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Regulatory Lag in Automobile Insurance

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ABSTRACT

This article investigates whether regulatory lag affects the character of unit prices (inverse loss ratios) in regulated private passenger automobile insurance markets. The analysis is developed in the context of a simple model of a regulated insurance market. The model predicts that regulatory lag will increase the intertemporal dependence of unit prices in regulated markets. The model is estimated using data from auto liability insurance markets in 35 states over the period 1975-1986. The estimation results support the view that regulatory lag is significant for most firms in most regulated states.

Most states that actively regulate automobile insurance rates do so on a prior approval basis. This system requires approval of rates by the state insurance commissioner prior to their introduction in the market. Many studies of regulation have postulated that regulatory lag in rate approval—that is, delays in the adjustment of insurance prices to changes in market conditions—will be an important byproduct of regulation (see Harrington, 1984a, for a comprehensive review of these studies). One potential source of regulatory lag is simply the time delay inherent in the review of proposed rates. Prior approval regulation could also delay firms' filing of rate changes since they must provide statistical justification for such filings. In addition, regulatory approval standards based on backward-looking accounting measures of profits (Cummins, 1990) could induce delays in price changes relative to their competitive market adjustment.

When prices are rising, regulatory lag will prevent prices from rising as rapidly as they otherwise would; in periods of falling prices, regulatory lag will cause prices to decline more slowly. This lag could significantly distort the market equilibrium. One concern is that regulatory lag will lead to rate inadequacy during periods of inflation, adversely affecting insurer profits and causing availability problems for consumers. Another hypothesis is that regulatory lag will increase the volatility of insurer profits (Witt and Miller,

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1981) by restricting the adjustment of prices. A related hypothesis is that regulatory lag will worsen profit cycles in insurance (Witt and Urrutia, 1983a; Cummins and Outreville, 1987; Outreville, 1990; Tennyson, 1991).

There is at least anecdotal evidence that rate regulation delays or discourages auto insurance price changes. Ippolito (1979) presents direct evidence of regulatory lag, at least in a few selected states. His study compares the frequency of price changes over the period 1965 through 1975 by large insurers in California (then unregulated), New Jersey, and Pennsylvania (regulated states). He documents price changes in the unregulated state approximately every 11 months, but in regulated states only once every 21 months. These results, if representative, suggest that prices in regulated states change only about half as frequently as prices in unregulated states.

Hiebert (1976) also finds evidence of regulatory lag using data on annual aggregate premiums and losses from five regulated states plus California over the period 1966 through 1971. Under assumptions about how loss expectations are formed from past loss experience, he develops an equation for predicting premium revenues as a function of lagged losses. Regulatory and other adjustment lags are then measured as the state-specific component of the residual from estimating this equation. He finds that premiums in three of the regulated states respond less to changes in expected losses than those in California. This is consistent with the existence of regulatory lag in those states, although the results are dependent on the assumptions about loss expectations.

More broadly based empirical studies have nevertheless failed to produce consistent evidence that regulatory lag affects market equilibrium. The hypothesis generally tested in the literature is that regulatory lag will lead to lower prices in periods of rising prices and to higher prices when the overall price trend is downward. Witt and Miller (1981) find support for this interpretation in a study of the effect of regulation on auto insurance loss ratios. They document that average loss ratios in regulated states were lower in the period 1971 through 1975 when loss ratios were increasing, but were higher during the 1975 to 1977 period of declining loss ratios. Because the loss ratio is the ratio of losses incurred to premium revenue earned, it can be thought of as the inverse of the average ex-post "price" of one dollar of insurance coverage (Frech and Samprone, 1980). Hence, these results support the regulatory lag hypothesis. Other studies comparing mean loss ratios across regulated and unregulated states have found no significant differences, however (see Harrington, 1984a).

Regression analyses of the regulatory lag hypothesis also yield mixed results. Harrington (1984b) estimates the effect of regulation on auto insurance loss ratios over the period 1976 through 1981, for each year separately. The results reveal no significant differences in regulatory effects for different years. Regulation raised loss ratios throughout the period, even though loss ratios first fell and then rose over the period. This result is at odds with the usual regulatory lag hypothesis. Using a systems approach to estimation, however, Weiss (1990) finds that regulation lowered loss ratios over the period 1980-

1984, a period of rising loss ratios.[1] This is consistent with the regulatory lag hypothesis.

One potential problem with these studies is their use of static measures of the effects of regulatory lag.[2] Regulatory lag is inherently a dynamic phenomenon. It will clearly slow price changes over time, but its effect on price levels at any point in time is less clear: Unless prices are initially at the same level across regulated and unregulated states, regulatory lag will have no predictable effect on relative price levels in a given year.

The purpose of this article is to overcome this ambiguity by returning to a dynamic approach to the study of regulatory lag. A simple model of a regulated insurance market is used to develop testable implications of the regulatory lag hypothesis. The model derives the effect of regulatory lag on the path of prices over time, and hence avoids the need to compare price levels across regulated and unregulated states. The model also predicts distinct effects of regulatory lag and regulatory price ceilings, allowing regulatory lag to be identified. The model appears in the next section. The following section discusses econometric methods and presents the empirical model used in the estimation. The subsequent section contains results from estimating the model using statewide data on private passenger auto liability insurance for the period 1972 through 1986. The final section discusses implications of the findings and possible avenues for further research.

The Theoretical Model

This section generates testable hypotheses about regulatory lag by examining the implications of such lags in an equilibrium model of a perfectly competitive insurance market. Insurers are assumed to offer the quantity of insurance that maximizes expected profits. Insurance output is defined as the volume of expected losses underwritten (Frech and Samprone, 1980). This implies that the price of one unit of insurance is the price of covering one dollar of expected loss, that is, the ratio of the total premium to expected losses, under a given insurance policy.

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[1] Weiss also finds that this effect diminished slightly over the period and interprets this as evidence that its cause is regulatory delays in rate changes, not persistent price-increasing effects of regulation.

[2] An exception is Venezian (1985), who compares the cyclical properties of different sublines of automobile insurance. His is not an effective test of regulatory lag, however, because it uses national aggregate data rather than separate data for regulated and unregulated states, and because it does not control for other factors (such as different loss payout profiles) which might cause cycle differences across sublines. A study more in keeping with this one is Outreville (1990), which examines regulatory effects on cycles using statewide data. He interprets the finding that regulated states are more likely to experience cycles as evidence of regulatory lag.

The Competitive Insurance Market

Each firm in the insurance industry is characterized by its cost function. The cost function for firm i is assumed to be approximated by

$$C_i(q_{it}) = F_i + (\alpha_i + .5c_iq_{it} + D_t)q_{it}, \tag{1}$$

where q_{it} = expected losses underwritten in period t (output);
F_i = the fixed costs of underwriting;
α_i and c_i = firm-specific underwriting cost parameters, with α_i representing that portion of variable costs which is linear in expected losses.[3]

The term $D_t = 1/(1+r_t)$ is a discount factor, where r_t is a measure of the total rate of return expected from financial assets invested in at time t. This term is included to recognize the time lag between contract signing and the payment of losses incurred by policyholders. An insurer receives premium revenue in return for a liability to pay future losses experienced by the consumer; in the intervening period between contract signing and loss occurrence, the premium revenue is invested in financial instruments. The appropriate valuation of expected loss payments is thus a discounted value, reflecting expectations about both interest rates and the average time to payment of losses. Both of these effects are incorporated in the term D_tq_{it}, by assuming r_t is a total (not annual) rate of return.

Under competition, insurer i will choose output q_{it} such that the market price equals marginal cost. Under this model's output characterization, each insurer sells a homogeneous product; hence, in equilibrium prices must be equal over all firms. Denoting this price by p_t, the first-order condition for a profit maximizing insurer is

$$p_t = D_t + \alpha_i + c_iq_{it}. \tag{2}$$

The aggregate supply relationship can be derived by summing equation (2) over all firms to yield:

$$Q_t = c^{-1}[p_t - D_t - \alpha], \tag{3}$$

where Q_t = aggregate output, and
α and c = appropriately weighted averages of α_i and c_i, over all firms.

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[3] The assumption that insurers face u-shaped average costs does not contradict studies which suggest that the industry is characterized by constant or increasing returns-to-scale (e.g., Cummins and VanDerhei, 1979; Joskow, 1973), since these studies compare average costs cross-sectionally over firms of different sizes. The model here allows for different average costs over firms, and, hence, for the possibility that costs vary systematically by size of firm; in fact, the model predicts that firms with lower average costs will be larger.

Market equilibrium will be determined by equating insurance supply and demand. For expositional simplicity, assume that aggregate demand for insurance can be approximated by a linear function of price,

$$Q_t^d = A - bp_t, \tag{4}$$

where Q_t^d = aggregate quantity demanded, and
A = a demand shift parameter.

Equating supply and demand, equilibrium aggregate insurance output is then

$$Q_t^* = (1+cb)^{-1}[A - bD_t - b\alpha], \tag{5a}$$

and the equilibrium market price is

$$p_t^* = (1+cb)^{-1}[cA + D_t + \alpha]. \tag{5b}$$

Equation (5b) tells us that the equilibrium price will depend positively on the level of demand (A), and cost factors (α) and negatively on the expected returns in financial markets (D_t being inversely related to interest rates).

Incorporating the Insurance "Cycle"

Equations (5a) and (5b) represent market equilibrium assuming that prices and quantities adjust instantaneously to changes in market conditions. In insurance markets, such instantaneous adjustment seems to be contradicted by the pervasiveness of the underwriting cycle, which is characterized by several years of stable or falling loss ratios alternating with periods of rapidly rising loss ratios. Recent research indicates that cycles are most likely due to supply rigidities in insurance markets. When market conditions change, these rigidities slow the adjustment of insurance supply to its new equilibrium level. This generates correlation of output between periods and leads to swings in insurance prices and profits.

No generally accepted view of the nature of these insurance supply rigidities has yet been presented. Several models of cycles are premised on supply rigidities due to capital market imperfections and consequent costs to adjusting insurance capacity (Winter, 1988; Doherty and Garven, 1991). Under incomplete risk diversification, an insurer's capacity to underwrite losses is limited by its stock of equity capital (policyholders' surplus), which acts as a buffer against unexpectedly high losses. When obtaining capital from external sources is costly, insurers will prefer to accumulate retained earnings to increase capacity rather than issue new equity, and they will be slow to retire equity when capacity is high. This will generate rigidities in insurance output and, hence, fluctuations in equilibrium prices and profits. Other authors argue instead that information lags and institutional features of insurance markets generate supply rigidities. Venezian (1985) shows that certain extrapolative ratemaking practices will lead to autocorrelated profits. Cummins and Outreville (1987) demonstrate that the practice of writing contracts of fixed length, and lags in the realization of losses, can lead to the correlation of profits over time, even under rational expectations.

Each of these theoretical models is capable of explaining the observed serial correlation of insurance loss ratios. Thus, the economic features of insurance profit "cycles" can be credibly explained as arising from some type of supply rigidity in the industry. Devising tests capable of distinguishing the competing hypotheses about supply rigidities is nevertheless difficult (see Tennyson, 1990). Each model generates similar predictions about the relationship between insurance loss ratios and losses or policyholders' surplus.

To avoid this controversy without ignoring the existence of cycles altogether, this analysis makes use of an empirical result first reported by Doherty and Kang (1988). They demonstrate that a partial-adjustment representation of insurance supply is adequate to explain the "cyclical" fluctuations in aggregate insurance data, once interest rate fluctuations are accounted for. The partial-adjustment framework simply assumes that actual output cannot adjust fully to its desired level in each period. Output supplied in each period is therefore a weighted average of desired output and the previous period's output. This representation of supply is consistent with the existence of supply rigidities and can be derived explicitly from dynamic optimization models of firms' output choices (Tennyson, 1990). Thus, although it is not a satisfactory theoretical model of cycles, it does allow the incorporation of cycles into the model while avoiding explicit modeling of supply rigidities.

The partial adjustment of output can be expressed by the relationship

$$Q_t^s = \gamma Q_t^* + (1-\gamma)Q^s_{t-1},\ \gamma \,\varepsilon\, (0,1), \tag{6}$$

where Q_t^s = actual industry output, and
Q_t^* = the equilibrium output that would occur in the absence of supply rigidities.

The parameter γ determines the degree to which output reflects current market conditions. A smaller value of γ indicates greater supply rigidities in the market, since actual output adjusts only the fraction γ of the way toward the true competitive equilibrium.

Substituting equation (5a) into equation (6) and using equation (4) yields the equilibrium market price process in the presence of supply rigidities:

$$p_t^e = (1-\gamma)p_{t-1} + \gamma(1+cb)^{-1}[cA + D_t + \alpha]. \tag{7}$$

Equation (7) indicates that the current market price is simply a weighted average of its lagged value and the equilibrium price in the absence of rigidities given by equation (5b). This dependence of price on its lagged value arises, of course, because of supply rigidities. The greater the supply rigidity (the smaller is γ), the greater is the intertemporal dependence of prices in equilibrium. As γ approaches one, so that actual output equals desired output in equation (6), the equilibrium price approaches that achieved under instantaneous adjustment (p_t^*) characterized in equation (5b).

The Regulated Insurance Market

Equation (7) indicates the path of prices in an unregulated insurance market in the presence of supply rigidities. This section considers the effect of regulatory lag on this equilibrium price path. In this model, regulatory constraints can be conveniently interpreted as constraining prices, even though rate regulation does not impose uniformity in policy premiums across firms. This is because the "price" measure used in the model represents the mark-up of premiums over expected losses, and the standard for regulatory approval is typically a rate-of-return or profitability constraint for the insurer.

Regulatory lag can therefore be easily incorporated into the model by defining a parameter $\lambda \in (0,1]$ such that

$$p_t^R = \lambda p_t^e + (1-\lambda)p^R_{t-1}, \tag{8a}$$

where p_t^R is the price under regulation. The most straightforward interpretation of this constraint is that regulation introduces a time delay to approval of price changes; λ is then the fraction of period t for which the new price is in effect. The parameter λ can also be interpreted as a limit imposed on the size of price changes: In this interpretation regulatory lag limits price changes in any one period to be λ percent of the amount which would otherwise occur. That is, equation (8a) can be rearranged to yield

$$p_t^R - p^R_{t-1} = \lambda(p_t^e - p^R_{t-1}). \tag{8b}$$

Equations (8a) and (8b) represent a model of pure regulatory lag: They predict that regulated prices in any given time period differ from those in the unregulated market, but long-run equilibrium prices and profits are unaffected. This is consistent with common perceptions about how regulatory lag affects insurance prices or underwriting profits (Harrington, 1984b). Such lags, however, are not likely to be the main objective of rate regulation. The model also must allow for regulatory constraints on profit levels. Recent empirical evidence indicates that rate regulation raises insurer loss ratios, that is, it lowers underwriting mark-ups (Harrington, 1984a, 1984b; Grabowski, Viscusi, and Evans, 1989).[4] This article thus considers only regulatory price (mark-up) ceilings, albeit allowing for the possibility that these are not binding.

In the most extreme case, in which the regulator imposes a fixed ceiling $p_t \leq p'$, the regulated price will be $p_t^R = \min(p_t^e, p')$ in any period. The path of prices in regulated states will then depend on the probability that the price constraint is binding in any given period. If the stochastic processes for the

[4] Historically, insurance profit constraints were often administered as price floors rather than ceilings, with firms required to demonstrate that rates were adequate to meet minimum "safe" standards of profitability. Under free entry and with large numbers of firms, however, competition will erode any supra-normal profits. Historical evidence on underwriting profits, presented in Hill (1979), suggests that this was the outcome in regulated insurance markets. Regulation failed to impose an effective profit floor, probably due to the difficulty of collusion over the large number of firms in the industry.

exogenous market variables are known, this could be determined as the probability that these variables take on values above some specific limit. Current regulatory standards, however, generally recognize the financial implications of the timing of insurance loss payments and, hence, recognize variations in expected investment returns (Cummins, 1990). Thus, a more realistic regulatory constraint is one that varies with market conditions.

A form for the constraint that is both particularly tractable and consistent with empirical studies of the effects of regulation on loss ratio levels (Harrington, 1984b) is $p_t^R = \eta p_t^e$, where $\eta \; \varepsilon \; (0,1]$ indexes the stringency of the regulatory price constraint. If η equals one, the regulated price ceiling is non-binding, in which case the regulated price is just the competitive market price; a value of η less than one imposes a binding regulatory constraint as a percentage of the unregulated price. This form of the constraint thus assumes that the intent of regulation is to keep prices in a fixed relation to their free market level and captures the idea that regulation lowers prices while still allowing them to vary with market conditions.[5]

Incorporating both regulatory price ceilings and regulatory lag into the model leads to a regulated price of

$$p_t^R = \lambda \eta p_t^e + (1-\lambda) p^R_{t-1}. \tag{9}$$

Then substitution of p_t^e from equation (7) yields the following equilibrium price process under regulation:

$$p_t^R = [\lambda\eta(1-\gamma)+(1-\lambda)]p_{t-1} + \eta\lambda\gamma(1+cb)^{-1}[cA + D_t + \alpha]. \tag{10}$$

The general nature of the equilibrium price process is unchanged by regulation: Price depends on lagged price and exogenous cost and demand variables. For values of λ equal to one (no regulatory lag) and η equal to one (no regulatory price ceiling) simultaneously, equation (10) becomes simply equation (7), the unregulated equilibrium price path in the presence of supply rigidities.

The Econometric Analysis

The basis of the test of the model is a comparison of equations (7) and (10); equation (7) represents the equilibrium price process in an unregulated market, while equation (10) denotes this process under rate regulation. The effect of regulation is to change the relative dependence of price on its lagged value and, hence, to change the dynamics of price over time. This is the key testable implication of the theoretical model.

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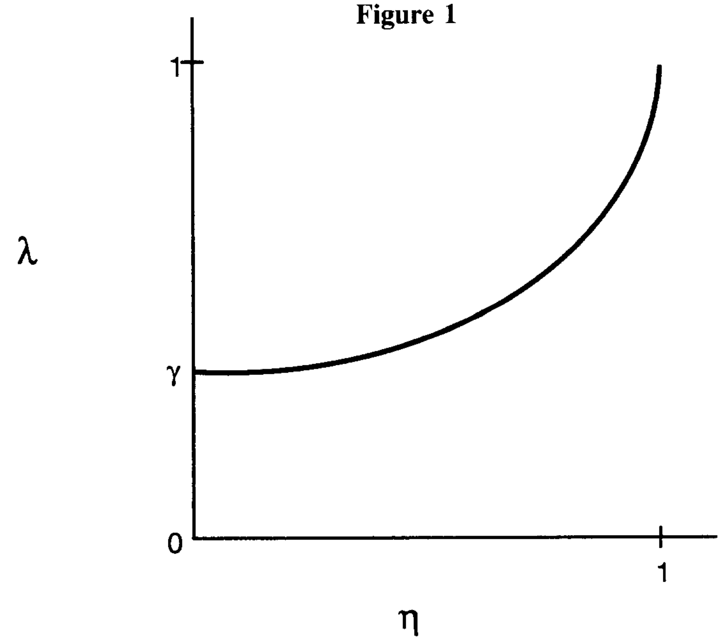
[5] For η strictly less than one, long-run equilibrium prices are lower under regulation than in the unregulated market. In the presence of assigned risk pools, however, restricting prices to lie below the competitive level need not imply shortages in the market. The regulatory constraint will simply affect the relative size of the voluntary and assigned risk pools. The fraction of insureds in assigned risk pools is only one to five percent larger in most regulated states (Grabowski, Viscusi, and Evans, 1989).

Testing for Regulatory Lag

The relative magnitudes of the coefficients on p_{t-1} in equations (7) and (10) depend on the values of the parameters γ, λ, and η. Holding γ fixed (since it is exogenous to regulation), we see from equation (10) that if the regulated price ceiling is not binding ($\eta = 1$), then the coefficient on p_{t-1} is simply $1-\lambda\gamma$; for all λ less than one this is greater than $1-\gamma$. Since the coefficient on p_{t-1} in equation (7) is simply $1-\gamma$, this implies that pure regulatory lag ($\lambda < 1$) increases the interdependence of prices across time periods. Alternatively, if λ equals one in equation (10)—that is, there is no regulatory lag—then the coefficient on p_{t-1} is $\eta(1-\gamma)$, which is less than $1-\gamma$ for all η less than one. Hence, regulatory price ceilings alone imply that p_t depends less on its lagged value.

More generally, for values of η and λ strictly between zero and one, the coefficient on p_{t-1} is greater under regulation the smaller is λ relative to η. That is, the more important is regulatory lag relative to regulatory profit ceilings, the more positive is the effect of regulation on the coefficient on p_{t-1}. Figure 1 illustrates this relationship. Holding γ fixed, the curve in Figure 1 depicts those λ and η values for which $[\lambda\eta(1-\gamma)+(1-\lambda] = 1-\gamma$. At points along this curve, rate regulation has no effect on price dynamics. At points above the curve, where regulatory lag is a relatively unimportant feature of regulation, rate regulation decreases the dependence of p_t on p_{t-1}. At points below the curve, regulatory lag is more important and results in an increased dependence of price on its lagged value.

Figure 1



This yields a simple test for regulatory lag in auto insurance markets: Regulatory lag increases the dependence of price on its lagged value. The empirical analysis therefore estimates the effect of regulation on the dependence of price on its lagged value; the regulatory lag hypothesis is supported if this effect is significantly positive. This approach improves on existing tests of regulatory lag in two ways. First, it yields an unambiguous prediction about the effect of regulatory lag on insurance prices by examining price dynamics rather than price levels. Second, it predicts separate and distinct effects of regulatory lag and regulatory price ceilings, thereby allowing regulatory lag to be identified.

Sample and Measurement Issues

The analysis uses annual statewide data on private passenger auto liability insurance, obtained from A. M. Best's Executive Data Service, for 35 states over the period 1972 through 1986.[6] The study uses data from only those states with consistent regulatory regimes throughout the sample period in order to avoid picking up effects of the deregulation transition and to minimize problems associated with the possible endogeneity of the deregulation decision.[7] Following Harrington (1984b), this study classifies a state as regulating the price of insurance if it maintains a system of prior approval regulation, state-made rates, or bureau rates; all other regulatory systems are classified as unregulated, or competitive, rating systems. Under this classification scheme, 19 states were regulated throughout the sample period, and 16 states were not regulated.[8] A complete listing of the regulatory classifications used is presented in Table 1.

Consistent with the definition of the price (mark-up) variable in the theoretical model, the empirical analysis uses the inverse of the average loss ratio by state as the dependent variable. Figure 2 depicts annual weighted average loss ratios in regulated and unregulated states over the sample period.

Several adjustments are required to transform the reported loss ratio into an economic measure. Without such adjustments, the correlation of profits between periods may be spuriously increased by accounting peculiarities in the insurance data and may bias estimation results if these features differ across regulated and unregulated states. These adjustments deal with two separate accounting issues. The first issue is the inclusion of revisions to past estimates of losses in the reported loss amount. Because of the time lag to loss occurrence, insurers set aside reserves for losses based on loss expectations. The reported level of losses incurred in each year includes changes to this

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[6] These data include premiums and losses for no-fault automobile insurance for states in which this is applicable.

[7] This is in keeping with the findings of Grabowski, Viscusi, and Evans (1989) that there were differences in the effect of regulation for those states that deregulated during the 1970s.

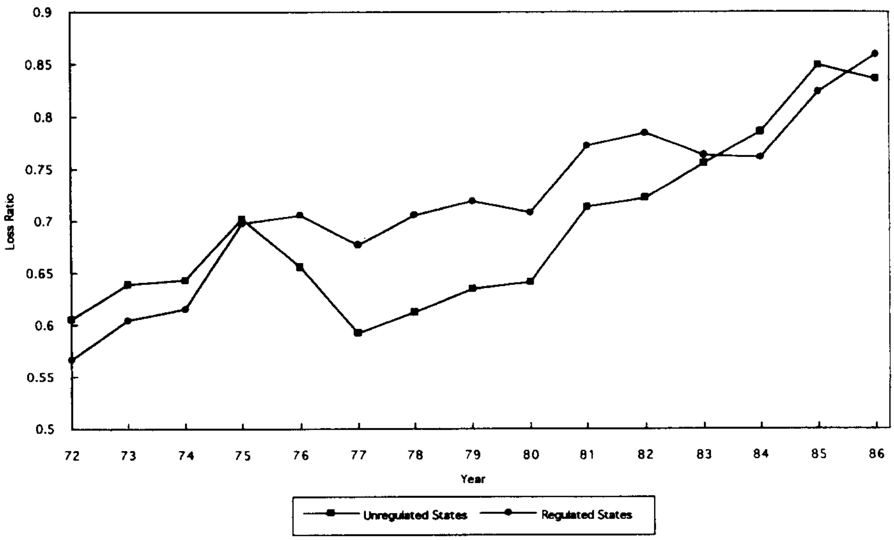
[8] Nevada was unregulated for the entire period but has been excluded from the analysis because it had no-fault auto insurance for 1976 to 1980 only. The importance of no-fault is discussed below.

Table 1
State Systems for Private Passenger Automobile Insurance

	Regulated	*Unregulated*
No-Fault	Connecticut	Colorado
	Delaware	Florida
	Kansas	Georgia
	Massachusetts	Michigan
	North Dakota	Minnesota
	New Jersey	Oregon
	Pennsylvania	
	South Carolina	
	Texas	
	Washington	
Tort	Iowa	Alabama
	Maine	California
	Mississippi	Idaho
	Nebraska	Illinois
	New Hampshire	Indiana
	North Carolina	Louisiana
	Oklahoma	Missouri
	Rhode Island	Montana
	West Virginia	Ohio
		Wisconsin

Source: Alliance of American Insurers (1987)

Figure 2
State Weighted Average Loss Ratios



reserve. One source of changes to the loss reserve is past unexpected inflation; the other source of changes is perceived changes in the population loss distribution. Regardless of the source, if actual loss experience differs from expected loss experience, adjustments will be made to the loss reserve. The data available do not report changes in the loss reserve. The extent of such changes is therefore proxied here by estimating the deviation in losses from recent trends. To calculate this deviation, the annual change in losses incurred is regressed on the lagged value of losses and time trends. The residual from this regression is then used as a proxy for changes in the loss reserve.

The second accounting issue is that losses incurred in a calendar year do not solely represent claims against contracts for which premium revenue is reported in that year (Cummins and Outreville, 1987). This is due to the convention that premiums are recorded as income only after the contract coverage period (which generates a liability) has passed. Hence, reported earned premiums represent a weighted average of premiums for contracts from current and previous years. For policies of one year or less, which are common in auto insurance, this implies that the inverse loss ratio reflects market conditions in both the current and previous year. To account for this dependence, both the current and lagged values of key exogenous variables are included in the empirical model.

The Empirical Model

Subject to these adjustments, the model to be estimated is of the form of equation (7), with a regulation dummy variable interacted with the key insurance market variables—the lagged price, the discount factor, and loss error variables. The interest rate proxy for the discount factor is the annual average three-month Treasury Bill rate less the realized inflation rate as measured by the Consumer Price Index.[9] Differences in the level of insurance demand (A) and underwriting costs (α) across states are controlled for by including state-specific effects in the estimating equation.[10]

The general form of the model to be estimated is thus

$$\begin{aligned} p_{st} &= a_1 p_{st-1} + a_2(\mu \text{Loss error}_{st} + (1-\mu)\text{Loss error}_{st-1}) + a_3(\mu r_t + (1-\mu) r_{t-1}) \\ &+ b_1 p_{st-1} \text{x} \text{Regulation}_s + b_2(\mu \text{Loss error}_{st} + (1-\mu)\text{Loss error}_{st-1}) \text{ x Regulation}_s \\ &+ b_3(\mu r_t + (1-\mu) r_{t-1}) \text{ x Regulation}_s + \omega_{st}, \end{aligned} \tag{11}$$

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[9] Although this interest rate does not accurately reflect the market portfolio of insurers, interest rates tend to move together (Fama, 1975), so its use should not be problematic as long as the composition of investments, and loss payout profiles of insurers, do not vary across regulated and unregulated states. Data on loss payout rates and investments are not available by state, and, hence, no adjustment could be made for the possibility of systematic differences.

[10] Preliminary analyses included per capita income, average hourly wage rates, and population density as proxies for cost and demand differences across states. Because these variables do not change much over time, however, they introduced significant multicollinearity into the model when estimated using the method discussed below.

where $s = 1, \ldots, N$ = state,
$t = 1, \ldots, T$ = year,
p_t (as in the theoretical model) = the price variable,
r_t (as in the theoretical model) = the interest rate proxy for the (inverse of the) discount factor,
ω is a random error term, and
the parameter μ = the fraction of reported premiums in year t which were written in that year, as discussed in the previous section.

The expected signs of the model parameters are as follows. A value of a_1 between zero and one is required to support the basic supply rigidities theory. The parameter a_2 is expected to be negative, because the price measure used in the equation is an ex-post measure, and the loss error variable approximates the deviation of realized losses from ex-ante expected losses. The parameter a_3 should also be negative, since unit prices should vary inversely with financial market returns. If regulatory lag is significant, b_1 will be positive. Because of the possibility of accounting differences across regulated and unregulated states, the values of b_2 and b_3 are more difficult to interpret, but should be positive if regulatory lag decreases the relative dependence of price on exogenous market factors.

The state-specific effects associated with omitted characteristics are captured in the error term. That is, $\omega_{st} = f_s + \varepsilon_{st}$, where f_s is a state-specific term, and ε_{st} is the usual random error term. The error terms f_s and ε_{st} are assumed to be independent, to have zero means and variances σ^2_f and $\sigma^2_{\varepsilon t}$, respectively. The error term ε_{st} is allowed to be serially correlated, because there seems no valid reason to suppose otherwise, given the presence of underwriting cycles.

Because the existence of state-specific effects in the presence of a lagged dependent variable in the model violates the conditions for consistent least squares estimates, the model is estimated on first-differences. This eliminates the state-specific effect (f_s) from the model. Letting L_{st} denote the loss error variable and REG_s the regulation dummy, the first-differenced model takes the form

$$\begin{aligned} p_{st}-p_{st-1} = {} & a_1(p_{st-1}-p_{st-2}) + a_2[\mu(L_{st}-L_{st-1})+(1-\mu)(L_{st-1}-L_{st-2})] \\ & + a_3[\mu(r_t-r_{t-1})+(1-\mu)(r_{t-1}-r_{t-2})] \\ & + b_1(p_{st-1}-p_{st-2}) \text{x} REG_s + b_2[\mu(L_{st}-L_{st-1})+(1-\mu)(L_{st-1}-L_{st-2})] \text{ x } REG_s \\ & + b_3[\mu(r_t-r_{t-1})+(1-\mu)(r_{t-1}-r_{t-2})] \text{ x } REG_s + \varepsilon_{st}-\varepsilon_{st-1}. \end{aligned} \tag{12}$$

The first-differencing procedure does not (nor is it intended to) eliminate serial correlation of the error term $\varepsilon_{st}-\varepsilon_{st-1}$. Moreover, the lagged price variable is still correlated with the error term. Following the suggestion in Hsaio (1986), the estimation uses the fact that the data form a system of simultaneous equations, one equation for each year. Hence, we have a cross-section of N observations (states) for each of T equations (years). The system can be estimated in two stages, first estimating each equation separately using instrumental variables methods to account for the endogeneity of the lagged price variable.[11] The equations for each year are then stacked and estimated as a system. This joint estimation approach allows for arbitrary forms of correlation of the errors across years.

The Estimation Results

Because the model includes lagged values and the estimation is undertaken on first-differences, the model can only be estimated for the period 1975 through 1986. Table 2 lists parameter estimates for the model using the entire sample of states; the first column of the table contains estimates for the basic model as given by equation (12). The model performs quite well: The explanatory variables exhibit the expected signs and most are statistically significant.

The effect of lagged price is positive, and between zero and one, as required by the theoretical model. Unexpected losses have the negative sign expected and are highly significant. Thus, some portion of the variation in the ex-post unit price of insurance is due to deviations in actual losses from expected losses. The interest rate variable also has a negative effect on price, demonstrating that underwriting mark-ups are indeed sensitive to returns in financial markets.[12] Most importantly, rate regulation has a significant positive effect on the lagged price. The results are thus consistent with the regulatory lag hypothesis.

Controlling for No-Fault Insurance

The estimated effects of regulation in the basic model may be biased due to uncontrolled-for differences in insurance compensation systems. While some states compensate accident victims through traditional tort-based liability, other states have enacted some type of no-fault compensation law. No-fault insurance compensates small bodily injury claims via first-party coverage and sometimes restricts individuals' rights to sue responsible parties for pain-and-suffering damages. The product being offered under no-fault insurance is thus a mix of first-party personal injury coverage and traditional bodily injury liability coverage and, hence, differs from the liability coverage offered in tort states (Witt and Urrutia, 1983b). Because first-party personal injury claims are generally paid more quickly than liability claims (Insurance Research Council, 1989), there could be natural differences in the time path of unit prices across states with no-fault and tort compensation systems. Sixteen states in the sample had no-fault auto insurance compensation systems and 19 had tort systems (see Table 1).

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[11] The instruments used include all available lags of the loss error variable and the interest rate. Note that the number of lagged variables in the instrument list differs for different years in the sample; instruments are formed separately for each year.

[12] Estimating each year separately in the first stage precludes including two interest rate variables in the equation. The compromise taken here is to include the second-difference of the interest rate. Similar results are obtained when the first-difference is used.

Table 2
Instrumental Variables Estimation of First-Differenced Model
Arbitrary Error Correlation
Dependent Variable = Differenced Inverse Loss Ratio (p_{st}-p_{st-1})

	Model 1	*Model 2*	*Model 3*
Price$_{t-1}$	0.0784* (0.0395)	0.1798* (0.0440)	0.2067 (0.1458)
Loss Error$_t$	-0.5020* (0.0269)	-0.4923* (0.0270)	-0.4829* (0.0257)
Loss Error$_{t-1}$	-0.2819* (0.0320)	-0.2790* (0.0318)	-0.3069* (0.0314)
Interest$_{t-t-1}$	-0.0078* (0.0011)	-0.0073* (0.0011)	-0.0067* (0.0011)
Price$_{t-1}$ x Regulation	0.1893* (0.0542)	0.2336* (0.0553)	0.2910* (0.0588)
Loss Error$_t$ x Regulation	0.0573 (0.0390)	0.0501 (0.0392)	0.0408 (0.0391)
Loss Error$_{t-1}$ x Regulation	0.1111* (0.0407)	0.1076* (0.0400)	0.1204* (0.0391)
Interest$_{t-t-1}$ x Regulation	0.0016 (0.0016)	0.0025 (0.0015)	0.0025 (0.0015)
Price$_{t-1}$ x No-Fault	---	-0.1781* (0.0478)	-0.2139* (0.0482)
Price$_{t-1}$ x Direct Writer Share	---	---	-0.0561 (0.2334)
System R^2	0.3812	0.3757	0.3827

Note: Standard errors are in parentheses. Variable definitions are provided in the Appendix.
* Statistically different from zero, 5 percent significance, 1-tailed test.

To allow for potential differences in the path of prices across compensation systems, the second column of Table 2 includes a no-fault dummy variable (equal to 1 in no-fault states) interacted with the lagged price term in addition to the regulatory interaction variables. As expected, the no-fault interaction variable is significantly negative, indicating smaller supply rigidities in no-fault states. In fact, the net dependence is very close to zero in no-fault states. The effect of regulation on the lagged price term is nevertheless significantly positive in this version of the model as well. Thus, regulatory lag appears to be significant regardless of the insurance compensation system.

Controlling for Market Structure

A potentially more important bias in the estimation is the implicit assumption that supply rigidities due to forces other than regulation (γ) do not vary across regulated and unregulated states. This may be invalid if regulation affects the structure of insurance markets. As Joskow (1973) first noted, there are significant differences in market structure between regulated and unregulated states. For example, in regulated states both the share of the market written by the largest firms and the average number of firms are significantly lower than in unregulated states.

In addition, the market share of direct writers is lower in regulated states. In my sample, the average market share of direct writers is 61.2 percent in unregulated states but only 51.6 percent in regulated states. Direct writers are firms that employ their own insurance sales force or contract with exclusive sales agents, as opposed to selling policies through independent agents who represent several insurers. Because they are more integrated, direct writers may face smaller supply rigidities (have a larger γ) than other firms. The price measure used in the analysis is aggregated across all firms in a state; to the extent that direct writers represent a lower share of the market in regulated states, the effects of regulation estimated above may be misleading. In particular, if direct writers face less severe supply rigidities, their smaller market share in regulated states could be the source of the apparent regulatory lag in the aggregate data.

Model 3 of Table 2 reports estimation of the model with the lagged price variable interacted with the direct writers' market share in the state, as well as the other interaction variables. Although this variable is not statistically significant, it does have the expected negative sign. The results also indicate that regulatory lag is still significant after controlling for direct writer market share differences. The more pronounced temporal interdependence of prices under regulation does not seem to be an artifact of the smaller market share of direct writers in regulated states, but rather is attributable to regulation itself.

Estimation of Subsamples

To check the robustness of the results reported above, the estimation is also undertaken for no-fault and tort states separately, and for agency and direct writers separately. The estimation of the overall average unit price equations

Table 3
Instrumental Variables Estimation of First-Differenced Model
Arbitrary Error Autocorrelation
Tort and No-Fault States Separately
Dependent Variable = Differenced Inverse Loss Ratio (p_{st}-p_{st-1})

	Tort States	*No-Fault States*
$Price_{t-1}$	0.2481*	0.0710*
	(0.0426)	(0.0271)
$Loss\ Error_t$	-0.8825*	-0.3394*
	(0.0356)	(0.0120)
$Loss\ Error_{t-1}$	-0.0413	-0.1720*
	(0.0400)	(0.0169)
$Interest_{t-t-1}$	0.0029*	-0.0104*
	(0.0007)	(0.0008)
$Price_{t-1}$	0.3049*	0.0338
x Regulation	(0.0642)	(0.0340)
$Loss\ Error_t$	0.0125	-0.0018
x Regulation	(0.0557)	(0.0192)
$Loss\ Error_{t-1}$	-0.0502	-0.0812*
x Regulation	(0.0562)	(0.0203)
$Interest_{t-t-1}$	0.0029*	-0.0010
x Regulation	(0.0010)	(0.0010)
System R^2	0.5429	0.3545

Note: Standard errors are in parentheses.
* Statistically different from zero, 5 percent significance, 1-tailed test.

for tort and no-fault states are presented in Table 3. There are significant differences in price dynamics for no-fault and tort states. As expected, the degree of price dependence over time is lower in no-fault states than in tort states. In addition, the (aggregate) effect of unexpected losses on the current price is greater in tort states, possibly reflecting longer claim settlement lags in tort states. Moreover, the effect of regulation on the lagged price variable is positive in both sets of states, although significant only in tort states. This is consistent with the results obtained for the full state sample.

Table 4 reports estimation of the model for agency writers and direct writers separately, for the subsamples of no-fault and tort states. For direct writers, the model performs well: Price dependence is lower in no-fault states, and the model parameters have the expected signs. Moreover, regulation has the expected positive effect on price dependence and is significant in no-fault states. The model performs less well for the agency writer subsample, especially in no-fault states. While regulatory lag is significant for agency writers in tort states, regulation decreases intertemporal price dependence in no-fault states, possibly reflecting greater stringency of price regulation in regulated no-fault states. Regulatory price ceilings should fall more heavily on less efficient agency writers (Joskow, 1973) and, hence, could outweigh regulatory lag effects for these firms. Note, however, that supply rigidities also appear to be greater in no-fault states than in tort states for agency writers. This is contrary to expectations; thus, the results could merely reflect the relative heterogeneity of the agency writer subsample.[13] While the average number of direct writers per state is 30 for regulated states and 35 for unregulated states in my sample, the average number of agency writers is 111 and 135 for regulated and unregulated states, respectively. In addition, many of the agency writers are small regional firms, especially in regulated states. Estimating average price paths for agency writers may thus be less reliable.

Discussion

This study provides support for the hypothesis that regulatory lag affects market equilibrium in auto insurance. Estimating a reduced form equation for the mark-up of premiums over losses (the inverse loss ratio), it is found that the dependence of this price on its past levels is significantly greater in regulated states. This is shown to be consistent with the view that regulatory lag prevents timely adjustment of insurance prices. These results are obtained even after controlling for other factors that might affect the time path of prices, such as the type of compensation system and the market structure.

In the aggregate, regulatory lag appears to have significant effects on market equilibrium. An important extension to this study would consider insurer responses to regulation. For example, it is often conjectured that insurers manipulate reported loss reserves in order to avoid the effects of stringent regulation or to obtain higher rate allowances from regulators. Since loss reserves affect the value of incurred losses, this could affect equilibrium price dynamics as measured by inverse loss ratios. Although such behavior has not been empirically demonstrated, studies of loss reserves indicate that loss reserving errors do not appear to be random (Weiss, 1985; Grace, 1990). In addition, Pauly, Kunreuther, and Kleindorfer (1986) find evidence to suggest that insurer service provision is lower in regulated states (as measured by expense ratios), presumably in response to profit ceilings or delays in rate changes. While it is not clear that service differences will affect price dynamics, they may do so if, for example, one component of service is

[13] The results may also reflect bias introduced by the relatively small sample size used in estimation of the subsamples. Recall that the first stage of the estimation uses only a cross section of states for a single year. In estimating tort and no-fault samples separately, this means that the first-stage regressions use only 19 and 16 observations, respectively.

Table 4
Instrumental Variables Estimation of First-Differenced Model
Arbitrary Error Autocorrelation
Direct and Agency Writers Separately
Dependent Variable = Inverse Loss Ratio (p_{st}-p_{st-1})

	Direct Writers		*Agency Writers*	
	No-Fault	*Tort*	*No-Fault*	*Tort*
Price_{t-1}	-0.0219	0.2963*	0.3853*	0.2792*
	(0.0379)	(0.0335)	(0.0398)	(0.0266)
Loss Error_t	-0.4902*	-0.5438*	-0.7715*	-0.5276*
	(0.0196)	(0.0213)	(0.0180)	(0.0269)
Loss Error_{t-1}	-0.2209*	-0.1870*	-0.3652*	-0.1525*
	(0.0235)	(0.0153)	(0.0287)	(0.0158)
Interest_{t-t-1}	0.0045*	-5.1E-5	-0.0033*	0.0114*
	(0.0013)	(0.0012)	(0.001)	(0.0010)
Price_{t-1} x Regulation	0.3631*	0.0614	-0.2053*	0.1486*
	(0.0505)	(0.0440)	(0.0472)	(0.0459)
Loss Error_t x Regulation	0.2942*	0.2469*	0.5147*	-0.2935*
	(0.0221)	(0.0246)	(0.0202)	(0.0331)
Loss Error_{t-1} x Regulation	0.2139*	0.1040*	0.1779*	-0.0101
	(0.0251)	(0.0193)	(0.0295)	(0.0283)
Interest_{t-t-1} x Regulation	-0.0044*	0.0094*	0.0015	-0.0049*
	(0.0017)	(0.0018)	(0.0013)	(0.0015)
System R^2	0.4953	0.3020	0.6707	0.6883

Note: Standard errors are in parentheses.
* Statistically different from zero, 5 percent significance, 1-tailed test.

timely payment of claims. Thus, changes in product quality in response to rate regulation could affect price dynamics.

This study controls for loss reserving differences across states by including deviations in losses from trends in the estimated model. Potential differences in services have been implicitly allowed for in the theoretical model by allowing cost functions to vary across firms.[14] Such differences are controlled

[14] Consumer demand does not, however, depend on service in this model; hence, it is not an equilibrium model of insurance products differentiated by service provision.

for in the estimation only by incorporating state-specific effects. More explicit consideration of these issues is beyond the scope of this article, but further investigation would enhance our understanding of the microeconomics of the regulatory lag identified here.

Another important consideration is that regulatory lag may be an intrinsic part of regulatory policy rather than an unintended consequence. Studies of other regulated industries have found that in periods of increasing input costs, rate-of-return regulation tends to be more stringent, lowering industry profit margins to curtail price inflation for consumers (Joskow, 1974). This could be the source of regulatory lag in the insurance industry as well. Particularly in the recent climate of high loss cost inflation (see e.g., Cummins and Tennyson, 1992), regulatory lag may be used as a tool to keep prices from rising too quickly. The analysis presented here cannot distinguish a deliberate policy of regulatory lag from other sources of such lags. More direct evidence on this point would provide useful insights into the regulatory process.

Appendix

Price = Earned Premiums/Incurred Losses
Source: Best's Executive Data Service, various years.

Interest Rate = Annual Average of 3-Month U.S. Treasury Bill Rate - CPI Annual Inflation Rate
Source: U.S. Statistical Abstract.

$$\text{Loss Error} = \frac{\text{(Annual Change in Losses - Predicted Change in Losses)}}{\text{Incurred Losses}}$$

Construction of Predicted Change in Losses

The variable of interest is the change in losses not attributable to changes in market share or the size of the market, or due to inflation in current losses. Hence, a variable to measure "unexpected" changes in losses incurred is constructed as the residual from the regression of the actual change in losses incurred on lagged losses, a time trend and a time-squared trend. The regressions for agency writers and direct writers also includes a control for market share. Predicted values were obtained for the time period 1974 through 1986, using the entire 35-state sample.

Predicted Loss Change Regressions

Statewide: $L_t\text{-}L_{t-1} = 8606.6 + 0.1694L_{t-1} - 5513\text{TIME} + 337.6\text{TIME}^2.$

Agency: $L_t\text{-}L_{t-1} = 4079.7 + 0.1324L_{t-1} - 1832.6\text{TIME} + 125.5\text{TIME}^2 + 1624.5\text{MKTSHARE}.$

Direct: $L_t\text{-}L_{t-1} = 6858.3 + 0.1703L_{t-1} - 3931.6\text{TIME} + 232.7\text{TIME}^2 + 2415.4\text{MKTSHARE}.$

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