Telecommunications Regulation and New Services: a Case Study at the State Level*

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Abstract

The effects that regulation has on the innovation and the introduction of new telecommunications services have not been previously quantified in the literature. This study compares state-regulated services in Indiana under rate of return regulation (RoRR) and under alternative regulation. The econometric model comprises an count process (for innovation) followed by a duration process with selection (for regulatory delay). Moving away from RoRR increased the rate of service creation to three times the old rate. Expected approval delays nearly disappear. A prediction exercise indicates that the firm would have introduced 12 times as many services to consumers if the alternative regulation had been in place the entire time.

Keywords: regulation, product innovation, telecommunications, count data, duration data, tobit model

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1 Introduction

Telecommunications regulation in the United States has changed rapidly in the past 15 years. Regulatory authorities in many states have moved away from rate of return regulation (RoRR) toward incentive regulation such as price caps. Most often, regulatory reform has focused on improving consumers' access to existing services: lower prices, availability to more households, higher service quality, improved efficiency in production, and so on. The economic literature also mirrors this emphasis. Most regulation studies take the existing product mix as given and look at "getting the prices right" or the incentive to reduce costs. Often overlooked—in both the regulatory arena and the economic literature—are the impacts on incentives to create and introduce new products. This study quantifies the impact of a regulatory reform in Indiana on the creation and introduction of new services.

The regulatory regime change of interest is a form of alternative regulation known as Opportunity Indiana, described more fully in the next section. In 1994, the Indiana regulatory commission switched from traditional RoRR of Ameritech Indiana[1] to a combination of price caps and market (i.e., unregulated) pricing.[2] Since RoRR limits the ability of the firm to retain as profit the economic benefit created by the service,[3] market pricing should increase the incentive to introduce a new product. Similarly, by controlling prices rather than profit levels, it is possible that a price caps regime would also see more new products than a rate of return regime.[4] This study finds that the incentives did increase under Opportunity Indiana: the creation of new services went up by 97% to 359%, depending on the type of service.

The literature on price caps and innovation focuses almost exclusively on cost reduction, known

[1] Ameritech is the Bell Operating Company in Indiana, with 62% of the local phone lines. Ameritech's operating company is named Indiana Bell, which does business as Ameritech Indiana.

[2] Rate of return regulation limits profits by setting prices to achieve a desired return on capital. Price cap regulation limits prices directly, without explicit reference to profits or costs. See Averch and Johnson (1962) and Acton and Vogelsang (1989) for characterizations of rate of return regulation and price caps, respectively.

[3] As long as the targeted rate is binding, that is.

[4] Although the comparison could go the other way, if a new product allows a firm to increase its capital base under rate of return regulation, which would increase its allowed profit. The literature does not contain a clear answer on this point.

as *process innovation*. Littlechild (1983) claimed that price caps increase the incentive to reduce costs. Theoretical comparisons of rate of return and price cap regulation mainly confirm Littlechild's claims (e.g., Cabral and Riordan, 1989). *Product innovation*—bringing new goods to market—may be put into the theoretical framework of process innovation: product innovation occurs when a product's cost is reduced from infinity. However, the process innovation literature typically focuses on small cost changes, and not all claims for process innovation necessarily apply to product innovation.[5] At least as important for product innovation as the distinction between price caps and RoRR is the treatment of new services. Are new services immediately placed in the rate base or the price cap basket, or is there a "grace period" to encourage innovation? Does regulatory approval delay differ markedly between traditional and alternative regulation? While these characteristics could be built into a theoretical model, ultimately whether price caps increase innovation in any particular setting is an empirical issue. In the present study, most of the new services were under market pricing rather than price caps. The comparison here, therefore, is not a clear-cut comparison of price caps and RoRR but of RoRR and a mixture of incentive regulation and deregulation.

Opportunity Indiana effectively increased the expected profit to the firm of introducing a new service for at least two reasons. First, as mentioned above most new services were allowed to be freely priced. RoRR, on the other hand, places constraints on prices. Second, there are long tariff approval delays under RoRR. associated with plan approval—over four months on average, a long time in the competitive and dynamic telecommunications industry. The delay reduces the present value of a proposed new service and (more importantly) allows competitors to beat the regulated firm to market. By contrast, there are few approval delays under Opportunity Indiana.

The effect of the alternative regulation on the number of services introduced can be formalized with a simple argument. Let π be the random variable representing the net present value of profit

[5] Bonanno and Haworth (1998) investigate the strategic differences between product and process innovation (although their concern is not with regulation).

from a service, distributed with density F_1 under RoRR and F_2 during Opportunity Indiana. In keeping with the discussion above assume that F_2 stochastically dominates[6] F_1, implying that expected profit is higher under Opportunity Indiana. Assuming the firm introduces every service in the population of potential services that has positive expected profit, then $1-F_i(0)$ is the fraction of the total population introduced under regime i. It follows directly from stochastic dominance that if the size of the population of potential services is the same under both regimes, the firm would introduce more new services under alternative regulation. If the population of potential services increases under lighter regulation, due to increased basic innovation by outside firms because of the more favorable prospects, then even more new services would be introduced under alternative regulation.

The only empirical study of which I am aware looking at product innovation and telecommunications regulation is Mueller (1993), who examines new service offerings by US West in Nebraska. The (non-econometric) evidence suggests that deregulation was successful at speeding new service introductions in Nebraska, compared with other states.[7] There are a few studies quantifying the impact of incentive regulation on process innovation. Taylor, Zarkadas and Zona (1992), Greenstein, McMaster and Spiller (1995), and Ai and Sappington (1998) all find that incentive regulation hastens the diffusion of digital infrastructure compared with other regulatory methods. Unlike these studies, which look at the means of providing a service, I look directly at the new services that consumers purchase. Unlike Hausman (1997), who estimates the impact of regulatory delay on cellular telephony and voice mail consumers, I am interested in the bigger picture of how regulation affects innovation in the entire class of telecommunications services and whether potential services are introduced.

The rest of this introduction previews the results. Ameritech greatly increased its rate of service introduction under Opportunity Indiana, responding to the new freedom to inaugurate and price

[6] I.e., $F_2(\pi) \leq F_1(\pi)$ for all π and the inequality is strict for some π.

[7] Mueller (1993) notes that US West may have merely shifted the order of introduction between states to provide a "success story" for deregulation.

services. The model estimates that moving away from RoRR increased the rate of service creation to 2 to 4.6 times the old rate, depending on the type of service. Overall, the estimated rate of service creation under Opportunity Indiana is 3.0 times the rate under the rate of return regime. Expected approval delay times were reduced to almost nothing (from 130 days before Opportunity Indiana to three days during). During Opportunity Indiana, proposed services had a higher probability of being approved as quickly as the law allowed, and the law allowed quicker introductions. Using the estimated model to project the innovation and introduction process under RoRR and Opportunity Indiana, I find that Ameritech would have introduced twelve times as many services to consumers during the study period if Opportunity Indiana had been in place the entire time.

Opportunity Indiana is described in the next section and the econometric model is introduced in section 3. The data, estimations, and specification tests are in section 4.1.1 for the innovation submodel, and in section 4.2 for the regulatory delay submodel. The benefit of Opportunity Indiana is measured in section 5. Section 6 contains discussion of the validity of the model for policy analysis and concluding remarks.

2 Opportunity Indiana

In Indiana, Ameritech offers basic local service, intrastate advanced services, and intrastate inter-LATA access services, which are regulated by the Indiana Utility Regulatory Commission (IURC). Mirroring regulation by the FCC, until the 1990s most state commissions placed the intrastate activities of incumbent local exchange companies under RoRR. Thereafter several states, Indiana included, began to experiment with alternative regulation.

This study examines new intrastate services introduced by Ameritech Indiana in its operating territory. New services may allow higher transmission capability (e.g., ISDN[8]), capability to trans-

[8] Integrated Services Digital Networks (although suspicious regulators sometimes refer to ISDN as "Improvements Subscribers Don't Need"). ISDN is composed of digital lines capable of circuit and packet transmission for voice and data communications at data rates up to 1.544 Mbps.

mit new kinds of data (e.g., video), or central office features such as caller ID or restricted access to 900 numbers. To offer these services, Ameritech typically makes use of technology developed by upstream firms and adapts it to work on their network and to offer the desired functionality.[9] For many of these services, Ameritech is the first to offer them (particularly to residential subscribers). For other services, substitutes may be available from other firms (particularly for business customers).[10]

Until 1994, Ameritech Indiana operated under traditional RoRR. On June 30, 1994, the IURC approved a stipulated agreement among various parties in response to an incentive regulation plan proposed by Ameritech.[11] The plan, known as *Opportunity Indiana*, gave Ameritech increased flexibility through 1997 to introduce and price services. Responding to Ameritech's concern about regulatory delay, the IURC streamlined the tariff approval process to allow quicker introduction of new services. Before Opportunity Indiana, approval of a new service required docketing and commission action. Under the new plan most services were approved automatically upon a day's notice. Also under the new plan, certain services related to basic local service were placed under price caps. Other services deemed to be competitive were moved to a separate tariff, the *Service Catalog*; pricing of Catalog services is essentially unregulated, under the presumption that actual or potential competition limits the prices Ameritech can charge. Opportunity Indiana also contains provisions for local service rate reductions, infrastructure and education investments, and a free subscription program in under-served areas (Indiana Utility Regulatory Commission, 1997*b*). After review of the first term, a modified version of Opportunity Indiana was approved in December 1997 for a second term.[12]

[9] There are elements of both innovation and diffusion to the service introduction process. For simplicity, I will use the term innovation.

[10] The wealth of technological options available in telecommunications means that there is usually something unique about any firm's service, however.

[11] *Order* in Cause No. 39705 (Indiana Utility Regulatory Commission, 1990).

[12] See Indiana Utility Regulatory Commission (1998).

3 A Model for Service Innovation and Introduction

In the course of introducing regulated telecommunications services to subscribers, a firm goes through two steps. The firm first creates a new end-user service, using technology that may be developed by other firms; I refer to this as *innovation*. After innovation, the services are not *introduced* to subscribers until they are approved by the regulator. I model innovation as a count process and the regulatory approval process as a duration process with selection. In the selection model, the regulator first selects whether a service is to be delayed, and if so, the length of the delay. The conditional means of the distributions are modeled as functions of economic, demographic, and regulatory covariates. The model is formally equivalent to an infinite-server queuing model.[13] This equivalence allows me to apply well-known queuing results, which will be useful for the projection exercise carried out in Section 5.

To construct the likelihood of the data for the arrival and regulatory approval processes, the marginal distributions and the nature of the correlation between the processes must be specified. Let the number of arrivals in period t be n_t, a realization of a count (non-negative integer) valued random variable, where $t = 1, \ldots, T$. Denote the probability density function (pdf) of n_t, conditional on covariates $\mathbf{w}_t$, parameter vector $\boldsymbol{\alpha}$, and a random effect u_{1t}, by $f(n_t|\mathbf{w}_t'\boldsymbol{\alpha}, u_{1t})$. The simplest such count process is the Poisson model with mean

$$\lambda_t = \exp\left(\mathbf{w}_t'\boldsymbol{\alpha} + u_{1t}\right) = e^{u_{1t}} \exp\left(\mathbf{w}_t'\boldsymbol{\alpha}\right) \tag{1}$$

and pdf

$$f(n_t|\mathbf{w}_t'\boldsymbol{\alpha}, u_{1t}) = \exp(-\lambda_t)\lambda_t^{n_t}/n_t! \tag{2}$$

The random effect u_{1t} is an unobserved heterogeneity term with variance τ_1^2; assume that $E(e^{u_{1t}}) = 1$ so that $E(n_t) = \exp\left(\mathbf{w}_t'\boldsymbol{\alpha}\right)$ and that the intercept coefficient in $\boldsymbol{\alpha}$ is identified. The inclusion of u_{1t} results in a generalized Poisson model that relaxes the equality of the mean and the variance

[13] See Prieger (1999) for details.

implied by the simple Poisson model and allows overdispersion (Cameron and Trivedi, 1998).[14]

For the regulatory delay part of the model, a standard duration model is not appropriate, for the following reasons. Both before and during Opportunity Indiana, there is a mandated minimum regulatory delay, say s. One can break a complete duration into the deterministic part s and a stochastic part t. In the data, s is 30 days under RoRR and one day under Opportunity Indiana. Many observations are not delayed beyond the minimum (i.e., $t = 0$). Such events have zero probability in any simple duration model. To accommodate these observations I adopt a selection model, in which the regulator selects certain services to be delayed beyond the minimum and then determines the length of the delay. Index the durations by i; each arrival has an associated stochastic duration (possibly of length zero), so that $i = 1, \ldots, N$, where $N \equiv \sum_{t=1}^{T} n_t$. Split i into index sets I_t so that $\{i|i \in I_t\}$ are the indices pertaining to period t. The econometric model I use is the Type 2 Tobit bivariate selection model of Amemiya (1985), with an added random effect u_{2t}. Let y_{1i} be the log duration of the stochastic part t_i of the ith regulatory delay, and let y_{2i} be an indicator for non-zero t_i. Assume that $\{y_{1i}, y_{2i}|i \in I_t\}$ are defined by

$$y_{1i} \equiv \log(t_i) = \mathbf{x}_i'\boldsymbol{\beta} + \sigma\varepsilon_{1i} + u_{2t}, \text{ observed if } y_{2i} = 1 \tag{3}$$

$$y_{2i}^* = \mathbf{z}_i'\boldsymbol{\gamma} + \rho_2\varepsilon_{1i} + \sqrt{1-\rho_2^2}\,\varepsilon_{2i} + u_{2t} \tag{4}$$

$$y_{2i} = 1\{y_i^* > 0\}, \tag{5}$$

where $(\varepsilon_{1i}, \varepsilon_{2i})$ are iid draws from a bivariate standard normal distribution, $(\mathbf{x}_i', \mathbf{z}_i')$ are covariates, $\sigma > 0$, $\rho_2 \in [-1, 1]$, and $(\boldsymbol{\beta}, \boldsymbol{\gamma})$ are parameters to be estimated, and $1\{\cdot\}$ is the indicator function. The disturbances $(\sigma\varepsilon_{1i}, \rho_2\varepsilon_{1i} + \sqrt{1-\rho_2^2}\varepsilon_{2i})$ in (3)–(4) are therefore mean zero bivariate normal with covariance matrix $\begin{bmatrix} \sigma & \rho_2\sigma \\ \rho_2\sigma & 1 \end{bmatrix}$ and correlation ρ_2.[15] Equations (4) and (5) compose a probit selection process. When $y_{2i} = 1$, the service is chosen to be delayed and delay t_i is drawn from the lognormal distribution given in (3). The model allows the random variable in the selection equation

[14] In particular, if $m_t \equiv \exp(\mathbf{w}_t'\boldsymbol{\alpha})$ and $\delta \equiv \exp(\tau_1^2)$, then $E(n_t|\mathbf{w}_t) = m_t$ and $Var(n_t|\mathbf{w}_t) = m_t[1 + m_t(\delta - 1)]$ (Anscombe, 1950).

[15] The variance of the selection equation disturbance is normalized to unity for identification of $\boldsymbol{\gamma}$.

to be correlated with observed regulatory delays through ρ_2 and u_{2t}. If the selection and duration processes are correlated, then applying the baseline model to the *observed* nonzero delays would lead to biased inference. Correlation through ρ_2 is spell-specific, and might have many causes in the regulatory context. For example, if the regulator mistakenly delays approval of a service it would normally approve, then realizes its error and quickly approves the service after a short delay, then the selection and delay variables will be negatively correlated.[16] Conversely, if unobserved factors make a specific service both more likely to be delayed and to have longer delays than observables account for, then there will be positive correlation. In (3) and (4), u_{2t} is a mixing term, common across y_{1i} and y_{2i}^* and across all i in period t. Correlation through u_{2t} might represent unobserved secular variables that influences regulatory delay.

Taken together, (u_{1t}, u_{2t}) represent unobserved period-specific heterogeneity. To complete the model, assume that (u_{1t}, u_{2t}) are defined by

$$\begin{aligned} u_{1t} &= \tau_1 (\eta_{1t} - \tau_1/2) \\ u_{2t} &= \tau_2 \left(\rho_1 \eta_{1t} + \sqrt{1-\rho_1^2}\, \eta_{2t} \right) \end{aligned}$$

where (η_{1t}, η_{2t}) are iid draws from a bivariate standard normal distribution, τ_1, $\tau_2 \geq 0$, and $\rho_1 \in [-1, 1]$. Furthermore, (η_{1t}, η_{2t}) are independent of all $(\varepsilon_{1i}, \varepsilon_{2i})$. Then (u_{1t}, u_{2t}) are bivariate normal with mean $\begin{bmatrix} -\tau_1^2/2 & 0 \end{bmatrix}'$, covariance $\begin{bmatrix} \tau_1^2 & \tau_1\tau_2\rho_1 \\ \tau_1\tau_2\rho_1 & \tau_2^2 \end{bmatrix}$, and correlation ρ_1. Thus ρ_1 is the key parameter governing correlation between the count and regulatory delay processes. If ρ_1 is positive, then departures from the means in the count and regulatory delay processes are positively correlated, as might happen if there is "regulatory congestion" due to the finite resources of the regulator. If ρ_1 is negative, then the count and regulatory delay processes are negatively correlated, as might happen if the regulator feels pressure to expedite service approvals when there are many new services created or if the firm submits more new services to the regulator when approval times are short. The definition of u_{1t} ensures that $E(e^{u_{1t}}) = 1$, as required above. All parameters are

[16] I thank Michael Katz for this observation.

identified as long as $n_t \geq 1$ for at least one period.[17]

The joint pdf for the data in period t, $\{n_t, y_{1i}, y_{2i} | i \in I_t\}$, conditional on (u_{1t}, u_{2t}), is then

$$f(\{n_t, y_{1i}, y_{2i} | i \in I_t\} | u_{1t}, u_{2t}) = f(n_t | u_{1t}) \cdot \prod_{i \in I_t} \left[F(y_{2i}^* | y_{2i} = 0, u_{2t})^{1(y_{2i}=0)} \cdot g(y_{1i} | u_{2t})^{1(y_{2i}=1)} \right] \quad (6)$$

where the dependence on the parameters is suppressed in the notation. In the expression above, the form of $f(n_t|u_{1t})$, the Poisson pdf, is given by (2). The probability of observing a stochastic duration of length zero, $F(y_{2i}^*|y_{2i} = 0, u_{2t})$, is given by

$$F(y_{2i}^* | y_{2i} = 0, u_{2t}) = \Phi\left(-\mathbf{z}_i'\boldsymbol{\gamma} - u_{2t}\right) \quad (7)$$

and the likelihood of the non-zero durations is

$$\begin{aligned} g(y_{1i}|u_{2t}) &= f(y_{1i}|u_{2t}) F(y_{2i}^*|y_{2i} = 1, u_{2t}) && (8) \\ &= \frac{1}{\sigma}\phi\left(\frac{y_{i1} - \mathbf{x}_i'\boldsymbol{\beta} - u_{2t}}{\sigma}\right)\Phi\left(\frac{\mathbf{z}_i'\boldsymbol{\gamma} + u_{2t} + \rho_2\left[y_{i1} - \mathbf{x}_i'\boldsymbol{\beta} - u_{2t}\right]/\sigma}{\sqrt{1-\rho_2^2}}\right) && (9) \end{aligned}$$

where ϕ is the pdf of a standard normal random variable, Φ is the cdf of the same, $f(y_{1i}|u_{2t})$ is the pdf of the durations conditional on observation and $F(y_{2i}^*|y_{2i} = 1, u_{2t})$ is the probability of observing a non-zero duration (Amemiya, 1985). Since (u_{1t}, u_{2t}) are not observed, one finds the unconditional joint pdf by integrating out (u_{1t}, u_{2t}):

$$f(n_t, \{y_{1i}, y_{2i} | i \in I_t\}) = E_{u_{1t},u_{2t}} f(n_t|u_{1t}) \cdot \prod_{i \in I_t} \left[F(y_{2i}^* | y_{2i} = 0, u_{2t})^{1(y_{2i}=0)} \cdot g(y_{1i}|u_{2t})^{1(y_{2i}=1)} \right] \quad (10)$$

where $E_{u_{1t},u_{2t}}$ denotes expectation taken over (u_{1t}, u_{2t}). The log likelihood for all time periods is

$$l(\boldsymbol{\alpha}, \boldsymbol{\beta}, \boldsymbol{\gamma}, \rho_1, \rho_2, \sigma, \tau_1, \tau_2) = \sum_{t=1}^{T} \log f(n_t, \{y_{1i}, y_{2i} | i \in I_t\}) \quad (11)$$

The expectation in (10) is a double integral that cannot be solved analytically, and numerical methods must be used. I use Gaussian quadrature, with fifty evaluation points in both dimensions.[18]

[17] If not, then the disturbance terms for y_{1i}^* are not separable and (σ, τ_2) enters g (see (6)) only through $\sigma^2 + \tau_2^2$. Strictly speaking, σ and τ_2 are identified even in such a case through the cross equation restriction that τ_2 enters the equations for both y_{1i}^* and y_{2i}^*, but without large sample sizes that restriction is likely to only weakly numerically identify σ and τ_2.

[18] See Quandt (1983, section 8.2) for a brief introduction (and further references) to Gaussian quadrature.

Estimation of this model therefore can be expensive when there are many covariates. Because of this cost, in the application below, I experiment with small numbers of covariates at a time.

Although the general model (11) is rather complicated, it is about the simplest model that allows correlation between the innovation and regulatory delay processes. The model contains several familiar models as special cases. When $\rho_1 = 0$, the count and Tobit models are independent and may be estimated separately with full efficiency. When $\tau_1 = 0$, the count model is the standard Poisson model with no accounting for overdispersion. When $\tau_1 > 0$ the count model is the generalized Poisson model. When $\rho_1 = 0$, these count models are independent of the regulatory approval process, and $\boldsymbol{\alpha}$ can be estimated consistently in either model by MLE based on (2) with u_1 identically equal to zero.[19] Similarly, the Tobit model when $\rho_2 > 0$ but $\tau_2 = 0$ (selection but no correlation with the arrival process) may be estimated by any of the usual methods for the Type II Tobit model (Amemiya, 1985). Estimating these restricted models is useful for hypothesis testing and to provide starting values of $(\boldsymbol{\alpha}, \boldsymbol{\beta}, \boldsymbol{\gamma}, \rho_2, \sigma, \tau_1)$ for the full model maximum likelihood routine.

4 Data and Estimation of the Model

Since the innovation and regulatory delay models are independent if $\tau_1 = 0$, I first discuss each separately in sections 4.1.1 and 4.2. In section 4.3 I discuss the results from joint estimation of the fully correlated model.

4.1 The Innovation Model

4.1.1 Data

I examine new intrastate non-access services introduced by Ameritech in its Indiana territory. In 1997, Ameritech provided information on all of its non-access tariff filings since July 1991 (these tariff filings are part of the public record). The data compose a balanced study period of three

[19] Even when $\tau_1 > 0$, MLE based on a simple Poisson model yields consistent estimates of $\boldsymbol{\alpha}$ as long as the conditional mean is correctly specified (Cameron and Trivedi, 1998, sec. 3.2.3).

Service Category	RoRR Period (7/91–6/94) Total	Yearly Rate	OI Period (7/94–6/97) Total	Yearly Rate	**Total**
A: Local Access, Centrex, and Central Office Services	3	1.0	14	4.67	17
B: All Other Services	6	2.0	12	4.0	18
Total	9	3.0	26	8.67	35

Table notes: counts are new services in Ameritech Indiana's Tariff numbers 7 and 8 before Opportunity Indiana, and in their Service Catalog during Opportunity Indiana. *RoRR* is rate of return regulation; *OI* is Opportunity Indiana.

Table 1: New Services Introduced Before and After Opportunity Indiana

years each before and after the start of Opportunity Indiana. Since no new service can be offered before being added to the tariff, the filings include all new services (see the Data Appendix for a complete list).

Ameritech filed 185 tariff revisions, of which 51 are for new services. The figure does not include promotional offerings (temporary price decreases for existing services) or access services. This figure includes 15 filings of new pricing options for existing services, which are filed as "new services" under IURC rules.[20] Because I want to count truly new services, and not new ways to buy existing services, optional pricing plans are removed from the data set. After removing another filing that contained no new technology,[21] 35 new service filings remain.

The new services are presented by category in Table 1. The services are grouped into two categories. The first category, labeled A, comprises Local (exchange access), Centrex, and Central Office services and features. The second category, labeled B, comprises Operator, Public Telephone, Wireless, Dedicated Communications, ISDN, Video, and Wide Area Telecommunications Services, and other miscellaneous services. These categories are determined from the tariff section in which the new service is included.

[20] This is not to say that new pricing options do not increase consumer welfare. If no existing pricing option is replaced by a new option, the new option *must* (weakly) increase welfare.

[21] The ISDN Data Usage filing (11/25/92) is merely a different price for a particular usage of technology already covered in the ISDN Prime, Direct, and Centrex filings from the same date.

4.1.2 Results

The dependent variable n_t for (2) is the number of services in each month in each service category. Separating the filings between categories allows the arrival process to differ between categories (i.e., allows category-specific fixed effects). The random effect u_{1t} is common to both category counts in month t.[22] The two categories and 72 months in the observation period yield 144 count observations.

The results of the estimation when the counts and durations are taken to be independent ($\rho_1 = \tau_2 = 0$) are in Table 2. The first two columns are for the *no covariates* model with no variables other than dummies for service category and Opportunity Indiana. The table includes the estimates for the regulatory delay models as well, although I defer discussion of those results until section 4.2. The category A coefficient (category B is the excluded dummy) is positive, meaning that under the rate of return regime Ameritech introduced services in the Local, Centrex, and Central Office category at a slower rate than Other services. The estimated coefficient of -0.693 for Local, Centrex, and Central Office services implies that the creation rate of such services is 50% that of Other services (which exactly matches the data in Table 1).[23] To interpret the coefficients in the generalized Poisson estimations, recall that the log of the conditional mean of the count process is linear in the covariates, so that positive coefficients imply greater counts.

Both of the Opportunity Indiana coefficients are positive (although only the first is statistically significant); there were more services in each category during the alternative regulatory regime. The coefficients imply that the mean arrival rate of services in the Local, Centrex, and Central

[22] For clarity of presentation, the notation in section 3 does not refect the panel aspect of the count data. If the category dimension is indexed by $j \in J$, then (1) and (2) would be replaced by

$$\lambda_{jt} = \exp\left(\mathbf{w}_{jt}'\boldsymbol{\alpha} + u_{1t}\right) = e^{u_{1t}} \exp\left(\mathbf{w}_{jt}'\boldsymbol{\alpha}\right)$$

and

$$f(\{n_{jt}\}_{j\in J} | \{\mathbf{w}_t'\boldsymbol{\alpha}\}_{j\in J}, u_{1t}) = \prod_{j\in J} \exp(-\lambda_{jt})\lambda_{jt}^{n_{jt}}/n_{jt}!$$

[23] The percentage change in arrivals due to an indicator x_j taking a value of 1 is $\%\Delta = e^{\beta_j} - 1$.

Variable	**No Covariates** coefficient	s.e.	**Covariates** coefficient	s.e.
Innovation Count Model				
intercept	-1.801^{***}	0.444	-2.607^{***}	0.470
category A services	-0.693	0.724	-0.691^{*}	0.385
Opp. Indiana: category A services	1.563^{**}	0.672	2.953^{***}	0.643
Opp. Indiana: category B services	0.716	0.543	2.108^{***}	0.663
population			-0.758	4.208
firm R&D (one year lag)			2.079^{***}	0.706
Regulatory Delay Duration Model				
intercept	4.255^{***}	0.272	4.209^{***}	0.286
Opportunity Indiana	-3.690^{***}	0.329	-3.647^{***}	0.348
revenue (Opp. Indiana period only)			-0.005	0.041
pages in tariff filing			0.136	0.243
Regulatory Delay Selection Model (Opp. Indiana period only)				
intercept	0.487^{*}	0.259	0.497	0.255
pages in tariff filing			0.668	0.462
Incidental Parameters				
ρ_1 (innovation and delay corr.)	0	fixed	0	fixed
ρ_2 (selection and duration corr.)	0.942^{***}	0.067	0.973^{***}	0.061
σ (duration dispersion)	0.817^{***}	0.126	0.822^{***}	0.132
τ_1 (innovation random effect s.d.)	0.816^{***}	0.267	0.559^{*}	0.345
τ_2 (delay random effect s.d.)	0	fixed	0	fixed
log likelihood	-125.08		-121.22	
observations (counts)	144		144	
observations (durations)	34		34	

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

Table notes: estimation is MLE based on (11) with ρ_1 and τ_2 fixed at zero, where the likelihood is evaluated via Gauss-Hermite quadrature with 50 weighting points. Significance levels are for one-tailed tests for τ and σ and two-tailed tests for all other parameters. Dependent variable for the count model is monthly services within each category; dependent variable for the duration model is log of the stochastic part of the regulatory delays (in days), as described in section 3; dependent variable for the selection equation is 1 if delayed, 0 if not. The selection model is estimated for the Opportunity Indiana (OI) period only; Pr(delay) before OI is taken to be 1. In each model, the conditional mean is linear in the covariates. Continuous variables are in logs. *Revenue* is the yearly forecast at time of filing, available only during OI.

Table 2: Estimation Results for the Generalized Poisson/Tobit II Model—Independent Version

Office category is 359% higher during Opportunity Indiana than before; that for Other services is 97% higher. The former is significant at the 1% level; the latter is not significant (although the two are jointly significant).[24] Thus Opportunity Indiana is correlated with a large increase in services created, especially Local, Centrex, and Central Office services.

Before putting weight on any interpretation of these results, it is necessary to test the *no covariates* model's assumptions. Consider first the possibility that the mean is incorrectly specified through the omission of relevant variables. In empirical studies of telecommunications innovation, economic and demographic variables often significantly affect innovation (e.g., Greenstein et al., 1995). If variables that in fact determine the mean are omitted, then all coefficient estimates will be inconsistent. We want to be assured that innovation did not jump up during Opportunity Indiana in response to changes in some causal variable other than the removal of the regulatory burden. Accordingly, the second and third columns of Table 2 (labeled *covariates*) report an expanded regression including other variables.[25]

In a short time series from a single state such as this one, most economic and demographic variables are highly correlated, so that multicollinearity precludes including all variables of interest. For example, population, income, and number of access lines in the territory should affect the demand for new services. Including them all, however, yields nonsensically large coefficients and huge standard errors. In the *covariates* regression reported in Table 2, I include log population in Indiana. The coefficient is not significant. The same result obtains if access lines or per-capita income are included instead of population. In each case the significance of the Opportunity Indiana variables is unchanged.

The next variable to check in the specification is a control for changes in the size of the population of potential new services. I add the log of real R&D spending by Ameritech, lagged one year. The estimated coefficient is positive and statistically significant at the 1% level.[26] The coefficient

[24] A Wald test statistic for joint significance is at the 0.96 quantile of a $\chi^2(2)$ distribution.

[25] In these estimations all new covariates are demeaned, so the intercepts are comparable across columns.

[26] Firm R&D figures are the portion of R&D spending by Ameritech allocated to Indiana Bell (FCC *Statistics of*

on R&D implies that every extra percent spent by the firm on R&D resulted in 2.08 percent more new services the following year, which works out to about \$1.5M in R&D spending (above the average) per extra new service (above the average). In exploratory work, neither industry R&D nor U.S. patent counts from relevant technological areas were significant.

Adding the extra variables in the *covariates* estimation does not steal the significance from the Opportunity Indiana indicators,[27] and in fact greatly increase the estimated impact (1816% more services in the category A and 723% more in category B).

A final area to explore is competition. To avoid the *competition effect* pitfall (Sappington and Weisman, 1996), which ascribes benefits to lighter regulation when in fact they are caused by increased competition, I control for the number of competitors in a separate estimation. What was the competitive environment in Indiana during the study period? For the services under consideration here, the closest form of competition would come from competing local exchange companies and competitive access providers (CAPs). Local competition was still virtually nonexistent by the end of 1997.[28] CAPs were making more progress in Indiana, however. The variable *CAPs* is the log number of firms providing high-speed bypass access to businesses.[29] The estimated coefficient on *CAPs* when it is added to the *covariates* regression is -0.645 (1.03), insignificant. Most importantly, the Opportunity Indiana coefficients are still positive and significant at the 5% level;[30] it does not appear that competition explains the increase in services.

It is never possible to test every potential covariate when considering omitted variables. However, to examine the effects of shocks apart from the regulatory variables in a general way, I also estimated the innovation model with the variables from the first column of 2 and a set of year

Communications Common Carriers), adjusted by the GDP deflator.

[27] A Wald test statistic for the joint significance of the Opportuntiy Indiana indicators has a p-value of 1.5E-4.

[28] Compared with the 2.4 million access lines owned by Ameritech Indiana, the company sold only 159 local service lines for resale and five lines as unbundled network elements to other companies as of December 1997 (Indiana Utility Regulatory Commission, 1998).

[29] Data are from Harris (1997, exhibit 1), the FCC (Kraushaar, 1991–1997), and the IURC *Annual Report 1996-1997*. The number of CAPs ranges from one in 1991 to 14 in 1997.

[30] The two Opportunity Indiana coefficients are 2.73*** (0.915) for category A and 1.89** (0.82) for category B. They are jointly significant.

dummies. These shocks might represent changes in the national or regional economy, competition, or demand that affect the firm's innovation. Even with the year dummies (none of which was significant),[31] the Opportunity Indiana coefficients are still negative, with values falling between the *covariates* and *no covariates* estimates in table 2. Both are significant.[32]

Finally, consider the generalized Poisson specification itself. The estimated dispersion parameter τ_1 is 0.8 in the *no covariates* estimation, which implies that the average ratio of the conditional variance to the conditional mean in the sample is 6.6 (this ratio would be unity in the simple Poisson model).[33] The estimated τ_1 drops to 0.56 in the *covariates* estimation, which implies that the average predicted coefficient of variation in the sample is 3.1. Apparently the added covariates account for some of the unexplained heterogeneity in the *no covariates* estimation. Both estimates of τ_1 are significant, although only marginally so in the *covariates* estimation. Since omitted variables in the conditional mean function can induce overdispersion, the marginal significance of τ_1 in the *covariates* estimation gives some assurance that other important variables are not omitted.

The predicted mean services per year from the *no covariates* and *covariates* models are given in Table 3 for the periods before and during Opportunity Indiana. The prediction is $T^{-1}\sum \exp(\mathbf{w}_t'\hat{\boldsymbol{\alpha}})$, the mean-in-sample of the predicted mean. The *no covariates* model estimates that under Opportunity Indiana the average number of new services increased by 5.8 per year compared with the RoRR period (a 196% increase), which is close to the actual observed increase of five and two-thirds per year.[34] The *covariates* model predictions are similar: the estimated mean yearly services from the *covariates* estimation is 3.00 for the rate of return era and 8.69 for the Opportunity Indiana era.

[31] An LR test of the model with year fixed effects vs. the model without fails to reject the simpler model (statistic is at the 0.818 quantile of a $\chi^2(6)$ random variable).

[32] The two Opportunity Indiana coefficients are 2.78^{***} (1.05) for category A and 1.96^{*} (1.00) for category B. They are jointly significant.

[33] The conditional variance of the generalized Poisson n_t is $\mu_t(1+\mu_t[\exp(\tau_1^2)-1])$, where the conditional mean μ_t is $\exp(\mathbf{w}_t'\hat{\boldsymbol{\alpha}})$.

[34] Unlike the standard Poisson model, when there are random effects ($\tau_1 > 0$) the mean-in-sample of the predicted mean calculated from the ML estimates is not identically equal to the sample average of the counts.

Period	**Predicted Mean Yearly Services** *No Covariates*	*Covariates*	**Actual Mean Yearly Services**
Rate of Return Regulation	2.97	3.00	3.00
Opportunity Indiana	8.78	8.69	8.67

Table 3: Predicted vs. Actual New Services

4.2 The Regulatory Delay Model

4.2.1 Data

Turn now to the regulatory delay submodel. The tariff data also contain the effective date of the filing, so that the approval delay can be calculated. Before Opportunity Indiana, a new service could not be introduced until its tariff filing had been actively approved by the IURC, and the commission required the filing to be docketed for at least 30 days before decision. Delay in excess of the 30 day minimum was caused by staff investigation of the tariff and the wait until the next commission meeting. After July 1994 under Opportunity Indiana, new services were generally presumed lawful (requiring no action by the IURC) and were introduced with one day of notice. The average delay for all services over the entire observation period was 39 days, the median was four days, and the maximum was 217 days.

The minimum mandated regulatory delays (30 days before Opportunity Indiana and one day during Opportunity Indiana) are treated as deterministic in the analysis (refer to discussion in section 3). Summary statistics on the remaining, stochastic part of the delay times are in Table 4. I refer to the total delay less the deterministic part as the *adjusted delay time*.

The delays during Opportunity Indiana were very short, averaging about a week.[35] In fact, if the largest adjusted delay time of 93 days is dropped, the next highest is a mere nine days and the average falls to three days. This outlier was not included in the estimations to avoid undue

[35] According to Ameritech, there were no approval delays at all during Opportunity Indiana, except when the company chose a later effective date for a filing than the minimum allowed (personal communication with Bruce Hazelett, Director of Regulatory Affairs, Ameritech Indiana). I remain agnostic about the cause of the apparent delays. Given that the "delays" (real or self-imposed) are so short, they do not affect the cost of regulation calculation much in section 5.

	Observations	Zeroes	For Non-Zero Observations Min.	Median	Mean	Max.
RoRR	9	0	13	96	102.6	187
Opportunity Indiana	26	7	1	2	7.6	93
Entire Period	35	7	1	3	38.1	187

Table notes: all units are days. *RoRR* is Rate of Return Regulation. The adjusted delay times are the total delay times less the minimum mandated approval periods (30 days for the RoRR regime, one day for Opportunity Indiana).

Table 4: Summary Statistics for the Adjusted Approval Delay Durations

influence on the results.

Limited characteristics of each filing are also available. The number of pages composing the filing can proxy the unobserved complexity of the filing. During Opportunity Indiana we also have the projected revenue from the service, which is submitted to the IURC with the filing.

4.2.2 Results

Section 3 introduces a bivariate selection model to handle the large number of zeroes in the adjusted delay times. Recall that correlation between the decisions is captured by the parameter ρ_2, which is zero if the decisions are uncorrelated. There are no zeroes in the data before Opportunity Indiana (i.e., the services are always delayed), so the selection equation is estimated only for the Opportunity Indiana period (i.e., before Opportunity Indiana the selection equation puts weight 1 on $y_2 = 1$).

The bottom part of Table 2 contains the estimated model for the case in which the duration and selection models are independent of the count model ($\tau_1 = 0$). The first two columns include no covariates, other than an indicator for Opportunity Indiana. The final two columns include the pages and revenue covariates. In both estimations, the indicator for Opportunity Indiana in the duration model is strongly significant. In fact, the indicator alone is responsible for reducing the delays nearly to zero: the percentage change due to the indicator is about -97.5% in each

	No Covariates		Covariates	
	RoRR	OI	RoRR	OI
Mandated Minimum Delay	30	1	30	1
Predicted Probability of Delay	1.0 (fixed)	0.69	1.0 (fixed)	0.67
Predicted Total Delay				
observed (conditional on $y_2 = 1$)	128.4	4.09	129.4	4.13
latent	128.4	3.45	129.4	3.44

Table notes: units are days. Last two rows are the sum of the mandated minimum delay and the predicted stochastic delay. *RoRR* is the rate of return regulation period; *OI* is the Opportunity Indiana period. Figures calculated as mean-in-sample using estimated coefficients from Table 2; see footnote 37 for details.

Table 5: Predicted Regulatory Approval Delays

estimation.[36] Before Opportunity Indiana all approvals were delayed; during Opportunity Indiana the chance of delay drops to 0.69 in the first estimation and 0.67 in the second (see Table 5). Neither tariff filing pages nor service revenues are significant in either the selection or the duration models. Pages has the expected sign: more pages (i.e., more complex filings) means that the probability of delay is higher and that delays are longer. Services with higher expected revenues create longer delays when delayed. In both estimations the selection and delay disturbances are highly correlated; ρ differs significantly from zero. This positive correlation might be caused by unobserved factors that lead the regulator both to delay approval with higher probability and to make the delay longer than average.

The estimated total delay, including the minimum approval period, is much shorter under Opportunity Indiana. Combining the selection and the delay results, expected total observed delay time fell from about 129 days under RoRR to about 4 days during Opportunity Indiana.[37] Table 5 lists the predicted delays from the two estimations.

[36] The formula in footnote 23 applies to this model as well.

[37] Predicted total *observed* delay before Opportunity Indiana is calculated as the mean-in-sample of the minimum delay plus $E(e^{y_1}|\mathbf{x}_i)$: i.e., the average of $30 + \exp(\mathbf{x}_i'\hat{\boldsymbol{\beta}} + \sigma^2/2)$ (e^{y_1} is lognormal) in the pre-OI sample. During Opportunity Indiana, the formula is the average of $\Pr(y_2 = 0|\mathbf{z}_i) \cdot 1 + \Pr(y_2 = 1|\mathbf{z}_i) \cdot E(e^{y_1}|y_2 = 1, \mathbf{x}_i, \mathbf{z}_i)$ in the OI sample, where $E(e^{y_1}|y_2 = 1, \mathbf{x}_i, \mathbf{z}_i)$ is approximated by $\exp\{E(y_1|y_2 = 1, \mathbf{x}_i, \mathbf{z}_i)\} \cdot [1 + V(y_1|y_2 = 1, \mathbf{x}_i, \mathbf{z}_i)/2]$, a second order approximation. The expressions for $E(y_1|y_2 = 1, \mathbf{x}, \mathbf{z})$ and $V(y_1|y_2 = 1, \mathbf{x}, \mathbf{z})$ are given in Greene (1993, p.707). Predicted total *latent* delay is the average of $\exp(\mathbf{x}_i'\hat{\boldsymbol{\beta}} + \sigma^2/2)$in the pre- or post-OI sample, as indicated.

4.3 The Jointly Correlated Model

Turn now to the full model with unrestricted correlation allowed between the count and duration models. The results from the unrestricted *no covariates* and *covariates* estimations are in Table 6. In these estimations, unlike those reported in Table 2, ρ_1 and τ_2 are allow to vary. In general, the coefficients of interest are close to those in Table 2. In particular, the Opportunity Indiana indicators are still significant.

An examination of the incidental parameters gives insight into the correlation among the models. The correlation between the count and duration models is negative, as measured by ρ_1 in the *no covariates* estimation. Such negative correlation means that a shock leading to more innovation, and therefore more tariffs submitted to the regulator, has the effect of reducing regulatory delay time. An interpretation may be that the IURC relaxes its scrutiny of tariffs as they stacked up in the regulatory inbox. The estimate of ρ_1 is not significant, however. Turning to the *covariates* estimation, $\hat{\rho}_1$ is again negative but no assessment of its significance can be made because the likelihood is maximized at the boundary $\rho_1 = -1$. This is likely due to the modest sample size and the numerical complexity of the estimation. LR tests of each unrestricted estimation against its restricted counterpart fail to reject the restricted model. Given this, it appears best to rely on the estimations in Table 2 for the prediction exercise in the next section.

One change between the restricted and unrestricted models is that σ, the duration dispersion parameter, drops from about 0.8 in the restricted model to about 0.6 in the unrestricted model. This is partly due to the extra heterogeneity entering the duration equation through τ_2; heterogeneity from σ is being partly replaced by heterogeneity from the monthly random effect, the strength of which is measured by τ_2.

Variable	No Covariates coefficient	s.e.	Covariates coefficient	s.e.
Innovation Count Model				
intercept	-1.781^{***}	0.452	-2.605^{***}	0.498
category A services	-0.693	0.761	-0.693	0.758
Opp. Indiana: category A services	1.523^{**}	0.707	2.936^{***}	0.800
Opp. Indiana: category B services	0.676	0.561	2.089^{***}	0.674
population			-0.155	1.406
firm R&D (one year lag)			2.198^{***}	2.198
Regulatory Delay Duration Model				
intercept	4.294^{***}	0.396	4.339^{***}	0.330
Opportunity Indiana	-3.585^{***}	0.364	-3.505^{***}	0.366
revenue (Opp. Indiana period only)			0.011	0.032
pages in tariff filing			0.200	0.244
Regulatory Delay Selection Model (Opp. Indiana period only)				
intercept	0.709^{**}	0.361	0.846^{***}	0.299
pages in tariff filing			0.791	0.533
Incidental Parameters				
ρ_1 (innovation and delay corr.)	-0.279	0.661	-1.000	fixed†
ρ_2 (selection and duration corr.)	0.941^{***}	0.092	0.966^{***}	0.075
σ (duration dispersion)	0.666^{***}	0.131	0.640^{***}	0.140
τ_1 (innovation random effect s.d.)	0.794^{***}	0.298	0.420^{**}	0.250
τ_2 (delay random effect s.d.)	0.472^{***}	0.203	0.541^{***}	0.199
log likelihood	-124.15		-119.50	
observations (counts)	144		144	
observations (durations)	34		34	

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

†The estimated likelihood is maximized when $\rho_1 = -1$; see text.

Table notes: see notes to table 2.

Table 6: Estimation Results for the Generalized Poisson/Tobit II Model—Joint Version

5 The Benefits of Opportunity Indiana

Now that we have estimated the parameters for the entire model, we can compare the overall effect of the regulatory change on service introductions in the two periods. This section uses the parameters estimated in the restricted, independent *covariates* model (columns three and four of Table 2). We can compare two counterfactual scenarios. In scenario 1, Ameritech remains under RoRR for the entire observation period (July 1991 to June 1997). In scenario 2, Ameritech is under Opportunity Indiana for the entire observation period. For each scenario, the covariates for the innovation model are set to their appropriate yearly values and the covariates for the regulatory delay models are set to sample averages. How many new services would we expect to have been introduced in each scenario? Comparing the two answers will measure the total effect that the regulatory regime had on innovation. These answers differ from a simple extrapolation of the predicted means in Table 3 because of the counterfactual value of the Opportunity Indiana dummy variable here and because here I am interested in services that are both created and approved, not just innovation (although the latter ends up making little difference).

Table 7 shows that in Scenario 1, the model predicts Ameritech creates and introduces about 11 new services. Under Opportunity Indiana, the company creates and introduces about 133 services—12 times as many as in scenario 1. In each scenario more services are created in the first part of the observation period because the coefficient on R&D is positive and observed R&D expenditure exhibits a downward trend. Approval delays have little effect in these calculations, because the delays in each scenario are relatively short compared to the observation period and because most services are created near the beginning of the observation period.[38] The average is 121.82 more services approved under Opportunity Indiana than under rate of return regulation; the standard

[38] Less that 0.1 service on average is still delayed at the end of the observation period in each scenario. Finding the services introduced net of those still pending approval at the end of the observation period requires a result from queuing theory. If arrivals are Poisson with mean λ and the CDF of the delays is G, then the services introduced (the output of the queuing system) at time t is a nonhomogeneous Poisson process with mean $\Lambda(t) = \lambda \int_0^t G(s)ds$ (Ross, 1983, p.39). The actual formula used is a modification of this result reflecting two extensions: the arrival process is generalized Poisson, and the arrival rate varies each year. Details of the calculation are available upon request.

	Scenario 1: Rate of Return	**Scenario 2:** Opportunity Indiana	**Difference**
Year 1	3.56	42.25	
Year 2	3.06	36.39	
Year 3	2.38	28.30	
Year 4	1.17	13.89	
Year 5	0.63	7.55	
Year 6	0.30	4.60	
Entire Period (sum)	11.10	132.98	121.88
standard deviation	*3.47*	*16.30*	*16.67*

Table notes: cell entries are the average number of services created in the given year that are introduced to subscribers (i.e., were approved by the regulator) by the end of the observation period. Prediction is based on the estimated *covariates* model from Table 2. See text and footnote 39 for details.

Table 7: Comparison of Predicted Services Introduced to Subscribers

deviation of 16.7 indicates that the difference is bounded well away from zero in probability.[39]

Exercises comparing counterfactual scenarios are only as good as the assumptions they rest upon. Given the difficulties inherent in out-of-sample prediction, it is perhaps best to view this exercise as illustrative rather than to place stress on the numerical results *per se*. The results from the exercise illustrate the two ways that incentive regulation benefits consumers of telecommunications services. First, many more new telecommunications services were created under the price caps and market pricing. It seems that the incentive to introduce new services increased substantially when the IURC replaced RoRR with Opportunity Indiana. Second, the reduced mandatory approval time and the lower probability of a service being delayed beyond the minimum means that services were available to customers sooner under price caps.

Even if Ameritech would have introduced twelve times as many services under Opportunity Indiana, welfare probably would not have risen proportionally. The services introduced during Opportunity Indiana were probably not of importance comparable to the services introduced under

[39] The standard deviation accounts for intrinsic variation only, not estimation error.

the rate of return regime. The incremental services introduced during Opportunity Indiana—those that would not have been introduced under the previous regime—most likely created less revenue for the BOCs since they were formerly deemed unprofitable to introduce. Accordingly, they were probably worth less to consumers as well. There is no way to estimate this difference in worth (to the company or the consumer) with the present data.[40]

6 The Validity of the Model for Policy Analysis

It appears that the price cap/deregulatory regime had a large impact on the number and timing of service introductions and therefore on consumer welfare. There are potential problems to consider when moving from correlation to causality, however. In addition to the competition effect pitfall discussed above, two other pitfalls may apply. The *demonstration effects* pitfall (Sappington and Weisman, 1996) is created when regulated carriers perform actions regulators desire, such as introducing new services, to "demonstrate" the success of a favorable regulatory regime. Unfortunately—from the standpoint of looking for a "natural experiment"—it is quite likely that the demonstration effect influenced Ameritech's actions. There was much discussion in the regulatory hearings about the potential for lighter regulation to spur innovation, both in 1994 before Opportunity Indiana and in 1997 when it was under review. Innovation was clearly one dimension of performance on which Ameritech was to be judged.

The *sequencing pitfall* applies if firms' actions are spurred by anticipated changes in the regulatory regime. This pitfall might apply to these data if, right before the switch to Opportunity Indiana, Ameritech held off on introducing new services in anticipation of the more favorable environment shortly coming. This pitfall is related to the demonstration effect; shifting a service introduction into Opportunity Indiana also has the advantage of making the relaxed regulation look more conducive to innovation. The evidence is inconclusive on whether sequencing effects are

[40] Service revenue data are available only for the Opportunity Indiana filings.

present or not. Ameritech submitted a spate of tariffs shortly after Opportunity Indiana began. When indicators for the six months right before the regime change and right after are included in the *no covariates* estimation, they have the signs that the sequencing pitfall predicts. However, when other covariates (population and R&D) are controlled for, neither sequencing effect indicator is signficant. In both estimations the positive signs remain on the Opportunity Indiana indicators.[41]

A simple test to see if the demonstration and sequencing effects unduly influence the data is to compare a prediction from the estimated model with some out-of-sample data. In particular, if the demonstration and sequencing effects caused Ameritech to introduce more services than Opportunity Indiana would have spurred apart from these incentives, then the rate of service introduction should drop off over time. The model predicts Ameritech would produce about 6.5 services on average in a nine month period under the lighter regulation (from Table 3). During the first nine months of 1999, the company introduced eight new services—exceeding the prediction, and well within the normal variation of a Poisson process.[42] So although the caution remains, the demonstration and sequencing effects do not appear to unduly bias the estimation.

No regulatory reform should be judged along a single dimension. Opportunity Indiana was a success at getting more services into the hands of consumers, but may not have performed as well by other criteria.[43] It would also be informative to measure the impact of the new services on consumer welfare—a task that awaits better data. Regardless, the evidence indicates that moving away from rate of return regulation spurs product innovation. This study is an initial step towards filling the lacuna in the literature concerning alternative regulation and product innovation.

[41] When added to the *no covariates* estimation with $\rho_1 = 0$, the indicator for the six months just before the regime change has coefficient -0.458 (0.973); the indicator for the six months just after the regime change has coefficient 1.150 (0.500). Adding these indicators also steals some significance from the Opportunity Indiana indicator for local, Centrex, and CO services: the p-value rises to 0.089. When added to the *covariates* estimation with $\rho_1 = 0$, the indicator for the six months before the change has coefficient -0.111 (0.972); the indicator for the six months after the change has coefficient -1.12 (0.708). Again, adding these indicators steals some significance from the Opportunity Indiana indicator for local, Centrex, and CO services: the p-value rises to 0.062.

[42] Eight is the 0.79-quantile of a Poisson random variable with mean 6.5.

[43] In particular, questions were raised about apparently declining service quality (Sword, 1999).

References

Acton, Jan Paul and Vogelsang, Ingo (1989), 'Introduction', *RAND Journal of Economics* **20**(3), 369–372.

Ai, Chunrong and Sappington, David E. M. (1998), The Impact of State Incentive Regulation on the U.S. Telecommunications Industry. Mimeo.

Amemiya, Takeshi (1985), *Advanced Econometrics*, Cambridge: Harvard University Press.

Anscombe, F. J. (1950), 'Sampling Theory of the Negative Binomial and Logarithmic Series Distributions', *Biometrika* **37**(3/4), 358–382.

Averch, Harvey and Johnson, L.L. (1962), 'Behavior of the Firm Under Regulatory Constraint', *American Economic Review* **52**(5), 1053–1069.

Bonanno, Giacomo and Haworth, Barry (1998), 'Intensity of Competition and the Choice Between Product and Process Innovation', *International Journal of Industrial Organization* **16**(4), 495–510.

Cabral, Luis M. B. and Riordan, Michael H. (1989), 'Incentives for Cost Reduction Under Price Cap Regulation', *Journal of Regulatory Economics* **1**(2), 93–102.

Cameron, A. Colin and Trivedi, Pravin K. (1998), *Regression Analysis of Count Data*, Econometric Society Monographs, 30, Cambridge: Cambridge University Press.

Greene, William H. (1993), *Econometric Analysis*, 2nd edn, New York: Macmillan Publishing Co.

Greenstein, Shane, McMaster, Susan and Spiller, Pablo T. (1995), 'The Effect of Incentive Regulation of Infrastructure Modernization: Local Exchage Companies' Deployment of Digital Technology', *Journal of Economics and Management Strategy* **4**(2), 187–236.

Harris, Robert G. (1997), 'Testimony in Support of Opportunity Indiana II Submitted by Ameritech Indiana to the IURC, Cause No. 39705'.

Hausman, Jerry (1997), 'Valuing the Effect of Regulation on New Services in Telecommunications', *Brookings Papers on Economic Activity* pp. 1–38. Microeconomics.

Indiana Utility Regulatory Commission (1990), 'Final Report, Cause No. 37905', September 19.

Indiana Utility Regulatory Commission (1997*a*), 'Annual Report, 1996–1997'.

Indiana Utility Regulatory Commission (1997*b*), 'Telephone Report to the Regulatory Flexibility Committee of the Indiana General Assembly', July 1.

Indiana Utility Regulatory Commission (1998), 'Telephone Report to the Regulatory Flexibility Committee of the Indiana General Assembly', July 1.

Kraushaar, Jonathan M. (1991–1997), 'Fiber Deployment Update'. Industry Analysis Division, Common Carrier Bureau, Federal Communications Commission.

Littlechild, Stephen C. (1983), 'Regulation of British Telecom's Profitability', Report to the Secretary of State, Department of Industry, London, February.

Mueller, Milton L. (1993), *Telephone Companies in Paradise: a Case Study in Telecommunications Deregulation*, New Brunswick: Transaction Publishers.

Prieger, James E. (1999), Regulation, Innovation, and the Introduction of New Telecommunications Services, PhD thesis, University of California, Berkeley.

Quandt, Richard E. (1983), Computational Problems and Methods, *in* Z. Griliches and M. Intriligator, eds, 'Handbook of Econometrics. Volume I', Amsterdam; New York: Elsevier, pp. 699–764.

Ross, Sheldon M. (1983), *Stochastic Processes*, Wiley Series in Probability and Mathematical Statistics. Probability and Mathematical Statistics., New York: John Wiley & Sons.

Sappington, David E.M. and Weisman, Dennis L. (1996), 'Potential Pitfalls in Empirical Investigations of the Effects of Incentive Regulation Plans in the Telecommunications Industry', *Information Economics and Policy* **8**, 125–140.

Sword, Doug (1999), 'Ameritech Repair is Slowest', *The Indianapolis Star* p. E1, May 31.

Taylor, William E., Zarkadas, Charles J. and Zona, J. Douglas (1992), Incentive Regulation and the Diffusion of New Technology in Telecommunications. NERA mimeo.

United States. Federal Communications Commission. Common Carrier Bureau (1987–1997), *Statistics of Communications Common Carriers*, Washington, D.C.: U.S. GPO.

Data Appendix

This appendix contains all the new services Ameritech introduced in Indiana during the period three years before and after the onset of Opportunity Indiana.

Service	Date of Filing	Date Effective	Approval Delay
Before Opportunity Indiana			
OPTINET 64 Kbps	11/14/91	02/05/92	83
Centrex Custom Calling Features	09/30/92	02/03/93	126
OPTINET DS1 384 Kbps	10/15/92	01/06/93	83
ISDN Prime	11/25/92	06/30/93	217
ISDN Direct	11/25/92	06/30/93	217
ISDN Centrex	11/25/92	06/30/93	217
Residence Custom Calling Type 2	03/29/93	09/08/93	163
Caller ID with Name	11/19/93	01/02/94	44
800 Directory Assistance	05/11/94	06/23/94	43
During Opportunity Indiana			
Billing Reports	07/27/94	07/28/94	1
Toll Restriction—Residential	09/08/94	09/13/94	5
Ameritech Area Wide Networking	09/15/94	09/19/94	4
Scan Alert	10/13/94	10/15/94	2
High Voltage Protection Service	10/28/94	10/30/94	2
128 & 256 Kbps (Fractional DS1)	10/28/94	10/31/94	3
ISDN Direct New Features	10/28/94	11/07/94	10
Ameritech ISDN Prime New Features	11/03/94	11/07/94	4
2-Way DID with Call Transfer	11/09/94	11/15/94	6

Service	Date of Filing	Date Effective	Approval Delay
Answer Supervision with Line Side Interface	11/14/94	11/15/94	1
Ameritech Digital Transport Service	01/17/95	01/18/95	1
Ameritech Advanced Video Service	02/03/95	02/06/95	3
Packet Switched Network Services	02/24/95	02/27/95	3
Business Call Forwarding—Temporary	03/03/95	03/06/95	3
Ameritech Intercept Referral Extension	03/16/95	03/20/95	4
Advanced Custom Calling Features—Pay Per Use	05/31/95	06/01/95	1
Ameritech Call Control	07/07/95	07/10/95	3
FlexLine	09/15/95	09/18/95	3
Prepaid Card	12/11/95	12/15/95	4
Information Call Completion	02/29/96	03/01/96	1
Calling Party Pays—Paging	04/09/96	04/10/96	1
Movers Call Forwarding	05/24/96	08/26/96	94
Ameritech Prime Number	06/04/96	06/06/96	2
Call Detail	06/18/96	06/19/96	1
Inmate Collect	06/28/96	07/01/96	3
3-Way Calling Pay Per Use	03/20/97	03/24/97	4