Regulatory Delay and the Timing of Product Innovation

James E. Prieger[1]
Department of Economics
University of California
One Shields Avenue
Davis, CA 95616-8578
jeprieger@ucdavis.edu

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Abstract

This paper examines how regulatory delay affects innovation by a regulated firm. When product introduction costs fall over time, an extra day of regulatory delay increases time to introduction by more than a day. When the uncertainty of regulatory delay increases, a risk neutral firm responds by moving up the product introduction. The model places testable restrictions on the empirical relationship between introduction delay and regulatory delay, and is consistent with data gathered from a large U.S. telecommunications provider.

JEL Codes: L51, L96

1 Introduction

The potential for economic regulation to distort the incentives of the firm to innovate and introduce new products is well known (e.g., Sweeney, 1981; Cabral and Riordan, 1989). Most of the literature examining regulation and innovation focuses on the impact of the type of regulatory regime (rate of return vs. incentive regulation, for example) or on the frequency of policy revision (the so-called "regulatory lag"). A little-explored avenue is the effect of regulatory delay on innovation.[1] Regulatory delay exists when the regulator does not allow the introduction of new products without regulatory review and approval. Regulated firms—for example, in the telecommunications, pharmaceutical, and banking industries[2]—often claim that regulatory delays are long, costly, and distort the incentives to introduce new products. Furthermore, uncertainty inherent in the regulatory process may further impact the firms' decisions. This paper explores the relationship between regulatory delay and the timing of innovation. "Innovation" in this context refers to the commercialization and introduction of new products to consumers (Iansiti and Lerner, 2002), rather than basic invention. The model I develop places testable restrictions on the empirical relationship between introduction delay and regulatory delay, and is compared with novel data from an incumbent local exchange telephone company in the Midwest.

There are regulator-side and firm-side components to the delay between technological feasibility of a product and its introduction to consumers. The regulator-side component is the time between the firm's submission of a new product to the regulator for approval and the granting of approval. I term this component *regulatory delay* (the term is not intended to be pejorative; delay may have social benefits). The firm-side component is the time between the first technologically feasible introduction date[3] and the submission of the product to the regulator. I term the firm's component *introduction delay.* Recent history in the telecommunications industry shows that introduction and

[1] Note that "regulatory delay" is a different concept than "regulatory lag". The former refers to delayed introduction of a new product, whereas the latter refers to the term of regulatory commitment.

[2] Examples of regulatory delay in telecommunications are presented in this article. In the pharmaceutical industry, regulatory delay comes from required FDA approval of new drugs. Regulatory delay in the banking industry came from line-of-business restrictions before deregulation.

[3] I.e., the first date at which the introduction costs are less than infinite.

regulatory delay tend to move in the same direction. In the data from the beginning of the 1990's from four Midwestern states examined here, a given new product tended to be introduced in different areas at very different times. For example, products were typically introduced in Ohio more than a year after availability in other states. By the end of the decade, product launches were more likely to be closer together among the states, and in some cases were simultaneous.

On the other side, many state regulatory streamlined their procedures for product introductions by regulated firms, leading to shorter regulatory delay and allowing products to reach the market sooner. This pattern also shows up in these data. I show formally that the profit-maximizing response of the firm is to reduce introduction delay. Regulatory delay reduces the opportunity cost of introduction delay for the firm by pushing the forgone profits from the new product farther into the future. When introduction costs fall over time, regulatory delay thereby induces the firm to postpone introducing the product. I also show that an increase in regulatory uncertainty causes a risk-neutral firm to introduce products sooner.

It is important to understand the determinants of introduction delay, because the impact on economic welfare can be large. For example, regulators may delay approval of a new product because of concerns about the price at which it will be offered to consumers. However, welfare gains from introducing a new product (the so-called Dupuit triangle) are typically an order of magnitude higher than any deadweight loss from a supra-competitive price (the Harberger triangle).[4] By empirically demonstrating with a novel dataset that regulatory delay increases introduction delay, this article sheds light on this important issue. The theoretical model I develop is closest to that of Braeutigam (1979), who also looks at a monopolist's incentive to innovate given a regulatory regime. Other literature focuses on adoption timing as entry deterrence or accommodation under different regulatory regimes (Riordan, 1992; Lyon and Huang, 1995), but does not explicitly consider regulatory delay. This paper leaves aside rivalry considerations, which were not important in local telephony in the time period studied. There are a few empirical studies of the impacts of regulatory

[4]Hausman (1997) quantifies of the high welfare cost of regulatory delay in telecommunications.

delay on innovation and product introduction. Prager (1989) finds that regulatory delay by public utility commissions raises the cost of capital for electricity firms considering constructing new plants. Gruber and Verboven (2001) show that regulatory delay in the granting of operating licenses to providers had a persistent effect on the evolution of the mobile telecommunications industry in Europe. Prieger (2001, 2002a, 2002b) finds that increased regulatory delay is associated with fewer new telecommunications products introduced in several different contexts. Hazlett and Ford (2001) highlight the potential for firms to use regulatory delay to raise rivals' entry costs.[5] Aside from case studies, systematic econometric studies documenting that regulatory delay influences the timing of innovation by the firm are rare.

The outline of the paper is as follows. In the next section, I introduce a basic model of a firm's decision of when to introduce a new product, first with fixed regulatory delay (Section 2.1) and then with uncertain delay (Section 2.2). Section 3 presents the testable implications derived from the signaling model and introduces the data from a large incumbent local exchange telephone company that are used to perform the tests. Testing of the predictions is carried out in Section 4. The results show that the model is generally consistent with the observed patterns of introduction delay and regulatory delay, although if the firm is risk neutral the model cannot explain the firm's response to regulatory uncertainty seen in the data. However, even a mild amount of risk aversion can bring the predicted response to uncertainty in line with the data. A final section concludes.

2 The Theoretical Model

2.1 The case of fixed regulatory delay

I now introduce a simple model of regulated product introduction. Let time $t = 0$ represent the point at which a firm can first feasibly introduce a given product. The firm chooses to submit

[5] In the another empirical study of regulatory delay, Sanyal (2003) asserts that patent approval delays detrimentally affect the incentive of firms to innovate. However, the dependent variable in the estimations performed is patent approvals and not applications. Approvals would decline as patent delays increase merely due to queuing, even if the application rate were unchanged.

the product to the regulator for approval at time $s \geq 0$, at which time it incurs fixed cost $F(s)$. F may include the cost of development, technology adoption, or regulatory filing. The length of introduction delay s will be referred to as the product introduction date.[6] Following Riordan (1992), fixed costs are assumed to be falling over time as exogenous technological advances lower the cost of adopting the new service. In dynamic industries such as telecommunications, it is realistic to assume that the fixed costs of introducing a new product fall over time. I assume $F'(t) < 0$ and $F''(t) > 0$, and all functions in the model are assumed to be continuous and twice differentiable. Falling fixed costs give the firm an incentive to delay introducing the product. The regulator approves the service after an examination period (i.e., regulatory delay) of length a. Firms cannot sell and consumers cannot purchase the good until time $s + a$, the introduction date. After introduction, the firm earns constant flow profit of π per unit time.[7] Note that π does not vary with any choice variable of the firm; lacking data on service prices, I focus on the timing variable s instead.[8] The firm's discount rate is r, so that its net present value of introduction at time s is:

$$\Pi(s, a) = -e^{-rs}F(s) + \int_{s+a}^{\infty} e^{-rt}\pi dt = e^{-rs}\left(-F(s) + e^{-ra}\frac{\pi}{r}\right) \tag{1}$$

The firm chooses optimal introduction date $s^* = \operatorname{argmax}_s \Pi$, which is defined by the first order condition:[9]

$$\frac{\partial \Pi(s, a)}{\partial s} = 0 \Rightarrow rF(s^*) - F'(s^*) = e^{-ra}\pi \tag{2}$$

The left side of equation (2) is the marginal benefit from postponing introduction (the reduction in fixed costs), and right side is the marginal cost from postponing introduction (the forgone profit). The higher the firm's profit flow π, the higher the firm's opportunity cost of delay.

Given the assumptions of the model, regulatory delay is unambiguously bad for the firm.

[6] Whether s represents true innovation or merely adoption of existing technology (diffusion), the structure of the model is the same.

[7] The timing of the model is similar to that of Braeutigam (1979).

[8] Many of the new telecommunications services introduced in the data are classed as "competitive" services and are allowed to be freely priced by the firm. In that case π is the flow profit resulting from the profit-maximizing price.

[9] To guarantee $s^* > 0$, assume $rF(0) - F'(0) > e^{-ra}\pi$. To guarantee finite s^*, assume that $\lim_{t\to\infty} rF(t) - F'(t) \leq 0$.

Proposition 1 $\partial\Pi/\partial a < 0$. *Longer regulatory delay lowers the firm's profit.*

From (1), $\partial\Pi/\partial a = -e^{-r(s+a)}\pi < 0$. There is no provision in this model for the firm to expect that regulatory delay will be beneficial. An example of beneficial delay is for the regulator to delay introduction until technical standards or coordination issues are resolved, which may reduce the firm's cost or increase demand for the service.[10] The model does not rule out that delay is beneficial *ex post*, merely that the firm expects delay to affect profit only through discounting.

Proposition 2 $\partial s^*/\partial a > 0$. *Longer regulatory delay induces the firm to introduce the product later.*

The proof is in the appendix. As regulatory delay increases (e.g., from a_L to a_H in Figure 1), the forgone profit is pushed farther into the future and its present value, which is the marginal cost of delay, falls. Since marginal benefit is decreasing, to re-equate marginal cost and marginal benefit later introduction dates are chosen by the firm. Thus there is a multiplier associated with regulatory delay: adding a day of regulatory delay, if anticipated by the firm, increases the time until introduction by more than a day.

Proposition 3 $\partial s^*/\partial\pi < 0$. *A higher opportunity cost of delay induces the firm to introduce the product earlier.*

The proof is in the appendix. The relevant picture is the same as Figure 1, where now the top marginal cost curve corresponds to a higher π and the bottom marginal cost curve corresponds to a lower π. At first this result might appear counterintuitive; if regulation is "bad for the firm" why would higher marginal costs of regulatory delay lead to *earlier* product introduction? The answer requires distinguishing between the direct and opportunity costs of regulation. Flow profit π creates the opportunity costs of regulation; as the forgone profit from delay increases, the firm

[10] The Federal Communications Commission, for example, delayed approval of high-definition television broadcasts for many years during the late 1980's and 1990's while it tested various technologies and chose a standard. The FCC apparently believed that consumers would ultimately benefit more from a high-quality product offered under a single standard, even if they had to wait an extra decade. If so, then increased demand may raise firms' profits.

introduces the product earlier to speed accrual of those profits. If the direct cost of the regulatory process is included as a constant in F, then an increase in direct cost would postpone introducing the product. This can be seen from Figure 1 by shifting the marginal benefit of delay curve up.

2.2 The case of stochastic regulatory delay

The data examined below reveal that there is substantial variation in regulatory delay a. The firm may therefore view regulatory delay as a random variable instead of a known quantity. Accordingly, split regulatory delay into a fixed component, $\bar{a} \geq 0$ and a stochastic component $\tilde{a} \geq -\bar{a}$. Structural delay $\bar{a}$ is non-random and represents the regulatory delay that the firm expects any service to go through. Structural delay includes average time to get on the regulator's docket, waiting for commissioner sessions, and mandatory examination periods. In the empirical models to follow, structural delay is taken to be unchanging during the course of a particular regulatory regime in a state, and exogenous when considering any single product or service. The part of regulatory delay that is random and varies among services is $\tilde{a}$. The regulator may set $\tilde{a} < 0$ and choose to expedite approval. Thus the regulator can choose total regulatory delay $a = \bar{a} + \tilde{a}$ to be any positive length. The role of $\bar{a}$ in the model is to provide a link to the empirical application, in which structural, service-inspecific delay clearly is a salient feature of the regulatory regimes examined.

Rather than explicitly modeling the choice of delay by the regulator, which requires specifying a regulatory objective function, here I assume the firm treats $\tilde{a}$ as a random variable with known distribution function G. In the working paper (Prieger, 2005) I explore a game theoretic model in which the firm views regulatory delay as a strategic choice variable of the regulator. Assuming the firm is risk neutral, the first order condition to the firm's decision problem with random regulatory delay is of the same form as (2) with $\bar{a} + \tilde{a}$ replacing a and the right side of the equation evaluated in expectation with respect to density G. Given that the random delay $\tilde{a}$ enters the first order condition nonlinearly, the variance of $\tilde{a}$ affects the firm's choice of introduction delay even under risk neutrality. In particular:

Proposition 4 *Given two densities G_1 and G_2 with the same mean for $\tilde{a}$, if G_1 second-order stochastically dominates G_2 and the firm is risk neutral, then s^* is lower under G_2 than G_1. Riskier regulatory delay induces the firm to introduce the product earlier.*

Being convex in $\tilde{a}$, the expectation of the right side of (2) is higher when $\tilde{a}$ is riskier.[11] Thus the effect of increasing regulatory risk on s^* is similar to the effect of decreasing a. Figure 1 can be re-interpreted as depicting the expected marginal cost of delay when regulatory risk is high (the top line) and low (the bottom). The firm views the impacts from a mean-preserving spread of regulatory delay asymmetrically. The firm places more weight on the possibility that regulatory delay becomes shorter, because longer delay times are discounted more heavily. Thus a risk-neutral firm views an increase in the riskiness of regulatory delay the same way it views a decrease in structural regulatory delay. This point was first made by Braeutigam (1979) in a different model of innovation by a regulated firm, and has not been empirically tested to my knowledge.[12]

If the firm is risk averse, however, it might choose to delay introducing new products more when regulatory riskiness increases. It is not hard to construct examples where adding risk aversion to the model induces the firm to respond to increased risk by choosing later introduction dates. In such examples, increasing s lowers profit but also makes it less sensitive to a, so that expected profit has less variance. If risk aversion is strong enough, when the riskiness of a increases the firm will increase s in response to flatten the profit function and take away some of the uncertainty created by the higher variance of regulatory delay. I return to this possibility in Section 4.

3 Data and Discussion of the Tests

The theoretical model places restrictions on the relationship between regulatory and introduction delay. First, from Proposition 2, the firm's introduction delay rises as structural regulatory delay

[11] This follows directly from the definition of second-order stochastic domination ($\succ_2$): $G \succ_2 G_2$ iff for every non--decreasing concave real function u, $\int u \; dG_1 \geq \int u \; dG_2$. The conclusion then follows from setting $u = -\exp(-r\tilde{a})$.

[12] The empirical literature on regulatory risk typically treats risk as the possibility of *ex post* regulatory interference, rather than uncertainty about the timing of project approval. Examples include studies of the effect of regulatory risk on the return to assets. A separate strand of literature looks at the impact of risk imposed on irreversible investment by open access regulation (with the deployment of DSL by an incumbent phone company being a leading example).

rises. This prediction applies to average behavior within a regulatory regime, and implies that introduction delay is longer in regimes with longer structural delay. It does not apply to service-specific regulatory delay, since introduction delay necessarily precedes regulatory delay, assuming that the regulator cannot commit to a policy a before the firm moves.[13] Second, Proposition 4 implies that riskier regulatory regimes lead to shorter product introduction delay by the firm. Assuming that a distribution for regulatory delay with higher variance is a mean-preserving spread of a lower-variance distribution,[14] Proposition 4 suggests that if firms are risk neutral the variance of regulatory delay will be negatively correlated with product introduction delay by the firm. In these regressions I will control for average regulatory delay to proxy preserving the mean.

Data were collected on regulatory and introduction dates for telecommunications services introduced in the 1990's by Ameritech in Illinois, Indiana, Ohio, and Wisconsin.[15] Ameritech also operates in Michigan, but since new services were effectively deregulated in Michigan, no tariff data are available. Ameritech, one of the Bell regional holding companies and later acquired by SBC (now AT&T), is the dominant local exchange company in each of these states, and its intrastate activities are regulated by the state commissions. Introduction of a new service required petitioning the public utility commission in each state; the service could not be offered to subscribers until regulatory approval was granted. Examples of the residential and business services in the data are new voice mail features, virtual networking services, and high-speed transmission services. The data cover the span 1991 through 1999, which comprises three regulatory periods.[16] In the first period, 1991 through mid 1994, Ameritech was under some form of rate of return regulation in each state. Following this first period, each state switched to some form of incentive regulation. After three years of the new regulation, in 1997 the regimes were reviewed in at least some of these

[13] Lack of commitment is a common assumption in regulatory games (outside of the mechanism design literature). See Spiegel and Spulber (1997) for a discussion of why regulatory commitment is not a realistic assumption.

[14] The assumption is required because when comparing distributions with the same mean, second order stochastic domination implies lower variance, but the converse does not always hold.

[15] The data are from the tariff filing logs of the company and the state commissions. Supplemental information was culled from the actual state tariffs where needed.

[16] The data for Ohio are complete only for years 1994-1999.

states.[17] Thus the regulators (or state legislatures) had three opportunities to set their policy concerning structural regulatory delay (i.e., to choose $\bar{a}$). Preliminary statistical work revealed that the latter two periods were indistinguishable in terms of average innovation and regulatory delay, and so I collapse the years 1994–1999 into a single period of incentive regulation in the empirical models and refer to it as Period 2 in the tables.

The first difficulty for the empirical investigation is measuring s, time between earliest potential and the actual product introduction ("introduction delay"). I take the date at which a service is first introduced in any of these states or in the FCC's access tariff to be $t = 0$, and then measure s for the other states relative to the first state's introduction date. This effectively underestimates true introduction delay: the true time 0 must be weakly before the observed first "product introduction" under this definition. If the unobserved delay before the first product introduction is constant in the sample, then the same offset would be added to all observations on s, the difference would be absorbed into the nonparametric baseline hazard of the empirical models in the next section, and the estimated marginal effect of regulatory delay would be the same. However, if the unobserved delay before the first product introduction is shorter in period 2 than in period 1, as I show the observed introduction delay to be, then the results will understate the true effect of shortening regulatory delay.

Applying the single-actor, single-market theoretical model requires the assumption that there are no strategic interactions among jurisdictions. To be included in the data set, a new service had to be introduced in at least two states. One hundred fourteen services were introduced in at least two states, generating 349 observations. Summary statistics for the observations on introduction delay are in Table 1. Regulatory delay, a, is measured as the time from the first tariff filing submission date to the approval date of the last tariff filing for the service.[18] Regulatory delay data is not available for Ohio. Summary statistics for regulatory delay are in Table 2.

Heterogeneity among regimes may induce correlation between regulatory and introduction delay.

[17] See Roycroft (1999) for more information on the regulatory regimes.

[18] Some services had multiple tariff filings and withdrawals before approval was granted.

For example, if profit opportunities are systematically higher in one regime than another, and the first regime also happens to have lower structural regulatory delay, then we may observe spurious positive correlation between s and $\bar{a}$. In the empirical application, therefore, I control for variables that affect average profit, cost, and consumers' surplus (size, density, and wealth of the market, etc.) in the regulatory regime. Controlling for differences in profit, cost, and consumers' surplus helps isolate the impact of regulatory delay on the timing of product introduction.

4 Empirical Results

The goal of the empirical work is to uncover relationships in the data between introduction delay and structural regulatory delay. The predictions of the theoretical model will also be tested.

Estimating how introduction delay varies with structural regulatory delay. Structural regulatory delay varies greatly between regulatory regimes (across time and states) in the data. The institutional changes that took place in 1994 in each state expedited approval for new services. Streamlining the regulatory approval process received special attention in the new incentive regulation plans. In Illinois, the legislature mandated that the regulatory commission evaluate whether an alternative regulatory plan would "reduce regulatory delay and costs over time".[19] Under the new regulation, termed Advantage Illinois, new services deemed competitive were allowed to be introduced on one day's notice, and many more services were classified as competitive after the regulatory change. In Indiana, all new services were allowed to be introduced on one day's notice under the "Opportunity Indiana" alternative regulatory plan, down from at least a month of regulatory delay before the new plan. In Ohio, the legislature explicitly noted that "Alternative methods [of regulation] may include, but are not limited to, methods that...minimize the costs and time expended in the regulatory process...."[20] In response, the commission effectively detariffed competitive services and allowed them to be introduced with essentially no regulatory scrutiny. In

[19] See § 220 Illinois Compiled Statutes, sec. 13-506.1.
[20] See Ohio Revised Code § 4927.04.

Wisconsin, the commission revised its procedures to ensure that approval for new services would be granted after 10 days unless suspended for investigation, down from about a month under rate of return regulation. The intent of the new regulation in each state was to ensure that structural regulatory delay be smaller in period 2 under the alternative regulatory schemes.

Another source of evidence is to examine the data themselves. Table 2 shows that the mean and median tariff approval delay dropped in Illinois, Indiana, and Wisconsin (no data are available for Ohio). Results from estimations lead to similar conclusions. I estimate the entire distribution of regulatory delay in the two periods via the Kaplan-Meier (Kalbfleisch and Prentice, 1980) non-parametric method (Figure 2). The figure presents the survival curves (defined to be $1 - CDF$) for regulatory delay and indicates that delay in period 2 stochastically dominates period 1 delay. Estimated means and medians from the curves are in Table 4. Mean and median regulatory delay is smaller in period 2 in each state. The confidence intervals for the median delay in periods 1 and 2 are non-overlapping in all states.

How did Ameritech respond to the reduced structural delay? From the raw data in Table 1 it is clear that mean and median introduction delay dropped substantially from period 1 to period 2. The estimated survival curves in Figure 3 reveal convincing evidence that period 2 introduction delay stochastically dominates period 1 delay. Estimated means and medians from the curves are in Table 3, and confirm the visual evidence from the curves: the mean and median introduction delay is smaller in each state in period 2. Although the confidence intervals for the median delay in the two periods overlap for Illinois and Ohio, if a slightly higher quantile is chosen, e.g. the 0.6 quantile (which corresponds to the ordinate 0.4 on the survival curves), the confidence intervals do not overlap for any state. Furthermore, a log-rank test rejects the hypothesis that the distributions from periods 1 and 2 are the same in Illinois and Ohio, and also jointly for the entire data.[21] Thus it appears that introduction delay fell in each state in period 2, confirming the prediction of the model that introduction delay moves in the same direction as structural regulatory delay.

[21] The test, also known as the Mantel-Haenszel test, has a p-value of 0.003 for Illinois, 1.6×10^{-6} for Ohio, and less than 10^{-11} when all states are grouped.

Since the nonparametric method does not allow covariates, I turn next to a semiparametric model to control for economic conditions and other factors that may change over time and influence the firm's behavior apart from the strategic considerations that I want to isolate. Estimates from a Cox proportional hazards model are in Table 5. In the Cox model, the hazard rate of the introduction delay durations is

$$\lambda(t, x_i) = \exp\left(x_i'\beta\right)\lambda_0(t), \tag{3}$$

where λ_0 is an arbitrary, unspecified baseline hazard and x_i is a vector of covariates for spell i.[22] The baseline hazard is left free to take any shape, so positive, negative, and non-monotonic duration dependence are all possible. Positive coefficients for β increase the hazard and therefore decrease mean duration. The first estimation replicates the finding from the survival curve estimation. When only fixed effects are included—state dummies, state-specific indicators for period 2 (*state:regulatory change*), and an indicator that delay is calculated from the federal access tariff—the coefficients for the regulatory change all indicate shorter delay times in period 2 (see Estimation I1 in Table 5). Also reported for each estimation are the χ^2 statistic for the joint significance of all coefficients and Grambsch and Therneau's (1994) $T(G)$ statistic. The latter is for a test of the proportional hazards assumption of the Cox model. In each estimation, the coefficients are jointly significant and the proportional hazards assumption is not rejected.

Estimation I2 is a more flexible specification, in which stratification by state replaces the state dummy variables, which allows the baseline hazard to vary in shape and level without restriction across states. The coefficients again indicate shorter delay times in period 2, although the coefficient for Illinois loses significance. This finding persists when state-level economic covariates are added in Estimation I3 to replace the state dummy variables and stratification. The new variables are log average household income, the log number of access lines (a measure of market size), the log land area, and log number of telecommunications industry patents.[23] Because Ameritech is not the

[22] The Cox (1972; 1975) model uses a $\sqrt{N}$-consistent partial likelihood method to estimate β.

[23] The Bell Operating Companies take out few patents themselves, and typically create new services from underlying technology patented by others. The variable used instead is the count of patents in the year (based on application date) from any company approved in the classes relevant to telecommunications services (359: optics, systems

incumbent in some rural areas, the first three variables are calculated only for Ameritech's operating territory in the state. All covariates are allowed to evolve over the course of a duration. Adding these variables does not remove the conclusion that introduction delay fell in period 2 except, again, for Illinois, for which the coefficient is smaller than before and statistically insignificant. Taken together, then, the evidence from all estimations indicates that introduction delay is positively associated with structural regulatory delay in accordance with Proposition 2, except possibly in Illinois.

Why is the evidence for positive correlation between introduction delay and structural regulatory delay so weak in Illinois? The construction of the data gives a clue to the answer. As explained in the previous section, introduction delay is measured as the time elapsing between the first introduction in any state or in the federal access tariff. The first introduction among these states is more often Illinois than not, given its large market size, and there is little that can be said about the firm's choice of s in such cases. As can be seen from the survival curves at $s = 0$ in the top left panel of Figure 3, the first introduction is very often in Illinois both before and after the regime change. Even if true introduction delay fell in Illinois markedly after the regime change, in these cases the data are not be able to capture it.

Turn now to the prediction from Proposition 4: riskier regulatory regimes lead to quicker product introductions by a risk-neutral firm. I test this proposition by including two new variables in the introduction delay estimations to proxy the expected mean and variance of regulatory delay. Recall that Proposition 4 is stated in terms of second-order stochastic domination, which is equivalent to mean-preserving spreads. Therefore expected mean delay is included as a control variable so that changes in the variance of regulatory delay can isolate changes in the riskiness of regulation. The expected mean regulatory delay for a service is constructed as the posterior mean duration calculated from all completed and ongoing regulatory delay durations at the time of tariff

(including communications) and elements; 370: multiplex communications; 379: telephonic communications; and 395: information processing system organization) and is compiled from data from Patent and Full-Text Image Database of the U.S. Patent and Trademark Office.

approval filing (see appendix for details). By taking an explicitly Bayesian approach to measuring average regulatory delay, the durations that are ongoing at the time of filing can be incorporated into the estimate, improving its accuracy. The expected variance of regulatory delay is measured by the sample variance of completed regulatory delay durations at the time of filing.[24] Both proxies are calculated from data only from the same state and regime. There are a smaller number of observations in these estimations because regulatory delay data are not available for Ohio.

The results for the estimations examining how introduction delay varies with the riskiness of regulation are in Table 6. For the sake of parsimony, the only economic covariate included is the number of access lines, given that the other economic controls were not significant in Estimation I3. In Estimation I4, the proxy for average regulatory delay enters linearly. The coefficients for the regulatory change (*state:regulatory change*) again all indicate shorter delay times in the latter regulatory regime (and the effect regains significance in Illinois), even though one would expect that most of the difference between regimes is already picked up by the proxy for average regulatory delay. The coefficients of interest in Table 6, however, are the state-specific proxies for the variance of regulatory delay (*state:regulatory uncertainty*). In contrast to the implication of Proposition 4, there is no evidence that increased regulatory riskiness leads to a higher hazard rate for introduction delay. In fact, all the coefficients show the opposite effect (although only one is marginally significant). To check the robustness of this unexpected finding, Estimation I5 in Table 6 controls for average regulatory delay with cubic p-splines (Eilers and Marx, 1996), which allow the variable to enter the hazard specification in a flexible way. It is especially important with duration data to control for the average of regulatory delay adequately when interested in the effect of its variance, since the mean and variance tend to move together. Thus if one does not appropriately control for average delay, the coefficients for riskiness might pick up some effects of the co-varying mean, leading to misleading inference. The other variables enter the specification of the hazard rate linearly as in Estimation I4. The results in Estimation I5 are the same as in I4,

[24] I did not incorporate ongoing delays into the calculation of the variance, because such Bayesian updating proved to be much more sensitive to the assumption of the prior distribution for the second moment than for the first.

and even stronger: increased riskiness leads to lower hazard rates for introduction delay, and the effect is statistically significant in Indiana and Wisconsin. Finally, to check for non-monotonicity in the effect of riskiness on the introduction delay hazard, I repeat specification I4 allowing riskiness itself to enter through p-splines. Figure 4 shows the partial effect of riskiness on the log hazard for introduction delay from this estimation. The inset box in the figure shows the entire range of the data, and the main graph shows the data omitting the top 3% of the variances, which are all from Wisconsin and much higher than the others. If riskiness enters the model linearly, then the partial effect would be linear in such a graph. The figure reveals that increasing riskiness decreases the hazard rate for introduction delay, at least up to the point where the few outliers from Wisconsin are reached. Thus the negative coefficients for riskiness found Estimations I4 and I5 do not appear to be artifacts of a linear constraint on the functional form.

Thus we are presented with a puzzle: why would the firm choose to delay introducing new products more when regulatory riskiness increases? As noted after Proposition 4, a potential answer is risk aversion. While it is traditional to assume that large corporations such as Ameritech are risk neutral, it takes little risk aversion to reverse the implication of Proposition 4. In an example where introduction costs F are exponentially declining and such that the firm just breaks even introducing a service immediately, a utility function with an Arrow-Pratt measure of 0.25 is enough to induce the firm to delay introduction further when the variance of regulatory risk increases.[25] Such utility exhibits little curvature in the relevant region of profit.[26] Thus it may be the case that even a small departure from risk neutrality by the firm is leading to an outcome inconsistent with Proposition 4.

[25] In this example, $F(s) = \exp(-s)$, $r = 10\%$, $\pi = 0.1$, the utility of wealth function is CARA, and b is lognormal with mean $\exp(.5)$ and scale parameter σ^2 rising from 1 to 2.

[26] In particular, consider the domain bounded by $\Pi(s^*, b_{0.025})$ and $\Pi(s^*, b_{0.975})$, where b_p is the pth quantile of b and b has the distribution from the previous footnote. This is the set covering 95% of the relevant range of profit outcomes for the firm. If one plots the utility function from the previous footnote on this domain, it is difficult to see any concavity.

Summary of the Empirical Results To conclude the empirical work, I summarize the results with a suite of hypothesis tests in Table 7. Panel A of the table reports results from the introduction delay estimations I1–I3. The tests are for $ds/d\bar{a} \leq 0$, which would violate Proposition 2. The test is implemented with the null hypothesis that the coefficients for the regime change are non-positive. Thus, under the null, the hazard rate for introduction delay stays the same or decreases (or, equivalently, introduction delay did not decrease) in the period when structural regulatory delay was lower. This hypothesis is soundly rejected in all states, with the possible exception of Illinois, for which the test rejects at the 5% level for I1 and at the 10% level for I2, but does not reject for I3. The hypothesis that the relevant coefficients are zero is also rejected at the 1% level when the states are tested jointly.

Panel B of Table 7 reports results from estimations I4–I5, where the tests are for the effect of regulatory uncertainty on introduction delay. Since the estimated coefficients on the regulatory uncertainty variables were negative in each state, the test is implemented with the null hypothesis that the coefficients are non-negative. Thus, under the null, the hazard rate for introduction delay stays the same or increases (or, equivalently, introduction delay did not increase) as the variance of regulatory delay increases. Rejecting the null violates Proposition 4. The hypothesis is rejected for Illinois in Estimation I4, and for Indiana, Wisconsin, and all coefficients jointly in Estimation I5. Thus the bulk of the evidence runs counter to Proposition 4. As discussed above, these results may stem from risk aversion of the part of the firm.

5 Conclusions

This paper presents a model that investigates the timing of product introduction, with particular attention paid to the role of regulatory delay. The model generates two main testable predictions, which are compared with results from a Bell Operating Company's operations in four states. The main theoretical result, which is supported empirically, is that the reduction in average regulatory delay in the Ameritech states contributed toward the speedier product introductions by the firm observed in the latter half of the 1990's. To the extent that regulatory delay still exists in these and other jurisdictions, the additional social cost of delay from distorting the incentive to create and introduce new products should be factored into regulators' and legislatures' social calculus. A second prediction, that a risk-neutral firm responds to increased regulatory risk by introducing products sooner, is not borne out by the data. A mild amount of risk aversion on the part of the firm, however, can rationalize the data with the model.

There are other applications and interesting extensions that deserve future attention. The theoretical model may also apply to other regulatory settings, such as the timing of patenting and patent approval, or of pharmaceutical development and regulatory approval. With minor modifications to the objective functions, the model may also apply to decision-making within a firm, where the agents are the R&D division and management, in place of the firm and the regulator, respectively. One extension would be to explicitly incorporate unregulated rivals into the model. The impacts of competition in the current model are tacit: competition may be seen as affecting the marginal cost of delay to the firm (through π) or the regulator's distribution of delay. Given that local telecommunications competition was just getting off the ground during the period studied, including competition in the model seems to be most useful for application to future data sets. Finally, a few papers in political economy attempt to peer inside the black box of regulatory delay in other contexts (Ando, 1999; Dwyer, Brooks and Marco, 1999). Exploring the political economy of regulatory delay in the telecommunications industry would be an interesting complement to the present work, where the regulator's objective function is not modeled.

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6 Appendix

Proof of Proposition 2 We can find $\partial s^*/\partial a$ by differentiation of (2), since (2) holds for all a: $\partial s^*/\partial a = re^{-ra}\pi/\left[F''(s^*) - rF'(s^*)\right] > 0$, where the inequality follows because $F' < 0$ and $F'' > 0$.

Proof of Proposition 3 The marginal cost of delay to the firm is the same whether the delay stems from the firm's or the regulator's choice. We can find $\partial s^*/\partial\pi$ by differentiation of (2), since (2) holds for all π: $\partial s^*/\partial\pi = -e^{-ra}/\left[F''(s^*) - rF'(s^*)\right] < 0$, where the inequality follows from the assumptions on F.

The calculation of expected regulatory delay The variable *average regulatory delay* in Estimations I4 and I5 is constructed as the posterior mean duration of regulatory delay calculated from all delay durations (completed and ongoing) at the time of filing for tariff approval for an observeration. The firm is assumed to believe that delay is exponentially distributed with rate parameter λ. At a particular point in time, let there be N current and past durations t_i for the firm to observe in the same regulatory regime, $M \leq N$ of which are completed. Let S_1 be the index set for the completed spells and S_2 be the same for the ongoing spells. The likelihood function for the data is then

$$p(t_1, \ldots, t_N|\lambda) = \left(\prod_{i \in S_1} \lambda \exp(-\lambda t_i)\right)\left(\prod_{i \in S_2} \exp(-\lambda t_i)\right) = \lambda^M \exp(-\lambda T)$$

where $T = \sum_{i=1}^{N} t_i$. The firm's expectations at the beginning of regulatory period r in state s are summarized by a prior distribution for λ, $p(\lambda)_{rs} \sim \mathcal{G}(\alpha, \mu_{rs})$, where $\mathcal{G}(\alpha, \mu)$ is the Gamma density:

$$\mathcal{G}(\alpha, \mu) = \mu^\alpha \lambda^{\alpha-1} \exp(-\mu\lambda)/\Gamma(\alpha)$$

The prior expected duration can be shown to be $\mu/(\alpha-1)!$. Using Bayes' rule, the posterior density for λ is distributed $\mathcal{G}(M+\alpha, T+\mu)$, and the posterior expected duration of regulatory delay is $(T+\mu)/(M+\alpha-1)$. I set $\alpha = 2$ for ease of interpretation, so that the posterior expected duration is an average of the spells with the prior contributing an "observation" of length μ (the prior expected duration). It remains to choose μ_{rs}. In each regulatory regime there was a modal delay, observed with great frequency, that reflected the "usual" regulatory delay. Choosing these values for μ, I set $\mu_{1,IL} = 45$, $\mu_{2,IL} = 30$, $\mu_{1,IN} = 90$, $\mu_{2,IN} = 1$, $\mu_{1,WI} = 45$, and $\mu_{2,WI} = 10$.MB/MCfig

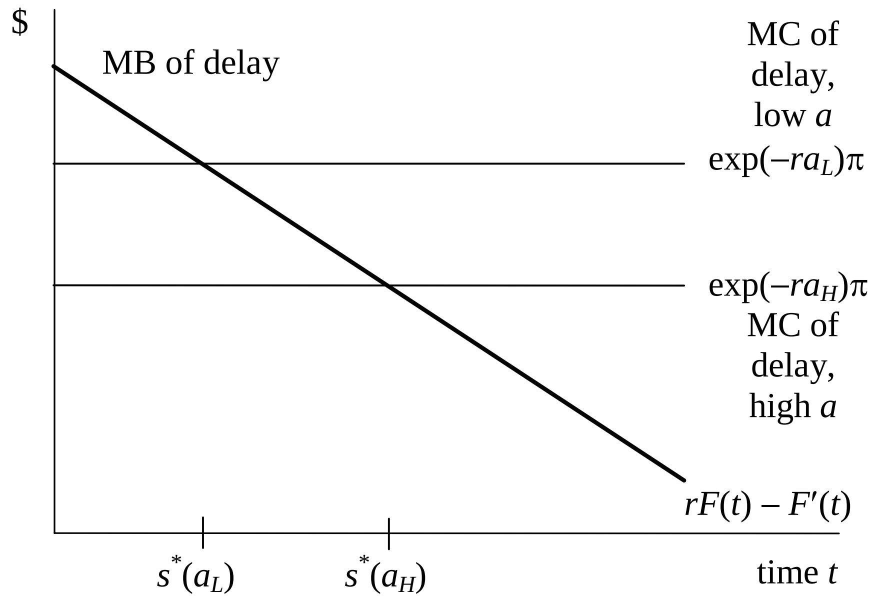

Figure 1: Determination of Optimal Innovation Date

State	**Sample** N	Product Introduction Delay Before Regulatory Change, Period 1 (1991–mid 1994) *min*	*mean*	*median*	*max*	N	Product Introduction Delay After Regulatory Change, Period 2 (mid 1994–1999) *min*	*mean*	*median*	*max*	N
IL	95	0	128	34	665	34	0	45	0	503	62
IN	77	0	457	199	2605	29	0	159	45	1,318	48
OH	65	361	1,267	1,235	2,518	8	0	98	26	1,071	62
WI	106	0	357	150	2,441	40	0	100	18	1,667	66
Total	349					111					238

Table notes: figures are in days. See text for calculation of product introduction delay s.

Table 1: Change in Product Introduction Delay Between Periods

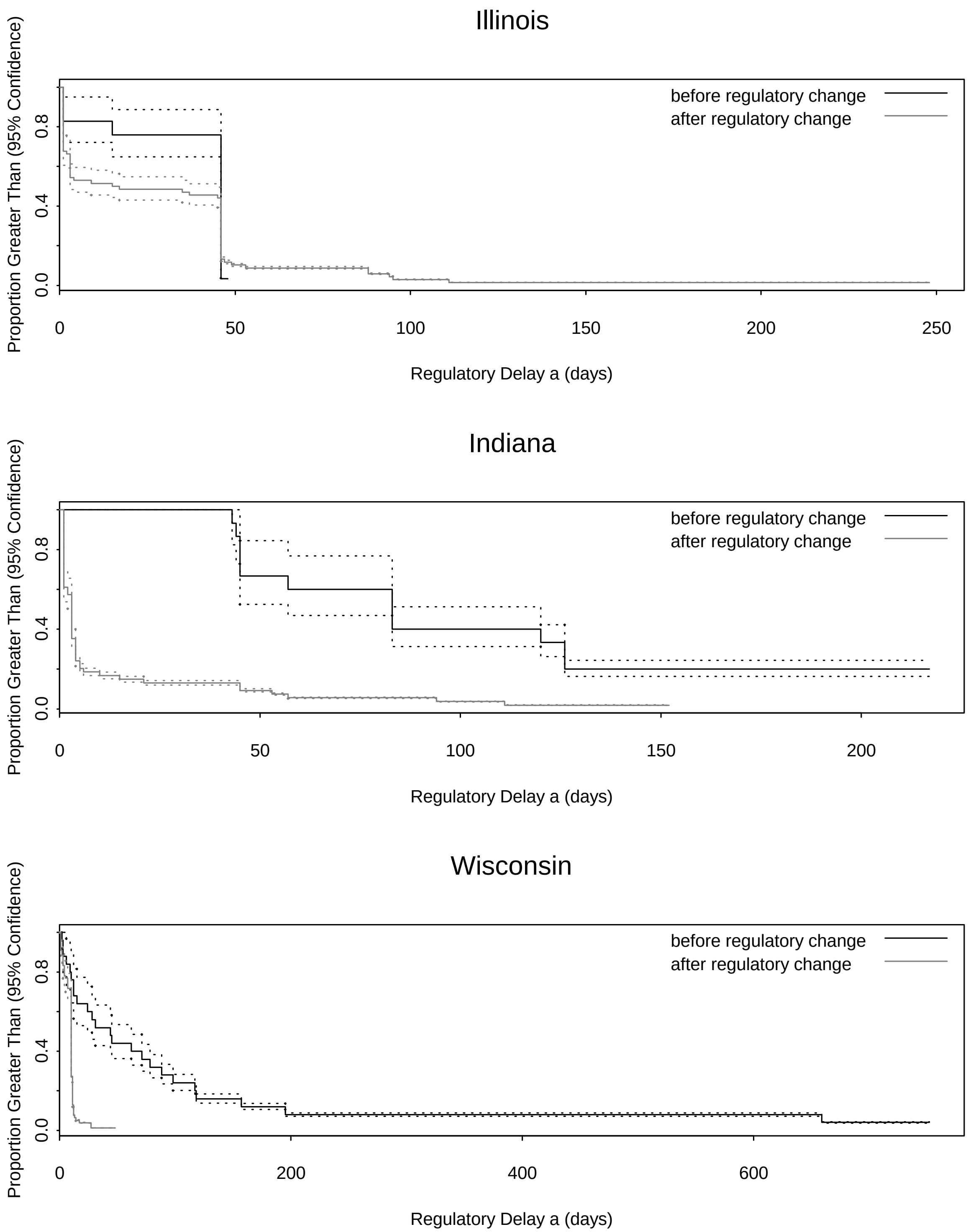

Figure 2: Nonparametric survival curves for regulatory delay a

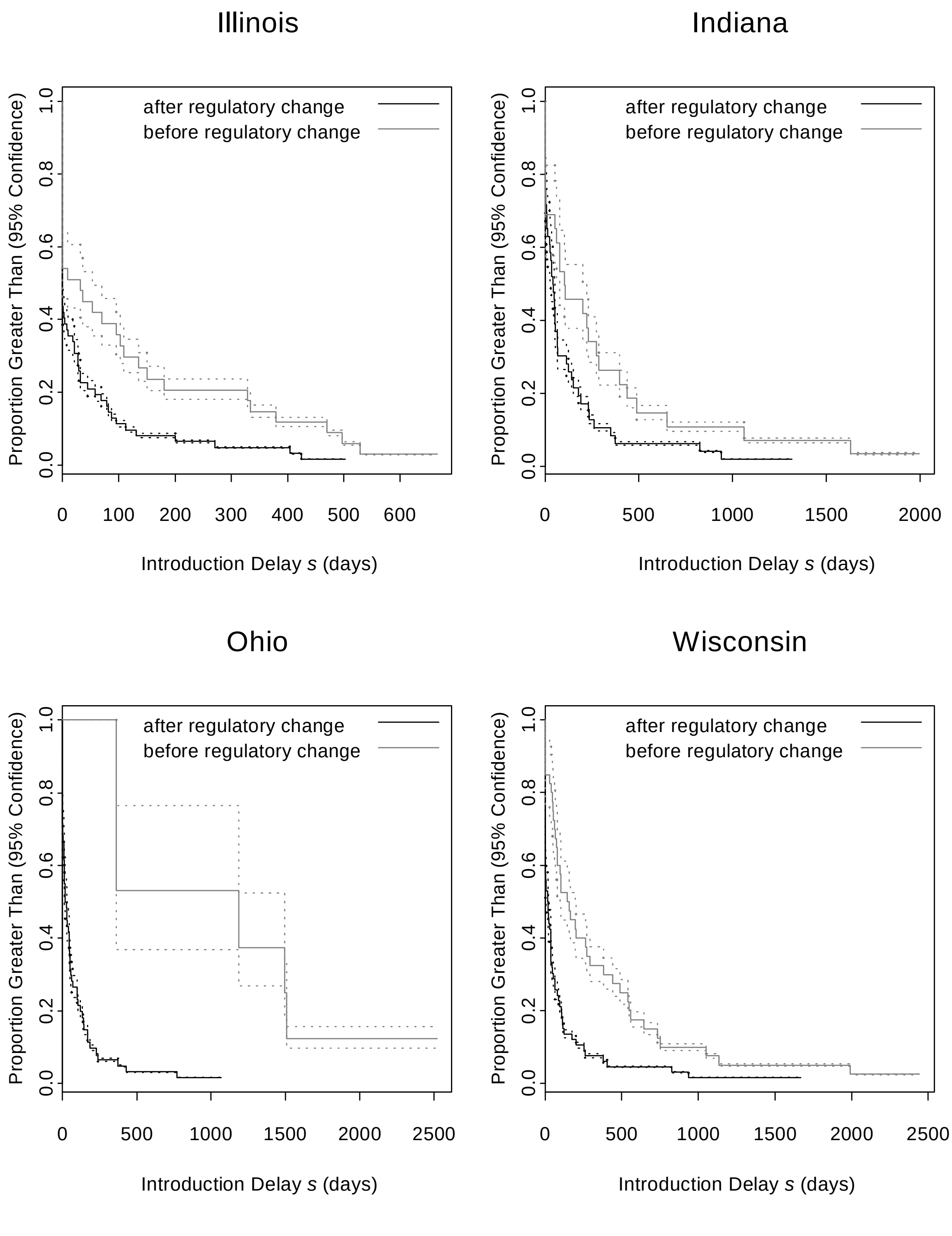

Figure 3: Nonparametric survival curves for introduction delay s

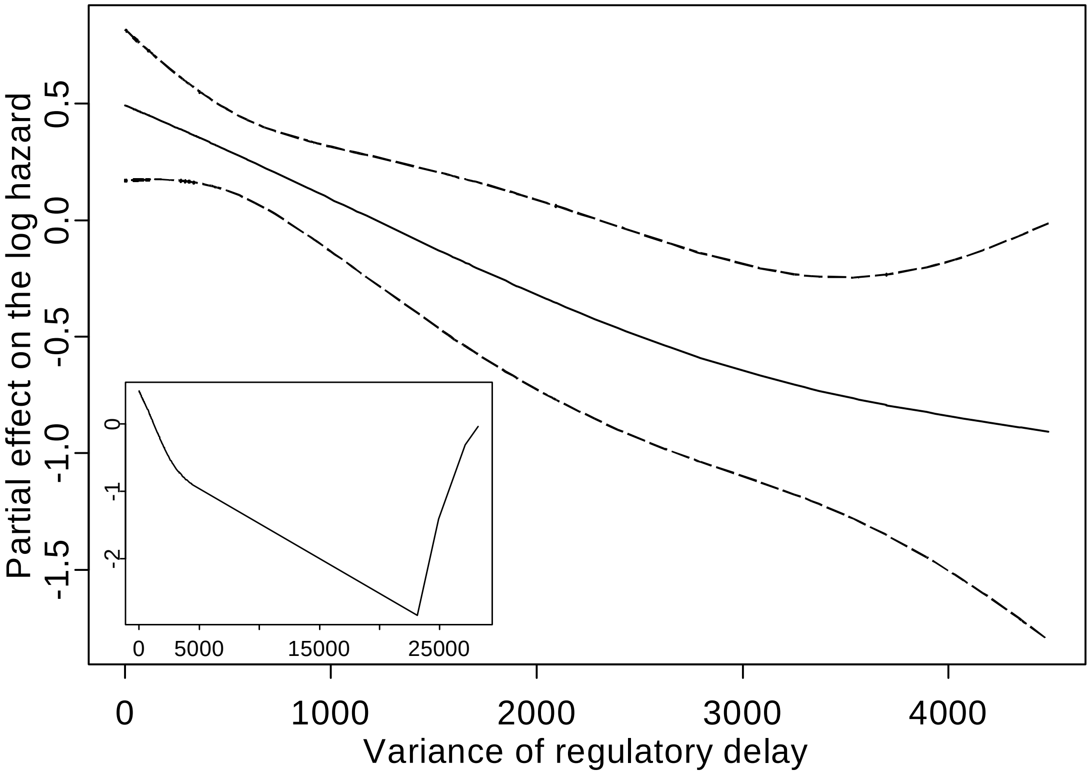

Figure 4: Effect of regulatory riskiness on introduction delay s

	Sample	**Regulatory Delay Before Regulatory Change** Period 1 (1991–mid 1994)					**Regulatory Delay After Regulatory Change** Period 2 (mid 1994–1999)				
State	*N*	*min*	*mean*	*median*	*max*	*N*	*min*	*mean*	*median*	*max*	*N*
IL	97	1	36	46	48	29	1	30	16	248	68
IN	69	43	103	83	217	15	1	13	3	152	54
WI	103	2	106	44	752	25	1	9	10	48	78
Total	269					69					200

Table notes: figures are in days. Regulatory delay data are not available for Ohio.

Table 2: Change in Regulatory Delay Between Periods

State	Period	Mean	(s.e.)	Median	Lower 95% conf. limit for median	Upper 95% conf. limit for median
IL	1	126.5	(32.4)	32	0	53
	2	45.0	(13.1)	0	0	0
IN	1	320.7	(97.6)	106	76	221
	2	133.5	(38.2)	42	28	53
OH	1	1,413.0	(309)	1,493	0	1,493
	2	87.3	(23.2)	19	13	32
WI	1	356.6	(81.6)	143	77	201
	2	99.5	(31.4)	19	3	23

Table notes: figures (in days) are based on survival curve estimates (see Figure 3). period 1 is 1991 to mid 1994, period 2 is thereafter.

Table 3: Estimated product introduction delay s

State	Period	Mean	(s.e.)	Median	Lower 95% conf. limit for median	Upper 95% conf. limit for median
IL	1	36.2	(3.3)	46	46	46
	2	30.1	(4.7)	17	3	45
IN	1	103.4	(16.4)	83	57	120
	2	13.0	(4.0)	3	3	3
WI	1	105.6	(36.9)	44	24	62
	2	9.5	(0.7)	10	10	10

Table notes: figures (in days) are based on survival curve estimates (see Figure 2). period 1 is 1991 to mid 1994, period 2 is thereafter. Regulatory delay data are not available for Ohio.

Table 4: Estimated regulatory delay a

	Cox Proportional Hazards Models for Product Introduction Delay					
	Estimation I1		Estimation I2		Estimation I3	
	coef.	*s.e.*	*coef.*	*s.e.*	*coef.*	*s.e.*
IL:reg change	0.358*	0.212	0.302	0.211	0.079	0.285
IN:reg change	0.689**	0.276	0.673**	0.279	0.709**	0.312
WI:reg change	1.334***	0.225	1.378***	0.222	0.848***	0.277
IN	−0.584**	0.288				
OH	0.149	0.203				
WI	−1.001***	0.251				
Federal tariff first	−0.690***	0.141	−0.737***	0.148	−0.603***	0.134
Household income					3.053	3.005
Access lines					1.260***	0.470
Area					0.606	0.510
Telecom patents					−0.392	0.835
Stratification	none		state		none	
N	349		349		349	
χ^2 statistic (d.o.f.)	81.9(7)	$p = 0.00$	62.2(4)	$p = 0.00$	83.5(8)	$p = 0.00$
$T(G)$ statistic (d.o.f.)	1.88(7)	$p = 0.97$	0.74(4)	$p = 0.99$	16.3(8)	$p = 0.04$
Log likelihood	−1621.1		−1167.6		−1617.6	

* = 10% level significance; ** = 5% level significance; *** = 1% level significance.

Table notes: The model incorporates time-varying covariates. S.e.'s are calculated with the Huber-White robust sandwich formula. Larger positive coefficients imply shorter delays. Excluded state dummy is Illinois. *Federal tariff first* is an indicator for introduction delays calculated from the initial date the service was filed in the Federal Access Tariff; other delays calculated from the date of the first filing in a state tariff (with first state's delay changed from 0 to 0.5). *Household income* is the average for Ameritech's territory within a state, in log real terms. *Access lines* is for Ameritech's subscribers in the state, logged. *Area* is for Ameritech's territory within a state, logged. *Telecom patents* is the log count of patents approved in the classes relevant to telecommunications services (359, 370, 379, and 395) in the calendar year, based on application date. χ^2 statistic is for the null hypothesis that all coefficients are zero. Figures in parentheses are degrees of freedom. $T(G)$ statistic is for a global test of the proportional hazards assumption and has a χ^2 distribution; rejection would indicate that the model is misspecified (test 4 of Grambsch and Therneau (1994)).

Table 5: Semiparametric estimation results for product introduction delay s

	Cox Proportional Hazards Models for Product Introduction Delay			
	Estimation I4		Estimation I5	
	coef.	s.e.	coef.	s.e.
IL:regulatory change	0.484*	0.277	0.200	0.292
IN:regulatory change	1.641**	0.713	1.164	1.410
WI:regulatory change	1.842***	0.647	1.137	1.384
Federal tariff first	−0.785***	0.189	−0.722***	0.192
Access lines	1.065***	0.270	0.869***	0.255
IL:regulatory uncertainty	−1.897*	1.040	−0.848	1.034
IN:regulatory uncertainty	−0.404	1.459	−4.800**	2.119
WI:regulatory uncertainty	−0.101	0.210	−0.327*	0.196
Average regulatory delay	0.017	0.016	p-splined†	
Stratification	none		none	
N	255		255	
χ^2 statistic (d.o.f.)	85.2(9)	$p = 0.00$	183.43(11.1)	$p = 0.00$
$T(G)$ statistic (d.o.f.)	1.77(9)	$p = 1.00$	5.57(20)	$p = 1.00$
Log likelihood	−1096.8		−1089.9	

* = 10% level significance; ** = 5% level significance; *** = 1% level significance.

† *Average regulatory delay* enters specification I5 flexibly through cubic p-splines (Eilers and Marx, 1996).

Table notes: *Regulatory uncertainty* is the variance of regulatory delay as described in the text, divided by $10,000$. See also notes to previous table.

Table 6: Semiparametric estimation results for product introduction delay s

Panel A: Evidence that introduction delay fell when structural regulatory delay fell			
	Estimation I1	Estimation I2	Estimation I3
	p-value	*p-value*	*p-value*
H_0: $\beta_j \leq 0$ vs. H_A: $\beta_j > 0$			
IL:reg change	0.045	0.076	0.391
IN:reg change	0.006	0.008	0.011
OH:reg change	*	*	*
WI:reg change	0.000	0.000	0.001
Joint test: H_0: $\boldsymbol{\beta} = 0$ vs. H_A: $\boldsymbol{\beta} \neq 0$	0.000	0.000	0.000

Panel B: Evidence that introduction delay is positively correlated with regulatory uncertainty		
	Estimation I4	Estimation I5
	p-value	*p-value*
H_0: $\beta_j \geq 0$ vs. H_A: $\beta_j < 0$		
IL:regulatory uncertainty	0.034	0.205
IN:regulatory uncertainty	0.391	0.012
WI:regulatory uncertainty	0.315	0.048
Joint test: H_0: $\boldsymbol{\beta} = 0$ vs. H_A: $\boldsymbol{\beta} \neq 0$	0.136	0.014

*None of the delay durations in Ohio are completed in the period before the change in regulatory regime, and so $\hat{\beta}$ for the OH:reg change variable would be $+\infty$ and the hypothesis test is moot.

Table notes: One-sided tests are computed with one-sided t tests. Joint tests are computed with Wald tests.

Table 7: Hypothesis Tests of Outcomes That Would Reject the Theoretical Model