A Model for Regulated Product Innovation and Introduction with Application to Telecommunications

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Abstract

This empirical examination compares product innovation under price caps and rate of return regulation (RoRR) in U.S. telecommunications. The econometric model comprises a count process (for product innovation) followed by a duration process (for regulatory delay). More services were created under price caps than under RoRR. The model may also be useful for other regulatory settings and issues.

Keywords: product innovation, regulation, telecommunications, Poisson, Weibull, count data, duration data.

Suggested Running Title: Regulated Product Innovation and Introduction

1 Introduction

Some proponents of price cap regulation claim that the price-capped firm has greater incentives to innovate than does the firm under rate of return regulation (RoRR).[1] The literature on price caps and innovation (e.g., Cabral and Riordan, 1989) focuses almost exclusively on cost reduction (process innovation), not product innovation. However, the surplus created by a new product can be much higher than the surplus from cost reductions for existing products. Notwithstanding, one finds few attempts to measure the effects of regulation on product innovation. In this letter, I present an econometric model for regulated product innovation and examine a regulatory shift from RoRR to price caps in U.S. telecommunications.

2 A Model for Regulated Service Innovation and Introduction

A regulated telecommunications firm goes through two steps in the course of introducing services to subscribers. The firm first creates a new end-user service, a countable event. This is the *innovation* step. After innovation, the service is not available to subscribers until it is approved by the regulator, a duration. This is the *introduction* step. Regulation may affect both steps, and if the two are correlated, joint estimation is required for correct inference. I present an econometric model that allows for correlated heterogeneity in the innovation and introduction steps. In the model, innovation follows a generalized Poisson event process. After creation, a new service immediately enters a generalized Weibull duration process that determines the length of the time until regulatory approval. The duration process may be correlated with the event process. Regulatory explanatory variables may affect the rates of the Poisson and Weibull processes.

In the next section, I derive the likelihood function for the observed innovation and introductions, from which MLE may be performed. In section 4, I examine the effects of switching from

[1] RoRR 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 RoRR and price caps, respectively.

RoRR to price caps on the innovation of Ameritech, a U.S. local exchange company (LEC). Variants of this model have been applied to other national (Prieger, 2001*a*) and state-level (Prieger, 2001*b*) data sets on regulated telecommunications services.

3 The Econometric Implementation

Let n_{tk} be the number of events (new services) in period t of type k, $t = 1, \ldots, T$, $k = 1, \ldots, K$. The k dimension covers different service categories. The length of each period is the same. Denote by $f(n_{tk}|u_{1t})$ the probability density function (pdf) of n_{tk}, conditional on type fixed effect α_k, covariate vector x_{tk}, coefficients β, and a random effect u_{1t}. A simple form for f is the generalized Poisson (Cameron and Trivedi, 1998) with pdf

$$f(n_{tk}|u_{1t}) = \exp(-\lambda_{tk} + n_{tk} \log(\lambda_{tk}))/n_{tk}!, \tag{1}$$

where the conditional mean is modeled $\lambda_{tk} \equiv \exp(\alpha_k + x'_{tk}\beta + u_{1t})$. The random effect u_{1t}, with $E(\exp(u_{1t})) = 1$ and $\mathrm{var}(u_{1t}) = \tau_1^2$, is an unobserved heterogeneity term.

Each counted event has an associated duration. The durations are assumed to follow a Weibull distribution with a hazard that depends on covariates and an unobserved heterogeneity term. The notation requires care because the number of durations does not match the number of periods (instead it matches the sum of all the counts) and because spells may begin in one period and end in another. Set aside the heterogeneity and the period-matching problem for the moment and focus on a single duration that spans multiple periods.

Let $y > 0$ be the duration and z be a vector of exogenous covariates that are fixed at the beginning of the spell. The hazard rate h depends on elapsed time, on z, and on a period-specific heterogeneity term u_{2t}. Thus, h changes discontinuously due to u_{2t}, with jumps at the beginnings of the periods. Label these points in $[0, y]$, endpoints included, as $(y_0, y_1, \ldots, y_J)$, and let h_j be the hazard rate in period j before the jth jump (set h_J equal to the hazard on $[y_{J-1}, y_J]$). Define the "completing spell" indicator d_j to be 1 for $j = J$ and 0 for $j < J$. The final d_J is also 0 if duration

y is right-censored. Define

$$g_j(y_j, y_{j-1}, d_j, z) = \begin{cases} \frac{1-F_j(y_j)}{1-F_j(y_{j-1})} \text{ if } d_j = 0 \\ \frac{f_j(y_j)}{1-F_j(y_{j-1})} \text{ if } d_j = 1 \end{cases} \tag{2}$$

for $j = 1, ..., J$, where F_j and f_j represent the Weibull cumulative density function (cdf) and pdf, respectively, under hazard h_j. A single event's duration, then, has pdf

$$f(y, z) = \prod_{j=1}^{J} g_j(y_j, y_{j-1}, d_j, z) \tag{3}$$

(Lancaster, 1990, sec. 2.3). For the Weibull model, f_j and F_j in (2) take the form

$$f_j(y) = \frac{1}{y\sigma}\left(\frac{y}{\mu_j}\right)^{1/\sigma-1} \exp\left(-\left[\left(\frac{y}{\mu_j}\right)^{1/\sigma}\right]\right) \tag{4}$$

$$F_j(y) = 1 - \exp\left(-\left[\left(\frac{y}{\mu_j}\right)^{1/\sigma}\right]\right), \tag{5}$$

where σ and the rate parameter μ_j are positive. The shape parameter σ captures duration dependence. When $\sigma < 1$, the hazard increases over the duration and there is positive duration dependence. When $\sigma > 1$, there is negative duration dependence. The Weibull mean is $\sigma\Gamma(\sigma)\mu_j$.

Turn now to the period-matching problem. Consider the entire sample of $M \equiv \sum_{t=1}^{T}\sum_{k=1}^{K} n_{tk}$ whole durations. Let the mth duration have J_m terms after splitting at the period changes, resulting in a split sample $((y_{mj})_{j=1}^{J_m})_{m=1}^{M}$. For simplicity of notation, relabel these with the single index i, resulting in $(y_i)_{i=1}^{N}$, where $N \equiv \sum_{m=1}^{M} J_m$, taking care to preserve the order of the series for the sake of the arguments of g in (2). Similarly, relabel the associated completing spell indicators d_i, rate parameters μ_i, likelihood terms g_i, and explanatory variables z_i, $i = 1, \ldots, N$. To match the split durations to the periods associate each i with an index set I_t so that $\{i|i \in I_t\}$ are the duration indices pertaining to period t.

Now I specify the functional form of the covariates and heterogeneity. For $i \in I_t$, the rate parameter μ_i is modeled as

$$\mu_i = \exp(z_i'\delta + u_{2t}). \tag{6}$$

The random effect u_{2t}, with $E(e^{u_{2t}}) = 1$ and $\text{var}(u_{1t}) = \tau_2^2$, is an unobserved heterogeneity term common across all split durations within a period. The inclusion of u_{2t} results in a mixture model that generalizes the standard Weibull model.

Taken together, (u_{1t}, u_{2t}) represent unobserved period-specific heterogeneity. Let (u_{1t}, u_{2t}) be iid draws from a bivariate normal distribution with mean $(-\tau_1^2/2, -\tau_2^2/2)$, variance (τ_1^2, τ_2^2), and correlation $\rho \in [-1, 1]$. If ρ is positive, then durations are longer in periods with many arrivals (congestion). If ρ is negative, then durations are shorter in periods with many arrivals.

The joint pdf for the data in period t, $\{n_{tk}, y_i | i \in I_t\}$, conditional on $\{y_{i-1}, u_{1t}, u_{2t} | i \in I_t\}$, is then

$$f(\{n_{tk}, y_i\} | u_{1t}, u_{2t}) = f(n_{tk} | u_{1t}) \prod_{i \in I_t} g_i(y_i, y_{i-1}, d_i, z_i | y_{i-1}, u_{2t}), \tag{7}$$

where the dependence on the parameters is suppressed in the notation. In the expression above, the form of $f(n_{tk}|u_{1t})$, the Poisson pdf, is given by (1). The form of g_i, the contribution of split duration i to the joint likelihood, is given by (2).

Since (u_{1t}, u_{2t}) are not observed, one finds the unconditional joint pdf by taking expectation over (u_{1t}, u_{2t}):

$$f(\{n_{tk}, y_i\}) = E_{u_{1t}, u_{2t}} f(\{n_{tk}, y_i\} | u_{1t}, u_{2t}). \tag{8}$$

The log likelihood for the sample is

$$\ln L(\beta, \delta, \sigma, \rho, \tau_1, \tau_2) = \sum_{t=1}^{T} \log f(\{n_{tk}, y_i\}) \tag{9}$$

The double integral in (8) requires numerical methods; I used Gauss-Hermite quadrature in the application.

The general model (9) contains several familiar models as special cases. When $\rho = 0$, the count and duration 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 $\tau_2 = 0$, the duration model is the standard homogeneous Weibull model.[2] Estimating these

[2] When, in addition, $\sigma = 1$, the duration model is exponential.

restricted models provides starting values for estimation of the full model.

4 Application to Telecommunications Regulation

In 1991, the U.S. Federal Communications Commission (FCC) switched from traditional RoRR of the major LECs to price caps. Since RoRR limits the ability of the firm to retain as profit the economic benefit created by the service, RoRR decreases the incentive to introduce a new product. Price caps place no direct restrictions on profit. Furthermore, under the FCC's price cap plan, a new service is allowed to be freely priced for a year before it is added to the price cap index, which may allow the firm to appropriate more of the surplus from a new product. Thus, a price caps regime may offer greater incentives to innovate than RoRR does. Whether price caps increase innovation in any particular setting is ultimately an empirical issue.

The data comprise the interstate access services introduced by Ameritech, one of the Bell operating companies, from July 1984 through March 1999.[3] The new services are summarized in Table 1. The count variable n_{tk} is the number of new services in a category in a calendar quarter. The categories (the k dimension of the model) are Database services, which enable services like toll-free calling; Digital Data services, which allow long-distance data transmission; High Capacity access services, which enable high-speed transmission of calls, and Other. I calculate regulatory delay—the y variable—as the time from submission of the tariff revision to approval. Regulatory approval times are broken out by regulatory classification (switched access, special access, and other) in Table 2.[4] Under price caps, Ameritech created more services than under RoRR. Regulatory delay was shorter.

Table 3 has the results from three estimations. In the first two, ρ is fixed at zero, so the count and duration models are independent. In the first estimation, only fixed effects are included.

[3] Data were culled from the FCC tariff filings. See Prieger (1999) for a detailed description of the data.

[4] Since switched access services face stricter regulation than special access services, regulatory classification provides a more germane set of indicator variables for the duration estimation than technological classification does.

The indicators for the price cap period in the innovation model are positive, reflecting that there was more innovation under price caps than under RoRR. The coefficients are significant for all categories except digital data. In the duration model, the price cap indicator is negative (but not significant), reflecting that regulatory delay was shorter under price caps.

In the second estimation, covariates are added to control for demand and supply factors. In the innovation model, the added variables are log real per capita income in Ameritech's territory, one-year-lagged log number of industry patents in the relevant technological areas,[5] and lagged log real R&D expenditure at the firm and industry level. In the regulatory delay model, I add the number of rate elements in the tariff filing, and the number of pages composing the filing. These proxy the unobservable complexity of the filing, which may affect regulatory delay. When these covariates are added, only the price cap indicator for high capacity services in the innovation model retains significance.

In the third estimation, correlation is allowed between the counts and the durations. The estimated price cap effects in the innovation model increase somewhat. The correlation parameter ρ is estimated to be negative and significant. This negative correlation suggests that a shock leading to more innovation, and therefore more tariffs submitted to the regulator, has the effect of reducing regulatory delay time. One possible interpretation: in periods of above-average innovation when the regulatory "inbox" begins to stack up, the FCC relaxes its scrutiny of new services and approves them quicker. Another interpretation is that the firm timed its innovation to match periods of rapid regulatory approval.

5 Conclusion

The results show that of the apparent increases in new service creation under the price cap regime, after controlling for demand factors (income) and supply factors (R&D and patents), only the effect

[5] All patents issued in classes 359, 370, 379, and 395 are included.

for high capacity services remains significant. The estimated coefficient implies a 301% change in new high capacity services from switching to price caps. Apparently the regulatory regime had a large impact on at least high capacity services—some of the most important access services. There are potential problems when moving from correlation to causality, however, that are beyond the scope of this letter.[6] Apart from the specific numerical results obtained, the main contribution of this work is methodological. The econometric model should prove useful for exploring other regulatory issues with other data.

[6] See Sappington and Weisman (1996) for a discussion of potential pitfall when moving from positive analysis to normative prescription in regulatory studies.

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	RoRR Period		Price Cap Period		
Service Category	Total	Yearly Rate	Total	Yearly Rate	**Total**
Database	4	0.62	13	1.57	17
Digital Data	3	0.46	4	0.48	7
High Capacity	5	0.77	31	3.75	36
Other	15	2.14	33	4	48
Total	27	4.15	81	9.79	108

Table notes RoRR is Rate of Return Regulation. Counts are new services in Ameritech's FCC Tariff no. 2.

Table 1: New Access Services Introduced by Ameritech

	Minimum	Mean	Median	Maximum
RoRR Period	1	110.33	59	771
Price Cap Period	5	53.79	45	349
Total	1	67.54	45	771

Table notes RoRR is Rate of Return Regulation. Durations are in days.

Table 2: Regulatory Approval Times of Ameritech's Proposed Services

	Independent Count and Duration Models				Joint Count and Duration Model	
	fixed effects only		*covariates*			
	coef.	s.e.	coef.	s.e.	coef.	s.e.
Innovation Count Model						
constant	0.84	(0.26)***	1.04	(0.36)***	0.92	(0.38)**
database	–1.32	(0.56)**	–1.32	(0.57)**	–1.32	(0.61)**
digital data	–1.61	(0.69)**	–1.61	(0.65)**	–1.61	(0.72)**
high capacity	–1.10	(0.49)**	–1.10	(0.53)**	–1.10	(0.52)**
price cap:database	1.15	(0.54)**	0.74	(0.68)	0.96	(0.76)
price cap:digital data	0.05	(0.78)	–0.36	(0.83)	–0.14	(0.87)
price cap:high capacity	1.59	(0.46)***	1.18	(0.62)*	1.39	(0.69)**
price cap:other	0.55	(0.31)*	0.14	(0.54)	0.36	(0.59)
income			0.43	(5.72)	2.08	(6.80)
industry patents$_{t-4}$			0.29	(0.37)	0.13	(0.44)
firm R&D$_{t-4}$			0.36	(0.27)	0.46	(0.35)
industry R&D$_{t-4}$			–0.11	(0.30)	–0.03	(0.33)
Regulatory Delay Duration Model						
constant	4.79	(0.30)***	4.91	(0.29)***	4.96	(0.30)***
switched access	0.35	(0.16)**	0.14	(0.19)	0.13	(0.19)
special access	–0.10	(0.16)	–0.24	(0.17)	–0.25	(0.17)
price cap	–0.40	(0.30)	–0.42	(0.27)	–0.45	(0.28)
rate elements			0.17	(0.07)**	0.17	(0.07)**
tariff pages			–0.02	(0.06)	–0.03	(0.05)
Incidental Parameters						
ρ	0.00	fixed	0.00	fixed	–0.98	(0.00)***
σ	0.52	(0.05)***	0.52	(0.05)***	0.67	(0.11)***
τ_1	3.9E-06	(0.24)	1.8E-05	(0.21)	0.13	(0.14)
τ_2	0.77	(0.12)***	0.68	(0.11)***	0.67	(0.11)***
log likelihood	–723.22		–717.38		–716.96	

* = 10% level significance; ** = 5% level significance; *** = 1% level significance; stars for σ, τ_1, and τ_2 based on one-sided tests.

Table notes Estimations based on 236 count observation (n_{tk}) and 196 duration observation (y_i), as described in the text. Periods t are quarters. Durations are in days. All continuous covariates are in logs. Dollar-denominated variables are adjusted by GDP deflator. Income data are from BEA *REIS*; patent data are from the USPTO; firm R&D data are from the FCC *Statistics of Communications Common Carriers*; industry R&D data are from the NSF *R&D in Industry*.

Table 3: Estimation Results