Does Rate Regulation Alter Underwriting Risk?

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Abstract: This research tests for a relationship between prior-approval rate regulation and differences in the mean and variance of individual insurer combined ratios in the private passenger auto line. A number of company-specific and market-specific control variables are also included to better isolate the effect of rate regulation. In addition to the traditional "prior-approval / open-competition" dichotomy for state rate regulation, the 1994 Conning & Company measure of regulatory stringency is included in the regression model to account for the effect of overall regulatory environment. Prior-approval laws showed no statistically significant relationship with either the mean or variance of the combined ratio, although an inverse relationship between the variability of year-to-year underwriting results and the Conning & Company regulatory freedom score was found.

The purpose of this paper is to examine the effect of state rate regulation on the profitability of individual insurers in the private passenger auto insurance market. There is a large body of research into the effect of rate regulation on statewide auto insurance markets. Researchers have looked at the relative cost of auto insurance in rate-regulated states versus non-rate-regulated states, the market shares of certain types of insurers, and the effect of rate regulation on the size of alternative (e.g., assigned risk) markets for auto insurance. However, relatively little research has been conducted on the effect of rate regulation on an individual company's operating results.

The effect of rate regulation on insurer profitability has been an issue during the development of the National Association of Insurance Commis-

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Journal of Insurance Issues, 1999, **22**, 1, pp. 26–50.

sioners (NAIC) risk-based capital formulas. The American Academy of Actuaries has taken the position that rate regulation increases insolvency risk for health insurers by increasing the volatility of underwriting results.[1] Based on the recommendation of technical advisors from the insurance industry, the NAIC's Life Risk-Based Capital formula assigns higher risk factors to certain lines of insurance because those lines are subject to rate regulation.[2] The Academy recommended that the NAIC's Health Organizations RBC formula include a twenty percent surcharge for business subject to rate regulation. This recommendation was subsequently rejected by the NAIC's Health Organizations Risk-Based Capital Working Group because of conflicting evidence about the effect of rate regulation on individual insurers.[3]

The effect of rate regulation on the profitability and stability of insurer earnings is an important public policy concern. If rate regulation reduces individual insurer profitability or increases the volatility of individual insurer earnings, solvency risk will be increased, all else held constant. If, on the other hand, rate regulation increases insurer profitability or reduces volatility, then rate regulation reduces insolvency risk. If rate regulation makes insurer profits more stable, increased risk due to reductions in profitability may be more than offset by reductions in the variability of those earnings. Put simply, an insurer with a combined ratio of 97.5 plus or minus 2 points is relatively safer from an insolvency standpoint than an insurer with a combined ratio of 96.0 plus or minus 5 points because the first insurer has a zero probability of losing surplus through its underwriting operations while the latter insurer has at least some non-zero probability of losing surplus.

Past research has concentrated on market-wide effects rather than on individual company effects. This research extends the literature by measuring the effect of rate regulation on the variability of individual companies. Also, unlike much of the prior research, this paper focuses on the overall underwriting results rather than just the loss portion of the premium dollar. Instability in underwriting expenses can be just as damaging to a company's overall profitability as instability in losses alone.

The remainder of this paper is organized into four sections. The following section discusses the general theory and summarizes some of the prior research. After that, a section is devoted to explaining the research design used to test for the effect of rate regulation on individual insurers. That section is followed by a general discussion of the results of the research, followed by a summary section.

THEORY AND PAST RESEARCH

Research into the effect of rate regulation has focused on auto insurance markets for a number of reasons. The auto insurance market is large and there is a great deal of readily available quantitative data. Auto insurance coverage is mandatory in many states and quasi-mandatory in most others. There are economic pressures that cause auto insurance coverage to be widespread as well. For example, most lenders will not finance the purchase of a car without proof of insurance coverage. Even when auto insurance coverage is not mandatory, there are no ready substitute products for auto insurance available. In contrast, customers for many other types of insurance have the option to substitute alternative products for insurance,[4] which makes it more difficult to measure the effect of regulation on insurance prices in those markets.

Harrington (1984) cites three general theories on the effect of rate regulation on profits in insurance markets: the regulatory-lag hypothesis, the excessive-rate hypothesis, and the consumer-pressure hypothesis.

- The *regulatory-lag hypothesis* is that regulatory delays in approving rate filings result in delays for companies trying to implement new rates and react to market changes. In the long run, there are no differences in the profit ratios between regulated and unregulated jurisdictions, but in the short run, rate regulation will exacerbate cyclical behavior, leading to more variability in underwriting results in rate-regulated states and, therefore, higher risk. If the regulatory-lag hypothesis holds true, then the long-run average underwriting profit should be unaffected by rate regulation, but the variability of those profits will be greater in states with rate regulation.

- The *excessive-rate hypothesis* assumes that regulators protect consumers against insolvency risk through minimum rate floors that reduce cutthroat competition. This means that, on average, insurers are forced to charge rates that are higher than they would otherwise be in the absence of rate regulation. If this hypothesis holds true, the average profit ratio in rate-regulated states will be higher than the average profit ratio in open-competition states.

- The *consumer-pressure hypothesis* holds that consumers pressure regulators to restrict prices to enhance affordability, so prices would be below the level that the competitive market would establish in rate-regulated states. If so, the average profit will be lower in rate-regulated states.

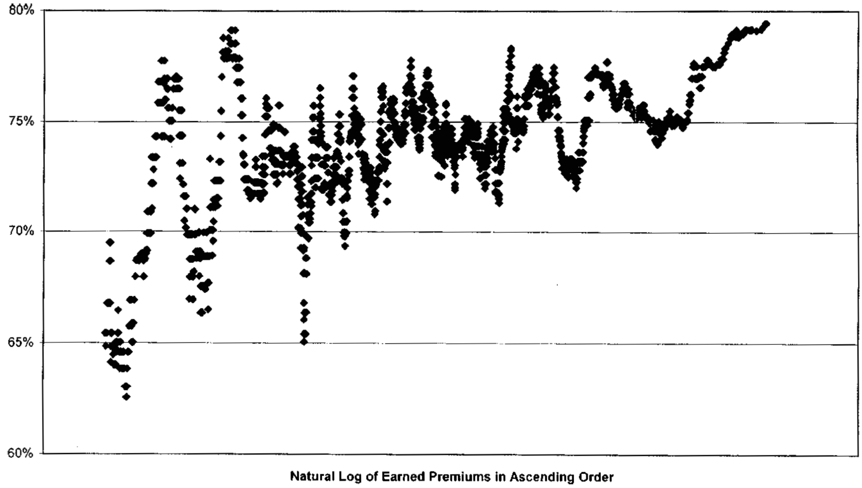

Fig. 1. Median trace of private passenger auto loss ratios, arranged in ascending premium order.

The empirical evidence as to the effect of rate regulation under these competing theories has been mixed, partially because of price measurement problems and partially because of competing definitions of "rate regulation." These differences have led different researchers to find at least some support for each of these competing theories.

Conflicting Definitions of "Price"

Most studies have used the inverse of the statewide aggregate loss ratio as the measure of price. There are several problems with using the statewide loss ratio with respect to analyzing the effect on individual insurers. First, the statewide loss ratio is affected by the mix of insurers doing business in a state. Loss ratios are positively correlated with premium volume, so larger companies tend to have higher loss ratios (Figure 1) and lower expense ratios (Figure 2), although the combined ratio (loss ratio plus expense ratio) tends to be relatively constant across insurer size (Figure 3). If the mix of large/small insurers differs from one state to the next, the state aggregate loss ratio will differ as well. Tennyson (1997) showed that overly stringent regulation discouraged participation by high-volume producers, which could lead to higher prices for consumers in the long run, with price defined in terms of premiums per dollar of losses.

Other research has shown that rate regulation is negatively correlated with direct writer market shares (Gron, 1995). Since direct writers are thought to be more efficient providers of insurance, and can thus maintain

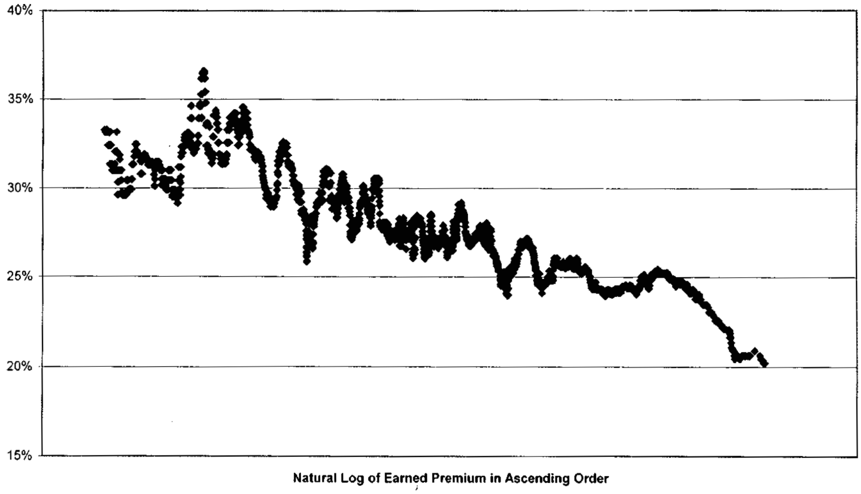

Fig. 2. Median trace of private passenger auto expense ratios, arranged in ascending premium order.

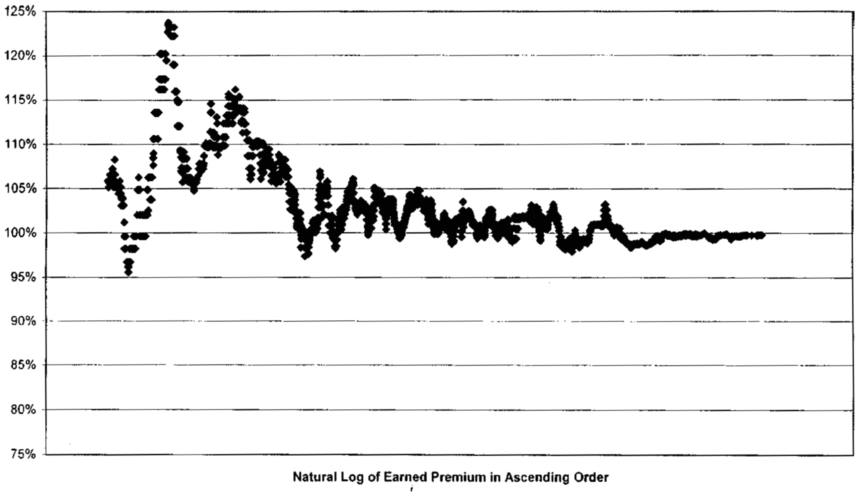

Fig. 3. Median trace of private passenger auto combined ratios, arranged in ascending premium order.

relatively higher loss ratios, the effect would be lower statewide loss ratios if the presence of rate regulation discouraged participation from direct writers.

Economic differences between states can also affect the statewide loss ratio. Outreville (1990) reported that rate regulation caused the auto insurance cycle to increase in intensity, but Tennyson (1991) suggested that those results were flawed because they did not account for systematic economic differences between states. Tennyson, using a multivariate approach to account for these economic differences, reported that prior-approval rate regulation did increase the variance of private passenger auto loss ratios, but did not affect cyclic extremes in that line. She also reported that the homeowners line of business did not show the same effect. Grabowski et al. (1989) reported that a significant amount of the total differences in auto insurance underwriting results between regulated and unregulated states could be attributed to just three states. Their work suggests that the results of other empirical studies might have been unduly influenced by the data from those three states.

There are significant differences between states' laws and markets that can affect either the absolute level of loss ratios or the variability of loss ratios. For example, researchers have used indicator variables to measure differences in loss ratios attributable to the presence of no-fault (e.g., Harrington, 1984; Grabowski et al., 1989; Eastman, 1994) and found that average loss ratios are higher in no-fault states. Eastman reported that the variance of underwriting results declined in no-fault states; the type of no-fault can be expected to affect insurer profits differently as well. Maroney et al. (1991) showed that the presence of a verbal threshold in a no-fault law was able to reduce costs, but that monetary thresholds proved to be a target rather than a limitation. Other differences include state laws on uninsured and underinsured motorist coverage, driver exclusions, minimum limits of liability, cancellation requirements and negligence laws (American Insurance Association, 1996).

The omission of insurer expenses from the calculation of "price" also can be misleading. Expense ratios differ between insurers (Figure 2). Some of those differences are size-related and reflect economies of scale, while some differences are quality-related. For example, the unallocated loss adjustment expenses is an expense load that varies from one insurer to the next. Unallocated loss adjustment expenses are the claims-adjusting expenses that cannot be allocated directly to a single claim. An insurance company such as Progressive, which sets a corporate goal of immediate claims service 24 hours a day, 7 days a week, will incur higher unallocated loss adjustment expenses than an otherwise identical company that handles claims only during office hours. Those additional unallocated loss adjustment expenses are built into the premium and reflect quality differences.

Commissions are another example of potential quality differences between products. Producer commissions are paid for services rendered by the insurance producer to both consumers and insurers. The 15 percent commission paid to an independent producer buys the consumer a wide range of knowledge about competing insurance products and prices that is not available when the consumer contacts a direct writer or exclusive producer. The additional cost paid in commissions can be more than offset by the total amount of premium dollars saved when the consumer finds a better value on his or her insurance through the independent producer. Also, since the producer acts as the front-line underwriter, part of those commission payments from the insurer are compensation from the insurer for those underwriting services, which can have a material effect on the loss portion of the total premium dollar as well. For example, many insurers include a profit-sharing bonus in their commission schedules to alleviate agency problems that arise when the producer is paid strictly for volume, so higher commission rates can signal tighter underwriting at the application stage of the insurance transaction.

The insurance premium is a mix of both expected losses and expected expenses in varying proportions. Consumer preferences play a role in the mixture and relative level of losses and expenses that make up the premium dollar. However, the total premium dollar still has to equal the total costs.[5] This suggests that the combined ratio (losses plus expenses divided by premium) is a more appropriate measure of price than the inverse loss ratio (premium per dollar of losses) when quality differences exist.

The Difficulty of Defining “Rate Regulation”

Much of the empirical research has categorized rate regulation according to the type of rate filing law in place in each state. Generally, prior-approval rating laws cause a state to be classified in the “rate regulation” group while file-and-use, use-and-file, and no-filing states are lumped into the “open-competition” group. Studies that have used this dichotomy (e.g., Harrington, 1984; Tennyson, 1991) have generally found empirical support for the consumer-pressure hypothesis. Yet the actual application of a rate filing law can differ markedly from the statutory definition. For example, a state may have a “file and use” statute but be “prior approval” in practice.[6] Additionally, many prior-approval states include a “deemer” provision that states that rate filings are automatically approved if the state does not take action within a given time period, typically 30 or 60 days. Arguably, a prior-approval state with a 30-day deemer clause is less restrictive than a state with a 60-day file-and-use statute because the insurer can implement the rate adjustment more quickly.

Other studies have used a measure of the overall stringency of regulation produced by Conning & Company as the measure of “rate regulation” (D’Arcy, 1982; Grabowski et al., 1989). The regulatory stringency measure used in this paper comes from a series of Conning & Company studies that use survey data from insurance companies to quantify a relative measure of the effect of state regulation. In the Conning & Company studies, insurance executives are asked to rank states by their level of stringency, and these surveys are used to develop scores. The scores encompass much more about a state than simple rate regulation, though. Executives are asked to assess each state on a scale of 1 to 5, taking into account "such factors as the regulatory climate, implementation of rating classifications and territories, setting adequate rate levels, cancellation and non-renewal of risks, and involuntary assignments" (Conning, 1994, p. 33).

The regulatory stringency measure tends to be correlated with the type of rate filing system in place in a state, although the correlation is not perfect. That is, rate regulation is incorporated into the Conning & Company score, along with a host of other regulatory factors. The focus of this study is the effect of prior-approval rate regulation, but regulatory stringency in other aspects of insurer operations is important as well. Therefore, both measures will be included in testing the effect on underwriting results.

RESEARCH DESIGN

Simple t-tests are used to determine whether the average company combined ratio in prior-approval states is different from the average in open-competition states. A regression model is used to test whether the presence of a prior-approval rating law affects the mean and standard deviation of the *growth rate* of individual companies' combined ratios after controlling for other risk factors.

Underwriting risk is a function of both internal and external forces. Market factors, such as the degree of competition, can affect both the average profitability and the variability of profits for companies in a particular market. Internal factors, such as market experience or size, also contribute to the variability of profits. These influences should be controlled for when trying to distinguish the effect of rate regulation on individual company results. If the typical combined ratio in prior-approval states is greater than 100 percent and / or greater than the typical combined ratio in non-prior-approval states, then the consumer pressure hypothesis is supported. If the reverse is true, then the excessive-rate hypothesis is supported. If combined ratios in prior-approval states have the same long-run average as the combined ratios in open-competition states but are more variable from one year to the next (i.e., the mean or variance of the growth rate is higher), then the regulatory-lag hypothesis is supported. The research plan is to calculate the combined ratio for each company in each state and then to measure for systematic differences in the mean and variance of the combined ratio.

The combined ratio is constructed from information contained in both the Insurance Expense Exhibit (IEE) and the Exhibit of Premiums and Losses (EPL) of the statutory annual statement for calendar years 1992 through 1997. The EPL provides direct premiums, losses, allocated loss adjustment expenses, commission and brokerage expenses, dividends and taxes, and licenses and fees by line of business for each state. These cost categories make up the bulk of the premium dollar but do not include items such as unallocated loss adjustment expenses and other general expenses, which are difficult to accurately allocate both by line and by state. However, the IEE does have an allocation on a by-line basis, and the by-line expense ratios are used to develop the rest of the combined ratio in each state.[7] Unallocated loss adjustment expense ratios, other acquisition and general expense ratios, and other income ratios for the auto liability and auto physical damage lines from the IEE are assumed to be a constant across all states for a given line of business. Because insurers generally sell liability and physical damage as bundled products[8] (that is, most insurers will not sell the physical damage coverage without also providing the liability coverage, although liability-only policies are generally acceptable), the combined ratio is calculated on private passenger auto insurance as a whole, rather than the individual liability and physical damage component pieces.

Underwriting results are generated for each company by state by calendar year as long as that company earned at least $250,000 of direct premiums in that state in a given year. The formula for computing the combined ratio is:

$$CR_i = \sum \Big[I_{L,i} + I_{P,i} + A_{L,i} + A_{P,i} + D_{L,i} + D_{P,i} + C_{L,i} + C(P,i) + T_{L,i} + T_{P,i} + \\ \left(\frac{O_L + G_L + U_L - OI_L}{E_L}\right) \times E_{L,i} + \left(\frac{O_P + G_P + U_p - OI_P}{E_P}\right) \times E_{P,i} \Big] / \sum (E_{L,i} + E_{P,i})$$

where

I	=	incurred losses
A	=	allocated loss adjustment expenses
D	=	dividends
C	=	commission and brokerage expense
T	=	taxes, licenses and fees
O	=	other acquisition expenses
G	=	general expenses
U	=	unallocated loss adjustment expenses
OI	=	other income
E	=	direct premiums earned.

The *L* subscript indicates liability business (including no-fault) and the *P* subscript indicates physical damage business. Each company's CR is calculated separately for each state *i* for calendar years 1992 through 1997, based on data contained in the annual statement filings to the NAIC. To mitigate measurement problems caused by outliers, the highest and lowest one percent of observations were discarded.

Not surprisingly, empirical analysis of the distribution of the combined ratio indicated a significant positive skew, even after elimination of the highest and lowest one percent. The combined ratio is constrained to a lower bound of zero[9] although the upper bound is theoretically infinity, which suggests that the combined ratio follows a nonsymmetrical distribution. The mean combined ratio for the entire sample was 1.0254 while the median was 1.0048, and the 2 percent difference was statistically significant. Evaluation of the spread of the distribution about the median showed that the range between the median and the 90th percentile was greater than the range between the median and the 10th percentile, another indicator of a skewed distribution. Similarly, the spread between the median and the 75th, 95th, and 99th percentiles was greater than the spread between the median and the 5th, 10th, and 25th percentiles, further evidence that the distribution was right-skewed. After applying the natural log transformation to CR, the skewness statistic improved sufficient to indicate a symmetrical distribution, suggesting that the lognormal may be a more appropriate assumption for the combined ratio results of individual companies.

A simple t test was used to measure the difference in the average combined ratio between prior-approval states and competitive rating states. The regulatory lag hypothesis holds that there are no long-run differences in average underwriting results, but that the underwriting results in prior-approval states will be more variable than in open-competition states. The next step was to test for differences in the variability of operating results from one year to the next.

The estimated standard deviation of the growth rate of the combined ratio was used to measure differences in underwriting risk, using a procedure outlined in Beckers (1980). If the combined ratio follows a lognormal distribution, then the ratio of the combined ratio in year t+1 to the combined ratio in year t is lognormal as well and the growth rate of the combined ratio will be normally distributed. If the growth rate of the combined ratio is highly variable from one year to the next, then an insurer would have to hold proportionately more surplus to absorb those fluctuations from year to year or else have a higher risk of becoming insolvent. Therefore, if prior-approval rate regulation increases the volatility of underwriting results without changing the mean underwriting result, solvency risk is increased and the regulatory-lag hypothesis is supported.

The growth rate of the combined ratio (LNDCR) is computed as

$$LNDCR = \ln\left(\frac{CR_{t+1}}{CR_t}\right) = \mu + \sigma dz .$$

Assuming that the long-run value of μ is zero, that term can be dropped from the equation. The realized growth rate is therefore the standard deviation σ times a random variable *dz*, assumed to be standard normal. The standard deviation of the growth rate of the combined ratio can be approximated by the absolute value of LNDCR as long as the combined ratio is lognormally distributed and the mean of LNDCR is both close to zero and small relative to the standard deviation.[10] Under these general conditions the absolute value of LNDCR is proportional in distribution to the standard deviation of LNDCR.

The standard deviation of the growth rate of the combined ratio is not constant for all insurers, though, and is not constant for individual insurers over time. The standard deviation is affected by volume (the law of large numbers), various internal company factors, market factors, and, presumably, regulatory factors. It can be rewritten as:

$$\sigma = \alpha\beta P^{\delta}$$

where α is the underlying "normal" standard deviation, β is a vector of company, market, and regulatory factors affecting the standard deviation, *P* is premium volume, and δ is the elasticity of the standard deviation with respect to volume. Premium volume should be an approximation of the number of policyholders in a given pool of insurance. The law of large numbers shows that the results for an insurance pool will become more stable as the pool size increases. Although the actual premiums will differ from one insured to the next and from one insurer to the next, volume will still be a good indicator of pool size.

The absolute value of LNDCR is proportional in distribution to σ, so the equation can be rewritten as:

$$|LNDR| \propto \alpha\beta P^{\delta}$$

$$\ln|LNDCR| \propto \ln\alpha + \ln\beta + \delta\ln P$$

$$LNABS = a + b_1X_1 + b_2X_2 + \ldots + b_nX_n + d\ln P + e$$

OLS regression is used to estimate the contribution of each of these various components from the company-specific vector of risk factors and from the regulatory / market vector of risk factors. The proportional constant will be absorbed into the intercept term a in the regression equation. The $b_1, b_2, \dots,$ b_n parameters are the factors for company, market and regulatory factors $X_1, X_2, \dots, X_n$ affecting underwriting risk. Although the error term e will not normally be distributed, the parameter estimates will follow a normal distribution if the sample size is sufficiently large.[11]

Explanatory Variables

The explanatory variables include a vector of company-specific variables, a vector of state-specific market and regulatory variables, and dummy variables for the individual calendar years. The variables and their hypothesized effect on underwriting variability (LNABS) are shown in Table 1. Note that there is no predicted effect on the mean growth rate of the combined ratio (LNDCR), which is expected to be zero over the long run. In the short run, LNDCR may be non-zero and change in concert with these or other non-specified variables. If the expected value of LNDCR is actually a function of these explanatory variables, then the parameter estimates for the LNABS regression will be biased.[12]

Company-Specific Explanatory Variables

The company-specific variables include LNDPE, POOLYES, NSACODE, and NR_STATE. The inverse relationship between volume and variability of underwriting results is well established. The natural log of direct premiums earned (LNDPE) is included to measure the elasticity of the standard deviation with respect to premium volume. POOLYES is a dummy variable that takes on a value of 1 if the company participates in an intercompany pooling arrangement for its private passenger auto business, 0 otherwise. Intercompany pooling arrangements allow individual companies to take on more risk in specific markets because the individual company results are combined with other companies, and then the total underwriting pie is sliced and distributed among the participants at fixed percentages. The presence of a pooling arrangement should allow individual companies to take on more direct risk, so POOLYES is expected to show a positive correlation with the variability of the combined ratio. Similarly, a company that operates in a number of different states enjoys a certain level of diversity and can therefore be expected to take on more risk in individual states. A single-state carrier, on the other hand, has to exercise caution because all of its eggs are in one basket. Therefore, the number of states that a company operates in (NR_STATE) should show a positive

Table 1. Independent Variables for LNABS and LNDCR Regression Equations

Variable Type	Variable Name	Description	Expected Relationship with LNABS
Company-Specific Variables	LNDPE	Natural log of earned premiums by state by year	–
	POOLYES	Intercompany pooling dummy (Yes=1; No=0)	+
	NSACODE	Nonstandard auto dummy (Yes=1; No=0)	–
	NR_STATE	Number of states where company writes PPA	+
State Market and Regulatory Variables	FOURFIRM	Four-firm concentration ratio for state	–
	NR_CO	Number of companies writing PPA in the state	–
	CAP_VEH	Per capita registered vehicles	–
	CAP_AGT	Per capita licensed producers	–
	PIP	Dummy for mandatory no-fault (Yes=1; No=0)	–
	ADDON	Dummy for add-on type of no-fault (Yes=1; No=0)	–
	RESID	Percentage share of residual market by year	+
	URBANPCT	Percent of population in urban areas	+/–
	FREEDOM	Conning & Company Score, 1994	?
	PRIAPP	Prior-approval rate regulation (Yes=1; No=0)	?
Time Dummies	CY92	Dummy for change period from 1992 to 1993	None
	CY93	Dummy for change period from 1993 to 1994	None
	CY94	Dummy for change period from 1994 to 1995	None

NOTE: The parameter estimate for each of these variables is expected to be zero with respect to LNDCR.

relationship with the variability of the individual state loss ratios. All of these variables were obtained from the NAIC database.

A dummy variable (NSACODE) was also included for companies that are predominantly nonstandard auto writers. Traditionally, nonstandard auto has been a market composed of high-risk drivers and low-limits policies. However, the nonstandard auto market has generated above-average profits over the past few years, leading to an increase in the number of competitors engaged in that market.[13] Nonstandard auto policies are generally issued at low policy limits, which should reduce the variability of the results for nonstandard companies. Additionally, since the nonstandard companies attempt to cherry-pick the better drivers out of the assigned risk pool, the results for the remaining assigned risk drivers should become more erratic, which can increase both the average loss ratio and the variability of the loss ratio for standard market companies. Companies were identified as nonstandard auto writers by reference to A.M. Best publications, state insurance department records, and/or miscellaneous trade press articles. This list is somewhat subjective, as the distinction between nonstandard and standard coverages has blurred somewhat over the years, and it included 191 out of a total of 1,119 auto insurers classified as nonstandard auto writers.

State-Specific Market and Regulatory Variables

A number of state-specific variables are included in the model. Per capita vehicles and per capita producers (CAP_VEH and CAP_AGT) are included as proxies for search costs. As shown by Eastman (1994), lower search costs lead to a reduction in the degree of price dispersion among insurers, which should decrease the variability of individual insurers' underwriting results. Search costs are therefore expected to be inversely related to the variability of underwriting results.

Two variables are also included to measure the effect of competition among insurers. The number of companies operating in a state (NR_CO) should be inversely related to the variability of LNDCR because the increased competition will also lead to more homogeneity in product pricing. Similarly, the four-firm concentration ratio (FOURFIRM) should be inversely related to variability for individual companies. The greater the degree of market concentration in the four largest firms, the more likely the remaining insurers are price takers rather than price makers. The number of companies in the market and the four-firm concentration ratios were calculated by state by year from the NAIC database.

The percentage of the population in urban areas (URBANPCT) is also included to test for differences between predominantly urban and predominantly rural state insurance markets. Insurers should have lower distribution costs in urbanized markets, and policyholder search costs should be lower as well, leading to less variability. On the other hand, some research has shown that urban areas are not as well served by insurers as suburban and rural areas are. There are arguments that could be made for reductions in underwriting uncertainty and arguments that can be made for increases in underwriting uncertainty, so the expected sign for URBANPCT is uncertain.

Researchers have also noted a difference in the variability of results in no-fault states. However, the effect probably differs by the type of no-fault in place. Some no-fault laws are mandatory, while others require that no-fault coverage be offered as an additional coverage. The expected relationship between mandatory no-fault (PIP) and the variability of underwriting results will be negative if no-fault actually reduces unpredictable tort claims. Add-on types of no-fault (ADDON) may also reduce tort claims, but should have less of an impact than PIP.

The number of drivers in the assigned risk pool or alternative market mechanism is expected to affect all companies in a particular state. Large pools of assigned risk drivers are indicative of noncompetitive markets. That is, a large pool of drivers that either cannot or choose not to purchase insurance in the private sector is prima facie evidence of dysfunctional markets. There may be an availability problem if there is not a sufficient number of insurers to make a market for these drivers, or there may be a pricing problem if assigned risk plan rates are set at artificially low levels relative to the fair market price. The percentage of drivers in the residual market, RESID, is expected to be positively related to the variability of underwriting results.

Type of Rating Law

A dummy variable for prior-approval rate regulation (PRIORAPP) is included to test whether the variability of underwriting results is affected by prior-approval rating laws. States that have a prior-approval rating law are coded as 1, 0 otherwise. A positive sign would lend support to the regulatory-lag hypothesis, while a negative sign would indicate that rate regulation dampens earnings volatility, as suggested by proponents of rate regulation (see Hsia and Reierson, 1987).

Although this research is intended to test the contention that prior-approval rate regulation, in and of itself, increases all companies' underwriting variability because of the regulator-induced lag in insurers' response time to changing market conditions, general regulatory climate is also of concern. For example, the underwriting results in a state with stringent cancellation and nonrenewal restrictions should be more variable, *ceteris paribus,* than an otherwise identical state with relatively few restrictions. A state with a liberal judicial system should generate higher variability than other states. To capture a measure of the general regulatory climate, the 1994 Conning & Company regulatory stringency score (FREEDOM) is used to test whether general regulatory climate affects the variability of underwriting results. As with prior-approval rate regulation, the overall effect could be to either increase or decrease the variability of underwriting results, so the sign remains uncertain.[14] The 1994 score is used for each period, although ideally a different score would be used for each year to pick up intertemporal differences and/or changes in regulatory policy. Unfortunately, Conning & Company does not conduct their survey annually, so yearly FREEDOM values are not available.

Time Dummies

Dummy variables were constructed to detect any systematic differences in the variability of underwriting results from year to year. The calendar-year dummy variables are CY92, CY93, CY94, and CY95.

DISCUSSION OF RESULTS

Table 2 shows the results of the t tests for equality of mean combined ratios in rate-regulated and open-competition states. The results indicated that there was no statistically significant difference in the mean combined ratio between states with prior-approval laws and states with open-competition laws. Separate tests were conducted for each calendar year as well, but the only calendar year that indicated a statistically significant difference was 1992. Insurers in Kansas, a prior-approval state, suffered a particularly poor year because of weather-related losses that year. When the results for Kansas insurers were dropped, the difference between prior-approval and open-competition states became statistically insignificant in 1992 as well. This finding reinforces the caution previously expressed by Grabowski et al. about the possibility that statistical results could be unduly influenced by a small number of states. Still, these results fail to support either the excessive-rate hypothesis or the consumer-pressure hypothesis as it applies to individual companies.

Interestingly, there was a positive correlation between the average state combined ratio over the entire six-year period and the FREEDOM score. That is, the combined ratio was higher (and profits relatively lower) in states that had more regulatory freedom. However, there are a number of alternative explanations that could be made for this relationship, including the assumption that the combined ratio follows a lognormal distribution.[15] This remains an area that deserves further study, but again, the focus of this paper is limited to the effect of prior-approval rating laws.

Regression results for LNDCR, the growth rate of the combined ratio, are shown in Table 3. The assumption was that all of the parameter estimates for the explanatory variables would be zero, and for the most part that held true. However, the explanatory variables NR_STATE and NR_CO, along with two calendar-year indicator variables, were statistically significantly different from zero. This could introduce some unintended systematic bias into the estimated standard deviation of the growth rate, LNABS. It is likely the statistically significant parameter estimates for

Table 2. T-test Results for Difference in Mean Combined Ratio by Prior-Approval Law

YEAR	PRIOR APPROVAL	N	Mean	Std. Deviation	Std. Error of the Mean	Mean Difference	Probability Under H0:μ1=μ2
92–97	NO	18,185	1.0300	0.2508	0.0019	0.0017	0.5551
	YES	15,612	1.0317	0.2671	0.0021		
92*	NO	3,043	1.0224	0.2534	0.0046	0.0222	0.0023
	YES	2,598	1.0447	0.2933	0.0058		
93	NO	3,043	1.0208	0.2558	0.0046	0.0015	0.8303
	YES	2,544	1.0223	0.2593	0.0051		
94	NO	3,001	1.0118	0.2362	0.0043	0.0047	0.4653
	YES	2,600	1.0071	0.2466	0.0048		
95	NO	3,079	1.0356	0.2464	0.0044	0.0050	0.4553
	YES	2,694	1.0405	0.2570	0.0050		
96	NO	3,114	1.0488	0.2559	0.0046	0.0042	0.5441
	YES	2,700	1.0446	0.2762	0.0053		
97	NO	2,905	1.0405	0.2547	0.0047	0.0105	0.1395
	YES	2,476	1.0300	0.2658	0.0053		

*Mean difference was not statistically significant when data for Kansas was omitted from the calculation for 1992.

NR_STATE and the NR_CO will disappear as more data years become available,[16] and the size of the deviations from zero was relatively small. However, caution should be exercised when interpreting the parameter estimates generated for these variables in the LNABS regression.

Table 4 shows the results for LNABS, the estimated standard deviation of the growth rate of the combined ratio. A number of statistically significant relationships were found. The primary determinant of underwriting risk, based on analysis of variance,[17] is premium volume, consistent with the law of large numbers: the greater the volume, the more predictable the results. As expected, insurers who operate within intercompany pooling arrangements (POOLYES) have more volatility in the underwriting results for their direct business. Since these companies' financial statements are

Table 3. Regression Results for LNDCR, the Growth Rate of the Combined Ratio

Variable	Parameter Estimate	Standard Error	T for H0: Parameter = 0	Prob > |T|
INTERCEPT	–0.0092	0.0388	–0.2380	0.8118
LNDPE	0.0000	0.0012	0.0330	0.9738
POOLYES	0.0032	0.0036	0.8960	0.3701
NSACODE	–0.0054	0.0047	–1.1360	0.2558
NR_STATE	0.0005	0.0001	4.2700	0.0001
FOURFIRM	–0.0037	0.0270	–0.1360	0.8919
NR_CO	–0.0001	0.0000	–2.0390	0.0415
CAP_VEH	0.0102	0.0202	0.5080	0.6117
CAP_AGT	–0.5810	0.5719	–1.0160	0.3097
PIP	–0.0060	0.0048	–1.2470	0.2124
ADDON	–0.0030	0.0047	–0.6370	0.5240
RESID	0.0003	0.0003	1.0060	0.3146
URBANPCT	–0.0207	0.0166	–1.2510	0.2110
FREEDOM	0.0036	0.0033	1.0670	0.2860
PRIORAPP	–0.0056	0.0044	–1.2670	0.2052
CY92	0.0005	0.0058	0.0870	0.9303
CY93	–0.0013	0.0058	–0.2200	0.8256
CY94	0.0325	0.0057	5.6640	0.0001
CY95	0.0156	0.0057	2.7520	0.0059
F Value	5.6640			
Prob > F	0.0001			
R^2	0.0040			
NOBS	25,696			

based primarily on the net results of the entire pool rather than direct results of the individual companies, pool participants can accept more volatility in their direct business. There may be an effect attributable to nonaffiliated reinsurance transactions as well, but the level of detail in the annual statement precludes testing for that kind of effect on a state-by-state basis. The number of states in which a company operates (NR_STATE) showed a statistically significant positive relationship with the variability of underwriting results, which is also indicative of an enhanced ability to spread risk geographically. Nonstandard auto insurers (NSACODE) show less volatility in their underwriting results, as expected.

Table 4. Regression Results for LNABS, the Estimated Standard Deviation of the Growth Rate of the Combined Ratio

Variable	Parameter Estimate	Standard Error	T for H0: Parameter = 0	Prob > |T|
INTERCEPT	2.3776	0.1598	14.8820	0.0001
LNDPE	–0.2908	0.0050	–57.9490	0.0001
POOLYES	0.0733	0.0149	4.9220	0.0001
NSACODE	–0.1434	0.0195	–7.3630	0.0001
NR_STATE	0.0022	0.0005	4.2480	0.0001
FOURFIRM	–0.1940	0.1112	–1.7450	0.0811
NR_CO	–0.0004	0.0002	–1.9850	0.0472
CAP_VEH	–0.2192	0.0830	–2.6410	0.0083
CAP_AGT	–3.4629	2.3552	–1.4700	0.1415
PIP	0.0652	0.0198	3.2920	0.0010
ADDON	0.0180	0.0195	0.9190	0.3583
RESID	0.0034	0.0011	2.9940	0.0028
URBANPCT	0.2787	0.0683	4.0830	0.0001
FREEDOM	–0.0329	0.0137	–2.4020	0.0163
PRIORAPP	0.0335	0.0181	1.8490	0.0645
CY92	0.0049	0.0240	0.2040	0.8384
CY93	–0.0473	0.0238	–1.9850	0.0471
CY94	–0.0192	0.0236	–0.8150	0.4148
CY95	–0.0203	0.0234	–0.8680	0.3855
F Value	212.196			
Prob > F	0.0001			
R^2	0.1295			
NOBS	25,696			

The state and market variables produced results generally as expected. The search cost proxies, CAP_AGT and CAP_VEH, showed a negative relationship with LNABS, although the CAP_VEH parameter estimate was not statistically significantly different from zero. The market concentration variables FOURFIRM and NR_CO also showed the expected sign, although FOURFIRM also was not statistically significantly different from zero. The urban percentage variable (URBANPCT) was positive and significant, so states with highly urbanized populations exhibited relatively greater underwriting risk.

No-fault (PIP) showed a positive relationship to the variability of the combined ratio, contrary to what one would expect if no-fault systems are effective in reducing uncertainty through elimination of tort claims. Some no-fault models, such as those employing a verbal threshold, have been shown to control costs better than those with monetary thresholds (Maroney et al., 1991). An enhancement to the model would be to segregate the types of no-fault laws more finely (e.g., verbal threshold versus monetary threshold). The parameter estimate for the add-on type of no-fault (ADDON) was not statistically significantly different from zero.

Although positive, the parameter estimate for the presence of a prior-approval rating law (PRIORAPP) was not statistically significant, which does not support the contention that prior-approval rating laws, in and of themselves, increase underwriting risk. However, the FREEDOM variable, which is a measure of overall regulatory policy in a state, did show a statistically significant inverse relationship to the variability of the combined ratio. Insurer underwriting results are more erratic in the presence of a strict regulatory climate.

Rate regulation is incorporated into the Conning & Company measure, so there is some degree of correlation between FREEDOM and PRIORAPP. Indeed, there appears to be some degree of correlation among several of the explanatory variables. The presence of collinearity among the explanatory variables can lead to a loss of power for the t tests of the parameter estimates and unstable signs for the parameter estimates. Collinearity diagnostics in SAS did not indicate model problems from multicollinearity, but those tests are limited in scope. Evaluation of the Pearson product-moment correlation coefficients among the explanatory variables in Table 5 showed several statistically significant relationships. For example, PIP is positively correlated with the URBANPCT and negatively correlated with CAP_VEH and FREEDOM, while FREEDOM is negatively correlated with the PRIORAPP, RESID, PIP, and URBANPCT.

Interestingly, there is also some degree of negative correlation between the FREEDOM variable and LNDPE, which suggests that there are relatively more low-volume insurers competing for business in those states with less stringent regulation. Analysis of variance shows that volume is the most significant contributor to the variability of underwriting results, so any analysis of differences between prior-approval and open-competition states that does not incorporate the effect of average insurer size could produce misleading results. It may be that regulation stifles competition, but it can also be argued that lack of competition encourages states to apply regulatory policies more stringently. Casual empiricism suggests the first explanation, but no proof is offered here in this paper to support either argument. We simply note that there is a higher degree of variability in

Table 5. Pearson Product-Moment Correlation Coefficients Between Explanatory Variables

	LNDPE	POOLYES	NSACODE	NR_STATE	FOURFIRM	NR_CO	URBANPCT	CAP_VEH	CAP_AGT	PIP	ADDON	RESID	FREEDOM	PRIAPP
LNDPE	1.00	–0.08	0.08	–0.02	–0.07	0.06	0.18	–0.16	–0.16	0.12	–0.02	0.14	–0.22	0.08
POOLYES	–0.08	1.00	–0.11	0.15	–0.01	0.00	0.00	–0.01	0.01	0.01	0.00	0.01	0.01	0.00
NSACODE	0.08	–0.11	1.00	–0.11	0.03	0.04	–0.02	0.06	0.00	–0.06	–0.03	–0.09	0.05	–0.02
NR_STATE	–0.02	0.15	–0.11	1.00	0.04	–0.13	–0.03	–0.01	0.06	0.02	0.02	0.04	0.01	0.00
FOURFIRM	–0.07	–0.01	0.03	0.04	1.00	–0.13	–0.02	–0.07	–0.12	–0.01	–0.10	–0.06	0.00	0.11
NR_CO	0.06	0.00	0.04	–0.13	–0.13	1.00	0.24	–0.06	–0.31	–0.12	–0.05	–0.30	0.06	–0.17
URBANPCT	0.18	0.00	–0.02	–0.03	–0.02	0.24	1.00	–0.40	–0.15	0.25	–0.02	0.10	–0.37	–0.01
CAP_VEH	–0.16	–0.01	0.06	–0.01	–0.07	–0.06	–0.40	1.00	0.09	–0.27	–0.01	–0.43	0.34	–0.16
CAP_AGT	–0.16	0.01	0.00	0.06	–0.12	–0.31	–0.15	0.09	1.00	–0.18	0.20	0.06	0.24	–0.09
PIP	0.12	0.01	–0.06	0.02	–0.01	–0.12	0.25	–0.27	–0.18	1.00	–0.33	0.17	–0.31	–0.04
ADDON	–0.02	0.00	–0.03	0.02	–0.10	–0.05	–0.02	–0.01	0.20	–0.33	1.00	0.13	0.01	0.00
RESID	0.14	0.01	–0.09	0.04	–0.06	–0.30	0.10	–0.43	0.06	0.17	0.13	1.00	–0.41	0.34
FREEDOM	–0.22	0.01	0.05	0.01	0.00	0.06	–0.37	0.34	0.24	–0.31	0.01	–0.41	1.00	–0.47
PRIAPP	0.08	0.00	–0.02	0.00	0.11	–0.17	–0.01	–0.16	–0.09	–0.04	0.00	0.34	–0.47	1.00

insurers' operating results in states that are perceived as having stringent regulatory policies.

SUMMARY

The main focus of this paper was to ascertain whether private passenger auto insurance underwriting risk for individual companies differs between prior-approval states and open-competition states. The *consumer-pressure hypothesis* holds that profits should be lower in prior-approval states relative to open-competition states. The *excessive-rate hypothesis* is that insurers earn a higher profit in prior-approval states relative to open-competition states. T-tests on the mean combined ratio of individual companies failed to show any statistically significant difference in these ratios between prior-approval states and open-competition states.

If the variability of underwriting results is exacerbated by rate regulation, as suggested by the *regulatory-lag hypothesis*, then insurance companies must hold a greater amount of capital as protection against insolvency. The results here do not support the hypothesis that rate regulation, *in and of itself*, alters underwriting risk. These results show that underwriting risk, as defined by the variability of the combined ratio from one year to the next, is no different in the presence of prior-approval rate regulation. However, these results also show that *overall regulatory climate* does indeed have some relationship to the underwriting risk experienced by insurers in each state's market. Conning & Company's regulatory stringency measure evaluates a state's overall regulatory climate and includes subjective evaluations of the effect of rate and form regulation, cancellation, and nonrenewal restrictions and other environmental factors. The relative degree of regulatory freedom had an inverse relationship to the variability of individual company underwriting results. Underwriting results are more stable, and thus underwriting risk is lower, in those states that insurers perceive to have less restrictive regulatory environments.

Proponents of rate regulation can argue that it is market instability that leads to more stringent regulation, rather than stringent regulation leading to greater market instability. The direction of any causal relationship is not formally tested here because of a lack of accurate data on the Conning scores over time. A logical extension of this research would be to apply this methodology to the results for individual states that have exhibited measurable swings in their degree of regulatory freedom to more closely evaluate and measure the cause-effect relationship.

NOTES

[1] American Academy of Actuaries Health Organizations Risk-Based Capital Simplification Task Force. *Final Report to the National Association of Insurance Commissioners Health Organizations Risk-Based Capital Working Group,* June 1996.

[2] NAIC (1994), p. 37.

[3] Minutes of the Health Organizations Risk-Based Capital (EX4) Working Group, June 1996.

[4] See Witt and Aird (1992) for a more detailed discussion.

[5] Investment income also plays a part, and investment income can be an integral part of the pricing of an insurance policy. However, the investment income potential available to insurers does not differ dramatically, and it shouldn't change materially from one state to the next.

[6] See American Insurance Association (1996) and other years for more details.

[7] This assumes that the general operating expenses are proportional to premiums. This may or may not be true in all instances, but these expenses are relatively minor and accurate allocation is problematic at best and arbitrary at worst. High premium volume states probably take up more resources than low premium volume states, so this assumption is reasonably accurate for the majority of insurers.

[8] The package pricing strategy differs from one insurer to the next. Some insurers intentionally pad the liability portion of the premium to make liability-only policies appear more expensive relative to full coverage policies. Other insurers take the opposite approach, factoring in higher investment income potential from the liability portion of the premium. However, internally, profitability is generally judged on the whole rather than on the individual pieces. That is, managers are judged on the overall results for the automobile line of business rather than separately on the physical damage and the liability components.

[9] Accounting conventions can generate a combined ratio less than zero. Earned premiums are measured as the beginning unearned premium reserve plus written premiums minus the ending unearned premium reserve. A company can generate negative premiums through a change in volume or in accounting treatment of unearned premiums. Similarly, losses are computed as paid losses plus the change in loss reserves, and that does generate negative losses on occasion. It is also theoretically possible through subrogration and salvage to recover more than the amount of losses paid, although not very likely. Absent accounting anomalies, though, in general incurred losses, incurred expenses, and earned premiums will be positive numbers.

[10] See Appendix in Beckers (1980) for the proof.

[11] This can be verified through simulation testing. The error terms are no longer N(0,1), though, so the parameter estimate for the intercept term, interpreted as the underlying "normal" standard deviation, must be increased by $2/\pi$ to obtain accurate estimates of the total standard deviation. This technique does require a large number of observations to assure that the parameter estimates for the explanatory variables are normally distributed.

[12] For example, consider the sample data set {–2, –1, 0, 1, 2} with sample mean 0, standard deviation 1.5811 and average absolute value of the mean-corrected sample equal to 1.20. If the true population mean is actually 0.1 rather than zero, the standard deviation would still be 1.5811 but the average absolute value of the mean deviated observations would be 1.22, which is slightly different. As long as the true mean is near zero and small relative to the standard deviation, the bias will be relatively harmless.

[13] See Covaleski, John, "Drive for Nonstandard Auto Profits Puts Agents at Risk," Best's Review—Property / Casualty Edition, October 1996.

[14] Most prior research has focused on the mean level of the loss ratio or profit ratio and hypothesized that regulation will increase or decrease the statewide mean. This research focuses on the change from one period to the next, which would be unaffected by the mean level. That is, it does not matter that the mean combined ratio is 95% in State A and 105% in

State B because the statistic of interest is the *change* in the combined ratio from one year to the next. The expected change for both State A and State B is zero in this instance.

[15] The expected value of a lognormal variate is $\mu + .5\sigma^2$ whereas the expected value of a normal variable is simply μ. If the mix of insurers differs from one state to the next, and smaller insurers have different variances, then there will be a difference attributable to insurer mix or size. The different mix could be caused by regulatory policies or it could be attributable to other factors.

[16] The variables POOLYES, NR_STATE, and FOURFIRM were statistically significant in an earlier regression, using data on growth rates for calendar years 1992 through 1995. With the addition of calendar year 1996 data, POOLYES and FOURFIRM became non-significant.

[17] Over 90 percent of the model sum of squares is contributed by LNDPE.

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