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Product Innovation, Signaling, and Endogenous Regulatory Delay

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Abstract This paper examines the determinants of the timing of a monopolistic firm's product innovation and regulatory approval, and proposes a signaling model with endogenous regulatory delay. Regulatory delay exerts a multiplier effect on total time to market, because when the firm expects the regulator to take longer to grant approval, the firm delays its product introduction. The firm can time its innovation to communicate its private information about the marginal cost of delay to the regulator. Successful signaling in the separating equilibrium leads the regulator to reduce regulatory delay. The implications of the model are consistent with data on innovation and regulatory delay in telecommunications markets in a few Midwest states in the U.S.

JEL Codes: L51, L96

Keywords: innovation delay, regulatory delay, telecommunications, Cox duration model

1 Introduction

Regulated firms often claim that regulatory review and approval of new products is costly and distorts the incentives to innovate. Examples of regulatory delay are FDA approval of new drugs and line-of-business restrictions in banking before deregulation. I explore the relationship between regulatory delay and the timing of the firm's innovation, a connection largely overlooked by the literature on regulation and innovation. I focus on the telecommunications industry. Delayed introduction of new telecommunications services can impose large costs on society (Hausman, 1997) and it is important to understand the determinants of delay. The model I develop recognizes that the firm may influence regulatory delay. The timing of innovation by the firm may reveal information about the cost of delay to the regulator, which might adjust regulatory delay in response. The implications of the model are consistent with data on innovation and regulatory delay in telecommunications markets in a few Midwest states.

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Both firms and regulators contribute to the delay between technological feasibility of a product and its introduction to consumers. The time between the first technologically feasible introduction date and the submission of the product to the regulator is *innovation delay*. The time between the firm's submission of a new product to the regulator for approval and the granting of approval is *regulatory delay*. The data examined here show that regulatory delay is positively correlated with innovation delay at the level of the individual product. There are many potential explanations of the correlation. The explanation developed here is that regulators delay a service based on the cost of so doing (other explanations are considered in Section 6). Whether the regulator has full or incomplete information regarding the cost of delay to the firm, there is a positive association between regulatory and innovation delay in equilibrium. In the asymmetric information model, the firm signals its private information to the regulator through its innovation delay.

In the model, the regulator trades off the benefit of reducing delay (quicker return on investment for the firm and earlier accrual of benefits for consumers) and the costs (loss of regulatory control, potentially lower quality of service, harm to competing firms, and the like). Examples of the cost to the regulator of quick introduction include the fallout from consumer confusion from introducing a product before agreement on a technical standard or loss of payoffs (perhaps intangible) from interest groups opposing the firm. These costs enter the model as a regulatory "taste for delay".[1] Instead of looking at the political economy of regulatory delay, I take the regulator's preferences as exogenous and merely note that given the lengthy regulatory delay exhibited in the data analyzed here, there must be a strong taste for delay.

The trade-off between the costs and benefits of regulatory delay depends in part on the cost that regulatory delay imposes on the firm. The firm is likely to know better the cost of delay than the regulator does. Unless the firm has the lowest possible delay cost, it would like to communicate its private information to the regulator. The firm can signal its cost of delay with an action that a firm with different cost cannot profitably mimic. A costly action available to the firm is innovation delay. In particular, innovating sooner than the full-information optimal delay signals the regulator that cost is high. Of course, the firm would also attempt to convince the regulator that the product is important through other forms of communication such as cost and demand reports. However, if the cost to the firm of sending such reports does not vary with the true type of the firm and verification is difficult, they may not be credible.

This article breaks new ground by jointly modeling the determination of regulatory and innovation delay. The earliest literature on regulation and the timing of innovation looked at a monopolist's incentive to innovate given a fixed regulatory regime (Braeutigam, 1979).[2] More recent work focuses on the strategic timing of innovation under different regulatory regimes (Riordan, 1992; Lyon and Huang, 1995), but does not explicitly consider regulatory delay. The empirical literature on how regulatory delay affects innovation and product introduction is small. It includes Gruber and Verboven (2001), Prieger (2001, 2002a, 2002b), and Hazlett and Ford (2001), all for the telecommunications industry, Prager (1989) for electricity generation, and Sanyal (2003) for the electronics and three other industries. In work related to the present study, Prieger (2007) shows that reduction in average regulatory delay contributed toward speedier product introductions. In that study, unlike the present one, regulatory delay is treated

[1] Sometimes delay stems from judicial rather than regulatory action. The arbiter of the Modified Final Judgement (MFJ) that split up AT&T, Judge Harold Greene, prohibited the dominant phone companies from offering information services even longer than the FCC wished. See chapter 12 of Brock (1994) for evidence that loss of regulatory control enters negatively into at least some objective functions.

[2] The literature review here and the description of the data in Section 4 draw on Prieger (2007).

as fixed or random. There are two differences between the extant literature and the present work. Previous studies all focus on aspects other than asymmetric information and signaling. Furthermore, in the empirical studies among the above, regulatory delay is an explanatory variable, rather than the dependent variable as modeled here. In one of the few other papers on signaling in a regulated environment, Spiegel and Wilkie (1996) consider a model in which firms signal to the capital market (not the regulator) through investment in a new technology. The model here thus offers a new role that signaling can play in regulated markets.

In the next section, I present the theoretical model. I give an example highlighting how the firm's and the regulator's strategies are formed in Section 3. In Section 4, I present the testable implications derived from the signaling model and introduce the data from a large incumbent local exchange telephone company. I test the predictions in Section 5 and show that the signaling model is consistent with the observed patterns of innovation delay and regulatory delay. I consider alternative explanations in Section 6, and conclude by discussing welfare implications and possible extensions to the model.

2 The Model

The firm's side of the model is an extension of that in Prieger (2007), to which I add heterogeneity of firms and endogenous regulatory delay. The firm first chooses its innovation date and then the regulator chooses an amount of additional delay. Time $t = 0$ is when a firm can first feasibly introduce a given product. Both the firm and the regulator observe t, which is required for innovation delay to function as a signal to the regulator. While in some settings the regulator may not know when products are technologically feasible, I argue below that the regulator can observe a proxy for t in the data examined here. The firm chooses to submit 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, adoption, or regulatory filing, and (as in Riordan (1992)) is a decreasing function of time. In dynamic industries such as telecommunications, it is realistic to assume that technological advances lower the cost of implementing the new service over time. I assume $F'(t) < 0$ and $F''(t) > 0$, and all functions in the model are assumed to be continuous and differentiable as needed. Falling fixed costs give the firm a reason to delay innovation.

The regulator approves the service after regulatory delay of length a. Regulatory delay may include time spent getting on the regulator's docket, waiting for a monthly review meeting, and mandatory examination periods. The firm offers the good for purchase at introduction time $s + a$. I assume that the regulator cannot commit to a policy a before the firm moves.[3] The firm earns flow profit of $\pi(\theta)$ per unit time after introduction, where θ is a parameter that may be private information of the firm.[4] There is a continuum of possible types, drawn from a compact set: $\theta \in [\theta^-, \theta^+] \subset \mathbb{R}$. I assume that $\pi'(\theta) > 0$, so larger θ might correspond to higher demand or to lower marginal costs. In order to focus on the strategic variable s, the determination of price (and therefore π) is taken to be unrelated to the introduction time in the model and is not modeled. The firm's net present value of introduction at time s, given discount rate is r, is:

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

[3] Lack of commitment is a common assumption in regulatory games (outside of the mechanism design literature). See Spiegel and Spulber (1997) for a justification.

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

I assume that for all firm types an interior maximum of Π in s exists for all a.[5] Since delay postpones the accrual of flow profit, and the firm's type is positively related to π, the firm's private information about θ can be interpreted as information about the firm's opportunity cost of delay.

Regulatory delay has no benefit for the firm in this model:

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

From (1), $\partial\Pi/\partial a = -e^{-r(s+a)}\pi(\theta) < 0$. Although one can conceive of situations in which regulatory delay can turn out to be beneficial *ex post*, the firm views regulatory delay as purely detrimental. Given that a firm could always choose on its own to delay introducing a product, it is hard to imagine realistic cases where regulatory delay enhances the profit of the firm.

Timing of the game The firm first chooses s. The regulator observes s, updates its beliefs about the firm's type θ, and chooses delay a. Because the regulator makes its decision after the firm's, the firm can signal its type to influence regulatory delay.

The regulator's objective function may represent either social welfare (the "benevolent dictator" framework) or the utility function of the regulator (the "economic theory of regulation" approach to regulation (Peltzman, 1976)). Utility at time s when the regulator believes the firm to be of type $\hat{\theta}$ is

$$U(\hat{\theta}, a) = W(a, \hat{\theta}) + V(a) \tag{2}$$

The first part of the utility function, W, comes from the profit of the firm and the consumers' surplus. In the simplest case, W is the sum of the present discounted value of total welfare. Other transformations of profit and consumers' surplus are allowed, but it is assumed that $\partial W/\partial a < 0$, $\partial W/\partial\hat{\theta} > 0$, $\partial^2 W/\partial a^2 > 0$, and $\partial^2 W/\partial a\partial\hat{\theta} < 0$. These assumptions are consistent with $W = \Pi + CS$, where CS is the present discounted value of a constant surplus flow that is increasing in θ. The firm's type affects CS at least indirectly because the monopoly prices charged are a function of θ. If θ represents a demand parameter, then θ will also have a direct impact on CS.

Crucial to the model is V, the benefit to the regulator from regulatory delay, with $V' > 0$ and $V'' < 0$. The interpretation of V varies with the interpretation of the regulator's objective. In a benevolent dictator setting, V may represent benefits not reflected in CS as defined above from higher quality or lower level of externalities that may result from regulatory delay. If delay represents the time taken by the firm to bring the product up to a regulatory quality standard, then longer delays may increase product quality. If delay represents time taken by the regulator to investigate safety or privacy concerns (e.g., caller ID or caller ID blocking), then longer delays may decrease externalities. In these cases, CS is read as surplus conditional on a fixed level of quality or externalities, and V subsumes all benefits of delay. In a political economy setting, V might represent a preference for exercising authority or direct or indirect payoffs to the regulator from the firm's rivals (although any such rivals are not modeled explicitly here). This "taste for delay" in the model, although *ad hoc*, is clearly realistic. Examination of the data below shows that regulatory delay is often quite lengthy in the real world, and therefore regulators must perceive there to be benefits of some sort to delay. For modeling purposes, all that is needed is a regulatory objective function that leads to non-zero regulatory delay.

[5] I assume $rF(0) - F'(0) > e^{-ra}\pi(\theta)$ for all θ and a to guarantee positive optimal innovation delay s^*, and that $\lim_{t\to\infty} rF(t) - F'(t) \leq 0$ to guarantee finite s^*.

For simplicity, the benefit to the regulator from regulatory delay is not a function of the firm's type. While the profitability of the proposed service may realistically affect V, especially if payoffs from the firm's rivals to the regulator are involved, making V depend explicitly on $\hat{\theta}$ complicates some proofs unnecessarily.[6] Finally, it is required that the concavity of V be great enough in magnitude so that $\partial^2 U/\partial a^2 < 0$. This assumption, for technical convenience, assures that the relevant single-crossing condition holds. Note finally that U is forward looking or "memoryless" in the sense that s does not affect U. This assumption means the regulator treats innovation delay as a bygone by the time its decision is to be made, and simplifies some of the results but is not intrinsic to the argument.

Solution concept I restrict focus in this Spence-type signaling game to cases of successful signaling: sequential separating equilibria. As will be shown, equilibrium in the model is unique and consists of pure strategies, and so I do not discuss mixed strategies here. Equilibrium consists of the firm's one-to-one strategy $\sigma(\theta)$ for s, and the regulator's strategy $\alpha(\hat{\theta}, s)$ for a, and the regulator's posterior beliefs $\hat{\theta}$ about θ such that

- $\sigma(\theta)$ maximizes $\Pi(\theta, s, a)$ given the firm's correct expectation that $a = \alpha(\theta, s)$,
- $\alpha(\hat{\theta})$ maximizes $U(\hat{\theta}, a)$ given the posterior beliefs and the regulator's correct expectation that $s = \sigma(\hat{\theta})$, and
- $\hat{\theta} = \theta$ on the equilibrium path.[7]

Mailath (1987) shows that when type space is continuous, a unique separating equilibrium exists if certain technical conditions are met, which I discuss below and in the appendix.

The regulator's strategy Because equilibrium is sequentially rational, we may use backward induction to solve the game. The regulator will choose a as

$$\alpha(\hat{\theta}) = \underset{a}{\operatorname{argmax}} U(\hat{\theta}, a) \tag{3}$$

when it believes the firm is type $\hat{\theta}$. Assuming an interior solution ($\alpha > 0$), the optimal choice of regulatory delay α equates the marginal benefit of delay for the regulator with the regulator's marginal cost of delay:

$$V'(\alpha) = -\left.\frac{\partial W}{\partial a}\right|_{a=\alpha} \tag{4}$$

Applying the implicit function theorem to equation 4 implies that

$$\frac{d\alpha}{d\hat{\theta}} = -\left.\frac{\partial^2 W}{\partial \hat{\theta} \partial a}\right/ \frac{\partial^2 U}{\partial a^2} \tag{5}$$

which, by the assumptions above, is negative. Figure 1 shows a typical case. The firm knows it will receive lower regulatory delay the higher the regulator thinks θ is. Thus, the worst belief the regulator can hold, from the firm's point of view, is that $\theta = \theta^-$. All firm types other than θ^- therefore wish to signal to the regulator to avoid the worst outcome, $\alpha(\theta^-)$.

[6] If V depends on $\hat{\theta}$, then as long as $\partial^2 U/\partial\hat{\theta}\partial a < 0$, no results change.

[7] More formally, strategies and beliefs are *sequentially rational* and *consistent* in the sense of Kreps and Wilson (1982). Sequential equilibrium in this game may impose more restrictions on play off the equilibrium path than does the more familiar perfect Bayesian equilibrium, because there are more than two types (see Thm. 8.2 of Fudenberg and Tirole (1991)).

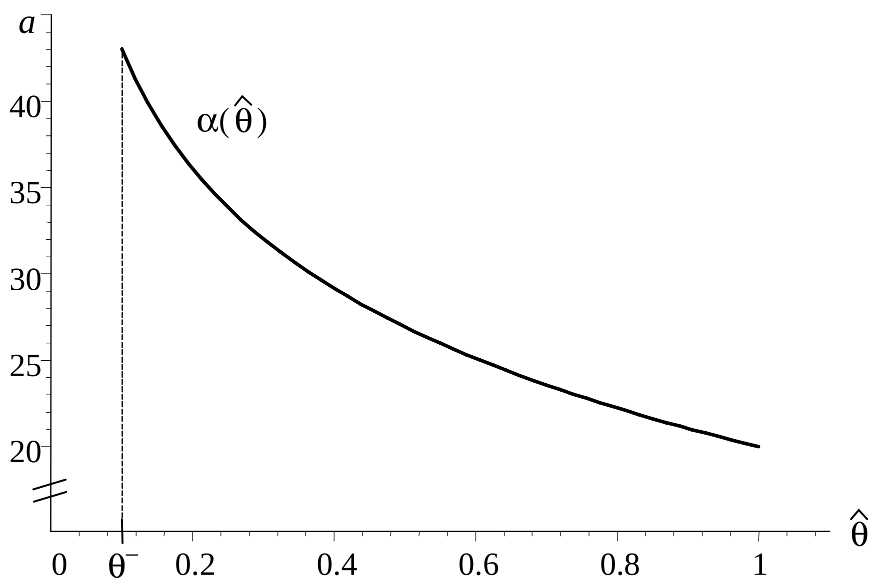

Fig. 1 The regulator's optimal strategy for regulatory delay

The firm's strategy Following Mailath (1987), define the firm's concentrated profit function as

$$\tilde{\Pi}(\theta, \hat{\theta}, s) = \Pi(\theta, s, \alpha(\hat{\theta})) \tag{6}$$

The function is concentrated in the sense that $\tilde{\Pi}$ incorporates the optimal action of the regulator. Given the assumptions, $\tilde{\Pi}$ satisfies a single crossing condition for θ:

$$\frac{\partial\tilde{\Pi}(\theta, \hat{\theta}, s)/\partial s}{\partial\tilde{\Pi}(\theta, \hat{\theta}, s)/\partial\hat{\theta}} \text{ is strictly monotonically decreasing in } \theta. \tag{7}$$

The single crossing condition for θ, proved in the appendix, means that strengthening the signal is costlier for firms with higher costs of delay. The condition implies that the firm's strategy will be monotone in θ.

If the firm's type were observable to the regulator, then the firm would choose its optimal innovation delay as

$$s^* = \tau(\theta) = \underset{s}{\operatorname{argmax}}\ \tilde{\Pi}(\theta, \theta, s) \tag{8}$$

The function τ is the full-information benchmark strategy for innovation delay.

Given that the firm's type is likely to be private information, the firm may wish to communicate its type to the regulator. Looking ahead to the regulator's policy $\alpha(\hat{\theta})$, the firm wishes to signal its type when doing so will cause the regulator to reduce regulatory delay from $\alpha(\theta^-)$. Thus, $\tau(\theta) = \sigma(\theta)$, where σ is the firm's signaling strategy for s, only at $\theta = \theta^-$. The type with the lowest cost of delay has no incentive to signal, which provides an initial value condition needed to solve for the equilibrium strategy below.

If $\sigma(\theta)$ is part of a separating equilibrium, it must be one-to-one and incentive compatible. Incentive compatibility requires that the firm maximize $\tilde{\Pi}$ recognizing that the regulator will correctly infer its type (in equilibrium): if the firm chooses delay s, the regulator will correctly believe that the firm's type is $\sigma^{-1}(s)$. Mathematically, incentive compatibility requires that

$$\sigma(\theta) = \operatorname{argmax}_{s \in \sigma([\theta^-, \theta^+])} \tilde{\Pi}(\theta, \sigma^{-1}(s), s) \tag{9}$$

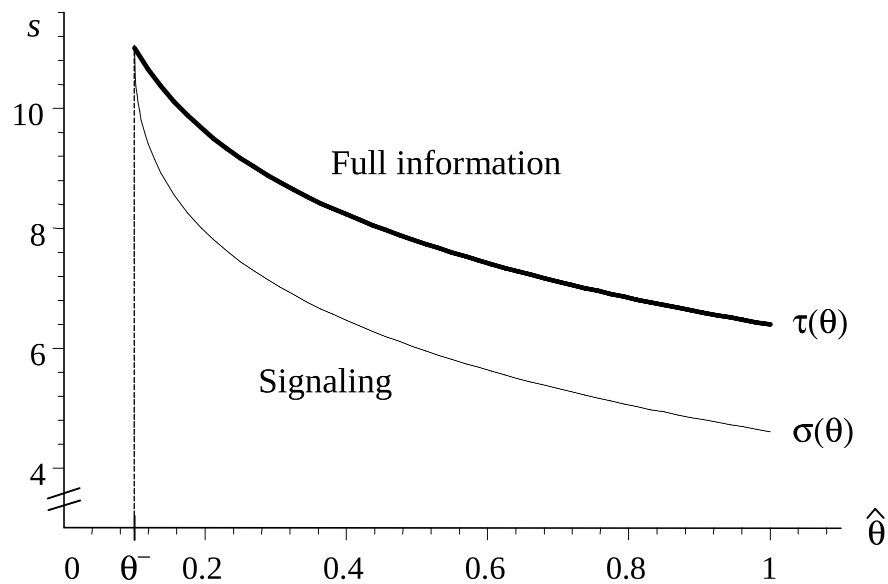

Fig. 2 The firm's optimal strategy for innovation delay

Mailath (1987) shows that under condition (7) and other regularity conditions satisfied here (see the appendix) a unique, continuous, strictly monotonic pure strategy $\sigma(\theta)$ exists. The firm's strategy may be found as the solution to an ordinary differential equation:

$$\frac{d\sigma}{d\theta} = - \left. \frac{\partial \tilde{\Pi}(\theta, \hat{\theta}, s)/\partial \hat{\theta}}{\partial \tilde{\Pi}(\theta, \hat{\theta}, s)/\partial s} \right|_{\hat{\theta}=\theta, s=\sigma} \tag{10}$$

$$\sigma(\theta^-) = \tau(\theta^-) \tag{11}$$

Given the assumptions of the model, $d\sigma/d\theta < 0$ and σ is below τ to the right of θ^- (see Proposition 4 in the appendix). In Figure 2, which shows a typical case, the firm's full-information strategy is the heavy line and the signaling strategy is the lighter line below. The economic interpretation of the firm's behavior at θ^-, where the slope of σ approaches infinity, is that types marginally higher than θ^- must decrease innovation delay a lot to differentiate themselves from the worst type. The derivative of σ, although still negative, is not as large for higher types. In all cases, however, the innovation delay chosen by the firm is less than that chosen in the full-information case. This is the cost of signaling for the firm. As one expects in a signaling model, the firm earns less profit compared to the full-information case. Note, however, that consumers receive the new service earlier when the firm has private information.

Since $d\sigma/d\theta < 0$, it follows that a higher opportunity cost of delay induces the firm to innovate earlier. This result is not caused by the need to signal; it can be shown (or seen from Figure 2) that $d\tau/d\theta < 0$ also. At first, this result might appear counterintuitive; if regulation is undesirable for the firm, why do higher marginal costs of regulatory delay lead to *earlier* innovation? The answer requires distinguishing between the direct and opportunity costs of regulation. Type θ measures the opportunity costs of regulation; as the forgone profit from delay increases, the firm innovates earlier to speed accrual of those profits. If the direct cost of the regulatory process were included as a constant in F, then an increase in direct cost would postpone innovation since the marginal benefit of delay is higher.

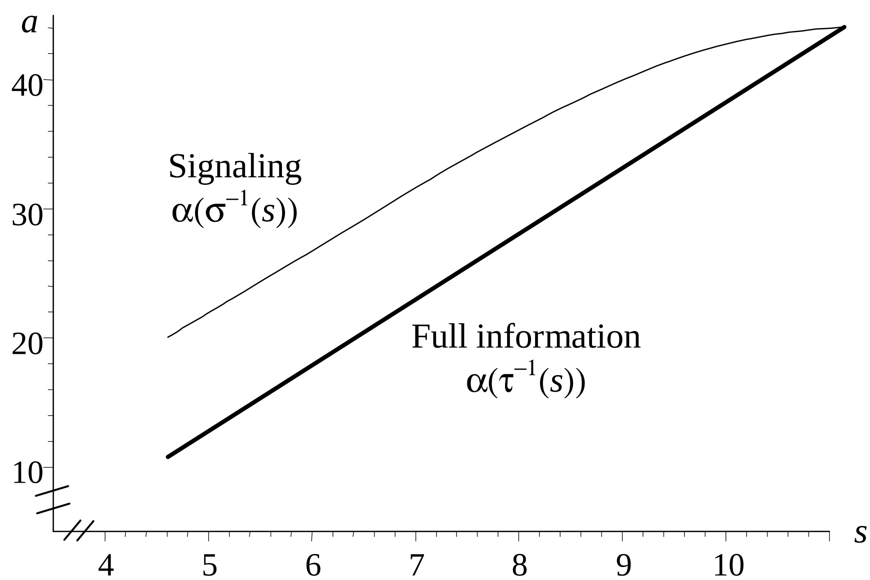

Fig. 3 The equilibrium relationship between regulatory and innovation delay

The central result is that innovation delay and regulatory delay move together in equilibrium:

Proposition 2 *$da/ds > 0$ in equilibrium. Regulatory delay is positively associated with innovation delay in the equilibrium of the signaling model.*

The proposition follows formally from the theory of supermodular games,[8] which shows that s and a are both decreasing in the firm's type, and which has already been shown above. Thus, in equilibrium, low types lead to high regulatory and innovation delay, and high types lead to low innovative and regulatory delay. The empirical implication is the following. Innovation delay and discretionary regulatory delay will be correlated in a sample of observations on the outcome of this one-shot game, since (ceteris paribus) s and a would move together.

In addition to positive association between innovation delay and regulatory delay, the model also predicts that the relationship is concave:

Proposition 3 *da/ds is decreasing in a neighborhood to the left of $\alpha(\theta^-)$ in equilibrium. Regulatory delay is concave in innovation delay in the equilibrium of the signaling model, at least in that region.*

The proof is in the appendix. The proposition states that marginal increases in innovation delay prompt diminishing marginal increases in regulatory delay, as can be readily seen in Figure 3. Concavity of regulatory delay is thus a necessary implication of signaling, and may be used to distinguish signaling from full-information behavior in an empirical investigation, since concavity need not hold in the full-information case.

[8] See, e.g., Topkis (1998), Lemma 4.2.2.

3 An Example

Consider the following example to illustrate the results of the signaling model. Assume these functional forms

$$W(a,\hat{\theta}) = \hat{\theta}\exp(-r\,[\bar{a}+a]) \tag{12}$$

$$V(a) = -\exp(-2ra)/2 \tag{13}$$

$$\pi(\theta) = r\theta \tag{14}$$

$$F(s) = e^{-s} \tag{15}$$

which satisfy the needed conditions given above. The parameter $\bar{a}$ in (12) represents a minimum level of delay required of all services. Thus, $\bar{a}$ represents structural, exogenous delay, which is fixed before the game begins.

The regulator's best response is found from (4) as

$$\alpha(\hat{\theta}) = \bar{a} - \frac{\ln\hat{\theta}}{r} \tag{16}$$

and the firm's best response under full information is found from (8) as

$$\tau(\theta) = 2r\bar{a} + \ln\frac{r+1}{r\theta^2} \tag{17}$$

When signaling is necessary in the separating equilibrium, (10) and (11) define the firm's strategy. This initial value problem does not have an analytic solution, but is readily solved by numerical methods. For example, set $\theta^- = 0.1$, $\theta^+ = 1$, $\bar{a} = 20$, and $r = 0.1$. The regulator's strategy for this case is the one depicted in Figure 1, the firm's full-information (the heavy line) and signaling strategies are the ones depicted in Figure 2, and the equilibrium relationship between regulatory and innovation delay is the one shown in Figure 3. The relationship between a and s for the full-information case is linear in this example:

$$\alpha(\tau^{-1}(s)) = \frac{s+\ln r - \ln(r+1)}{2r} \tag{18}$$

For the signaling case, in the function $\alpha(\sigma^{-1}(s))$ a given regulatory delay results in longer regulatory delay, compared with the full-information case. This follows directly from Figure 2, because regulatory delay depends only on the type of the firm, and a given type signals with an s shorter than it would like to absent signaling. Note that the sign of the relationship between regulatory and innovation delay is positive in both the full-information and signaling cases. However, Proposition 3 applies only to the signaling model: $\alpha(\sigma^{-1}(s))$ is concave but $\alpha(\tau^{-1}(s))$ is not. Therefore, the sign of correlation between regulatory and innovation delay observed in data from a regulated industry may be used to distinguish strategic from non-strategic or random behavior, but concavity must be tested to distinguish between the full-information and signaling cases.

4 Data and Discussion of the Tests

The theoretical model places restrictions on the relationship between regulatory and innovation delay. First, from Proposition (2) regulatory delay rises with innovation delay. Proposition 3 further asserts that regulatory delay is concave in innovation delay. Note that tests of these predictions are non-parametric, in the sense that the tests depend only

State	Sample N	Innovation Delay *min*	*mean*	*median*	*max*	Regulatory Delay *min*	*mean*	*median*	*max*
IL	95	0	82	1	824	1	33	46	248
IN	69	0	206	53	1,996	1	33	3	217
WI	103	0	191	39	2,441	1	33	10	752
Total	267	0	156	32	2,441	1	33	10	752

Table notes: figures are in days. See text for calculation of innovation delay.

Table 1 Summary statistics for delay

on the sign or shape of the correlation between the observed regulatory and innovation delay, and no specific functional forms need be assumed for π, W, or V.

The data are the innovation and introduction dates for every telecommunications service introduced 1991-1999 by Ameritech in Illinois, Indiana, and Wisconsin.[9] These data are a subset of those used by Prieger (2007) for other purposes. Ameritech, one of the Bell regional holding companies and later acquired by SBC (now AT&T), is the dominant local service provider in each of these states, and its intrastate activities are regulated by the state commissions. The firm could offer the new services to subscribers until the state public utility commission granted regulatory approval. Examples of the residential and business services in the data are new voice mail features, virtual networking services, and high-speed transmission services. From 1991 until mid 1994, Ameritech was under rate of return regulation in Illinois and Indiana, while in Wisconsin the firm was allowed some earnings and pricing flexibility (Roycroft, 1999). In 1994, each state switched to price cap regulation.

I take the date at which Ameritech first introduced a service in any of these states or in the FCC's access tariff to be $t = 0$, and then measure innovation delay s for the other states relative to the first state's innovation date. This may underestimate true innovation delay: the true time 0 must be weakly before the observed first "innovation" under this definition. However, time elapsed before the first tariff cannot serve as a signal because the regulator does not observe it. Thus, my measurement of s corresponds to the useful part (for signaling) of innovation delay, and therefore corresponds to s as used in the model.[10] The data include new services introduced in at least two states. Regulatory delay a for a service is the time from the submission of the first tariff filing to the approval of the last filing.[11]

Summary statistics for regulatory delay are in Table 1. The modal value for innovation delay is zero, and in most states delay of both types is right skewed, which affects the choice of empirical specification described in the next section. Prieger (2007) more extensively documents that innovation delay varied greatly among states and regulatory regimes, which provides good variation in this variable for the estimations here. The difference between Prieger (2007) and the present study is the focus: in the former, innovation delay is the dependent variable and the expected regulatory delay or regulatory uncertainty of the regime are the key regressors; in the present work, service-specific regulatory delay is the dependent variable and innovation delay is the explanatory variable of interest.

[9] The data are from the tariffs and the tariff filing logs of the company and the state commissions.

[10] A caveat is that Ameritech also operated in Michigan. New services were effectively deregulated in Michigan and were not tariffed. It is unclear whether regulators in the other states observed introductions in Michigan.

[11] Some services had multiple tariff filings and withdrawals before approval was granted in a state.

We may observe spurious positive correlation between innovation delay and discretionary regulatory delay if services differ in the complexity of implementation. If so, the firm may delay filing for approval as it works out technical issues, and the regulator may delay approval as it reviews the complex issues raised. Correlation between s and a would be positive, but not for any reason coming from the theoretical model. To control for the complexity of a new service offering, I use as a proxy the rank of the introduction of a service among the various states.[12] The notion of learning by doing motivates the use of the order of introduction to reveal complexity of product implementation. Ameritech gains experience each time it introduces a particular service in another state. Thus, the first introduction may be the most complex. Similarly, regulators in subsequent states can learn from the experience of regulators in previous states, as they examine the issues that were raised and their resolution during previous approval processes. Thus, the complexity of the regulatory approval process should also decrease in subsequent states. Of course, the rank of a state in the order of introduction is not unrelated to innovation delay: longer delay in a state increases the likelihood that introduction is later than in other states. However, in estimations including the rank of introduction as a control, I already control directly for innovation delay. The rank therefore communicates extra information about complexity not captured by innovation delay. The idea: given two distinct services with equal innovation delay, regulatory approval is more complex on average for a novel service than for a service already introduced elsewhere in neighboring jurisdictions.

Finally, note that the condition of concavity implied by Proposition 3 is necessary for the signaling model but not the full-information model (as may be seen from Figure 3). Thus, rejecting concavity of regulatory delay in innovation delay would reject the signaling model but not the full-information model.

5 Results of the Empirical Tests

The goal of the empirical work is to test the predictions from the model. To examine how regulatory delay a responds to innovation delay s, I perform semiparametric Cox estimations for regulatory delay where innovation delay for the service is included as a regressor. The hazard rate of the duration of innovation delay in the Cox model is

$$\lambda(t, x_i) = \exp(x_i'\beta)\, \lambda_0(t), \tag{19}$$

where λ_0 is an arbitrary, unspecified baseline hazard and x_i is a vector of regressors for duration i.[13] Positive coefficients for β increase the hazard and therefore decrease mean duration. By including innovation delay as a regressor, I assume it is exogenous or predetermined with respect to regulatory delay. The assumption of exogeneity may not hold if behind-the-scenes negotiations between the firm and the regulator jointly determine observed regulatory and innovation delay. For example, the firm could agree not to submit a product for approval until the regulator gathers information on the relevant issues, shortening observed regulatory delay. However, such maneuvering would induce negative correlation between regulatory and innovation delay, the opposite of what I find. I return to the issue of endogeneity below.

I first allow s to enter the hazard specification in a flexibly parametric way with cubic p-splines (Eilers and Marx, 1996). Figure 4 shows the partial effect of introduction delay

[12] Tariff filings from Ohio, another Ameritech state, are included in the calculation of the order statistic (and the innovation delay), and so the variable ranges from one to four. Ohio is not included in the main data set because regulatory delay data are unavailable.

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

on estimated mean regulatory delay.[14] The graph omits the top 10% of the outlying introduction delays (20 observations), since the data are sparse there and the standard error band is wide. The figure reveals that increasing introduction delay increases the hazard rate for regulatory delay, at least through 90% of the data range. Furthermore, the relationship between innovation and regulatory delay is concave, as predicted by Proposition 3. The concavity is not an artifact of the specification, because innovation delay enters the hazard specification effectively nonparametrically in the estimation.

To assess statistical significance, I turn now to Cox models in which innovation delay enters the hazard in a simpler fashion. Based on the previous analysis, in the rest of the Cox models introduction delay enters the regression function non-linearly. I include an indicator variable for no innovation delay, because 35% of the observations are zero. For positive innovation delay, I include a two-part spline with the knot at the median innovation delay.[15] Estimation 1 in Table 2 includes innovation delay, along with state-specific indicators for the years of price cap regulation. The latter remove differences in average regulatory delay across states and regimes. The estimation is stratified by state, which allows the baseline hazard to vary in shape and level without restriction across states. Standard errors are clustered to allow for arbitrary correlation among the observations in different states for the same service. The χ^2 statistic for the significance of the regressors is very large, and a test of the Cox model's assumed proportional hazards in (19) does not indicate any misspecification (see table notes for details).

Proposition 2 states that there must be a positive relationship between a and s. The coefficients for the innovation delay spline are both negative in all specifications in Table 2, implying that the hazard rate for regulatory delay decreases (and average delay increases) as innovation delay increases. For estimation 1, the estimates imply that if innovation delay rises from its mean value by one standard deviation, then regulatory delay increases by 7 days. The coefficients are not individually significant, but in the next section, I find better evidence for the hypotheses in question with tests that are more powerful. The signs of the innovation delay coefficients are robust to various changes in the controls used. In estimation 2 in Table 2, the stratification by state is replaced with state-specific political economy variables used in other studies of regulatory change (Donald and Sappington, 1997): the log annual budget of the regulatory authority, an indicator for Republican control of both houses of the state legislature and a Republican governor (*Republican*), and the average value of *Republican* from 1984 up to the previous year (*Republican history*).[16] These variables are the same for all observations within a particular state and year (although they can change for durations that span years), and help control for possible omitted factors due to the political and regulatory climate that affect regulatory delay. These three coefficients are jointly significant, but there is evidence the Cox model is misspecified. In any event, the coefficients for innovation delay change little from estimation 1.

In estimation 3, the first specification is repeated including the observation's rank in the order of introduction of that particular service across the states. As discussed above, the rank proxies the unobserved complexity of the service. One expects that holding the length of innovation delay constant, services introduced in later states are less complex and should be approved quicker. This is indeed what estimation 3 reveals

[14] Figure 4 is based on a penalized Cox regression where indicators for the incentive regulation periods in each state are also included.

[15] If a second knot is added at the third quartile, the slope coefficient for the third piece of the spline does not differ significantly from the second piece.

[16] The other political economy variable used in Donald and Sappington (1997), an indicator for elected public utility commissioners, can not be used here because commissioners are not elected in any state. Other variables I explored included the size of the PUC staff and the political composition of the legislature and governor's office separately; none of these was significant.

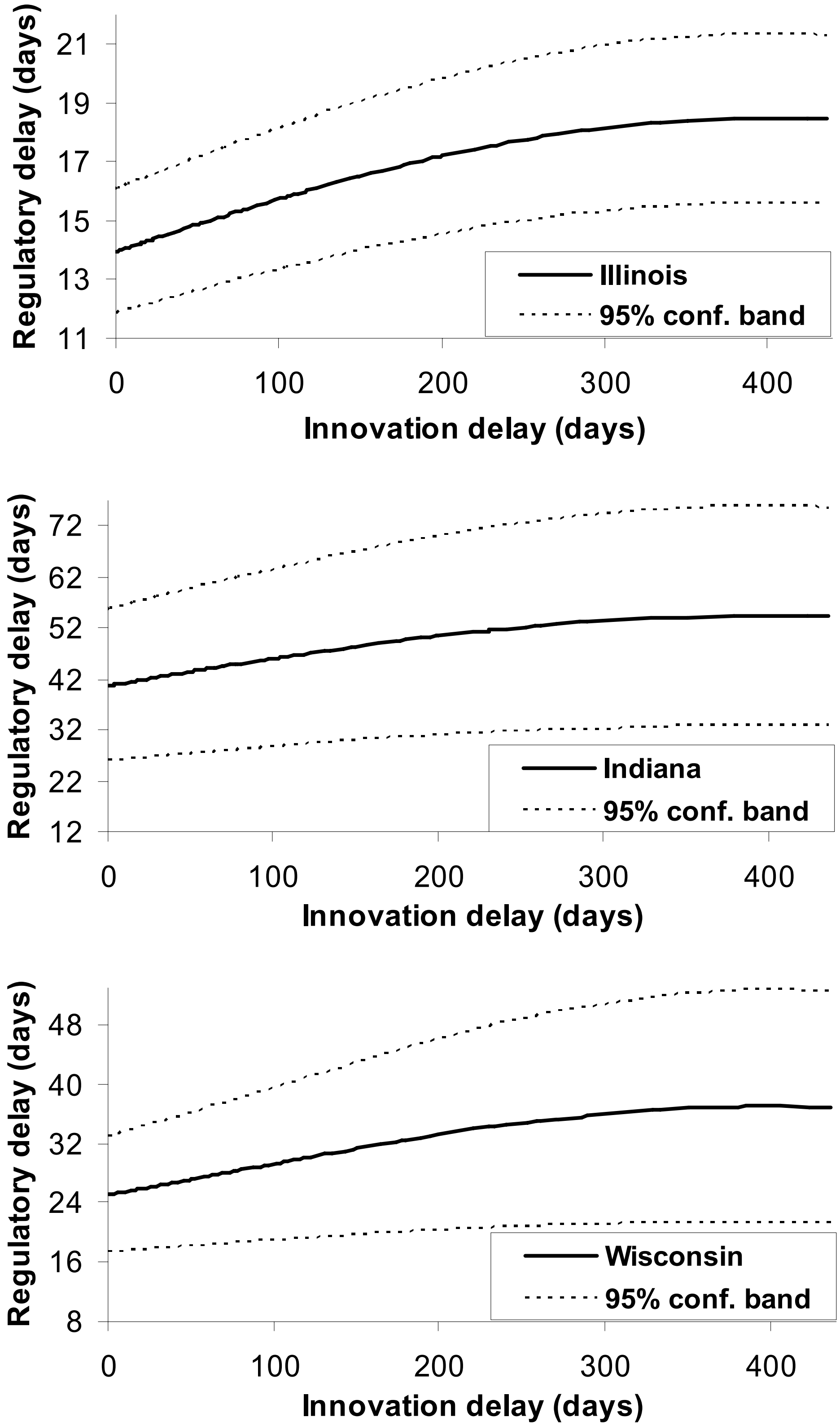

Fig. 4 Effect of innovation delay on regulatory delay (cubic p-splines)

	Cox Proportional Hazards Models for Regulatory Delay					
	Estimation 1		Estimation 2		Estimation 3	
	coef.	*s.e.*	*coef.*	*s.e.*	*coef.*	*s.e.*
Innovation delay = 0	−0.449	0.339	−0.528	0.392	−0.344	0.516
∈ [1,median] (×1000)	−4.510	10.40	−4.160	11.90	−7.560	11.60
> median (×1000)	−0.379*	0.224	−0.380	0.266	−0.486**	0.192
IL:incentive reg	0.212	0.190	0.224	0.243	0.168	0.203
IN:incentive reg	1.669***	0.301	1.958***	0.357	1.420***	0.326
WI:incentive reg	1.855***	0.348	1.484***	0.195	1.566***	0.396
Order: second					0.149	0.380
Order: third					0.450	0.388
Order: fourth					1.047***	0.382
PUC budget			1.574***	0.583		
Republican			−0.090	0.161		
Republican history			0.662	0.649		
Stratification	by state		none		by state	
N	267		267		246	
χ^2 statistic (d.o.f.)	55.7 (6)	$p = 0.00$	81.9 (9)	$p = 0.00$	120.6 (9)	$p = 0.00$
$T(G)$ statistic (d.o.f.)	4.85 (6)	$p = 0.56$	17.60 (9)	$p = 0.04$	9.3 (9)	$p = 0.41$
Log likelihood	−914.6		−1194.2		−821.1	

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

Table notes: The model incorporates time-varying covariates. Larger positive coefficients imply shorter delays. *PUC budget* is the log budget of the state public utility commission. *Republican* is an indicator for a Republican governor and majority in both houses of the state legislature. *Republican history* is the average value of *Republican* from 1984 to the previous year of the observation. 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 indicates the model is misspecified (test 4 of Grambsch and Therneau (1994)). Std. errors are robust to clustering on services.

Table 2 Semiparametric estimation results for regulatory delay a

in Table 2: the hazard rate increases monotonically with the introduction order of the service, implying that the average regulatory delay time decreases as the rank increases. More importantly, for the purposes of testing the theoretical model, the coefficients on the innovation delay variables remain negative. Taken together, the evidence from all estimations suggests that greater innovation delay is positively associated with greater discretionary regulatory delay, in accordance with the prediction of the model. If order is added to estimation 2, similar results obtain.

The relationship between regulatory and innovation delay is concave (i.e., a piecewise linear approximation of concavity) in all estimations. Figure 5 plots expected regulatory delay as a function of innovation delay, using the coefficients estimated in models 1–3.[17] Thus, the data are in accord with the prediction from Proposition 3. Concavity is a necessary implication of the signaling model, but nonconcavity is not a necessary implication of the full-information model in general (although it is for the example in Section 3), so that finding concavity suggests but does not prove that the firm is signaling.

I return now to the potential endogeneity of innovation delay in these estimations. Although innovation delay fully precedes regulatory delay by definition, there may be

[17] Mean regulatory delay is calculated as the average across the sample of the predicted mean durations. Predicted durations are computed from the estimated survival curves using actual covariates and the counterfactual innovation delay shown on the x-axis in the figure.

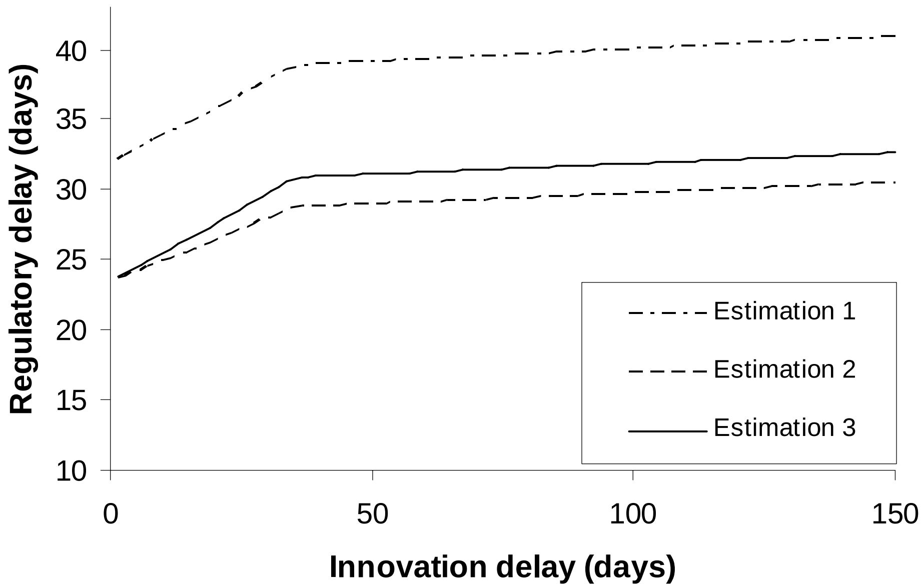

Fig. 5 Effect of innovation delay on regulatory delay (linear splines)

unobserved factors that contribute jointly to the determination of regulatory and innovation delay. If so, the relationship between regulatory and innovation delay is not causal. For example, in the full-information version of the model above, positive correlation is induced between regulatory and innovation delay only through the unobserved (to the econometrician) cost of delay. Endogeneity of covariates in a Cox model raises some subtle statistical issues and there is no analog to instrumental variables (IV),[18] so I investigate the potential endogeneity of regulatory delay with linear OLS and IV models. I instrument for innovation delay with an indicator variable that is one if delay is calculated from the federal access tariff (FAT) and zero if delay is calculated from the first state tariff. Prieger (2007) shows that FAT is highly significant with a large effect in estimations with innovation delay as the dependent variable, so the instrument is relevant. The exclusion restriction for IV requires that FAT be uncorrelated with the unobserved determinant of regulatory delay. FAT indicates that the service appears in the federal tariff before in a state tariff, and there is no reason to expect FAT to be correlated with regulatory delay after controlling for innovation delay.

The exclusion restriction is not formally testable here, because the IV model is just identified with the single instrument. However, if FAT is added to the estimations in Table 2, it is never significant, even at the 20% level. Further informal evidence for the exclusion restriction comes from assessing how balanced the other observed covariates from the regulatory delay model are with respect to FAT. Altonji, Elder and Taber (2000) explore using the relationship between an instrument and the observables as a guide to how much correlation there may be with the unobservable error in the outcome equation. In the ideal case where the instrument is randomly assigned, the instrument would show no correlation with either the other covariates or the error term. Two-sample tests performed independently for the covariates *order*, *PUC budget*, *Republican*, and *Republican history*, where the samples are split according to the value of FAT for the observation,

[18] If introduction delay is endogenous, the Cox estimations are still consistent for β in the hazard rate in (19). However, the hazard rate would no longer be interpretable as the probability of regulatory approval conditional on the covariates (Lancaster, 1990, sec. 2.3 and 9.2.).

generally failed to show any significant differences in means.[19] I also examined the distribution of the martingale residuals from Estimation 3, separated by value of *FAT*. Testing reveals no significant difference between the distributions of the residuals for the two values of *FAT*—further corroboration that *FAT* is not correlated with unobserved determinants of regulatory delay.[20] *FAT* should also be correlated highly enough with innovation delay to avoid the problems for inference that weak instruments can cause. The partial F statistic for *FAT* in the first stage regression of innovation delay on all the exogenous variables in the model is 16.2 (p-value$= 0.0001$), which is not in the region that the literature on weak instruments suggests may lead to non-negligible finite sample bias and artificially small standard errors in the second stage estimation (Staiger and Stock, 1997).

The results from the OLS and IV estimations of regulatory delay on innovation delay are in Table 3. Given that only one instrument is available, innovation delay enters the mean linearly, without the spline used in Estimations 1–3. Furthermore, since there is no additional instrument for the indicator for zero innovation delay, those observations are dropped. In the OLS estimation, the coefficients have signs in accord with Estimation 3, the analogous specifications from the Cox models.[21] Innovation delay is associated with longer regulatory delay, albeit only at the 10% significance level (probably due to the smaller sample size and linear constraint on how delay enters the mean). In the IV estimation, the coefficient for innovation delay switches sign but is insignificant. The purpose of the IV estimation is to test for the endogeneity of innovation delay. The bottom of Table 3 shows test statistics and p-values for the Hausman tests. The tests do not come close to rejecting the null hypothesis that innovation delay is exogenous in the OLS estimation, whether performed on the delay coefficient only or all coefficients. The conclusion, then, is that there is no evidence for the endogeneity of innovation delay in the regulation delay estimations.

To conclude the test of the predictions of the theoretical model, I summarize the evidence presented with a suite of hypothesis tests. Table 4 reports results from estimations 1–3, where the tests are for $da/ds \leq 0$, which would violate Proposition 2. The test is implemented for individual coefficients with the null hypothesis that the coefficients for the spline in innovation delay are non-negative. Thus, under the null, the hazard rate for endogenous regulation delay stays the same or increases (or, equivalently, regulation delay does not increase) as innovation delay increases. Although all the spline coefficients are negative, the hypothesis is not rejected for β_1, the spline coefficient for innovation delay less than the median. However, it is rejected for β_2, the coefficient for longer delay. Thus, the best evidence that innovation and regulation delay move together is found when innovation delay is longer than usual. Innovation delays are much more spread out in the upper quartiles (e.g., the median is 35 days but the third quartile is 145 days); perhaps unusually long delays signal more effectively because they catch the eye of the regulator more. I also report joint tests at the bottom of the table. The hypothesis that the innovation delay coefficients are zero is rejected at the 10% level in four out of six cases (three out of six cases at the 5% level) in joint tests.

[19] The one exception, curiously, is *Republican*, for which the test returns a p-value of 0.043. The tests are performed via a Welch two-sample t-test without assuming the variances of the two sample are equal. The test for *Republican* is not significant at the 5% level if equal variances are assumed.

[20] The residuals were tested with a two-sample difference-in-means t-test and a two-sample Kolmogorov-Smirnov test for equality of the distributions; neither test rejected the null that the residuals have the same distribution for the two values of *FAT*.

[21] The Cox models are for the hazard rate and the OLS model is for mean duration, so a negative coefficient in Table 2 accords with a positive coefficient in Table 3.

	Linear Models for Regulatory Delay			
	OLS		IV	
	coef.	*s.e.*	*coef.*	*s.e.*
Constant	44.24***	11.09	49.01***	12.33
Introduction delay (×1000)	11.88*	6.691	−11.21	22.42
IL:reg change	−9.155	10.88	−12.06	11.61
IN:reg change	−86.20***	13.33	−87.22***	13.87
WI:reg change	−34.56***	9.576	−33.30***	10.01
Indiana	61.75***	15.15	63.72***	15.84
Wisconsin	11.30	11.73	11.25	12.18
Order: second	−10.77	9.649	−10.99	10.02
Order: third	−18.64*	10.25	−17.73*	10.68
Order: fourth	−26.20**	13.33	−18.98	15.36
N	162		162	
R^2	0.325		0.272	
F statistic	8.140	$p = 0.00$	7.250	$p = 0.00$
(d.o.f)	$(9, 152)$		$(9, 152)$	
Hausman test statistic 1 (d.o.f.)			1.16 (9)	$p = 0.99$
Hausman test statistic 2 (d.o.f.)			1.16 (1)	$p = 0.28$

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

Table notes: sample is restricted to observations with nonzero introduction delay. The IV estimation uses Federal Access Tariff as an instrument for introduction delay. *Hausman test statistic 1* tests all coefficients except the constant. *Hausman test statistic 2* tests only the coefficient for introduction delay. *D.o.f.* is the degrees of freedom of the statistics.

Table 3 OLS and IV estimates for regulatory delay a

6 Alternative Explanations

The empirical work shows that discretionary regulatory delay is positively correlated with introduction delay at the service-specific level, in accord with the theoretical model presented here. In this section, I discuss a few competing explanations.

Let us first dispense with some conceptions of the interaction between the regulator and the firm that would lead to the opposite finding. One idea is that before the firm officially submits a service to the regulator, there is unofficial communication and bargaining between the parties, so that the firm submits only "pre-approved" services to the regulator. If so, then the division of total delay time $s + a$ into innovation and regulatory delay is arbitrary. In this case, however, there would either be no correlation

	Cox Models for Regulatory Delay		
	Estimation 1	Estimation 2	Estimation 3
Hypothesis	*p-value*	*p-value*	*p-value*
H_0: $\beta_j \geq 0$ vs. H_A: $\beta_j < 0$			
β_1 for innovation delay $\in$ [1,median]	0.333	0.364	0.258
β_2 for innovation delay $>$ median	0.046	0.077	0.006
Joint test: H_0: $\boldsymbol{\beta} = 0$ vs. H_A: $\boldsymbol{\beta} \neq 0$			
β_1 and β_2	0.169	0.278	0.014
β_1, β_2, and β_0 for innovation delay $= 0$	0.055	0.049	0.014

Table 4 Results of hypothesis tests

between innovation and regulatory delay (if regulatory delay were pro-forma and varied only due to the vagaries of the regulatory calendar) or negative correlation (if longer s assured shorter a).

Another possibility is that the firm delays submitting a service to the regulator while it performs "due diligence" tasks regarding cost-based pricing or other matters. If the regulator does not know the propensity of the firm to introduce over-priced services, but believes that firms that are more honest find it less costly to perform due diligence, then a Spence-type signaling model would lead to honest firms taking more time to introduce products than dishonest firms. The regulator would respond by taking more time to examine services with little innovation delay, and the result would be negative correlation between innovation and regulatory delay, the opposite of what the data show.

Other explanations could rationalize the empirical finding of positive correlation. First, there may be unobserved service-specific factors that are positively correlated with both types of delay. A candidate for this explanation is a service such as caller ID. When phone companies first introduced caller ID, consumer privacy advocates objected, which led to extended discussions about how the service would be offered. The privacy concern, an unobserved factor to the econometrician, may have delayed introduction of the service as the phone companies reprogrammed the service to meet the objections of the advocates, and also may have delayed regulatory approval due to lengthier regulatory hearings. Caller ID, however, is not a typical service in this respect. Few telecommunications services garner any attention from the public before they are introduced. Furthermore, if such unobserved factors were an important part of the regulatory approval process, then innovation delay would be endogenous in the estimations for regulatory delay. The statistical tests gave no evidence of such endogeneity. However, given that the exact power of the tests cannot be known, unobservables remain a potential explanation.

There may be many other explanations for the correlation between innovation and regulatory delay seen in the data. However, external evidence suggests that regulators are becoming increasingly attuned to the costs of regulatory delay, so it is realistic to assume that a firm would wish to communicate such costs to the regulator. Over the last few decades, regulatory commissions (in some cases prodded by state legislatures) have placed more emphasis on the benefits from new products. The older breed of regulatory official, accustomed to tight regulatory control and a stable industry, viewed new products with suspicion. As one regulator put it, "...regulation of telecommunications remain essential to protect the public from deleterious consequences of innovation..." (Oppenheim, 1991, p.310). Contrast this view with the more recent goals adopted by regulators in the Ameritech region to "...facilitate the introduction of innovative new services in this competitive marketplace." (PSC of Wisconsin, 1998, p.47) This change of attitude about the importance of new products to consumers and firms may make regulators willing recipients of signals about the cost of delay.

7 Conclusions

This paper shows there is a robust positive relationship between the time until submission of a new product to the regulator and subsequent regulatory delay. The model developed here suggests that the correlation may stem from the firm's and the regulator's optimal choice of timing in response to the cost of regulatory delay. Whether under complete or asymmetric information, with signaling taking place in the latter case, the model predicts that shorter innovation delay will be rewarded with shorter regulatory delay.

How much does the signaling matter? Consider the prediction from estimation 1 for the effect of changing innovation delay from one day to 2,441 days, the maximum in the data: regulatory delay increases 215% (by 69 days), about one standard deviation (s.d. = 71.1). The predicted percentage increases in regulatory delay from estimations 2 and 3 are even greater (293% and 237%, respectively). Thus, innovation delay has the potential to shift regulatory delay a significant amount. While the welfare cost of the delay cannot be estimated without price and quantity data, demand for voice mail may provide a comparison. Hausman (1997) estimates that the compensating variation from voice mail service averaged about $5 per subscriber in 1994, which was 8.2% of the average telephone expenditure (FCC, 1996). Voice mail undoubtedly has more subscribers than most of the services in the present data. However, if the surplus per subscriber for these services were similar, then while the absolute welfare cost of the additional regulatory delay would be small, it would not be insignificant for affected subscribers relative to their telecommunications budget.

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. In this setting consumer surplus would not enter management's objective function. In each of these settings, there is asymmetric information and the possibility of signaling and learning over time.

Some interesting extensions to the model deserve future attention. Taking the single-firm, single-regulator theoretical model to the data required the assumption that there are no strategic interactions among jurisdictions. However, many firms operate in multiple states. In the current formulation, actions undertaken in one jurisdiction have no signaling value to regulators in the other jurisdictions. A logical next step for the model is to expand the signaling game to include multiple receivers of the firm's multiple signals.

Another extension would be to incorporate unregulated rivals into the model explicitly. The only impact of competition in the current model is indirect: it may affect the marginal cost of delay to the firm (θ) or the regulator's benefits of delay (V). 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, exploring the political economy of regulatory delay in the telecommunications industry would be an interesting complement to the present work, in which the regulator's taste for delay is assumed but not derived.

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8 Appendix

Proof of Proposition 3 The limit of da/ds as $\theta \downarrow \theta^-$ is given by $\lim_{\theta\downarrow\theta^-} (d\alpha/d\theta)\,/\,(d\sigma/d\theta)$. Since $\lim_{\theta\downarrow\theta^-} d\sigma/d\theta = -\infty$ by the discussion after equation (11), it follows that da/ds, which is positive by Proposition 2 tends to zero as θ falls to θ^-. An increasing function with a vanishing derivative at θ^- must be concave at least in a neighborhood around θ^-.

Statement and proof of Proposition 4

Proposition 4 *$\sigma(\theta) < \tau(\theta)$ for $\theta \in (\theta^-, \theta^+)$. Innovation delay is smaller under signaling than the full-information case.*

From (11) we know the proposition holds in a neighborhood to the right θ^-, so if it does not hold for all types σ crosses τ from below at a type $\bar{\theta}$. But then pick a type $\theta' > \bar{\theta}$ such that $\tau(\theta') \in \sigma([\theta^-, \theta^+])$, and consider a deviation by a firm of type θ' to $\tau(\theta')$. Then the regulator infers (incorrectly) that the firm is of type θ'', where $\theta'' = \sigma^{-1}(\tau(\theta'))$. By definition of τ it must be that $\tilde{\Pi}(\theta', \theta', \sigma(\theta')) < \tilde{\Pi}(\theta', \theta', \tau(\theta'))$. Also, because $\partial\tilde{\Pi}/\partial\hat{\theta} > 0$, we have $\tilde{\Pi}(\theta', \theta', \tau(\theta')) < \tilde{\Pi}(\theta', \theta'', \tau(\theta'))$. Thus, the firm of type θ' does better to play $\tau(\theta')$, and deviation to $\tau(\theta^-)$ cannot be credibly punished. Thus, σ cannot cross τ.

The Mailath conditions In addition to the assumption that $\tilde{\Pi}$ is C^2, the following conditions are required to make use of Mailath's (1987) results:

1. Belief monotonicity: $\partial\tilde{\Pi}(\theta, \hat{\theta}, s)/\partial\hat{\theta} \neq 0$. Here, $\partial\tilde{\Pi}/\partial\hat{\theta} = \frac{\partial \Pi}{\partial a}\frac{da}{d\alpha}\frac{d\alpha}{d\hat{\theta}}$. Prop. 1 implies that $\partial\Pi/\partial a \neq 0$, $da/da = 1$, and equation (5) implies that $d\alpha/d\hat{\theta} \neq 0$, so $\partial\tilde{\Pi}/\partial\hat{\theta} \neq 0$.
2. Type monotonicity: $\frac{\partial^2\tilde{\Pi}}{\partial s\partial\theta} \neq 0$. Here, $\frac{\partial^2\tilde{\Pi}}{\partial s\partial\theta} = -e^{-r(s+a)}\pi'(\theta) < 0$.
3. Requirements of the full-information strategy:
   (a) Existence and uniqueness: $\partial\tilde{\Pi}(\theta, \theta, s)/\partial s = 0$ has unique solution in s, which maximizes $\tilde{\Pi}(\theta, \theta, s)$. Here, $\frac{\partial\tilde{\Pi}}{\partial s} = 0$ implies
   $$rF(s) - F'(s) = e^{-ra}\pi(\theta) \tag{20}$$
   Under the assumptions on F in footnote (5), a unique solution exists.
   (b) "Strict" quasiconcavity. $\frac{\partial^2\tilde{\Pi}(\theta,\theta,\tau(\theta))}{\partial s^2} < 0$. Evaluated at $(\theta, \theta, \tau(\theta))$, and using equation (20), we have $\frac{\partial^2\tilde{\Pi}}{\partial s^2} = e^{-rs}\left(rF'(s) - F''(s)\right) < 0$.
4. Boundedness. For all $(\theta, s) \in [\theta^-, \theta^+] \times \mathbb{R}$ there exists a $k > 0$ $\forall$ such that $\frac{\partial^2\tilde{\Pi}(\theta,\theta,s)}{\partial s^2} \geq 0 \Rightarrow \left|\frac{\partial\tilde{\Pi}(\theta,\theta,s)}{\partial s}\right| > k$. In this application this condition does not hold, because (due to the exponential terms in s) as $s \to \infty$, $\partial\tilde{\Pi}(\theta, \theta, s)/\partial s \to 0$. However, this condition is sufficient but not necessary, and is stronger than needed. The condition is used in Mailath (1987) to bound the set $\mathbb{S} = \{s \in \mathbb{R} | \exists\theta$ such that $\tilde{\Pi}(\theta, \theta, s) \geq \tilde{\Pi}(\theta, \theta^-, \tau(\theta^-)\}$ for arbitrary τ. The inequality condition may be written as
   $$e^{-r\left(s-\tau(\theta^-)\right)} \geq \frac{\left(-rF(\tau(\theta^-)) + e^{-r\alpha(\theta^-)}\pi(\theta)\right)}{\left(-rF(s) + e^{-r\alpha(\theta)}\pi(\theta)\right)} \tag{21}$$
   As $s \to \infty$, the left side of inequality (21) goes to zero for arbitrary τ. As $s \to \infty$, the numerator on the right side of (21) is unaffected, the denominator has $F(s) \to 0$ and $e^{-r\alpha(\theta)}\pi(\theta) > 0$, and so the right side is a positive number bounded away from zero. Thus, $\mathbb{S}$ is bounded above as required.
5. Initial condition: equation (11). Assume not: suppose σ is one-to-one and incentive compatible but that $\sigma(\theta^-) \neq \tau(\theta^-)$. Consider a deviation by the firm of type θ^- to $\tau(\theta^-)$. If $\tau(\theta^-) = \sigma(\theta')$ for some $\theta' \in [\theta^-, \theta^+]$, then the regulator will infer (incorrectly) that the firm is of type θ'. By definition of τ it must be that $\tilde{\Pi}(\theta^-, \theta^-, \sigma(\theta^-))$ $< \tilde{\Pi}(\theta^-, \theta^-, \tau(\theta^-))$. Also, because $\partial\tilde{\Pi}/\partial\hat{\theta} > 0$, we have $\tilde{\Pi}(\theta^-, \theta^-, \tau(\theta^-)) < \tilde{\Pi}(\theta^-, \theta', \tau(\theta^-))$. Thus, the firm does better to play $\tau(\theta^-)$. On the other hand, if $\tau(\theta^-) \notin \sigma([\theta^-, \theta^+])$, then a sensible refinement such as the intuitive criterion can ensure that the regulator holds the pessimistic belief that the firm's type is θ^-. If so, then (by definition

of τ) the firm does no worse playing $\tau(\theta^-)$ than by playing $\sigma(\theta^-)$. In either case, then, deviation to $\tau(\theta^-)$ cannot be credibly punished. Thus, it must be that $\sigma(\theta^-) = \tau(\theta^-)$.

6. Single crossing condition: $\frac{\partial \tilde{\Pi}(\theta,\hat{\theta},s)/\partial s}{\partial \tilde{\Pi}(\theta,\hat{\theta},s)/\partial \hat{\theta}}$ is monotonic in θ. We have

$$\frac{\partial \tilde{\Pi}(\theta,\hat{\theta},s)/\partial s}{\partial \tilde{\Pi}(\theta,\hat{\theta},s)/\partial \hat{\theta}} = \frac{rF(s) - F'(s) - e^{-ra}\pi(\theta)}{e^{-ra}\pi(\theta)\ \frac{\partial^2 W}{\partial \hat{\theta}\partial a} \Big/ \frac{\partial^2 U}{\partial a^2}} \tag{22}$$

Since neither W nor U depend on θ, the relevant terms are $\left[rF(s) - F'(s) - e^{-ra}\pi(\theta)\right]/\pi(\theta)$, which has derivative in θ of $-\left(rF(s) - F'(s)\right)\pi'(\theta)/\pi(\theta)^2$. This derivative has the sign of $-rF(s) + F'(s) < 0$. So the single crossing condition is satisfied, as asserted in (7).