presence of rising costs, the countrywide size of the workers' compensation insurance residual mar-

ket increased sharply, with the residual market share of premiums growing to over 50% in a number of states (see fig. 1).[6]

Growth in residual markets and expected operating deficits on residual market business increased voluntary market rate levels needed by insurers to cover expected costs for the overall workers' compensation insurance market in a given state. In the long run, expected residual market deficits require a cross subsidy from the voluntary to the residual market to prevent widespread exit, and linkage between deficits and voluntary market supply makes chronic cross subsidies feasible. In the short run, regulatory constraints on voluntary market rates may partially shift the incidence of residual market subsidies to insurer capital. Exit will be slow, and insurers are therefore vulnerable to expropriation, because exit requires insurers to write off much of their investment in a given state (Harrington 1992). Higher voluntary market rates to finance residual market deficits encourage more low-risk businesses to self-insure, further reducing the size of the voluntary market. Escalating growth in the residual market and inability to shift residual market deficits to a shrinking voluntary market in the 1980s caused a virtual collapse of the workers' compensation insurance market in a few states (e.g., Maine and later Rhode Island).

Related Work

As we explained in the introduction, Kaestner and Carroll (1997) and Barkume and Ruser (1998) provide mixed evidence of whether the statutory form of regulation is related to levels of workers' compensation injuries and costs. Several studies examine the relationship between the type of workers' compensation insurance regulation (prior approval, loss-cost systems, competitive rating) and premium levels and premium-cost margins to test hypotheses of industry versus consumer capture of the regulatory process (e.g., Appel, McMurray, and Mulvaney 1992; Klein et al. 1993; Carroll and Kaestner 1995; Schmidle 1995). Conclusions about the effects of rate regulation differ, presumably in part because of heterogeneity in the types of regulation and because the effects of a particular type of regulation differ depending on its implementation and the cost environment during the sample period.

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6. Reduction in regulated rate levels relative to the average cost of providing coverage for a class during the 1980s also led to lower dividend payments and fewer downward deviations and schedule rating credits. Concern with affordability of coverage to employers also played a role in the decision by some states to shift to prospective loss cost systems, in which the NCCI files prospective loss costs only on behalf of insurers, with individual insurers filing their own profit and expense factors. Loss cost systems often were advocated by regulators and other parties as a means of increasing competition, or at least fostering the impression of increased competition by reducing reliance on advisory organization expense and profit loadings in filed rates.

These studies generally use a broad definition of competitive rating that encompasses states that have adopted loss-cost systems (see n. 6) or allow insurers to deviate from rates filed by rate advisory organizations, even though loss costs or individual insurer rates remain subject to prior regulatory approval.[7] By the early 1990s, however, few states had true competitive rating systems that expressly relied on competition to control rates and permitted insurers to alter rates without close regulatory scrutiny. Furthermore, even if voluntary market rates are not regulated, regulation of residual market rates can suppress rates for many employers.

Several studies provide evidence that higher workers' compensation benefits increase the ratio of loss costs to payroll, particularly in firms subject to little or no experience rating of premiums (e.g., Ruser 1985, 1993; Chelius and Kavanaugh 1988; Fortin 1992; and Meyer, Viscusi, and Durbin 1995). The results of several studies suggest that workers' compensation insurance experience rating increases employer incentives to control costs (Ruser 1985; Bruce and Atkins 1993; and Hyatt and Kralj 1995). An implication of this is that suppression of rate levels to which experience rating is applied will reduce incentives for safety (see below).

Butler (1994) argues that workers' compensation loss growth largely reflects statutory benefit increases, declines in waiting periods, and changes in workforce demography. He provides evidence that cost growth in part reflects higher propensity to report claims rather than a change in workers' or firms' risk-taking behavior (also see Krueger 1990; and Butler and Worrall 1991). We noted earlier that increased claims reporting might explain part of any increase in loss growth due to rate suppression.

Carroll (1994) analyzes price regulation's effects on the estimated proportion of total workers' compensation paid losses that is self-insured across states during 1980–87. The study provides some evidence that higher ratios of premiums to losses increase the self-insurance proportion, implying that suppression of the statewide premium-loss margin might discourage self-insurance.[8] Butler and Worrall (1993) analyze the proportion of paid losses represented by self-insurance for selected years during the period 1954–82 as a function of the size distribution of firms. They provide evidence that the probability of self-insuring is a convex function of firm size. They do not explore possible effects of price regulation.

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7. Barkume and Ruser (1998) distinguish states that retained some type of prior approval from those that did not.

8. No relationship is found between residual market share and the self-insurance variable.

III.   Hypotheses

Rate suppression may inflate loss growth through several channels that are not mutually exclusive. Workers' compensation involves multiparty accidents subject to strict statutory liability for employers. Optimal loss control requires care by employers, employees, and insurers. Insurance that is accurately experience rated does not distort incentives for optimal investment in injury prevention (Shavell 1982). By preventing insurers from charging rates that reflect expected loss costs plus a competitive expense and profit margin, rate suppression is expected to dull loss control incentives for all three parties.

Subsidies to High-Risk Behavior

Regulation may constrain the incentive effects of experience rating in two main ways. First, it could limit percentage-experience-modification factors applied to class rates. Second, it could reduce the class rates to which the factors are applied, thus producing a smaller absolute debit for relatively poor experience or absolute credit for good experience. Because experience-rating systems generally have been uniform across the states for which the NCCI serves as a rating advisory organization and changed little until the 1990s, this second effect may be more likely than the first.

   Regulatory constraints on experience-rating debits and credits reduce employer incentives to invest in loss control and to require such investments by employees. The firm bears the costs of loss control by employers and employees, but benefits are diffused across other insured firms unless premiums are experience rated to reflect the change in expected losses. Postinjury moral hazard—overuse of medical care, delay in return to work—also will likely increase in response to rate suppression, leading to increased duration and cost per claim.

   In addition to effects related to experience rating, the tendency of rate suppression to lower rates for the highest risks in each class and increase rates or lower dividends for better risks also acts as a subsidy to high-risk activity. The prevalence of relatively high-risk firms within a state will tend to increase over time, and incentives for safety, conditional on the type of activity, are undermined. While incentives for safety might possibly increase for lower-risk firms that end up financing part or all of the subsidy to high-risk firms, the net effect is plausibly to reduce safety on average.

Possible Effects on Insurer Behavior

In some cases, regulation might allow a large part of increases in expected losses to pass through but constrain the markup for insurer expenses and return on equity to a level that is inadequate to cover optimal insurer investment in loss control. If so, insurers might reduce expendi-

tures on loss control, even if higher losses result, because part of the increase in expected losses is passed through whereas the loss-control expense comes entirely out of profit.[9] Whether rate regulation in practice places greater constraints on expense and profit loadings than on allowances for expected losses is uncertain. In the 1980s, many regulators failed to approve requests for rate increases based on projected trends in losses as well as requests for expense and profit loadings. To the extent that rate regulation is neutral in disallowing either expenses or loss costs, the distorting effects of rate suppression on insurer investments in loss control are reduced. Even in this case, however, some increase in losses due to reduced investments in loss control by insurers might arise because part of any increase in losses will be shared with employers through the operation of experience rating.[10]

Testable Implications and the Effects of Self-Insurance

The preceding discussion implies that rate suppression will increase loss growth by distorting the incentives of employers, employees, and insurers. We test this key prediction using data on insured loss growth at the state aggregate level and two measures of rate suppression, controlling for exogenous factors that could influence loss growth. Given available data and measures of rate suppression, it is not feasible for us to distinguish the possible effects of subsidies on employer, employee, and insurer behavior. The preceding discussion emphasizes rate regulation's possible effects on loss growth due to distorting effects on employers, employees, and insurers. Growth in insured losses may also be affected if regulation alters the proportion and mix of firms that self-

9. The incidence of any suboptimal investments in loss control by insurers is probably divided between employers and employees. Both theory and evidence indicate that the costs of workers' compensation premiums are borne by employees in the long run (see, e.g., Viscusi and Moore 1987; Gruber and Krueger 1990). Reductions in insurer loss control efforts that lead to higher costs per claim, and higher total premiums may be passed on to employer policyholders in the form of lower dividend payments and/or higher renewal premiums, including lower experience-rating credits.

10. As noted earlier, most residual markets are administered by assigning employers to a number of servicing carriers that receive a fixed proportion of premiums to service the policy (determine the premium, issue the policy, provide loss control advice, settle claims, etc.). Residual market deficits are then allocated to insurers in the state in proportion to their premium volume in the voluntary market. With this structure, the costs of most forms of loss control are fully internalized to the servicing carrier, but it ultimately bears only part of any resulting reduction in claim costs. A possible result is that incentives for investment in loss control (and perhaps premium collection) will be suboptimal in the residual market unless counteracted by other factors. The disincentive to invest in loss control compared to voluntary market coverage will be mitigated if not eliminated by a number of influences, including: (1) the extent to which it is costly for servicing carriers to distinguish residual market employers and treat them differently from voluntary market employers, (2) the magnitude of the voluntary market share of servicing carriers and thus their participation in deficits, and (3) the degree of regulatory monitoring and monitoring by other insurers through the NCCI and other means.

insure as opposed to buying commercial insurance. Suppression of the average margin between rates and costs might be expected to reduce self-insurance (e.g., Carroll 1994), at least in the short run before the possible cost-increasing effects of rate suppression reduce the attractiveness of commercial insurance despite the suppression of rates compared to costs. Any short-run reduction in self-insurance would tend to increase growth in insured losses relative to payroll if higher-risk employers are most likely to buy commercial insurance in response to rate constraints.

But in the long run, rate suppression is both likely to increase the percent of total payroll that is self-insured and increase the average risk level in the commercially insured sector. Regulation-induced cross subsidies from low- to high-risk firms encourage more low-risk firms to self-insure, while subsidies to high-risk firms reduce their incentive to self-insure. The resulting self-selection will increase the ratio of insured losses to insured payroll. There was substantial concern during the late 1980s and into the 1990s that growth in self-insurance was reducing the voluntary market premium base for assessing residual market deficits.

One method of controlling for loss growth due to changes in self-insurance would be to analyze data for a common sample of insured employers over time. Because we were unable to obtain such data, we use the indirect approach of controlling for growth in the estimated proportion of payroll that is self-insured.

IV. Methodology and Data

Empirical Framework

We assume that log growth in the expected loss per $100 of payroll, $E(G_{jt})$, depends on fixed time and state effects, regulation, and exogenous time and state varying factors:

$$E(G_{jt}) \equiv \ln\left(\frac{E(L_{jt})/P_{jt}}{L_{jt-1}/P_{jt-1}}\right) = \alpha_j + \alpha_t + \beta'X_{jt} + \gamma'R_{jt} + u_{jt}, \quad (1)$$

where for state j and year t, L_{jt} is insured losses, P_{jt} is insured payroll, α_j represents a fixed state effect, α_t represents a fixed time effect, X_{jt} is a vector of exogenous factors that influence expected loss growth, R_{jt} is a vector of variables measuring price regulation, β and γ are parameter vectors, and u_{jt} is a disturbance term that reflects idiosyncratic variation in expected loss growth.

Realized loss growth can be expressed as $G_{jt} \equiv (1 + \epsilon_{jt}) E(G_{jt})$,

where ϵ_{jt} is the percentage forecast error. Because $\ln(1 + \epsilon_{jt}) \approx \epsilon_{jt}$ for relatively small ϵ_{jt}, realized cost growth can be expressed as:

$$G_{jt} \approx \alpha_j + \alpha_t + \beta'X_{jt} + \gamma'R_{jt} + u_{jt} + \epsilon_{jt}. \qquad (2)$$

This equation is the basis for our empirical tests of the null hypothesis that realized loss growth is unrelated to our measures of rate suppression (i.e., $\gamma = 0$), versus the alternative hypothesis that rate suppression increases loss growth.

Data and Sample

We obtained by-state data from the NCCI for a variety of characteristics of the workers' compensation market, including residual market premiums and share of total premiums; filed and approved rate changes for the voluntary and residual markets; insured payroll, indemnity, and medical incurred losses (on a policy-year basis); average weekly wages; and the predicted growth in loss costs due solely to changes in benefit provisions.[11] The loss data are NCCI projections of ultimate costs using first report data and first-to-ultimate report loss development factors based on historical experience. The last available year for the payroll and loss data was 1990. The first year of filed and approved rate data was 1983. Because we use one-year lags of many of the variables in the regression models (see below), we analyze loss growth for policy periods from 1984 through 1990. This period encompasses the 1980s cost surge and precedes the subsequent enactment in many states of significant benefit reforms and changes in residual market pricing programs.

The NCCI sources did not include data for states with monopoly state insurers and a number of states with independent rating organizations and/or residual market pools. Many of the states for which some data were reported had missing loss and/or payroll data for one or more years. We excluded states with 3 or more years of missing data or 2 consecutive years of missing data. We also exclude states with sizable state insurers that compete with private insurers. These criteria produced a sample of 24 jurisdictions (23 states and the District of Columbia, hereafter ''24 states'') with reasonably complete data for residual market share, filed versus approved rates, payroll, and incurred losses. Reasonably complete data were not available for several of the largest states (e.g., California, Texas, New York, Michigan, and Pennsylvania).[12]

11. Residual market premiums and market shares were obtained from annual issues of the NCCI's *Management Summary*. Average weekly wages and predicted loss growth due to benefit changes were obtained from the NCCI's *Annual Statistical Bulletin*. The remaining data were contained in unpublished exhibits purchased from the NCCI.

12. The 24-state sample includes: Alabama, Alaska, Arkansas, Connecticut, Washington, D.C., Florida, Georgia, Illinois, Indiana, Iowa, Kansas, Mississippi, Missouri, Nebraska, New Hampshire, New Mexico, North Carolina, Rhode Island, South Carolina,

Measures of Loss Growth and Regulation

Loss growth. We used the loss and insured payroll data to calculate the average insured loss per $100 of insured payroll each year for total losses (indemnity plus medical).[13] For states with missing data for 1 or 2 (nonconsecutive) policy years, we used the geometric growth in losses and payroll to interpolate missing values.[14] We measure loss growth as the log growth rate in loss per $100 of payroll, that is, the natural logarithm of the variable in year t to its value in year $t-1$. The log transformation produces a continuously compounded annual growth rate and reduces skewness in the growth rates due to random loss variation.

Regulation. According to our main hypothesis, loss growth should be positively related to (current and prior) rate suppression. One approach to testing this hypothesis would be to use indicator variables for different types of rating laws. However, as we explained earlier, there is considerable heterogeneity in rating laws and their implementation, and very few states had true competitive rating systems for the voluntary market during our sample period. Moreover, material suppression of residual market rates can occur even if voluntary market rates are nominally unregulated, with regulated residual market rates acting as de facto constraints on voluntary market rates.[15]

South Dakota, Tennessee, Vermont, Virginia, and Wisconsin. The sample excludes Maine, a state that is well known for the collapse of its voluntary market in the mid-1980s. The primary data sources included data for Wisconsin except for residual market share. We obtained residual market share data from the Wisconsin residual market and included this state. Maine had extreme values for many of the variables that were available, and it enacted substantial and hotly contested benefit reforms and fundamentally altered the operation of its residual and voluntary markets with its Fresh Start legislation in the middle of the sample period. This program allowed assessments for residual market deficits against employers and contingent assessments on insurers depending on ''good faith'' efforts to depopulate the residual market. We experimented with including Maine in the sample using available data. The explanatory power of the regressions declined substantially, and the coefficients declined and standard errors increased for most of the explanatory variables, including the regulation variables.

13. Separate analysis of indemnity loss growth produced similar implications. The signs of the coefficients were similar for medical loss growth, but the equations had considerably less explanatory power.

14. The policy years are not strictly contemporaneous, as they begin in different months for different states; however, virtually all of the policy years began in the first 6 months of the calender year. For 1983, some of the policy periods differed from 12 months for some states. The payroll data for these cases were annualized for the purpose of calculating the self-insured payroll measures discussed later.

15. Based on information reported by Klein (1992) and in the NCCI's *Annual Statistical Bulletin* as well as information provided to us by the American Insurance Association, only three states in our sample (Illinois, Kentucky, and Vermont) and during our sample period had rating laws that explicitly relied on competition to control voluntary market rates without regulatory prior approval. The small number of states with true competitive rating in the voluntary market precludes meaningful segmentation of the sample and estimation of models separately for these and the remaining states. It also precludes a before-and-after analysis of the effects of deregulation. Consistent with rate suppression despite

Rather than rely on categorical measures of type of regulation, we employ two continuous measures of the stringency of rate regulation (see the appendix for details): (1) one plus the filed rate increase divided by one plus the approved rate increase and (2) residual market share of premiums (at voluntary-market-rate level). We include one-year lagged values of each of these variables in our loss-growth model. We do not estimate an explicit lag structure given limited time series observations and correlations in the variables over time and between the variables. However, each measure (and especially residual market share) reflects the cumulative effects of rate suppression over a period of years.

The filed rate versus approved rate variable measures the extent to which regulators failed to approve rates requested by the industry during a period of rapid loss growth. Residual market share reflects the overall effects of price regulation on the willingness of insurers to provide coverage voluntarily. In principle, the relationship between these measures of rate suppression has implications for the allocation of the cost between insurers and policyholders in the voluntary market. When the filed-rate versus approved-rate variable is held fixed, increases in residual market share should be associated with greater cross-subsidies among policyholders. Conversely, when the residual market share is held fixed, an increase in the filed-rate versus approved-rate variable should be associated with voluntary market insureds bearing less of the expected residual market deficit in the form of higher rates and thus with lower insurer profits (and perhaps lower expenditures on loss control). However, the high correlation between these two measures impedes disentangling these effects.

Control Variables

Total losses divided by payroll in \$100s in year $t - 1$ is included in the loss-growth model to allow for diminishing growth as loss levels increase, which could arise from possible effects of higher loss levels on incentives for loss control and from any tendency toward mean reversion in loss levels over time.[16] The models also include contemporaneous and lagged 1- and 2-year estimated growth in losses due to benefit law changes using data reported in the NCCI's *Annual Statistical Bulletin*.[17] Contemporaneous changes in benefit provisions obviously should directly affect loss growth. Lagged values of the NCCI esti-

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''competitive rating'' for the voluntary market, two of these states (Kentucky and Vermont) had large residual markets and deficits during our sample period.

16. Similar results were obtained using the lagged value of premiums per \$100 of payroll.

17. According to the NCCI, the ''theoretical monetary costs are determined in accordance with standard procedures for estimating the effect of changes in benefit provisions . . . as adopted by the National Association of Insurance Commissioners'' (NCCI 1993, p. 62).

mated effects on losses are included to control for noncontemporaneous influences, such as delayed responses in firm behavior or in incentives to file claims. [18] In addition to the direct effects, changes in benefits also could be indirectly related to loss growth. In particular, benefit changes will likely be correlated with the regulatory and loss-growth environment across states, with states with high losses and loss growth ultimately more likely to reduce benefits (or likely to increase benefits more slowly) than states without substantial cost pressure. If so, any positive correlation between loss growth and benefit increases that otherwise would occur will be reduced.

We include contemporaneous growth in the average weekly wage to allow for a possible nonproportional relationship between loss growth and payroll growth. Such a relationship might arise, for example, from a possible relation between wage growth and economic growth. We do not make a strong prediction concerning this variable's sign. We also include the lagged ratio of medical losses to total losses, again without making any prediction concerning the sign on this variable, although rapid inflation in medical costs during the sample period might imply a positive relation to loss growth.

We also analyze models that include an estimate of the growth in the proportion of payroll that is self-insured to control for the self-selection that could affect insured loss growth if rate suppression induces relatively low-risk employers to self-insure. As discussed in Section III, if restrictive regulation causes relatively more low-risk employers to self-insure than high-risk employers, the resultant increase in average expected loss as a proportion of payroll for insured employers will increase insured losses relative to payroll. Given available data, we cannot employ rigorous methods for controlling for sample selectivity. Assuming that any selection effect is associated with self-insurance growth, we can test for whether regulation affects loss growth after controlling for estimated growth in self-insurance, which we define as the log growth rate of the ratio of estimated uninsured payroll to total payroll in a state (see appendix). We can also provide evidence of whether regulation is related to our measure of log growth in self-insurance.

We include fixed year and state effects in the model to allow for the effects of omitted influences that could give rise to state-invariant effects or, more important, time-invariant differences in mean loss-growth rates across states that are not captured by our other variables. Our analysis of loss-growth rates rather than loss levels and the inclusion of state effects should substantially reduce concern with bias from

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18. See Butler (1994) for a brief review of the literature. Several studies have obtained estimated elasticities of claim frequency with respect to benefit levels greater than one.

possible omitted factors that cause differences in loss growth across states.

Estimation

We estimate the models using least squares with White's heteroscedasticity consistent standard errors to construct quasi *t*-values. We also use maximum likelihood estimation assuming normally distributed disturbances and multiplicative heteroscedasticity (Harvey 1976). Specifically, the log of disturbance variance is assumed to be a linear function of insured payroll (i.e., $\delta_0 + \delta_1$ Payroll). Compared to simply weighting the data by the inverse of payroll, this procedure does not assume a strictly proportional relationship between payroll and disturbance variance. This is desirable because the disturbance in equation (2) reflects both random fluctuations in realized loss growth (ϵ_{jt}), with variance inversely related to the magnitude of insured payroll via the law of large numbers, and model error (u_{jt}), with variance that need not be related to payroll.[19] The maximum likelihood estimator is more efficient than least squares under the assumed error structure.

V.   Empirical Results

Sample means, standard deviations, and selected percentile values for the variables used in the analysis are shown in table 1. The variables each exhibit substantial variation within the panel. Table 2 shows bivariate correlation coefficients between the variables. The results indicate relatively high absolute correlations among the regulatory variables, relatively high correlations between the regulatory variables and loss per $100 of payroll, and small positive correlations between the regulation variables and loss growth.[20]

Estimates of Loss Growth Equations

Tables 3 and 4 show least squares and maximum likelihood results for the loss-growth equations. Least squares estimates with heteroscedasticity-consistent *t*-values are shown in table 3; maximum likelihood estimates are shown in table 4. Equations (1), (2), and (3) in each table

19. In particular, weighting by inverse payroll assumes a constant, negative unitary elasticity of disturbance variance with respect to payroll. Our multiplicative model allows the elasticity to increase as payroll increases (i.e., to become smaller in absolute value). Compared to the alternative approach of assuming that disturbance variance is a linear, rather than exponential, function of payroll, the multiplicative specification avoids the potential problem of negative variance estimates for some observations. Qualitatively similar results to those reported were obtained weighting by the inverse of payroll.

20. The relatively high positive correlation between growth in self-insured share of payroll and growth in average weekly wage could reflect the use of wage growth in the estimation of growth in self-insured share of payroll (see appendix).

TABLE 1   **Summary Statistics for Loss Growth, Regulatory, and Control Variables: Annual Data from 24 States during 1984–90**

Variable	Mean	Standard Deviation	Minimum	10%	25%	50%	75%	90%	Maximum
Growth in losses$_t$ / payroll$_t$	.073	.109	−.365	−.065	.028	.094	.133	.193	.398
Filed rate$_{t-1}$ / approved rate$_{t-1}$	1.057	.085	.941	1.000	1.000	1.027	1.091	1.157	1.687
Residual market share$_{t-1}$	.152	.100	.020	.037	.073	.134	.217	.268	.653
Losses$_{t-1}$ / payroll$_{t-1}$ (× 100)	1.778	.900	.528	.915	1.188	1.533	2.097	2.985	5.065
Medical losses$_{t-1}$ / losses$_{t-1}$	.431	.097	.191	.328	.375	.435	.491	.547	.602
Growth in benefits$_t$	.009	.034	−.306	.000	.002	.006	.015	.038	.102
Growth in average weekly wage$_t$	.046	.026	−.046	.011	.031	.049	.063	.076	.121
Growth in self-insured share of payroll$_t$	.010	.032	−.129	−.030	−.009	.007	.033	.048	.142

NOTE.—All growth variables are defined as the natural logarithm of the ratio of the variable in year t to its value in year $t − 1$ (i.e., as $\ln(X_t / X_{t-1})$).

TABLE 2   Correlation Matrix for Loss Growth, Regulatory, and Control Variables: Annual Data from 24 States during 1984–90

Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
1. Growth in losses$_t$ / payroll$_t$	1.00							
2. Filed rate$_{t-1}$ / approved rate$_{t-1}$	.03	1.00						
3. Residual market share$_{t-1}$	.05	.41	1.00					
4. Losses$_{t-1}$ / payroll$_{t-1}$ ($\times$ 100)	−.13	.32	.52	1.00				
5. Medical losses$_{t-1}$ / losses$_{t-1}$	.05	−.15	−.12	−.38	1.00			
6. Growth in benefits$_t$	.38	−.08	−.05	−.26	.13	1.00		
7. Growth in average weekly wage$_t$	−.00	.13	.09	−.03	−.14	.16	1.00	
8. Growth in self-insured share of payroll$_t$	.01	.15	.24	.27	−.18	−.01	.50	1.00

TABLE 3 Fixed Effects Least Squares Regression Estimates for Loss Growth Equations: 24 States, 1984–90

Variable	Equation (1)	Equation (2)	Equation (3)	Equation (4)	Equation (5)	Equation (6)
Filed rate$_{t-1}$ / approved rate$_{t-1}$	.246 (3.67)	⋯	.208 (2.80)	.248 (3.79)	⋯	.223 (3.00)
Residual market share$_{t-1}$	⋯	.427 (1.91)	.256 (1.13)	⋯	.356 (1.65)	.167 (.77)
Losses$_{t-1}$ / payroll$_{t-1}$ (× 100)	−.156 (5.46)	−.176 (5.05)	−.175 (5.14)	−.159 (5.82)	−.173 (5.28)	−.171 (5.39)
Medical losses$_{t-1}$ / losses$_{t-1}$	−.093 (.33)	−.285 (.91)	−.201 (.66)	.073 (.27)	−.107 (.37)	−.006 (.02)
Growth in benefits$_t$	1.299 (5.42)	1.191 (5.39)	1.233 (5.45)	1.313 (6.11)	1.222 (6.00)	1.269 (6.12)
Growth in benefits$_{t-1}$	.522 (1.97)	.357 (1.28)	.492 (1.84)	.455 (1.69)	.298 (1.05)	.438 (1.61)
Growth in benefits$_{t-2}$	.155 (.73)	.172 (.75)	.165 (.76)	.149 (.69)	.164 (.70)	.156 (.71)
Growth in average weekly wage$_t$	−.345 (1.04)	−.289 (.86)	−.321 (.98)	−.75 (2.06)	−.654 (1.83)	−.714 (2.03)
Growth in self-insured share of payroll$_t$	⋯	⋯	⋯	.674 (2.51)	.602 (2.34)	.643 (2.46)
$F_{(2, n-k)}$ for coefficients on regulatory variables (significance level)	⋯	⋯	8.041 (.001)	⋯	⋯	8.245 (.000)
Adjusted R^2	.42	.41	.42	.42	.42	.44

NOTE.—Loss growth is the natural logarithm of the ratio of total losses per $100 of insured payroll in year t to the value in year $t − 1$. All equations also included year and state dummy variables. Absolute heteroscedasticity-consistent t-values are in parentheses. The F-statistic tests the joint hypothesis that the coefficients on both measures of rate suppression equal zero.

TABLE 4   Maximum Likelihood Estimates of Loss Growth Equations with Fixed Effects and Multiplicative Heteroscedasticity: 24 State Sample, 1984–90

Variable	Equation (1)	Equation (2)	Equation (3)	Equation (4)	Equation (5)	Equation (6)
Filed rate $_{t-1}$ / approved rate $_{t-1}$	.183 (2.24)	...	.136 (1.60)	.184 (2.30)	...	.145 (1.74)
Residual market share$_{t-1}$	...	.421 (2.36)	.334 (1.80)	...	.366 (2.08)	.270 (1.47)
Losses$_{t-1}$ / payroll$_{t-1}$ (× 100)	−.155 (6.95)	−.181 (7.17)	−.178 (7.05)	−.155 (7.12)	−.178 (7.14)	−.174 (7.00)
Medical losses$_{t-1}$ / losses$_{t-1}$	−.183 (.71)	−.406 (1.53)	−.338 (1.27)	.008 (.03)	−.208 (.76)	−.130 (.48)
Growth in benefits$_t$	1.077 (7.76)	.976 (7.23)	1.016 (7.35)	1.107 (8.09)	1.010 (7.58)	1.054 (7.70)
Growth in benefits$_{t-1}$	.085 (.32)	−.018 (.07)	.046 (.18)	.043 (.16)	−.061 (.82)	.010 (.04)
Growth in benefits$_{t-2}$	.057 (.33)	.072 (.42)	.058 (.35)	.047 (.29)	.060 (.36)	.047 (.29)
Growth in average weekly wage$_t$	−.422 (1.61)	−.383 (1.48)	−.402 (1.55)	−.786 (2.74)	−.713 (2.49)	−.743 (2.60)
Growth in self-insured share of payroll$_t$	...	...	...	.638 (2.82)	.577 (2.53)	.593 (2.62)
Variance intercept (× 10)	.096 (5.84)	.103 (5.84)	.098 (5.84)	.090 (5.84)	.099 (5.84)	.093 (5.84)
Variance slope	−.036 (5.15)	−.040 (5.71)	−.038 (5.42)	−.035 (5.06)	−.040 (5.69)	−.037 (5.34)
$\chi^2_{(2)}$ for coefficients on regulatory variables (significance level)	...	...	7.94 (.02)	...	...	7.12 (.03)
Log likelihood	209.12	209.46	210.69	213.00	212.60	214.04

NOTE.—Loss growth is the natural logarithm of the ratio of total losses per $100 of insured payroll in year t to the value in year $t - 1$. All equations also include year and state dummy variables. The disturbance is assumed to be characterized by muliplicative heteroscedasticity as a function of insured payroll in the state ($\sigma^2_{jt} = \exp(\delta_0 + \delta_1$ Payroll$_{jt}$). Absolute t-values are in parentheses below coefficients. The χ^2-statistic is for a likelihood ratio test of the joint hypothesis that the coefficients on both measures of rate suppression equal zero.

exclude the estimated growth in self-insured share of payroll, whereas equations (4), (5), and (6) include this variable.

The overall results provide compelling support for rejecting the null hypothesis of no relation between the measures of rate suppression and loss growth. In the equations that include a single measure of rate suppression, either the (lagged) filed-rate versus approved-rate variable or (lagged) residual market share is always positively and significantly related to loss growth (*t*-value exceeds two). The least-squares results suggest a more reliable relation between loss growth and the filed versus approved variable than between loss growth and residual market share. However, the maximum likelihood (quasi) *t*-values for the coefficient on residual market share also exceed two in equations (2) and (5), which omit the filed versus approved variable. Not surprisingly given the correlation between the regulatory measures, the *t*-values decline when both regulatory variables are included. However, an *F*-test for the least squares estimates and a likelihood ratio test for maximum likelihood estimates reject the null hypothesis that the coefficients on both regulatory variables are jointly zero at the .03 significance level or lower in all cases. The maximum likelihood results provide strong evidence that disturbance variance is inversely related to payroll.

The positive relation between loss growth and the rate suppression measures is robust to including the self-insured growth variable. The coefficients on self-insured share growth are positive and significant for each specification and estimation method. This result is consistent with the hypothesis that adverse selection reflected in self-insured share growth increased insured loss growth.[21] The coefficient estimates for residual market share (but not those for the filed versus approved variable) decline when the self-insured growth variable is included, as would be expected if some of the growth in loss costs reflects a selection effect. Consistent with regulatory-induced adverse selection, regressions (not shown) of the estimated growth in the self-insured share of payroll on the regulatory variables and a vector of controls, including fixed period and state effects, indicate a significantly positive relation between self-insured share growth and the regulation variables. Thus, our overall analysis of these data suggest that rate suppression increased loss growth both by inducing adverse selection and distorting incentives for loss control.

The point estimates for the regulatory variables indicate economically meaningful effects on loss growth. For example, the maximum

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21. We obtained similar results when we estimated the model using instrumental variables to allow for the possibility that loss growth and growth in self-insured share of payroll are jointly determined. We used variables indicating whether group self-insurance was permitted and the lagged estimated share of payroll represented by self-insurance as identifying exogenous variables.

likelihood estimate of the coefficient on the filed versus approved vari-
able from equation (4) in table 4 implies a 1.8% increase in loss growth
for a 10% increase in the lagged ratio of the filed rate versus approved
rate. Similarly, equation (5) in table 4 implies a 3.7% increase in loss
growth for a 10% increase in lagged residual market share.

The estimated coefficients for the control variables generally have
the expected signs and are often statistically significant. The coefficient
for lagged losses divided by payroll is negative and highly significant
in each equation, consistent with some mean reversion in loss growth.
As expected, the estimates for contemporaneous benefit growth are
positive and significant, with higher values and some evidence of
lagged responses to benefit growth for the least squares estimates but
not the maximum-likelihood estimates.[22] The lagged ratio of medical
losses to total losses is not significant. The coefficients for growth in
average weekly wage are negative, with absolute t-values greater than
two when the self-insurance growth variable is included. Not surpris-
ingly, the results (not shown) also indicate significant state and time
effects.

Robustness

Two additional procedures were used to analyze the robustness of our
main findings to possible influential observations. First, we obtained
similar results using a rough ''trimming'' procedure in which the 10%
of the observations with the highest squared residuals based on the full
sample least squares estimates were excluded from the sample.[23] Sec-
ond, we examined the sensitivity of the results to omitting different
states. The estimates obtained from single-state deletion were robust,
with the exception of two small states. Excluding Rhode Island, a state
with well-known cost problems and accompanying rate suppression,
produced positive but insignificant coefficient estimates for the regula-
tory variables. Conversely, excluding South Dakota produced a mate-
rial increase in the coefficient estimates and t-values for the regulatory
variables. Exclusion of both Rhode Island and South Dakota produced
results with the same implications as those reported in tables 3 and 4.

VI.   Conclusions

The results of our analysis of state aggregate loss growth in workers'
compensation insurance are consistent with the hypothesis that rate

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22. A contemporaneous and lagged response of loss growth to benefit growth with total
elasticity greater than one would not be unexpected given other evidence of a positive
effect of benefit growth on claim filings and duration (e.g., Ruser 1985; and Meyer, Viscusi,
and Durbin 1995; also see Butler and Worrall 1991).
23. See Knez and Ready (1998) for a related approach and discussion in a different
context.

suppression increased loss growth during the sample period by distorting incentives for loss control. We find evidence of a positive and significant relationship between loss growth and the lagged values of two measures of rate suppression (filed rate increase versus approved rate increase and residual market share). This positive relationship persists when growth in the estimated proportion of payroll that is self-insured is included in the models to control for the selection effect that arises if rate suppression increases the proportion of lower-risk employers that self-insure. Our results imply that binding constraints on price increases in the presence of rapid loss growth led to both adverse selection and increases in insured loss growth apart from any selection effect.

These results clearly are relevant to current debate and proposals for deregulating property-liability insurance rates in states that retain such regulation. Several states recently have deregulated rates for large, commercial insurance buyers while retaining rate regulation for workers' compensation and personal lines. Our results imply that deregulating voluntary market insurance rates and avoiding material subsidies to the residual market would temper future loss growth by improving incentives for loss control.

Appendix

Measures of Regulation and Self-Insurance Growth

Filed rate versus approved rate.   After making several adjustments to the data described below, we calculated separate ratios of one plus the requested (filed) rate increase to one plus the approved rate increase for increases that occurred in a given year for the voluntary and residual markets. (If the state had uniform rates for the voluntary and residual markets the same ratio was used for each market.) We then calculated a weighted average of the voluntary and residual-market ratios using residual and voluntary market shares in the prior year as weights. For example, if the filed rate increase was 20% for the voluntary market and 30% for the residual market, if an increase of 10% was approved for both markets, and if the residual market share in the prior year was 25%, the calculated filed rate versus approved rate would equal 1.11 (0.75 × 1.2 / 1.1 + 0.25 × 1.3 / 1.1). We did not attempt to include separate filed rate versus approved variables for the residual and voluntary markets in the models given correlations between the variables. It also would be necessary to weight each variable by its respective market share, further increasing the bivariate correlation between the variables and making it unlikely that we could reliably estimate separate coefficients. Klein et al. (1993) used a filed versus approved rate variable in their analysis of the effects of regulation on loss ratios.

All of the filing information was carefully reviewed and adjustments made to the reported filed increases to reflect factors mentioned in footnotes to the data that would reduce comparability between the original filings and approved filings. One adjustment involved reducing voluntary market filings to reflect the approval

of an ''offset factor'' for the voluntary market if the approved residual market filing included approval of a new pricing program that would reduce the necessary rate increase for the voluntary market. This adjustment was made because original voluntary market filings probably were gross of this amount in many cases. This adjustment had relatively little effect on the calculated ratios. Some states had multiple filings in a year. Subsequent filings usually were made to reflect the effects of changes in benefits rules and usually were approved as filed. In a few cases, the second filing clearly was an attempt to significantly increase rates following an earlier filing that was approved for less than the filed amount. In these instances we calculate the filed request for the year as the product of the earlier approved amount and the subsequent filed request. We did not adjust for within-year variation in timing of the earlier approval and subsequent request.

The most important adjustment involved the treatment of years in which no filing for an increase was made. In some states with restrictive regulation, a filing for a rate increase may not be made in a given year because of the knowledge that no agreement will be reached with regulators. Using a value of one for the filed rate versus approved rate would clearly be inappropriate in these cases. To reduce this potential bias, we used the ratio of the filed-to-approved rate change for the previous filing if this ratio was greater than or equal to 1.05 and if the ratio for the first subsequent filing was greater than or equal to 1.15. This procedure assumes that the difference between filed and approved rate increases in the prior year persists until another filing is approved.

While the filed rate versus approved rate variable should be highly correlated with rate suppression, it is an imperfect quantitative measure of the degree to which market prices are suppressed by rate regulation for at least two additional reasons. First, conventional actuarial procedures may allow the NCCI some flexibility in determining the filed rate increase. The possibility exists that the NCCI might submit higher requests, other things being equal, in states with more severe rate suppression, hoping to achieve whatever rate increase is possible. Second, the use of schedule rating, dividends, and other competitive pricing programs will be less prevalent when regulation suppresses class-rates gross of such adjustments. In both cases, however, the resulting error in the lagged filed rate versus approved rate variable as a measure of regulatory suppression of average market rates should be positively correlated with the true but unobservable value.[24] As a result, the filed rate versus approved rate variable should still increase with the degree of rate suppression, which makes it a useful albeit imperfect proxy.

Residual market share.   The residual market share of statewide direct premi-

24. To see this, define rate suppression as the average net rate (gross rate less price adjustments such as schedule rating and dividends) that would arise under competition less the average net rate under regulation. The average filed rate can be viewed as the average net rate under competition plus the expected average competitive price adjustment plus any strategic bias. The average approved rate is the average net rate under regulation plus the expected average price adjustment under regulation. The average filed rate less the average approved rate therefore equals the difference in net rates under competition versus regulation plus the expected difference in the average price adjustment plus any strategic bias. The expected difference in the average price adjustment under competition versus regulation and any bias should both be positively related to the amount of rate suppression (the difference in average net rates).

ums is highly correlated with the extent of any cross-subsidies from the voluntary to the residual market due to price regulation. It is also affected by the extent of any overall rate suppression (both current and in prior years). In addition, the residual market share of premiums reflects differences in the average rate level between the voluntary and residual markets. Other factors held constant, states with a greater rate differential between the residual and voluntary markets will have larger residual market shares of premiums just due to the rate level effect. To control for this in the regression analysis and to better approximate the residual market share of payroll (for which we did not have state aggregate data), we estimated the residual market share of premiums at voluntary market rate level using the voluntary and residual market rate differential. Specifically, the residual market share at voluntary rate level is calculated with residual market premiums deflated by the one plus the proportionate rate differential between the voluntary and residual market. The rate differential was calculated using the history of filed and approved rate increases for the voluntary and residual markets. The residual market share at voluntary rate level is highly correlated with the unadjusted share (correlation = 0.98) during our sample period because voluntary and residual market rate differentials generally were of modest size. Preliminary analysis indicated that results of estimating our loss growth models did not differ materially for the two measures.

Self-insurance. Accurate measures of the proportion of total payroll or insured losses represented by self-insurance are not available. Previous analyses of factors that influence levels of self-insurance by Butler and Worrall (1993) and Carroll (1994) used estimates of the proportion of total paid losses that represents losses paid by self-insurers. The estimates of losses paid by self-insurers and state data on losses paid by commercial insurers and state funds used in these studies are published annually in the *Social Security Bulletin* in articles written by William Nelson (e.g., Nelson 1993) and earlier by Daniel Price. We collected these data for our sample period and 1991–93 (given that changes in paid losses will lag changes in incurred losses and payroll). Prior to 1992, many of the estimates for losses paid by self-insurers had few significant digits and changed little over time. Later reports note that improvements in data collection were made in 1992. As a result, the estimates of losses paid by self-insurers increased sharply for many states and decreased sharply for a few others. These large changes when better data were obtained led us to conclude that the paid loss estimates in the late 1980s were unreliable.

As an alternative to using the estimates of losses paid by self-insurers, we employ estimates of the proportion of total payroll in a state represented by employers that self-insure (or who are not subject to workers' compensation statutes). We obtained data on total nonfarm employment by state and multiplied the employment figures by the statewide average annual wage (52 × average weekly wage) to get estimates of total nonfarm payroll. We then estimated the proportion of payroll that is self-insured (or not insured) as (nonfarm payroll − insured payroll) / (nonfarm payroll) and calculated (log) growth rates for this ratio. Note that these growth rates also include changes in payroll that are not insured due to, for example, employers not being subject to workers' compensation law. An equivalent procedure (with signs reversed) would have been to conduct the analysis with estimated growth in insured payroll.

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