# REGULATION, INNOVATION, AND THE INTRODUCTION OF NEW TELECOMMUNICATIONS SERVICES*

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**Abstract**

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I examine the effects of FCC regulation on the innovation and introduction of advanced telecommunications services in the U.S. An interim of lighter regulation provides an "experiment" to test the regulatory regime's impact. The econometric model comprises an arrival process (for service innovation) followed by a duration process (for regulatory delay). The number of services the firms created during the interim is 60–99% higher than the model predicts they would have created if the stricter regulation had still been in place. Overall, firms would have introduced 62% more services to consumers during the study period if the regulation had not been in place.

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# 1 Introduction

Innovation and new goods are important forces driving economic growth and increasing the welfare of consumers. In regulated industries such as telecommunications, however, firms attempting to introduce new services often must contend with restrictions designed more to protect the status quo than to promote innovation. Despite the prevalence of such restrictions, little effort has been devoted to measuring the effects of regulation on product innovation. Since Joskow and Rose (1989) noted that such studies were virtually non-existent, the dearth of empirical work on regulation and product innovation has largely remained. This study examines the effect of a particular regulatory regime on the creation and introduction of new telecommunications services.

The literature looking at the effects of regulation on innovation focuses almost exclusively on cost reduction, or *process innovation* (Taylor, Zarkadas and Zona, 1992; Greenstein, McMaster and Spiller, 1995; Ai and Sappington, 1998). Unlike these studies, which look at the means of providing a service, I look directly at creating new goods—*product innovation*. My contribution is twofold. First, I quantify the effects of regulation in a real-world regime, uncovering the extent to which regulation reduces innovation and delays introduction of new services. Second, I develop an econometric model suitable for studying the introduction of regulated goods. The model comprises an arrival process for innovation followed by a correlated duration process for the regulatory delay that drives a wedge between innovation and introduction to consumers.

The data for the study are information services (such as fax, voice mail, and audiotex information services) introduced by the Bell Operating Companies[1] (BOCs) and AT&T. These services require access to the local telephone network to function. For example, voice mail service for a household must intercept and reroute incoming calls. Because these carriers were the sole suppliers of the necessary network elements, the Federal Communications Commission (FCC) regulated information services. The approval process for a new service begins when the carrier submits a plan to the FCC offering competitors "comparably efficient interconnection" (CEI) to the network.

The CEI plan theoretically allows rivals to offer similar services, although the BOCs complain that the process merely raises the cost of introducing a service. For an interim around 1992–1995, such detailed plans and time-consuming approvals were not required. During that time, the FCC allowed the carriers to freely introduce new enhanced services. Section 2 describes the CEI regime in more detail.

Carriers likely found it more attractive to introduce new services in the interim of lighter regulation (hereafter the "interim"). The CEI requirements reduced the expected profit of introducing a new service for at least three reasons. First, there are substantial direct costs of preparing a CEI plan, which require great amounts of technical and legal staff time. One BOC claims that the CEI plan requirement is "the most significant" regulatory burden imposed on BOCs' enhanced service operations.[2] Second, the plan reveals information to potential competitors before the service is introduced, allowing competitors to preempt the BOCs.[3] Third, there are long delays associated with plan approval—over 200 days on average, a long time in the rapidly evolving telecommunications industry. Once proposed, the plans typically went through several rounds of public comment and rebuttals, and in six cases, the FCC requested changes to the plans. The delay reduces the present value of a proposed new service and allows competitors to beat the regulated firm to market. The econometric model for innovation and introduction is presented in section 3.

To see if the release from regulation spurred more innovation, I estimate a prediction interval for the interim using data from the regulated periods (section 4.1). The actual number of new services created during the interim lies outside the "regulated-conditions" prediction interval, evidence that the behavior of the firm changed when released from the regulation. This approach requires that the observed count from the interim have a small probability of being generated under the regulated conditions—a test of structural change.

The analysis in section 4.1 bears out that some otherwise profitable services are not financially viable under the CEI regime. The number of services the firms created during the interim is 60–99% higher than the model predicts they would have created if the stricter regulation had still been in place. In section 4.2, I look at the magnitude and determinants of product introduction delay caused by the regulatory review process.

Using the estimated model to simulate the innovation and introduction process with and without the CEI regulation and regulatory delay, I find that firms would have introduced 62% more services to consumers during the study period if the regulation had not been in place (section 5). This projection exercise requires stronger assumptions than does the prediction interval exercise in section 4.1, because it assumes a parametric form for innovation during the unregulated interim. The final section, 6, discusses some qualifications and alternative explanations for the increased innovation.

## 2 The Comparably Efficient Interconnection Regime

Before 1986, regulated telecommunications carriers wishing to offer enhanced services were required to form separate subsidiaries (so-called "structural separation"). Through the *Computer III* series of orders (1986–1989), the FCC allowed integrated provision of enhanced services.[4] *Computer III* established a long-term goal, Open Network Architecture (ONA), and a short-term plan, CEI, to open the dominant firms' networks to competitors. To gain the FCC's approval to offer enhanced services, a carrier had to fulfill two requirements. First, for each service it had to develop a CEI plan to abide by certain safeguards. The CEI plan allowed competitors to interconnect to the incumbent's network to offer a similar service. Each proposed service necessitated a new CEI plan. Second, the carrier had to develop a longer range ONA plan to open up the rest of their network by offering all of the individual "building blocks" of the network to all competitors. The FCC stated that after it approved a carrier's ONA plan, the requirement of structural separations would be lifted altogether, and individual CEI plans would no longer be required.

The BOCs began to submit CEI plans to the FCC in 1987. In June 1990, a court case (*California I*) required the FCC to disallow development and introduction of new enhanced services (although carriers were allowed to continue to file for CEI waivers). The FCC resumed the CEI regime in February 1992.[5] In 1992 and 1993, the BOCs individually received approval of their ONA plans and the freedom to offer enhanced services without filing CEI plans.

In 1994, another court case (*California III*) forced the FCC to reinstate the CEI plan requirements. The FCC required the BOCs to file retroactive CEI plans for services introduced after the lifting of structural separation. From the retroactive CEI filings in January 1995, one can thus enumerate the new services during the unregulated interim. Thereafter, the CEI plan requirement remained in effect for the BOCs until 1999. In February 1999, the FCC discontinued pre-approval and required only that the BOCs post CEI plans on the Internet. To avoid contamination of the data by anticipation effects, I use data only through 1997 for the innovation estimations.

The time line in Figure 1 shows when enhanced services were being developed and the activity is observable through CEI filings (the black areas), when services were being created but the activity is not observable in the public record until a later date (the gray areas), and when development and introduction are disallowed (the white areas). The extent to which the firms anticipated these regime changes affects the interpretation of the data; I explore this issue in section 6.

## 3 A Model for Regulated 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; I refer to this as *innovation*. After innovation, the services are not *introduced* to subscribers until the regulator approves them. Modeling this progression requires a statistical framework that incorporates both steps; I use an arrival process followed by a duration process. The arrival process generates the new services, and the duration process determines the delay between innovation and introduction to consumers.

The model is a generalized Poisson event process followed by a generalized Weibull duration model. In particular, arrivals (innovation) during a period follow a heterogeneous Poisson distribution, which generates count data. After innovation, a new service immediately enters a heterogeneous Weibull model that determines the length of time until regulatory approval, generating duration data. The hazard rate of the Weibull process is determined from time-varying covariates. The conditional means of both distributions are modeled as exponential functions of economic, demographic, and regulatory explanatory panel data. A bivariate unobserved heterogeneity term admits correlation between the counts and the durations.

The likelihood for the data is first found conditional on the heterogeneity terms. The unobserved heterogeneity terms may then be integrated out, resulting in the unconditional likelihood. The model nests several simpler models, including independent standard Poisson and Weibull models, thus allowing specification testing. The model is formally equivalent to an infinite-server queuing model. The equivalence of the model to a queuing system allows adaptation of a result from the queuing literature for the projection exercise carried out in Section 5.

To construct the likelihood, the marginal distributions and the nature of the correlation between the count and duration processes must be specified. Let the number of arrivals in period $t$ from firm $k$ and of type $\ell$ be $n_{tk\ell}$, a realization of a count (non-negative integer) valued random variable. In the data, $k = 1, \ldots, 8$ represents AT&T and the seven Bell Operating Companies, and $\ell = 1, 2, 3$ represents the type of CEI filing (plans, amendments, and waivers; these are described in section 4.1.1). The length of period $t$, $s_t > 0$, is usually 0.25 (one quarter year) but may be less if the quarter is incompletely observed. Denote by $f(n_{tk\ell}|u_{1t})$ the probability density function (pdf) of $n_{tk\ell}$, conditional on firm and type fixed effects $\alpha_{1k}$ and $\alpha_{2\ell}$, covariates $\mathbf{x}_t$, coefficient vector $\boldsymbol{\beta}$, and a random effect $u_{1t}$ (the notation suppresses explicit dependence on $\alpha_{1k}$, $\alpha_{2\ell}$, $\mathbf{x}_t$, and $\boldsymbol{\beta}$ for simplicity). The count model is the Poisson with rate and pdf

$$\lambda_{tk\ell} = \exp\left(\alpha_{1k} + \alpha_{2\ell} + \mathbf{x}_t'\boldsymbol{\beta} + u_{1t}\right) \equiv \lambda^0_{tk\ell} \exp(u_{1t}) \tag{1}$$

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

respectively. Functions of current period durations are not allowed in $\mathbf{x}_t$, although functions of past durations are allowed.

The random effect $u_{1t}$, for simplicity taken to be common across firms and types within a period, is an unobserved heterogeneity term with variance $\tau_1^2 \geq 0$. Assume that $E(e^{u_{1t}}) = 1$, so that $E(n_{tk\ell}) = s_t \lambda^0_{tk\ell} = s_t \exp(\alpha_{1k} + \alpha_{2\ell} + \mathbf{x}_t'\boldsymbol{\beta})$. The inclusion of $u_{1t}$ results in a generalized Poisson model that relaxes the equality of the mean and the variance implied by the simple Poisson model and allows overdispersion (Cameron and Trivedi, 1998). The random effect also allows correlation with the duration part of the model, as explained below.

Associated with each counted event (innovation) is a duration (regulatory delay). The durations are assumed to follow a Weibull distribution with a hazard that depends on time-varying covariates and an unobserved heterogeneity term. The details become cumbersome because the number of durations does not match the number of periods (instead it matches the sum of all the counts), the durations may begin in one period and end in another, the hazard rate may change each period, and the observations may be censored. The details of this part of the model are available in the working paper (Prieger, 2000). Here I merely describe the characteristics of the Weibull distribution and specify the rate parameter.

The rate parameter $\mu_{it}$ and hazard rate $h_{it}$ for an ongoing duration in period $t$ (which is part of the duration of the $i$th arrival) is modeled as

$$\mu_{it} = \exp(\gamma_{1k} + \gamma_{2\ell} + \mathbf{z}_{it}'\boldsymbol{\delta} + u_{2t}) \equiv \mu^0_{it} \exp(u_{2t}) \tag{2}$$

$$h_{it}(y_i|u_{2t}) = (\sigma y_i)^{-1} (\mu_{it} y_i)^{1/\sigma}$$

respectively, where $\gamma_{1k}$ and $\gamma_{2\ell}$ are firm and filing type fixed effects as in the count model, and $\sigma > 0$ is a shape parameter. Spell-specific (i.e., $i$-specific) attributes may be included in $\mathbf{z}_{it}$, as may be functions of current-period counts $n_{tk\ell}$.[6] Parallel to the count model above, the random effect $u_{2t}$ is common across firms, types, and all ongoing durations within a period; it is an unobserved heterogeneity term with variance $\tau_2^2 \geq 0$. Assume that $E(e^{u_{2t}}) = 1$. The inclusion of $u_{2t}$ results in a mixture model that generalizes the standard Weibull model and allows correlation with the duration part of the model.

The Weibull distribution allows *duration dependence*—a nonautonomous hazard rate—through $\sigma$. Values of $\sigma$ less than one (underdispersion) yield a hazard rate that increases over the duration; this is known as *positive duration dependence* (the likelihood of the spell ending in the next instant increases as time goes on). Values of $\sigma$ greater than one (overdispersion) yield a decreasing hazard rate and *negative duration dependence.* The Weibull mean is $\sigma\Gamma(\sigma)\mu$ when $\mu$ is constant.

Taken together, $(u_{1t}, u_{2t})$ represent unobserved period-specific heterogeneity. To complete the model and to specify the correlation between the counts and durations, let $(u_{1t}, u_{2t})$ be defined by

$$\begin{aligned} u_{1t} &= \tau_1 (\eta_{1t} - \tau_1/2) \\ u_{2t} &= \tau_2 \left(\rho\eta_{1t} + \sqrt{1-\rho^2}\eta_{2t} - \tau_2/2\right) \end{aligned} \tag{3}$$

where $\eta_{1t}$ and $\eta_{2t}$ are iid draws from the standard normal distribution, $\tau_1,\ \tau_2 \geq 0$, and $\rho \in [-1, 1]$. Then $(u_{1t}, u_{2t})$ are bivariate normal with variance $(\tau_1^2, \tau_2^2)$ and correlation $\rho$. Thus, $\rho$ is the key parameter governing correlation between the count and regulatory delay processes. If $\rho$ 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 (beyond that which may be captured by included covariates). If $\rho$ 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 non-zero mean of $(u_{1t}, u_{2t})$ ensures that $E(e^{u_{1t}}) = E(e^{u_{2t}}) = 1$, as required above.

The working paper shows that the joint pdf for the count and duration data in period $t$ may now be found, first conditional on $(u_{1t}, u_{2t})$ and then unconditionally. The latter likelihood, required for MLE, involves a bivariate expectation—a double integral that cannot be solved analytically, requiring numerical methods.[7]

The model is about the simplest one that allows correlation between the innovation and regulatory delay processes. The model 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. Estimation of the independent models is numerically easier than of the joint model because the likelihood involves only unidimensional integrals. 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. When, in addition, $\sigma = 1$, the duration model is exponential. Estimating these restricted models is useful for hypothesis testing and to provide starting values for estimation of the full model.

# 4 Data and Estimation of the Model

Since the innovation and regulatory delay models are independent if $\rho = 0$, I first discuss each separately in sections 4.1 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 collected data on all CEI plan filings and waiver requests from 1987 through 1997 from the *FCC Record* (see Table 1 for a partial list). The CEI plans, plan amendments, and waivers provide a complete record of new enhanced services that AT&T and the BOCs wished to introduce on an unseparated basis during that time. Plans are the firm's proposal to meet the CEI safeguards for a new information service. Amendments are modifications to existing plans that are required when significant new functionality is added to a service. Waivers are requests for permission to offer a new service without meeting all of the required CEI safeguards, usually because of technological limitations. During the interim around 1993 to 1995, when CEI plans were not required, there is no way to track the introduction date of new services. However, after the *Computer III* remand, CEI plans for all services introduced during the interim were filed *en masse*, so one can count new services in retrospect. This data set therefore encompasses all enhanced services offered by AT&T and the BOCs.

The new services are summarized in Table 2.[8] In all, 74 new enhanced services were introduced via CEI plans or waivers, and 27 were introduced during the interim, for a total of 101 services during the observation period.

To apply the methodology developed in Section 3, I first create the period-firm-type counts $n_{tk\ell}$ (784 observations). The periods are the quarters between the start of 1987 and the end of 1997 that at least partially overlap with the "innovation observed" part of the timeline in Figure 1. The CEI filing types are plans, amendments, and waivers. Fixed effects are estimated by two sets of dummy variables for carriers and filing types. Because the observations come from two disjoint periods (before and after the interim), I include a dummy variable for the remand period (1995-1997).

### 4.1.2 Estimation and Prediction

The model is estimated with the service innovation data from when the CEI regime was in place. The results of the simple Poisson model (i.e., $\tau_1 = 0$) estimation with only the firm, type, and remand fixed effects are in Table 3, column one. Positive coefficients increase the number of arrivals. The indicator for the remand period is not significant, which indicates that the pace of innovation was comparable in the two regulated periods.

The fitted model characterizes the BOCs' new service creation process under regulation. If the FCC never released the BOCs from the CEI filing requirements, how many new services would the BOCs have submitted during the interim? If in fact the reduced burden of regulation during the interim accelerated the introduction of new services, then we should expect that the model's prediction will be significantly lower than the actual number of services introduced. This is indeed the case.

The predictor is formed from the predicted yearly mean of the arrival process, $\hat{\lambda}^0_{tk\ell}$ from (1).

To calculate the predicted number of new services from the BOCs during the interim, calculate $\hat{\lambda}^0_{tk\ell}$ using the estimated coefficients, multiply by the length of the interim for firm $k$, and sum over $k$ and $\ell$.[9] In the homogeneous Poisson model without time-varying regressors,[10] such a predictor reproduces the means in the raw data. The advantage of the model over the raw data is that the prediction variance can be estimated.

The resulting prediction, prediction interval, and $p$-value for the actual number of new services is in the first row of Table 4.[11] The prediction is the expected number of new services during the interim under the counter-factual assumption that the CEI regime was still in place. The interim was between 1.5 and 2.6 years long, differing for each firm because the FCC released them from the CEI requirements on different dates. If the FCC had not lifted the structural separations requirement, we would expect to have seen 16.3 new plans and amendments introduced by the BOCs during the interim. In fact, we saw 26, about 60% more than expected and outside the 95% prediction interval. The $p$-value of 26 services is 0.02; thus, it is highly likely that the incentives to innovate did change during the interim. The predictors from the other models in the next section bolster this conclusion.

#### 4.1.3 Specification Checking

Before placing weight on the conclusions from the prediction exercise, I test the assumptions underlying the *fixed effects* model. Note that because the interim occurs in the middle of my observation period, the model already avoids a common pitfall of applied regulation studies. When a deregulated period generates the second half of the observations, such "before-and-after" studies typically cannot separate the regulatory effects from secular trends in innovation (Sappington and Weisman, 1996).

Consider the possibility that there is unobserved heterogeneity (i.e., that $\tau_1 > 0$). In the second column of Table 3, $\tau_1$ is a free parameter. The coefficient estimates do not change much, but $\tau_1$ differs significantly from zero (as it does in the following two generalized Poisson estimations).[12]

Given this evidence of heterogeneity, I re-do the prediction exercise with the heterogeneous fixed effect model to make sure that the prediction interval still excludes 26 services (the actual number introduced) after incorporating the additional variance. The resulting prediction, prediction interval, and $p$-value is in the second row of Table 4. Twenty-six services still has a low $p$-value (0.02), providing corroboration for the earlier evidence that innovation changed during the interim.

Next, consider the possibility that the mean is incorrectly specified through the omission of relevant variables. The first variable to check is expected regulatory delay time. As noted in the introduction, the long delays between innovation and introduction may be one reason that the regulation hampered innovation. An expected regulatory delay variable was constructed based on recent past experience with regulatory delay.[13] Expected regulatory delay is added to the next two estimations: the *congestion* and *economic variables* estimations in Table 3. In both estimations, longer expected delay leads to less innovation. The *congestion* model implies that if expected delay were to rise by 152 days above the sample average, one fewer service would be added to the total yearly innovation (all firms and plan types).[14] The same calculation based on the *economic variables* model shows that there would be one fewer service created if expected delay rose by only 83 days. Conversely, if regulatory delay disappeared altogether, there would be from three (*congestion*) to six (*economic variables*) new services per year. The congestion effect is statistically significant only in the second estimation, so the conclusions here remain tentative.

The next variables to check are economic and demographic variables that might affect the demand for new services. The *economic variables* estimation in Table 3 includes measures of consumer income (log per-capita real income in the BOC's territory) and BOC market size (log number of access lines) (all added variables are demeaned).[15] These variables are not significant,[16] although access lines does have the expected sign: more lines increase new services.

Next, I check controls for the changing size of the population of potential new services. I include the log of contemporaneous real R&D spending by the carriers and by the whole U.S. telecommunications industry. The estimated coefficients in the *economic variables* model are statistically insignificant.[17] The coefficient on firm R&D, 0.275, implies that (at sample averages) about $50M above average R&D spending would generate a new service.

Competition to the BOCs might also influence service creation, although the effect could go either way.[18] The first control for competition is the number of competitive access providers (CAPs) in the firm's territory. CAPs offer access services to high-volume customers in urban areas, and may also offer information services. The other control is the contribution of the communications industry (net of the BOCs' contribution) to real gross state product within the BOC's territory.[19] This variable captures a broad range of competitive activity that the BOCs face. Neither coefficient is significant.[20]

A linear time trend, to capture any other omitted variable such as competitive pressure or the expected profitability of introducing new services, is also insignificant. Finally, a test of the joint hypothesis that all of the coefficients in the *economic variables* estimation that are not in the *congestion* estimation are zero fails to reject.[21]

Viewed as a whole, the evidence strongly indicates that omitted variables are not a problem for the *fixed effects* specification. Nevertheless, I re-do the prediction exercise based on the *congestion* and *economic variables* models; the results are in rows three and four of Table 4. The predictors are calculated using the actual values of the covariates during the interim period (and therefore do not merely reproduce the sample mean of the dependent variable). Thus, if it is the covariates alone that cause more innovation during the interim, it would be reflected in the prediction and prediction interval. In fact, the covariates result in a smaller prediction than before. The resulting $p$-values for 26 services are even lower than for the previous predictions: further evidence that innovation changed during the interim.

One question remains: how sensitive are the predictions to the chosen parametric forms for the hazard? Here I re-do the prediction exercise using coefficient estimates from a Cox (1975) semiparametric proportional hazards model, using the *fixed effects* regressors. The dependent variable in this estimation is the interarrival times of innovation events. Flexibly modeling the interarrival process leads to a count distribution that can exhibit under- or overdispersion of any form. The coefficients are reported in the final column of Table 3, and are close to the Poisson model. The 26 services actually introduced by the BOCs still lie outside of the 95% prediction interval from this model (final row of Table 4).[22] The results from the Cox model corroborate the other predictions: there is a very small chance that innovation did not change during the interim.

## 4.2 The Regulatory Delay Model

### 4.2.1 Data

Turn now to the regulatory delay submodel. The data set also contains the time to approval or withdrawal for each CEI filing. Most observations end with approval by the FCC, a few with withdrawal by the carrier.[23] The approval delays were sizable: of the 68 spells ending in final approval, the average was 204 days and the longest (AT&T SPECS waiver) was 22 months.[24] The longest ongoing delays (as of December 1997) are two CEI plans (Fast Packet and Internet Access) Ameritech first submitted in March 1995. Splitting the durations into sub-spells, to match with time-varying covariates, results in sample size of 245.

Explanatory variables include firm and plan fixed effects, whether a plan is a "me-too" filing,[25] and whether the plan was refiled. One expects that delays will be longer if the CEI plan raises complex issues. I proxy plan complexity by counting the number of pages in the *FCC Record* reporting the approval, because the report includes a discussion of the issues. The number of competitors that are affected by the plan may also affect the time to approval; I count the number of interested parties submitting comments for the public record.

### 4.2.2 Estimation and Specification Checking

The results for the *fixed effects* Weibull model ($\tau_2 = 0$) are in Table 5, column one. The parameter $\sigma$ is estimated to be 0.613 and differs significantly from unity, rejecting the exponential model.

The data exhibit positive duration dependence, implying that the longer the delay lasts, the more likely it is to end the next day (conditional on lasting as long as it has). The positive duration dependence may have a structural origin. The FCC has a legal obligation to act on filings in a timely matter, and the pressure may mount to approve filings if they have been delayed a long time.

To interpret the coefficients, recall from (2) that positive coefficients increase mean delay.[26] The estimates indicate that me-too status garners 41% shorter delays. Amendments are approved 9% quicker and waivers 16% slower than plans. Refiling a plan not surprisingly increases delay, by 48%. The average predicted mean duration is 226 days.

The heterogeneity parameter $\tau_2$ is allowed to vary in the second column of Table 5. The results and likelihood from the generalized Weibull model differ little from the standard Weibull model in column one.

I also consider whether there is congestion in the regulatory approval process for CEI filings. Does approval slow down when there are many plans awaiting regulatory approval? It is of interest to separate the impact of plans already sitting on the regulator's desk when a new plan is filed, versus the impact of plans filed afterward.[27] I add two congestion variables: the number of services (from any firm) awaiting approval at the start of the spell and the number of new arrivals (from any firm) during the length of the spell.[28]

The results of the *congestion* estimation are in Table 5, column three. Both congestion coefficients are negative, a surprising result, although only the number of new filings during the spell is significant. Neither effect is large, however. An extra filing above average is estimated to decrease the mean approval duration by less than a week. An interpretation may be as follows. When services are stacked up in the regulatory inbox, the FCC relaxes its scrutiny of new plans a bit and approves them marginally quicker. Holding that effect constant, any filings arriving afterward speed up approval of a plan already in the inbox, perhaps because regulatory resources need to be shifted to the new arrivals.

Turn finally to the other covariates (the *covariates* estimation in Table 5).[29] *FCC Record* pages appears to be a weak proxy for plan complexity; the coefficient is positive but not significant. An extra page in the *Record* above the average increases mean delay by 4 days. The number of commenters is marginally significant; an extra reply commenter above the average increases mean delay by only 1.2 days, indicating that opponents have limited power to slow down their rival's plan approvals.

### 4.3 The Jointly Correlated Model

The full model allows correlation between the count and duration models. In these estimations, unlike those reported in Tables 3 and 5, $\rho$ is allowed to vary. In general, the coefficients are close to those from the independent estimations and the results are not reported here (see (Prieger, 2000) for the estimation results).

In a *fixed effects* estimation, including all the variables from the *fixed effects* estimations in Tables 3 and 5, the correlation between the count and duration models (as measured by $\rho$) is negative. 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. This indirect congestion effect is similar to the direct congestion effects in the independent regulatory delay *congestion* estimation, where congestion covariates elicited this effect. Since these congestion covariates are omitted here, the direct effect shows up as negative unexplained correlation. Once the *congestion* variables (those from the *congestion* estimations in Tables 3 and 5) are included in the joint estimation, the residual correlation $\rho$ turns positive. Thus, net of the direct congestion effects in the model, there may be additional, unobserved factors that cause both increased innovation and longer delays.

Although $\rho$ is significant in both joint estimations, LR tests of each unrestricted estimation against its restricted counterpart fail to reject the restricted models. Given this, and given the added complexity that correlation adds to simulating the joint model, I rely on the independent estimations for the comparison exercise in the next section.

## 5 The Costs of the CEI Regime

The predictions in section 4.1 have already provided evidence that innovation during the interim was 60–99% higher than would have been the case had regulation still been in place. In this section, I include the impact of regulatory delay and ask a different question. Now that I have estimated the parameters for the entire model, I can compare the overall effect on service introductions (not just innovation) in the two periods. In Scenario 1, the interim period of lighter regulation never happens, and the CEI regime is in place for the entire observation period (1987–1997). In Scenario 2, the CEI regime never happens, and the interim conditions are in effect for the entire observation period. How many new services would one expect the carriers to introduce in each scenario? Comparing the two answers will measure the total effect that the regulatory regime had on innovation.

Scenario 1 employs the generalized Poisson *fixed effects* model for the arrival process (column two of Table 3) and the Weibull *fixed effects* model for the delay process (column one of Table 5). For Scenario 2, the projection is calculated by MLE of the simple Poisson fixed effects model using the interim data (the only model that is identified with the interim counts). The Scenario 2 mean is 12.75 services created per year. Regulatory delay is zero for scenario 2. Table 6 shows that in Scenario 1, the BOCs introduce a projected 92 services. At the end of the period, 5.7 services are still awaiting regulatory approval.[30] In Scenario 2, the BOCs introduce a projected 140 services. Thus in Scenario 2, the firms create 52% more services and are able to introduce 62% more services than in Scenario 1. The average is 53.6 more services approved without the CEI regulation, and the standard deviation is small enough to assure that the difference is positive with high probability.[31]

In the steady state, the relevant comparison is between the innovation rates. As the observation period in the projection exercise grows, the difference between innovation and introduction goes away (in percentage terms). This is because (absent changes in covariates) the number of services created is proportional to the length of the observation period, while the absolute difference between innovation and introduction (*services not approved* in Table 6) is constant. Nevertheless, looking at introductions is important. There is a wedge between innovation and introduction caused by regulatory delay. That wedge represents services that consumers could be enjoying but are not—an opportunity cost of regulation. For example, AT&T proposed to offer voice mail in the 1970's but was prevented by the FCC from doing so until after divestiture in 1984.

Exercises comparing counterfactual scenarios are only as good as the assumptions upon which they rest. Given the difficulties with such exercises, it is perhaps best to view this exercise as illustrative rather than to stress the numerical results *per se*. The results from the exercise illustrate the two ways that relaxed regulation benefits consumers of telecommunications services. First, carriers introduced many more new telecommunications information services when the CEI regime was suspended. The evidence is strong that the onerous requirements of the CEI regime prevented many services from being introduced. Second, suspending the CEI regime removed the long approval delays for CEI plans, meaning that each new service became available to consumers much sooner. Adding the CEI delay to the wait associated with tariffing any underlying basic services pushes the time to introduction out even farther. The innovation *congestion* estimation indicates that the removal of these long introduction delays induced the BOCs' to introduce more services.

## 6 Discussion and Conclusions

One would like to move beyond the mere counts and measure changes in consumer welfare. This is not possible with the present data, any more than it is with most of the literature on patent counts. However, it is possible to do a "back of the envelope" calculation for voice messaging, one of the CEI services introduced by all the BOCs. One recent study (Hausman, 1997) estimates the consumer welfare gained from the introduction of voice mail by local telephone companies, finding that the structural prohibition against enhanced services cost consumers about \$1.2 billion per year for voice messaging. Hausman's methodology is not undisputed[32] and I cannot extrapolate from this figure to other services (voice mail is one of the most successful CEI services). However, if the figure is valid, then the CEI regime cost consumers \$690 million to \$910 million in lost welfare from delayed voice mail availability, depending on which predicted regulatory approval delay in Table 5 is used.

Even if the BOCs would have introduced more services if the regime had not been in place, I do not claim that welfare would have risen proportionally. Services introduced during the interim were probably not of comparable importance to the services introduced under the CEI regime. The incremental services introduced during the interim—those that firms would not have introduced under the CEI regime—most likely created less revenue for the BOCs since they were formerly deemed unprofitable to introduce. Accordingly, they may have been worth less to consumers as well. There is no way to estimate this difference in worth with the present data.

The analysis implicitly assumes that the services introduced are truly new; the welfare impact of the regulation is blunted if they are not. For some of these services, consumers face substitutes offered by competitors. For example, Internet services offered by BOCs are included in these data, and consumers have many choices of Internet service providers. The case of Internet services is the exception rather than the rule, however. In general, the CEI services introduced by BOCs are not identical to any previously offered by competitors. Because the BOCs seamlessly integrate CEI services with their networks, and therefore with the subscriber's local telephone service, these services can offer features that competitors previously could not. It is very likely that even if some competitor already offered a related service, the BOC's offering was highly differentiated. In fact, one of the reasons the FCC instituted the CEI regime was because these services are unique, and because competitors feared the BOCs would therefore have an unfair advantage.

A final question concerns the nature of the experiment. The estimation would stumble into the *sequencing pitfall* (identified by Sappington and Weisman (1996)) if firms' actions are spurred by anticipated changes in the regulatory regime. This pitfall might apply to these data if, right before the ONA plans were approved, the BOCs held off on creating new services in anticipation of the upcoming reduced introduction costs. On the other end of the interim, the sequencing effect would also imply that fewer services would be introduced shortly after the regime was reinstated, if firms hurried services into the marketplace just before the regime change. Three facts provide evidence against this pitfall. First, some BOCs submitted CEI plans as shortly as a few weeks before and after the interim. Second, although there are only three new services during the period from *Computer III* reinstatement to the ONA plan approvals (see Figure 1), that number is well within the normal variation of the estimated innovation process.[33] Third, a dummy variable for the periods just before and after the interim in the innovation regression is not significant.

Against the demonstrated costs to consumers of the CEI regime, what was there to balance? If regulators maximize social welfare, then they must have believed that the benefits of the CEI regime to consumers were large indeed. However, it appears that when the FCC began the CEI regulation, it did not consider the opportunity cost of the regulation: it compared the CEI regime to disallowing enhanced services altogether, instead of to a regime with even less regulation. I emphasize that this study is not a cost-benefits analysis; without more data on competitors the benefits are impossible to evaluate. However, the benefits of the regime appear to be slight: few requests for interconnection were ever made under the CEI regime,[34] and competitors have shown little interest in the CEI plans in recent years.[35] The FCC intended the CEI regime to be a short-term solution until the stricter ONA regime was fully in place. It took twelve years for the burdens of the CEI regime to be significantly reduced—a caution to those who would set up "temporary" regulation.

Of the three innovation-discouraging effects mentioned in the introduction, the estimations are able to disentangle only the expected regulatory delay effect. Recall from section 4.1.3 that the estimated impact of reducing expected regulatory delay to nil is 3 to 6 new services created per year. Given that the projection exercise in section 5 estimated the innovation difference between the regulated and unregulated scenarios to be about 4.3 services per year, it appears that the expected delay effect accounts for much of the reduction in innovation under the CEI regime. One interesting extension to the present study would be to attempt to disentangle the other two effects. As of February 1999, the BOCs are relieved from pre-filing CEI plans, but are still required to make the plans public. Therefore the effect of information revelation still exists under the present rules, while the effect from the direct costs of preparing the plan is reduced[36] and the effect from approval delay is eliminated. It will be interesting to see if the rule change spurs the creation of new services.

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# NOTES

[1] The BOCs are Ameritech, Bell Atlantic, BellSouth, NYNEX, Pacific Bell, Southwestern Bell Telephone, and US West. Bell Atlantic merged with NYNEX in 1996 and Southwestern Bell merged with Pacific Bell in 1997; the firms continued to issue separate CEI plans.

[2] BellSouth *Comments* in the Computer III Further Remand Proceedings, CC Dkt No. 95-20, 1998.

[3] MCI contested Ameritech's CEI plan for Personal Access Service while it added the functionality described in the plan to its competing service [FCC *Report and Order* (1999, footnote 66)].

[4] See Prieger (1999) for the legal citations for this section.

[5] Around this time, the FCC relieved AT&T from most CEI requirements.

[6] Strictly speaking, if $\rho$ in (3) is non-zero, then $n_{tk\ell}$ is endogenous. However, the usual likelihood expression assuming exogeneity is still a *partial likelihood* (Lancaster, 1990, sec. 9.2.11). Because estimates resulting from maximizing a partial likelihood behave like MLE, including consistency of the parameter and s.e. estimates (Lancaster, 1990, sec.9.2.2), I will not further emphasize this distinction.

[7] I use Gaussian quadrature, with 50 evaluation points in both dimensions.

[8] Using CEI filings as evidence of new service creation requires care. I deem an amended plan to represent a new service only if it introduced new features or functionality. Similarly, filings that reduce functionality are not included. Finally, if a carrier submitted two new plans because two previously distinct technologies had converged, only one of the new plans was counted.

[9] Only plans and amendments are used in the prediction exercise, both for the prediction and the actual number of services introduced. Waivers are observed during the interim and do not need to be predicted. AT&T is not included in the prediction because the firm's data ends after 1991.

[10] One change is made to the models in Table 3 when forming the predictions. Since the interim period is neither in the initial Computer III period nor in the subsequent remand period, I do not include the remand indicator in the estimations for the predictions. Given that the remand coefficient is small, this has little impact on the predictions.

[11] The prediction interval accounts for variation from two sources: estimation error and the intrinsic variation of the Poisson random variable. The bootstrap prediction interval (adapted from Stine (1985)) covers 95% of a sample of size 9,999 of $y^* + \epsilon_i$, where $y^*$ is the prediction and $\epsilon_i = y_i - y_i^*$ is the prediction error. The $y_i$ are pseudo-random deviates drawn from a Poisson distribution with mean $y^*$. The $y_i^*$ are bootstrapped predictions.

[12] $\hat{\tau}_1$ is half-normal under the hypothesis that $\tau_1 = 0$, and so the correct $p$-value is based on a one-sided critical value.

[13] The expected regulatory delay variable is constructed assuming the firms believe regulatory delay is exponentially distributed (with rate $\kappa$, which has a Gamma prior). The expectation is updated via Bayes' rule given observed completed and ongoing delays within the past year.

[14] The yearly mean of the arrival process is $\lambda^0_{tk\ell}$; the elasticity of the mean with respect to $x$ is $\beta x$. When $x = \log(w)$, then the elasticity of the mean with respect to $w$ is $\beta$.

[15] Since AT&T does not have a territory like the BOCs, AT&T's observations are dropped. I also drop the BOC fixed effects because they are highly correlated with the added covariates.

[16] A Wald test statistic for the two variables is at the 0.70 quantile of a $\chi^2(2)$ distribution. When population is added, it is also insignificant.

[17] See notes to Table 3 for data sources. A Wald test statistic with joint null hypothesis that each coefficient is 0 is at the 0.88 quantile of a $\chi^2(2)$ distribution.

[18] The *replacement effect* (Arrow, 1962) indicates that a monopolist that innovates merely replaces one stream of rents with another, so that the incremental profit from innovation is not as great

as for a competitive firm. The *efficiency effect* (Gilbert and Newbery, 1982) indicates that a new service is worth more to a monopolist than to an entrant who then gains only duopoly profit. Which effect predominates is an empirical matter.

[19] A BOC's contribution to GSP is estimated as (operating revenue – [operating expenses less depreciation]) using FCC *SOCC* data. This approximates BEA methodology for gross product.

[20] A Wald test statistic with joint null hypothesis that each coefficient is 0 is at the 0.31 quantile of a $\chi^2(2)$ distribution.

[21] A Wald test statistic is at the 0.51 quantile of a $\chi^2(7)$ distribution.

[22] The Cox model estimates the effect of regressors on an unspecified baseline hazard, which I next estimate via Breslow's (1974) nonparametric method. The prediction is the sum of (interim $\text{length}_k$/mean survival $\text{time}_{k\ell}$) for each firm-plan type. This predictor is the asymptotic mean of the compound counting process generated from the renewal processes associated with the estimated survival functions. The prediction may be biased upward due to censoring. The bootstrap prediction interval covers 95% of a sample of $y^* + \epsilon_i$, where $y^*$ is the prediction and $\epsilon_i = y_i - y_i^*$ is the prediction error. The $y_i$ are pseudo-random deviates (1,000,000 draws) drawn from a compound counting process with the interarrival times from the Cox/Breslow estimates. The $y_i^*$ were generated by drawing pseudo-random deviates from the estimated asymptotic distribution of the mean interarrival time and forming the corresponding counts.

[23] In some cases the FCC requires a carrier to modify and resubmit a plan. In such cases the approval delay is calculated from initial submission to final approval. Withdrawn plans are counted as censored observations.

[24] In the approval delay analysis I exclude the exceptional February 1995 *en masse* filing.

[25] A *me-too* filing is a CEI plan that is substantially similar to a previously approved plan filed by another carrier. The FCC rules expedite approval for such plans.

[26] When the regressors are in logs, the coefficients are elasticities.

[27] More physical measures of regulatory capacity—FCC total and CEI-related personnel—have little variation in the observation period and are not significant when added to the estimations.

[28] These are treated as time-varying covariates. Refer to note 3 concerning endogenous regressors.

[29] There is a sample selection issue with the *covariates* estimation. *FCC Record* pages and the number of commenters are available only for observations appearing in the *Record*, which excludes 14 withdrawn and unapproved plans. These observations may not be a random subsample. The impact of sample selection bias is not likely to be severe; none of the coefficients included in both the *congestion* and the *covariates* estimations changes sign.

[30] Finding the services introduced requires a result from queuing theory. Details of the calculations in this section are available upon request.

[31] The standard deviations in the table account for intrinsic (not estimation) variation.

[32] See comments by Pakes (1997); see also Bresnahan (1997) for comments on another application of similar econometric methodology.

[33] Three is the 0.13 quantile of a Poisson distribution with mean 6.3, the estimated mean for that time period from the *fixed effects* model; no reasonably sized test would reject the hypothesis that the mean of the arrival process was constant during that time.

[34] Vogelsang and Mitchell (1997), ch.6; GTE (1998) also reported to the FCC that the number of requests for interconnection was “modest.”

[35] FCC, *Computer III Further Remand Proceedings FNPRM* (January 29, 1998) at 64.

[36] Without the possibility of the FCC requiring changes to the CEI plan before the service is available, the firm’s expected costs of plan preparation drops.

| CEI Filing | Date Filed | Date Approved or Withdrawn |
|---|---|---|
| Interactive Audiotex Svcs | 3/23/89 | 9/21/89 |
| Enh Protocol Processing Svcs | 11/6/89 | 5/24/90 |
| Alarm Monitoring | 3/13/95 | 10/31/95 |
| Fast Packet Data Svcs | 3/13/95 | ongoing* |
| Fax Store & Forward | 3/13/95 | 10/31/95 |
| Interactive Voice Response Svcs | 3/13/95 | 10/31/95 |
| Internet Access Svcs | 3/13/95 | ongoing* |
| Voice Mail Messaging | 3/13/95 | 10/31/95 |
| Message Delivery Svc | 6/11/95 | 12/15/95 |
| Personal Access Svc (PAS) | 9/1/95 | 6/4/97 |
| Voice Mail Messaging (minor amendment) | 8/22/96 | 10/28/96 |
| Fax Store & Forward (amended) | 9/27/96 | 12/4/96 |
| Reverse Search (waiver) | 10/25/96 | 3/24/97 |
| Voice Mail Messaging (supplement to minor amendment) | 10/28/96 | 11/14/96 |
| Payphone Svc | 11/27/96 | 4/15/97 |
| Electronic Vaulting Svc | 2/27/97 | 12/31/97 |

*As of end of 1997.

Table note: contains the CEI plans, amended plans, and waiver requests for Ameritech as a sample of the types of services in the data set. For the full list of services from all the BOCs and AT&T, see Prieger (2000).

**Table 1: New Telecommunications Information Services Introduced Through CEI Plans, Amendments, and Waivers by Ameritech**

| Carrier | Initial CEI Regime (1987–1992/3) | Structural Separation Lifted (1992/3–Jan 1995) | Remand—Back to CEI Regime (1995–1997) | Total |
|---|---|---|---|---|
| Ameritech | 2 | 6 | 6 | 14 |
| Bell Atlantic | 11 | 1 | 4 | 16 |
| BellSouth | 7 | 2 | 2 | 11 |
| NYNEX | 3 | 2 | 3 | 8 |
| PacBell | 8 | 6 | 0 | 14 |
| SWBT | 4 | 3 | 8 | 15 |
| U S West | 4 | 7† | 1 | 12 |
| AT&T* | 11 | – | – | 11 |
| *Total* | *50* | *27* | *24*‡ | *101* |

* After 1991, the FCC treated enhanced services from AT&T differently than those from the BOCs. To preserve comparability, new services by AT&T after 1991 are not included in the data set.

† Includes a waiver filed during the interim.

‡ Excludes an FCC-mandated payphone plan for each BOC in 1997. The services were not new.

**Table 2: New Telecommunications Information Services Introduced Through CEI Plans, Amendments, and Waivers**

| | Poisson Model | Generalized Poisson Model | | | Cox Model |
|---|---|---|---|---|---|
| | *Fixed Effects* | *Fixed Effects* | *Congestion* | *Economic Variables* | *Fixed Effects* |
| intercept | 0.494 | 0.461 | 0.371 | −1.209** | |
| | (0.317) | (0.343) | (0.365) | (0.573) | |
| amendment | −1.056*** | −1.056*** | −1.056*** | −1.149*** | −1.072*** |
| | (0.318) | (0.305) | (0.313) | (0.317) | (0.314) |
| waiver | −1.664*** | −1.653*** | −1.626*** | −1.731*** | −1.729*** |
| | (0.339) | (0.340) | (0.351) | (0.380) | (0.367) |
| remand period | −0.067 | −0.061 | 0.072 | 2.607* | −0.529* |
| (1995-1997) | (0.270) | (0.299) | (0.361) | (1.358) | (0.285) |
| expected regulatory | | | −0.331 | −0.601* | |
| delay | | | (0.287) | (0.337) | |
| per cap income | | | | −0.407 | |
| | | | | (1.042) | |
| access lines | | | | 1.511 | |
| | | | | (1.155) | |
| firm R&D | | | | 0.275 | |
| | | | | (0.196) | |
| industry R&D | | | | −0.749 | |
| | | | | (0.617) | |
| CAPs | | | | −0.044 | |
| | | | | (0.205) | |
| communications GSP | | | | −0.640 | |
| | | | | (1.041) | |
| time trend | | | | −0.073 | |
| | | | | (0.046) | |
| Ameritech | −0.804* | −0.787* | −0.769 | | −0.714 |
| | (0.483) | (0.476) | (0.492) | | (0.507) |
| Bell Atlantic | −0.173 | −0.157 | −0.139 | | −0.262 |
| | (0.413) | (0.446) | (0.442) | | (0.432) |
| BellSouth | −0.812* | −0.784* | −0.752 | | −0.748 |
| | (0.461) | (0.475) | (0.481) | | (0.473) |
| NYNEX | −1.152** | −1.126* | −1.105** | | −1.166** |
| | (0.535) | (0.580) | (0.556) | | (0.561) |
| PacBell | −0.913* | −0.885* | −0.856* | | −0.980* |
| | (0.478) | (0.487) | (0.496) | | (0.507) |
| SWBT | −0.444 | −0.421 | −0.401 | | −0.287 |
| | (0.436) | (0.470) | (0.463) | | (0.449) |
| USWest | −1.089** | −1.073* | −1.055* | | −0.934* |
| | (0.536) | (0.577) | (0.552) | | (0.534) |
| $\tau_1$ (heterogeneity | 0.000 | 0.379** | 0.378** | 0.445** | |
| parameter) | (fixed) | (0.192) | (0.193) | (0.203) | |
| log likelihood | −236.26 | −235.51 | −230.91 | −199.63 | |
| observations | 784 | 784 | 784 | 726 | 137 |

* = 10% level significance; ** = 5% level significance; *** = 1% level significance; stars for $\tau_1$ based on one-sided test.
Table notes: Asy. s.e. in parentheses. Dependent variable is the number of new CEI filings of a given type per quarter except for the Cox model, for which it is the interarrival times of the filings. The log of the conditional mean of the dependent variable is linear in the covariates. Excluded carrier dummy is AT&T. All covariates are in logs, and adjusted by GDP deflator where appropriate. Income data are from BEA *REIS*, firm R&D data from the FCC *Statistics of Communications Common Carriers*, industry R&D data from the NSF *R&D in Industry*, Competitive Access Providers (CAPs) from NPRG (1997) and (Kraushaar, 1991–1997), and communications industry GSP (net of BOC contributions) from BEA *Gross Product by Industry*.

**Table 3: Estimation Results for the Innovation Model (Independent Version)**

| | *Prediction* (Assuming No Structural Change)* | *95% Prediction Interval†* | *P-value of Actual Number of Interim Services‡* |
|---|---|---|---|
| **Poisson Model** | 16.30 | (7.5, 25.6) | 0.0197 |
| **Generalized Poisson Model** | | | |
| Fixed Effects | 16.13 | (7.6, 25.8) | 0.0225 |
| Congestion | 13.35 | (5.7, 21.9) | 0.0024 |
| Economic Variables | 13.09 | (6.4, 22.9) | 0.0048 |
| **Cox Model** | 15.01 | (7.8, 23.1) | 0.0048 |

*Excludes waivers. Waiver requests were permitted and observed during the interim, and thus do not need to be predicted.

†95% centered prediction interval based on bootstrap prediction errors (see notes 11 and 22).

‡One-sided $p$-value of 26 services.

Table notes: predictions are based on the models estimated with data from the regulated era. A prediction that differs from the actual number of services created is evidence that innovation changed when the firms were released from regulation. Predictions are calculated using actual covariates during interim period.

**Table 4: Predictions and Prediction Intervals for the Innovation of Information Services During the Interim Period**

| Variable | Weibull Model<br>*Fixed Effects* | Generalized Weibull Model<br>*Fixed Effects* | <br>*Congestion* | <br>*Covariates* |
|---|---|---|---|---|
| intercept | −0.655*** | −0.403 | −0.349 | −0.464** |
| | (0.239) | (0.303) | (0.280) | (0.228) |
| BOC (not AT&T) | 0.318 | 0.242 | 0.288 | 0.041 |
| | (0.251) | (0.257) | (0.257) | (0.227) |
| amendment | −0.095 | −0.341 | −0.437* | −0.674*** |
| | (0.249) | (0.290) | (0.258) | (0.231) |
| waiver | 0.144 | 0.044 | 0.017 | 0.262 |
| | (0.238) | (0.255) | (0.240) | (0.195) |
| me-too filing | −0.522** | −0.504** | −0.442* | −0.363** |
| | (0.229) | (0.238) | (0.231) | (0.177) |
| refiled | 0.389 | 0.354 | 0.357 | 0.255 |
| | (0.332) | (0.302) | (0.299) | (0.248) |
| filings awaiting approval | | | −0.107 | −0.131 |
| at start of spell | | | (0.140) | (0.113) |
| filings arriving during | | | −0.288** | −0.133 |
| spell | | | (0.119) | (0.102) |
| pages in *FCC Record* | | | | 0.123 |
| (in logs) | | | | (0.104) |
| commenters | | | | 0.046* |
| (in logs) | | | | (0.025) |
| $\sigma$ (Weibull shape | 0.613*** | 0.570*** | 0.554*** | 0.426*** |
| parameter) | (0.060) | (0.067) | (0.065) | (0.049) |
| $\tau_2$ (heterogeneity | 0.000 | 0.323*** | 0.235** | 0.103 |
| parameter) | (fixed) | (0.131) | (0.128) | (0.129) |
| predicted mean duration (years) | 0.620 | 0.703 | 0.757 | 0.585 |
| (ave. in sample) | | | | |
| log likelihood | −19.79 | −18.59 | −16.30 | 7.34 |
| observations | 245 | 245 | 245 | 189 |

* = 10% level significance; ** = 5% level significance; *** = 1% level significance; stars for $\sigma$ and $\tau_2$ based on one-sided test.

Table notes: asy. s.e. in parentheses. The conditional mean of log dependent variable is linear in the regressors.

**Table 5: Estimation Results for the Regulatory Approval Delay Model (Independent Version)**

| | **Scenario 1:** *CEI Regime Conditions* | **Scenario 2:** *Interim Conditions* | **Difference** |
|---|---|---|---|
| Observation Period (years) | 11.00 | 11.00 | |
| Projected Services Created | 92.46 | 140.30 | |
| Projected Services Not Approved | 5.74 | 0.00 | |
| Projected Services Introduced | 86.72 | 140.30 | 53.58 |
| *standard deviation* | *10.69* | *11.84* | *15.96* |

Table notes: Services Not Approved are arrivals still awaiting regulatory approval at end of period. See text and notes 30 and 31 for details.

**Table 6: Projections for the Innovation and Introduction of Information Services During the Entire Observation Period**