Estimating Demand for Local Telephone Service with Asymmetric Information and Optional Calling Plans

Eugenio J. Miravete[†⋆]

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Abstract

In this paper, I study the theoretical and econometric implications of agents' uncertainty concerning their future consumption when a monopolist offers them either a unique, mandatory nonlinear tariff or a choice in advance from a menu of optional two–part tariffs. Agents' uncertainty is resolved through individual and privately known shocks to their types. In such a situation the principal may screen agents according to their *ex ante* or *ex post* type, by offering either a menu of optional tariffs or a standard nonlinear schedule. The theoretical implications of the model are used to evaluate a tariff experiment run by South Central Bell in two cities in Kentucky in 1986. The empirical approach explicitly accounts for the existence of informational asymmetries between local telephone users and the monopolist, leading to different, nested, econometric specifications under symmetric and asymmetric information. The empirical evidence suggests that there exists a significant asymmetry of information between consumers and the monopolist under both tariff regimes. All expected welfare components failed to increase with the introduction of optional tariffs for the estimated value of the parameters. JEL: D42, D82, L96.

Keywords: Optional Tariffs; Asymmetric Information; Type Shocks.

[†] Department of Economics, University of Pennsylvania, McNeil Building / 3718 Locust Walk, Philadelphia, PA 19104-6297; and CEPR, London, UK. Phone: 215-898-1505. Fax: 215-573-2057. E–mail: *miravete@ssc.upenn.edu*; http://www.ssc.upenn.edu/~miravete

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1 Introduction

Over the past fifteen years, the use of optional calling plans for pricing telephone services has grown in popularity in the United States and elsewhere. The introduction of optional local measured tariffs was originally envisioned as playing two roles. First, optional local measured tariffs were regarded as an instrument designed to facilitate universal service. Second, such local tariffs were also viewed as enabling a reduction in the economic distortion generated by the use of untimed local service, which was available at zero marginal charge in the U.S. despite a low but generally positive marginal cost of service. Today, optional calling plans are not restricted to local service, and interexchange carriers have used them intensively as marketing strategies in a competitive environment since the divestiture of AT&T.

The critical feature that distinguishes optional calling plans from standard nonlinear tariffs is the existence of a time lag between tariff choice and the consumption decision. Telephone pricing becomes a two–stage problem if optional tariffs are used. At the beginning of the billing period, consumers choose the tariff plan that the local monopolist will apply to pricing their future consumption. Later, given their tariff choice, they decide on telephone usage. The standard theory of nonlinear pricing neglects the two–stage nature of choice under optional tariffs and its implications. Similarly, empirical studies of telecommunications demand have treated tariff choice and usage as simultaneous decisions.

The objective of this paper is threefold: first, to develop a model that explicitly accounts for the effects of uncertain future consumption on the choice of tariffs, addressing both the consumer's choice of tariff and the subsequent telephone usage decision; second, to show that under this additional source of uncertainty for consumers, the monopolist may discriminate among them by offering a menu of optional calling plans; and finally, to develop an analytical structure to test the relevance of this additional source of uncertainty together with the importance of asymmetric information in a principal–agent relationship.

Optional tariffs are also common in other industries. In some cases, as in the subscription of insurance policies or personal health and retirement plans, the choice of an option embodies insurance against unknown states of nature. However, very often, the total amount paid under any of the options is so low relative to a consumers' income that risk aversion is likely to be unimportant. This is certainly the case for local telephone service, but also for car rentals (free mileage and full tank options), fees for electronic transfer of funds between financial institutions using the FED's network, weekly or monthly passes for public transport systems, advance subscription to spectacles, or the choice of a Sunday brunch instead of the *à la carte* menu. Therefore I assume risk neutrality on the part of consumers with respect to monthly bill variation.

In all these cases, the option chosen *ex ante* need not to be the one that minimizes expenditures *ex post* every billing period. For local telephone service, it is often observed that many consumers would have paid less had they chosen a different option. In particular, it is commonly claimed that telephone users with low calling profiles generally choose the more expensive flat rate service in larger proportion than users with intensive telephone demand patterns who had chosen the measured service option and who end up paying more than the fixed, flat rate tariff.[1] I argue that these "mistakes" may be explained by stochastic elements that affect consumer types so that each consumer chooses *ex ante* the tariff with the highest option value given the conjectures about their future, individually–relevant, state of nature.

To capture this idea, I distinguish between *ex ante* and *ex post* consumer types. The model explicitly deals with the two–stage decision process by assuming that consumers' *ex post* types have two dimensions: an *ex ante* type and a shock. A consumer knows only her *ex ante* type when she chooses among tariff options. In the interim between the tariff choice and the usage decision, each consumer receives an individually and privately known shock. The magnitude and sign of the shock incorporates any increase in the consumer's information set that is relevant for her usage decision, which is determined by her *ex post* type. Both *ex ante* and *ex post* types are always private information for consumers. The monopolist knows the distribution of *ex ante* types. Furthermore, at the time of choosing

[1] Using the data examined below, Miravete (2000) has shown that this common belief is not supported empirically. The reported evidence also rules out risk aversion as a sensible explanation of the observed tariff choices.

among plans, the monopolist and consumers share the same prior on the distribution of shocks. I show that the monopolist is able to design a menu of self–selecting optional tariffs for this case when there exists a stochastic component in consumers' demand. Therefore, the monopolist may discriminate among consumers according to their *ex ante* distribution by offering a menu of optional calling plans (*ex ante* tariff), or according to their *ex post* distribution, through a fully nonlinear schedule (*ex post* tariff). The monopolist decision does not constrain consumers to use the same information set to decide on consumption as in Lewis and Sappington (1994), but it affects consumers' purchases through the *ex ante* choice of tariff since such choice defines their *ex post* nonlinear budget sets.

I analyze the demand for residential local telephone service in two cities of Kentucky in the fall of 1986. These data are suitable to test the two different stochastic structures of the model. In Bowling Green all customers were placed on mandatory measured service and a standard nonlinear pricing model is estimated. The "choice" of marginal tariff and usage are simultaneous decisions and estimation of monthly expenditures on local telephone service identifies the parameter indexing the distribution of the *ex post* type (unobservable individual heterogeneity related to usage). In Louisville consumers chose between a flat tariff and local measured service. Therefore, the distinction between the tariff choice and usage decision is relevant, and an econometric specification based on the *ex ante* tariff is estimated. In this case, the choice of tariff plan identifies the distribution of the *ex ante* type (heterogeneity related to choice) and the subsequent usage decision identifies the distribution of the type shock.

The model is deliberately constructed to provide closed–form solutions for both pricing problems. Closed–form solutions to the theoretical model identify nonlinear restrictions on the estimated regression coefficients of monthly payments for local telephone service. I find that the hypothesis of symmetric information is always rejected both under mandatory measured and optional two–part tariffs. This result calls into question empirical work on telecommunications demand where the asymmetry of information between consumers and the monopolist has commonly been neglected. Finally, the model provides with a useful framework to evaluate the welfare effects of the introduction of optional calling plans.

Using the estimates of the structural parameters I conclude that the optional nonlinear tariff offered in Louisville failed to increase expected profits and consumer surplus.

The work of Clay, Sibley, and Srinagesh (1992) is close to the present theoretical approach. However, they do not show that the monopolist may discriminate among *ex ante* consumer types by offering a menu of *ex ante* tariffs. This result is proven by Miravete (1996) who works with a continuum of types. In addition he also shows that the private and social expected desirability of optional tariff plans depends on the relative variance of components of consumer types. Courty and Li (2000) develop a similar sequential–screening mechanism when consumers have unit demands. Panzar and Sibley (1978) and Spulber (1992) extend the basic nonlinear pricing model so that type changes are involved. In contrast with my model, contingent tariffs are such that the shock is common to every consumer and also observed by both consumers and the monopolist. Few empirical papers have explicitly incorporated asymmetries of information in estimating demand, *e.g.*, Bousquet and Ivaldi (1997), Feinstein and Wolak (1991), Ivaldi and Martimort (1994), and Wolak (1994). The closest to my model is the work of Hobson and Spady (1988). Using a pooled sample as in the present paper, they obtain an estimate of the distribution of an asymmetric information parameter, but this is done without a structural treatment of the two–stage decision process under optional tariffs, and imposing an *ad hoc* symmetric distribution of shocks, which in fact may critically condition the results obtained.

The paper is organized as follows. In Section 2, I describe the tariff experiment undertaken by South Central Bell in two cities of Kentucky in 1986 and discuss statistical differences between these two cities. In Section 3, I characterize the optimal, *ex post* nonlinear tariff for measured telephone service and optional calling plans. In Section 4, I present the econometric model, point out the links between the features of the experiment and the theoretical approach of this paper, and present the estimates of the determinants of monthly payments under symmetric and asymmetric information. In Section 5, I evaluate the welfare effects of introducing optional two–part tariffs in Louisville after estimating the structural parameters of the model. Finally, Section 6 concludes.

2 South Central Bell's Experiment

In November 1984, the Kentucky Public Service Commission (KPSC) established Administrative Case No. 285 to study the economic feasibility of providing local–measured–service telephone rates. The commission had placed a moratorium on South Central Bell Telephone Company's (SCB) optional–measured service a few months earlier. The reason, argued at the time, was that KPSC had not decided whether optional local–measured–service rate was "proper." Consequently SCB carried out an extensive experiment in the second half of 1986 in two Kentucky cities to provide the commission with evidence in favor of extending optional local–measured–service.

In spring, when all customers in Kentucky were on flat rate, SCB collected demographic and economic information for about 5,000 households in Bowling Green and Louisville. The local exchange carrier also collected monthly data on usage (number and duration of calls classified by time of the day, day of the week, and distance) for a period of three months. New tariffs were introduced in July 1st, and after a three month adjustment period, SCB collected another three months of billing and usage data for residential customers. These are the data used below.

The experiment defined two different scenarios in the second half of 1986. In Bowling Green all customers were on mandatory–measured service. The tariff included a fixed fee of \$8 per month and a \$21.50 bill cap. This was an explicit requirement of the KPSC who wanted to minimize the impact of the experiment on the expenses of individual customers. There was a setup charge of 1 cent and a duration charge of 1 cent per minute in peak time. Peak period was from 8 a.m. to 8 p.m. on weekdays. Off–peak charges had a 50% discount on both setup and duration. In Louisville consumers had a choice between unlimited calls at a cost of \$18.70 per month, or a measured service option with a monthly fee of \$14.02. The measured service tariff in Louisville included a \$5 allowance and distinguished setup, duration, peak periods, and distance. The tariff differentiated among three periods: peak was from 8 a.m. to 5 p.m. on weekdays; shoulder was between 5 p.m. to 11 p.m. on weekdays and Sundays; and off–peak was any other time. For distance band A, measured charges in peak periods were 2 cents for setup and 2 cents per minute. These charges had a 35% discount in shoulder time and 60% discount in off–peak time. For distance band B, setup charges remained equal but duration always had a 100% surcharge relative to the corresponding tariff for band A.

The situation in Bowling Green mimics the *ex post* tariff formulation of Section 3.4. Consumers were not given a choice of tariffs, and because of the nonlinearity of the tariff schedule, the marginal tariff of each customer depends on her telephone usage; *i.e.,* number of calls, duration, and temporal calling patterns according to the measured tariff in Bowling Green. After deleting observations with missing values, the sample includes 6,445 observations for Bowling Green. Note that 553 households, about 9% of the sample, fell above the cap, and therefore faced a zero marginal tariff. The existence of a binding tariff cap forces me to control for a censored endogenous variable.

The Louisville scenario resembles the case of *ex ante* tariffs in Section 3.5. Customers in this exchange faced a two–stage decision problem. At the beginning of the billing period they chose between alternative tariffs, one with a positive marginal charge and another with zero marginal charge. The consumption decision was taken over the billing period once the tariff plan had been chosen. The sample size is 5,576, of which 29% (1,615) selected measured service and 71% (3,961) chose flat rate. Tariff choice in Louisville was not between two two–part tariffs, but rather between two nonlinear tariff schedules defined on an abstract aggregate measure of usage. Customers who chose the measured option were not committed to a particular *ex post* positive marginal rate, but to the whole multidimensional schedule. The actual *ex post* marginal rate defined upon the aggregate usage unit was fixed only if customers did not exceed the allowance.

2.1 Data Description

The data set includes 12,021 observations on monthly payments for residential customers, 6,445 in Bowling Green and 5,576 in Louisville. In addition to monthly payments, the data set also contains several dummy variables from the original recorded data. These data are described in Table 1. Only active users were considered. Households that did not make any calls over the three months when the data were collected have been excluded from the sample. Observations with missing values for variables used in the analysis have also been excluded. In most cases the number of observations excluded is small and the results do not vary significantly. There is one exception. There are 1,026 households in Bowling Green (16%) and 1,036 in Louisville (19%) that did not report their income. Because of the importance of this variable and the large number of missing values, I have recoded these observations and included a dummy variable, DINCOME, to control for non–response and thus avoid selection–bias problems. Missing observations are recoded at the yearly average expected income of $19,851.[2]

Table 1 presents descriptive statistics and compares household characteristics in Bowling Green and Louisville, as well as households on measured and flat rate in Louisville. Consumers in Bowling Green and Louisville belong to populations with quite different characteristics. In general, households in Bowling Green are larger, and they have higher income, lower average age (more teenagers), include a larger proportion of married couples, and are geographically more mobile. There is also a larger proportion of college graduates and people that makes use of the telephone for charity purposes. By contrast, households in Louisville are more likely to receive benefits, include blacks or single and males among their members. There is another difference between the two exchanges that is not captured by Table 1. By the end of the 1980's Louisville had a population over 250,000 while Bowling Green barely reached 50,000. It is well documented that the size of the exchange generates important network externalities that lead to higher local telephone demand in more populous exchanges [Taylor (1994, §7.1)]. As for the choice among measured and flat rate service, note that on average, those who chose the measured service ended up paying $2.07 above the cost of the flat tariff option. This is an indication that errors in predicting future usage are not only common but also opposed to the common belief that most mistakes are made by those who choose the flat tariff. Estimates of the structural parameters in Section 4 are consistent with this result.

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[2] This estimate is close to the income per household reported by the 1990 U.S. Census of Population and Housing for Bowling Green ($20,043) and Louisville ($20,141). This estimate is the mean of a displaced gamma distribution used to fit the empirical frequency of the reported income categories. See the Appendix posted on the web page of this journal for a detailed description of the construction of a continuous income variable from the reported categorical data.

3 Ex Ante and Ex Post Tariffs

Consider the following game: A monopolist without capacity constraints offers customers either the choice between an *ex ante* tariff plan or an *ex post* tariff plan. In the case of the *ex ante* tariff (Louisville) a representative consumer chooses the tariff plan before she knows her *ex post* type completely, and the monopolist commits to bill her according to the plan chosen. When the state of nature is revealed and the consumer privately learns her *ex post* type, individual usage is determined by maximizing her utility subject to the billing system she has chosen. In the case of the *ex post* tariff (Bowling Green), consumers choose simultaneously the usage and the marginal tariff conditional on their *ex post* type. The *ex ante* type θ_1 is always the private information of the consumer, while the monopolist only knows its distribution. However, the monopolist and consumers share the same prior on the distribution of the shock θ_2. Since the monopolist knows the distribution of both components of the *ex post* type, he can compute the optimal nonlinear schedules using either the distribution of the *ex ante* type θ_1 or the distribution of the *ex post* type θ. In this section I make explicit the choices of specific functional form for telephone demand and distribution of types that allows me to obtain close–form solutions to both problems, as well as to identify the role of asymmetry of information. The features of this model will later guide the empirical analysis of this paper using cross–section data.

3.1 Demand

Although calling plans may be complicated, most of them consist of monthly payments of a fixed fee plus a constant marginal rate times usage in excess of the plan allowance. This modeling choice is less restrictive than it might look at first sight. For the *ex post* pricing case, a continuum of two–part tariffs characterizes the fully nonlinear tariff schedule. The argument is less straightforward for the *ex ante* pricing problem. It is true that if optional tariffs are only two–part tariffs the monopolist does not make use of potentially profitable incentives to screen consumers according to their *ex post* type, within each tariff option, once their type shock is realized. But the present approach characterizes the lower envelope of the tariff in closed form, and this happens to be the same for optional two–part tariffs and fully nonlinear options.[3] Thus, I assume that the monopolist's tariff plans consist of two–part tariffs defined over some aggregate measure of telephone usage, x:

$$T(x) = A + px. \tag{1}$$

In order to obtain a closed–form solution to the nonlinear pricing model, I specify the following linear demand for telephone service:

$$x(p, \theta) = \theta_0 + \theta - p, \tag{2}$$

where θ represents the consumer's *ex post* type, and θ_0 is assumed to be a parameter large enough to ensure that consumption equation (2) is always non–negative so that the monopolist serves all the market.[4] In addition, the share of local telephone bills over total household income is low enough to assume that the marginal utility of income is constant.[5] This particular demand specification leads to the following consumer surplus net of fixed fee payments for a customer of type θ:

$$V(p, A, \theta) = \frac{(\theta_0 + \theta - p)^2}{2} - A. \tag{3}$$

3.2 Costs

The monopolist uses a constant returns–to–scale technology. His cost function is given by

$$c(x) = K + cx, \tag{4}$$

which seems adequate for an industry characterized with very high fixed costs and very low marginal costs. The cost function is defined on usage which is an aggregate of duration, distance, peak periods, and setup charges.

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[3] See Miravete (2001b, §3). There are some other minor justifications to following this approach. Two–part tariffs are commonly used in many industries because of their simplicity and low monitoring costs. Furthermore, Miravete (2001b, §6) also evaluates that the increase in expected profits from implementing a menu of fully nonlinear options relative to a menu of two–part tariffs in Louisville is only of about 4%.

[4] This assumption is reasonable for basic telephone service since universal service is encouraged by the regulator, and because according to the 1990 U.S. Census of Population and Housing 89.33% of the households in Bowling Green and 92.07% of the households in Louisville subscribed local telephone service.

[5] In the present study the average share of telephone expenditure ranges from 1.6% in Bowling Green to 2.8% in Louisville. See Miravete (2000a, §3.3) for an explicit test of this hypothesis.

3.3 Beta Distributed Types

In principle, there are two ways to deal empirically with asymmetry of information in the presence of nonlinear pricing. If panel data were available, the econometrician could estimate price elasticity conditional on observable demographics while avoiding functional form assumptions. Estimates would serve to characterize the optimal tariff options that the monopolist should offer. If only pooled data are available, the econometrician has to resort to some functional specification for the distribution of unobservable characteristics of consumers. This approach, very common in the empirical auction literature [Donald and Paarsch (1996); Laffont, Ossard, and Vuong (1995); Paarsch (1992, 1997)], estimates the parameters of the model together with those that select one among a family of distributions of the unobservable, individual heterogeneity parameter. Thus, I assume that offered options are optimal, and characterize the distribution of asymmetric information parameters that is consistent with such tariff options.

I consider a two–stage model to address the decision process with optional tariffs. I define the *ex post* type θ so that it has two components: the *ex ante* type θ_1, already known at the time of the tariff choice, and the shock θ_2 which is learned in the interim between the tariff choice and the consumption decision. The idea is that the tariff choice depends on conjectures about the future, individual specific, state of the world, while consumption depends on the realized state for each individual. For analytical convenience I assume that the *ex ante* type, *ex post* type, and the shock are related as follows:

$$\theta = \theta_1 \theta_2. \tag{5}$$

The present approach not only requires that *ex ante* and *ex post* types are related, but also that their distributions are related as well. The distribution of the *ex post* type is the composition distribution of those of the *ex ante* type and the shock. To keep the model tractable, the distribution of the *ex ante* type is assumed to be a particular specification of the beta distribution:

$$\theta_1 \sim \beta\left[1, \frac{1}{\lambda_1}\right] \text{ on } \Theta_1 = [0, 1]; \quad \lambda_1 > 0. \tag{6}$$

A shock, ε, is assumed to be independently distributed over the unit interval with the following beta distribution:

$$\varepsilon \sim \beta\left[1+\frac{1}{\lambda_1}, \frac{1}{\lambda} - \frac{1}{\lambda_1}\right] \text{ on } 0 \leq \varepsilon \leq 1; \quad 1+\frac{1}{\lambda_1} > 0; \quad \frac{1}{\lambda} - \frac{1}{\lambda_1} > 0. \tag{7}$$

The independence assumption is critical to obtain closed–form solutions for the present model as discussed in Miravete (2001b, §4.3). The multiplicative structure of equation (5) captures the observed fact that higher telephone demand consumers are also those with higher variability of calling patterns. This observation may cause the heteroscedastic errors of the usage equations that I find at the estimation stage. Thus, the theoretical model is well suited to deal with this empirical feature of the data. I normalize the type shock as follows:

$$\theta_2 = 1 + \varepsilon - \mu_\varepsilon \text{ on } \Theta_2 = [1-\mu_\varepsilon, 2-\mu_\varepsilon]; \quad \text{and } 0 \leq \mu_\varepsilon = \frac{\lambda(1+\lambda_1)}{\lambda_1(1+\lambda)} \leq 1. \tag{8}$$

The normalization of the shock support and the linearity of the *ex post* type in θ_1 and θ_2 capture the idea that consumer's actual consumption equals her expectation when the realized shock equals its mean, *i.e.*, $E_2[\theta] = \theta_1\mu_2 = \theta_1$. From these assumptions it follows that the *ex post* type is also beta distributed [Kotlarski (1962)]:

$$\theta \sim \beta\left[1, \frac{1}{\lambda}\right] \text{ on } \Theta = [0, \overline{\theta}] = [0, 2-\mu_\varepsilon]; \quad \lambda > 0. \tag{9}$$

The beta is a very flexible distribution that, under the present parameterization allows me to find a closed–form solution for the optimal nonlinear tariff problem. Furthermore, parameters λ, and λ_1 directly identify the hazard rate of the distributions of θ and θ_1 respectively, which play a critical role in characterizing the relative power of *ex ante vs. ex post* pricing. The beta distributions of θ or θ_1 are beta distributions of the first kind with parameters 1 and λ^{-1} (also known as the Burr distribution of type XII). They mimic the exponential pattern of telephone usage behavior [Bousquet and Ivaldi (1997, §2.1)], can be integrated analytically, and are characterized by an inverse hazard rate proportional to λ and λ_1 respectively. Thus, the pricing problems are well defined as long as λ and λ_1 are positive. Observe that (7) also requires that $\lambda < \lambda_1$, *i.e.*, the hazard rate dominance of the distributions of types is also given by the relative magnitude of these parameters.

This feature proves critical to rank the power of alternative pricing mechanisms and to interpret the welfare implications of results. In addition, the beta distribution is defined on a compact support, so that the intercept of consumers' demands (consumers' *ex post* type) may be constrained to take only positive values.[6]

Observe that it is not possible to identify separately θ_1 and θ_2 in equation (5) in the case of Bowling Green because in that local exchange consumers are only offered a single nonlinear tariff. There, unobserved consumer characteristics determine different usage levels for customers with identical observable demographics. This heterogeneity of usage conditional on observable characteristics allows me to identify λ in Bowling Green. Identification of θ_1 and θ_2 is possible in Louisville because of the sequential decision process that their residents face. The heterogeneous choice of tariff plans conditional on identical demographics allows me to estimate λ_1 while different posterior usage decisions conditional on identical demographics and tariff choice enables me to estimate λ.

3.4 Ex Post Nonlinear Tariff: Bowling Green

In Bowling Green, consumers pay according to their consumption at the end of the billing period. Consumers optimally choose the usage level given their *ex post* type and the monopolist's tariff. Consumers are offered a single nonlinear tariff. The choice of a particular two–part tariff is dual to the *ex post* choice of usage if the tariff is concave. However, there is not any "choice of tariff plan" in the sense explained before and the problem becomes standard. The monopolist designs an optimal direct *ex post* mechanism $\{\hat{p}(\theta), \hat{A}(\theta)\}$ that induces consumers to truthfully reveal their *ex post* type. By making this mechanism incentive compatible (IC) the monopolist maximizes his expected profits while

[6] It is worth mentioning some of the properties of these distributions as the proposed test of asymmetric information will rely on the particular values of λ and λ_1. When $\lambda \to 0$, the distribution of θ becomes degenerate at 0 and $\mu_\varepsilon \to 0$. When $\lambda_1 \to 0$, the distributions of θ_1 and θ become degenerate at 0 since $\lambda < \lambda_1$ and $\mu_\varepsilon \to 1$. Similarly, when $\lambda_1 \to \infty$, $\mu_\varepsilon \to \lambda/(1+\lambda)$, but when $\lambda \to \infty$, $\mu_\varepsilon \to 1$. Finally, when $\lambda \to \lambda_1$, $\mu_\varepsilon \to 1$. The support of the distributions of θ_1 and θ_2 is always the unit interval while the support of θ expands or contracts depending on the variance of the shock. In all cases where Θ is the unit interval, the distribution of θ_2 becomes degenerate at $\theta_2 = 1$. See footnote 10 for a discussion on testing whether λ and λ_1 are significantly different from zero.

minimizing consumers' informational rents. Incentive compatibility in the usage/tariff choice decision requires

$$A'(\theta) = -(\theta_0 + \theta - p(\theta))p'(\theta), \tag{10}$$

By maximizing expected profits with respect to $F(\theta)$, we can characterize the two–part tariff effectively faced by any consumer type θ as follows:[7]

$$\hat{p}(\theta) = c + \lambda(\overline{\theta} - \theta), \tag{11a}$$

$$\hat{A}(\theta) = \frac{\lambda(1+\lambda)\theta^2}{2}. \tag{11b}$$

I assume that all consumers are served by the monopolist. The non–negativity constraint of purchases requires that $\hat{p}(\theta) \leq \theta_0$, but since $\hat{p}'(\theta) = -\lambda < 0$, it suffices that the constraint is binding for $\theta = 0$. Thus, for the *ex post* pricing problem, the "sufficiently large" θ_0 is found combining (2) and (11a):

$$\theta_0 = c + \lambda\overline{\theta}, \tag{12}$$

which has already been used in deriving (11b). Thus, $\hat{A}(0) = 0$ and $V(0) = 0$.

I can also write in closed form the optimal purchase for a consumer of type θ who is offered the *ex post* tariff (11),

$$\hat{x}(\theta) = \theta_0 + \theta - \hat{p}(\theta) = (1+\lambda)\theta. \tag{13}$$

The following proposition summarizes the standard properties of this mechanism:

PROPOSITION 1: *The ex post pricing solution is such that:*

(a) *Necessary conditions to solve the monopolist's expected profit maximization, are also sufficient,*

(b) *If $\{\hat{p}(\theta), \hat{A}(\theta)\}$ is an IC mechanism, it is almost everywhere differentiable. Consumers with higher valuations pay lower marginal tariffs but higher fixed fees,*

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[7] Derivation of results presented in the remaining of Section 3 are detailed in the Appendix 2 on the web page of this journal.

(c) Consumers with higher valuations purchase larger quantities if $\lambda > 0$,

(d) The marginal willingness to pay of any consumer exceeds marginal cost, except for the highest ex post consumer type,

(e) The mechanism leads to a fully separating equilibrium so that the ex post tariff may be implemented by a continuum of self–selecting two–part tariffs if $\lambda > 0$.

An important feature of this model is the link between parameter λ, the magnitude of asymmetric information, and the optimal price distortion that the monopolist has to introduce for each type θ in order to ensure incentive compatibility. As equations (11)$-$(13) show, the marginal tariff and optimal purchase are linear in θ while the fixed fee and the optimal *ex post* tariff are quadratic in θ. Thus, the solution can be considered a general quadratic approximation where the degree of concavity is determined by the value of λ. The empirical analysis identifies the magnitude of the information on pricing. Since the *ex post* tariff is concave whenever $\lambda > 0$, the monopolist offers price discounts to consumers. Only the highest *ex post* consumer type is efficiently priced while the others are priced above marginal cost in order to reduce their informational rents and ensure that truth telling characterizes a separating equilibrium (with respect to *ex post* types). Any increase in the value of λ means that there is a smaller proportion of infra–marginal consumers' *ex post* types, which leads to higher prices for every consumer (except for $\overline{\theta}$) in order to reduce consumers' informational rents and induce them to truthfully reveal their *ex post* type. Consequently, the *ex post* tariff will be characterized with more important price distortions the higher the value of λ.

PROPOSITION 2: *Price margin is increasing in* λ *when* $\theta \sim \beta[1, \lambda^{-1}]$ *on* $[0, \overline{\theta}]$.

3.5 Ex Ante Nonlinear Tariff: Louisville

In Louisville consumers first choose one from a number of exclusive tariff plans. At that time consumers are uncertain about their future consumption but know their *ex ante* type. The monopolist commits to apply the consumer's chosen tariff to her future usage, and so consumers choose their tariff in order to maximize their expected consumer surplus.

As in Section 3.4, both the choice of tariff plan and the usage decision may be written as communication games. The IC constraint for consumers' choice of plan is:

$$A'(\theta_1) = -\int_{\Theta_2} [(\theta_0 + \theta_1\theta_2 - p(\theta_1))p'(\theta_1)]\, dF_2(\theta_2) = -(\theta_0 + \theta_1 - p(\theta_1))p'(\theta_1). \tag{14}$$

When the uncertainty concerning θ_2 is resolved, consumers decide how much to purchase from the monopolist. This consumption decision may also be represented by a second communication game. The IC constraint for consumers' usage decision is:

$$\theta_0 + \theta_1\theta_2 - x(\theta_1\theta_2) = p(\theta_1). \tag{15}$$

Observe that the first IC constraint holds only in expectation. It is possible that for particular realizations of the shock consumers could have been better off *ex post* had they chosen a different tariff plan *ex ante*. However, they never violate any IC constraint. Each consumer maximizes her expected utility when she chooses the tariff plan, and when she learns her *ex post* type, her consumption maximizes her utility even though the same consumption could have been achieved at lower cost using a different tariff plan [Train, Ben–Akiva, and Atherton (1989)]. Therefore, by maximizing expected profits with respect to $F_1(\theta_1)$, we can characterize the optimal two–part tariff chosen by a consumer of *ex ante* type θ_1 after making use of the non-negativity constraint $\theta_0 = c + \lambda_1$:

$$\tilde{p}(\theta_1) = c + \lambda_1(1 - \theta_1), \tag{16a}$$

$$\tilde{A}(\theta_1) = \frac{\lambda_1(1+\lambda_1)\theta_1^2}{2}. \tag{16b}$$

The formal similarity with the *ex post* case allows me to summarize the relevant features of this mechanism in the following proposition.

PROPOSITION 3: *The ex ante pricing solution is such that:*

(a) Necessary conditions to solve the monopolist's expected profit maximization, are also sufficient,

(b) If $\{\tilde{p}(\theta_1), \tilde{A}(\theta_1)\}$ is an IC mechanism, it is almost everywhere differentiable. Consumers with higher valuations choose plans with lower marginal tariffs but higher fixed fees,

(c) *A higher ex ante valuation induces larger purchases of the good, independently of the shock, if* $\lambda_1 > 0$,

(d) *The marginal willingness to pay of any consumer exceeds marginal cost, except for the highest ex post consumer type,*

(e) *The mechanism leads to a fully separating equilibrium with respect to* θ_1 *so that the mathematical lower envelope of the ex ante tariff is concave if* $\lambda_1 > 0$.

Besides the well known results, the most important consequence of this proposition is that the monopolist can still screen consumers through the design of optional tariffs because it is possible to design a globally concave mechanism with a fully separating solution over θ_1. Furthermore, note that the monopolist cannot profit from the information that the consumer reveals to him when she chooses a particular plan. The choice of tariff plan is made at the beginning of the billing period and the monopolist cannot change plans in the interim. The monopolist's commitment is explained by the institutional legal framework, or on the basis of a repeated relationship with consumers.[8]

As before, I can also compute the optimal consumption and payments for each tariff plan $\{\tilde{p}_i(\theta_1), \tilde{A}_i(\theta_1)\}$ according to the chosen tariff $\tilde{T}_i$. The optimal consumption of an *ex post* type θ with *ex ante* type θ_1 under *ex ante* tariff billing is:

$$\tilde{x}(\theta) = \tilde{x}(\theta_1, \theta_2) = (\theta_2 + \lambda_1)\theta_1. \tag{17}$$

Observe that while the optimal outcome functions of the *ex ante* mechanism depends only on θ_1, consumption and total payments are contingent on the realization of the individual shock. It is also straightforward to show that the optimal marginal tariff $\tilde{p}(\theta)$ is linear in θ_1, and the tariff function $\tilde{T}(\theta)$ is quadratic in θ_1. In addition, the concavity of the *ex ante* schedule is directly related to the magnitude of the asymmetry of information relative to the distribution of consumers' *ex ante* types:

[8] I should also mention that the monopolist does not further profit from screening consumer's type shock by means of fully nonlinear options instead of just two–part tariffs. Miravete (2001b, §6) shows that the increase of profits from screening consumers' type shocks is very small. In the present paper, optional two–part tariffs are a deliberate choice to obtain solutions in closed form. The empirical analysis will be able to estimate λ_1, which determines the shape of the mathematical lower envelope of the *ex ante* tariff, and which happens to be common for a menu of optional two–part tariffs, and also for fully nonlinear options.

PROPOSITION 4: *Price margin is increasing in* λ_1 *when* $\theta_1 \sim \beta[1, \lambda_1^{-1}]$ *on* $[0, \overline{\theta}_1]$.

This characterization of tariffs provides an interesting interpretation to compare *ex ante* and *ex post* nonlinear schedules. The values of λ and λ_1 capture the degree of asymmetric information between consumers and the monopolist. Propositions 2 and 4 show that the more important is the asymmetry, the larger is the mark–up that the monopolist charges for each unit sold in order to reduce consumers informational rents. This result is obtained whenever θ_1 dominates in hazard rate to θ [Maskin and Riley (1984, §4)]. Appendix 1 shows that this is the case when $\lambda \leq \lambda_1$. But for the distribution of the shock to be properly defined it must be the case that the variance of θ_2, $\mathcal{V}(\theta_2)$, is positive which requires just the opposite, *i.e.*, that $\lambda \leq \lambda_1$. Therefore, the mark–up for any unit sold should be larger in the case of the *ex ante* tariff since the monopolist screens consumers only with respect to the *ex ante* type.

3.6 Testable Hypotheses

In this section I present testable hypotheses based on structural relationships implied by the theoretical model. The estimation of monthly payments for local telephone service will include the effect of several demographic and economic characteristics. Accounting for observable demographic differences does not mean that first degree price discrimination (individual pricing) is allowed, but rather that by making the analysis conditional on the demographic profile of consumers, parameters λ and λ_1 identify the magnitude of whatever additional individual heterogeneity may exists. The effect of individual characteristics on telephone demands is introduced in this model by re–scaling consumers' types. Therefore, different intercepts of consumers' demands are partially explained by their observable characteristics.

Let $\mathbf{W} = (w_1, \ldots, w_m)$ denote the set of consumers' observable characteristics both for the monopolist and the econometrician. Let $\boldsymbol{\psi} = (\psi_1, \ldots, \psi_m)^\top$ denote the vector of associated parameters. In general, the intercept may be any function of individual characteristics, but for simplicity, I shall use a linear specification. The effect of this procedure is that individual demographic characteristics shift the intercept of consumer demands. But there are not particular reasons that justify this choice over other functional specifications. However in Section 4.3 I show that the data are consistent with this redefinition of the type. For the case of an *ex post* tariff this is:

$$\breve{\theta} = \theta + \mathbf{W}\boldsymbol{\psi} \ \sim \ \beta\left[1, \frac{1}{\lambda}\right] \quad \text{on } [\mathbf{W}\boldsymbol{\psi}, \overline{\theta} + \mathbf{W}\boldsymbol{\psi}], \tag{18}$$

so that the distributions of $\breve{\theta}$ and θ are beta distributions but with different supports. In the case of an *ex ante* tariff, θ_1 is re–scaled to $\breve{\theta}_1$ and θ_2 is left as a shock with the same normalized distribution support so that $E_2[\breve{\theta}] = \breve{\theta}_1$.

The existence of asymmetric information can be tested. If unobserved heterogeneity is not relevant for pricing decisions, the monopolist then maximizes the sum of consumer surplus and profits:

$$\arg\max_x \left[(\theta_0 + \theta)x - \frac{1}{2}x^2 - cx - K\right] = \theta_0 + \theta - c, \tag{19}$$

but this is the solution for $\hat{x}(\theta)$ in equation (13) for the incomplete information case when $\lambda = 0$. This is equivalent to the beta distribution being degenerate at $\underline{\theta} = 0$. The monopolist will not charge anything different from a two–part tariff with a marginal charge equal to marginal cost. The two–part tariff will implicitly account for the distribution of demographics in the population (the average intercept defines the average consumer surplus for any given marginal charge). The introduction of additional markups to reduce informational rents are thus unnecessary. Therefore, testing $\lambda = 0$ will address the importance of asymmetric information for the design of the optimal *ex post* nonlinear tariff. Similarly, when $\lambda_1 = 0$, the solution $\tilde{x}(\theta)$ is exactly the solution under symmetric information concerning consumers' *ex ante* type. Therefore testing $\lambda_1 = 0$ will suffice to test the importance of asymmetric information for the design of optional calling plans.

Finally, I can also address whether type variation is relevant for the design of optional tariffs. The measure of consumers' uncertainty is the variance of the shock.

Proposition 5: *If $\lambda = \lambda_1$ the ex ante and ex post tariffs are the same, and*

$$\mathcal{V}(\theta_2) = \frac{\left(\frac{1}{\lambda} - \frac{1}{\lambda_1}\right)\left(1 + \frac{1}{\lambda_1}\right)}{\left(1 + \frac{1}{\lambda}\right)^2\left(1 + \frac{2}{\lambda}\right)} = 0. \tag{20}$$

If this hypothesis is not rejected, tariff plans offered by the monopolist at the beginning of the billing period may be considered two–part tariffs whose lower envelope is actually the *ex post* nonlinear schedule. In such a case, the monopolist is interested in offering optional tariffs because he may benefit from locking–in his customers. The smallest deviation from the expected consumption will push them above the tariff's lower envelope. On the contrary, if the hypothesis is rejected, the lower envelopes differ, and the monopolist accounts for consumers' uncertainty in the design of the optional tariff plans. Because of Proposition 5, it suffices to test whether $\lambda = \lambda_1$. If the test rejects this hypothesis, testing $\lambda < \lambda_1$ will confirm whether the hazard rate dominance prediction of the model holds, and consequently that the relative ordering of markups is supported by the data.

4 Estimation under Asymmetric Information

Many empirical studies have estimated telecommunications demands, and in particular the demand for telephone services after the breakup of AT&T [Kling and Van Der Ploeg (1990), Mitchell and Vogelsang (1991, §12.2.2), and Wolak (1993)]. Modeling telephone demand is difficult because of the heterogeneity of the service that distinguishes access, usage, peak periods, and sometimes distance [Park, Wetzel, and Mitchell (1983), Mackie–Mason and Lawson (1993)]. The existence of several options [Train, McFadden, and Ben–Akiva (1987)] makes the treatment of demand even more complex because it leads to nonlinearities of the tariff schedule, and the marginal tariff paid by consumers is determined endogenously. The approach of Hobson and Spady (1988) only accounts for simultaneous choice of usage and class of service and neglect the two–stage feature of the decision process and its informational implications. MacKie–Mason and Lawson (1993), control for tariff choice, endogenous marginal tariffs, and multiple dimensions of telephone pricing, although they ignore non–observed consumer's heterogeneity, as the estimation becomes intractable.

In this section I first present the econometric implications of the existence of asymmetric information according to the model of Section 3. Because of the features of the closed–form solution of the theoretical model and the specific properties of the beta distribution, I obtain an econometric specification that, under the symmetric information hypothesis, is a nested version of the model under asymmetric information. Later, I present the estimates and finally several specification tests that study whether the econometric and structural theoretical restrictions are sustained by the data. The results reported in this section indicate that the choice of tariff and usage decisions cannot be considered simultaneous, and that any reasonable estimation of demand under nonlinear pricing should not ignore the effect of individual unobservable heterogeneity.

4.1 Econometric Approach

The theoretical model provides us with a closed–form solution for the optimal tariff functions under two different regimes. Conditioning on the observable demographics as in (18), these solutions become:

$$\hat{T}(\breve{\theta}) = \hat{A}(\breve{\theta}) + \hat{p}(\breve{\theta})\hat{x}(\breve{\theta}), \tag{21a}$$

$$\tilde{T}(\breve{\theta}) = \tilde{A}(\breve{\theta}_1) + \tilde{p}(\breve{\theta}_1)\tilde{x}(\breve{\theta}). \tag{21b}$$

The model predictions can be compared with actual total payments of a sample of customers. These closed–form solutions of the theoretical model allows me to describe the demand for aggregated telephone services without addressing aggregation explicitly. Aggregate usage is an abstract measure, as are the structural parameters, but the estimates are useful for testing the importance of the asymmetry of information in the optimal design of optional tariffs. The following econometric specification decomposes the total bill payment into two factors: the average consumer type conditional on demographics and the unobservable type. Thus,

$$T = E_{\Theta}[\hat{T}(\breve{\theta})] + \theta, \tag{22a}$$

$$T = E_{\Theta_1}[\tilde{T}(\breve{\theta})] + \theta_1. \tag{22b}$$

To continue, first consider Bowling Green. Using the properties of the beta distribution and the identification condition $\overline{\theta} = 1$, it can be shown that expected monthly payment for customers below the cap, conditional on individual characteristics $\mathbf{W}$, is

$$E_\Theta[\hat{T}(\breve{\theta}) \mid \theta \leq \theta^\star] = (1+\lambda)\mu^\star \left[c + \lambda - \frac{\lambda^2}{1+2\lambda}\right]$$

$$+ (1+\lambda)\left[c + \lambda - \lambda\mu^\star\right] \sum_{i=1}^{m} \psi_i w_i - \frac{\lambda(1+\lambda)}{2} \sum_{i=1}^{m}\sum_{j=1}^{m} \psi_i w_i \psi_j w_j, \qquad (23)$$

where $\mu^\star = E(\theta \mid \theta \leq \theta^\star)$ is derived in Appendix 1. Thus, the tariff payment is a linear function of m individual characteristics and their cross–products:

$$T = \gamma_0 + \sum_{i=1}^{m} \gamma_i w_i + \sum_{i=1}^{m}\sum_{j=1}^{m} \gamma_{ij} w_i w_j + \theta. \qquad (24)$$

The existence of a binding tariff cap of \$21.50 in Bowling Green requires the estimation of a selection model whose selection rule is defined by the beta distribution of θ. The existence of individual heterogeneity makes that some households with similar characteristics are above and some below the tariff cap, thus identifying λ:

$$T = \gamma_0 + \sum_{i=1}^{m} \gamma_i w_i + \sum_{i=1}^{m}\sum_{j=1}^{m} \gamma_{ij} w_i w_j + \theta \quad \text{if } \theta \leq \theta^\star; \quad T = 21.50 \text{ otherwise}, \qquad (25a)$$

$$I^\star(\theta) = \theta^\star - \theta \quad ; \quad \theta \sim \beta\left(1, \frac{1}{\lambda}\right) \text{ on } [0,1]. \qquad (25b)$$

Next, consider the case of Louisville. Using again the properties of the beta distribution and the identification condition $\overline{\theta} = 1$, it can be shown that the expected monthly payment for customers on optional measured service, conditional on individual characteristics $\mathbf{W}$, is

$$E_\Theta[\hat{T}(\breve{\theta}) \mid \theta_1 \leq \theta_1^\star] = \left[(c+\lambda_1)\mu^\circ - \left(\lambda_1(c+\lambda_1) - \frac{\lambda_1^2(1+\lambda_1)}{1+2\lambda_1}\right)\mu_1^\star\right]$$

$$+ (1+\lambda_1)\left[c + \lambda_1 - \lambda_1\mu_1^\star\right] \sum_{i=1}^{m} \psi_i w_i - \frac{\lambda_1(1+\lambda_1)}{2} \sum_{i=1}^{m}\sum_{j=1}^{m} \psi_i w_i \psi_j w_j, \quad (26)$$

where $\mu_1^\star = E(\theta_1 \mid \theta_1 \leq \theta_1^\star)$, and $\mu^\circ = E(\theta \mid \theta_1 \leq \theta_1^\star) = E(\theta \mid \theta \leq \theta_1^\star\overline{\theta}_2)$. Therefore, the tariff payment is again a linear function of m individual characteristics and their cross–products. Similar to the case of Bowling Green, I deal with the two–stage decision process by estimating a selection model. Here, the existence of unobservable individual heterogeneity drives the choice among tariff options, which in this case identifies λ_1:

$$T = \gamma_0 + \sum_{i=1}^{m} \gamma_i w_i + \sum_{i=1}^{m}\sum_{j=1}^{m} \gamma_{ij} w_i w_j + \theta_1 \quad \text{if } \theta_1 \leq \theta_1^\star; \quad T = 18.70 \text{ otherwise}, \tag{27a}$$

$$I^\star(\theta_1) = \theta_1^\star - \theta_1 \qquad ; \qquad \theta_1 \sim \beta\left(1, \frac{1}{\lambda_1}\right) \text{ on } [0,1]. \tag{27b}$$

Equation (25) is the econometric specification of the model for Bowling Green under asymmetric information and mandatory measured service. A structural implication of the econometric specification (23) is that when $\lambda = 0$, the expected monthly payment should also be a linear function of demographic characteristics. Therefore, testing $\gamma_{ij} = 0$, $\forall\ i,j$ should be understood as a specification test of the proposed model. Given the similar structure, the same reasoning applies to the econometric specification (27) for Louisville, although in this latter case, the monthly allowance makes the estimation slightly more involved. Monthly income of consumers on optional measured service is increased by $5 (virtual income) and monthly payment neglects the allowance to estimate demand in the presence of nonlinear budget sets [Hausman (1985)]. Vector $\mathbf{W}$ includes m variables that explain monthly payments for customers under the tariff cap in Bowling Green, and on measured service in Louisville.

Straightforward application of the theoretical model to the empirical analysis is not possible for two reasons. First, the beta distributions of θ and θ_1 are defined on a compact support while the disturbance of the selection rules (25b) and (27b) should be defined on $\mathbb{R}$. The second problem is also of an econometric nature since the standard selection model makes use of the bivariate normal assumption to correct for sample selection in a common two–stage estimation procedure. I devote the rest of this subsection to deal with this technical problems of the estimation. Given the similitude of the models, I will focus in the case of Louisville.

First, whenever $I^\star(\theta_1) > 0$ in the selection equation (27b), consumers choose the flat tariff option. Obviously, the value of θ_1 needed to result in the choice of the flat tariff option is specific to each individual. I therefore need to relate the cutoff $\theta_1^\star$ to the individual characteristics that are available both for the monopolist and the econometrician. Let define, without loss of generality, the following index function:

$$\theta_1^\star = \Im(\mathbf{X}\boldsymbol{\zeta_1}) = \frac{\exp(\mathbf{X}\boldsymbol{\zeta_1})}{1+\exp(\mathbf{X}\boldsymbol{\zeta_1})} : \mathbb{R}^n \to [0,1], \tag{28}$$

where $\mathbf{X} \subseteq \mathbf{W}$ denotes the set of variables that explain the choice of tariff. Observe that the probability of choosing the flat tariff option can be written as:

$$\Pr[I^\star(\theta_1) \geq 0 \mid \mathbf{X}] = \Pr[\theta_1 \leq \Im(\mathbf{X}\boldsymbol{\zeta_1})] = \Pr\left[z_1 = \ln\left(\frac{\theta_1}{1-\theta_1}\right) \leq \ln\left(\frac{\Im(\mathbf{X}\boldsymbol{\zeta_1})}{1-\Im(\mathbf{X}\boldsymbol{\zeta_1})}\right)\right]. \tag{29}$$

The advantage of using the monotone transformation z_1 instead of θ_1 is that z_1 is distributed on $\mathbb{R}$ instead of on a bounded support. While θ_1 has a Burr type XII distribution defined on a closed support the distribution of z_1 is an exponential generalized version of the beta distribution $\beta(1, \lambda_1^{-1})$ known as Burr type II distribution (see Appendix 1) that admit a closed–form representation, that depends exclusively on the same parameter λ_1 of the theoretical model, ant that is defined over the whole real line. Therefore:

$$I^\star(z_1) = \mathbf{X}\boldsymbol{\zeta_1} - z_1 \qquad ; \qquad \frac{\exp(z_1)}{1+\exp(z_1)} \sim \beta\left(1, \frac{1}{\lambda_1}\right). \tag{30}$$

Second, equations (25) and (27) are estimated by a modified version of Heckman's (1979) two–stage method for correction of selectivity bias which provides consistent estimates for γ. The modification makes use of Lee's (1983) transformation to normality of non–normal disturbances in the selectivity equation. In the second stage I estimate the following specification by ordinary least squares:

$$T = \gamma_0 + \sum_{i=1}^{m} \gamma_i w_i + \sum_{i=1}^{m}\sum_{j=1}^{m} \gamma_{ij} w_i w_j - \sigma_1 \rho \frac{\phi\{\Phi^{-1}[F(\mathbf{X}\boldsymbol{\zeta_1})]\}}{F(\mathbf{X}\boldsymbol{\zeta_1})} + \eta, \tag{31}$$

where $\phi(\cdot)$ and $\Phi(\cdot)$ account respectively for the standard normal density and distribution functions, and where ρ is the correlation coefficient of disturbance z_1 and $\Phi^{-1}[F(\mathbf{X}\boldsymbol{\zeta_1})]$ [Maddala (1983, §9.4)], so that $E[\eta \mid I = 1, \mathbf{W}, \mathbf{Z_1}] = 0$. Finally,

$$F(\mathbf{X}\boldsymbol{\zeta_1}) = 1 - [1 + \exp(\mathbf{X}\boldsymbol{\zeta_1})]^{\frac{1}{\lambda_1}}. \tag{32}$$

4.2 Estimates

I consider the three most important demographic characteristics for telephone demand: income, size of the household, and number of teenagers. These three variables define six cross–products variables that will be used to test for mispecification of the model. The rest of variables will be used as demographic dummies.

Tables 2 and 3 present estimates of equations (25) and (27). The first refers to the sample of customers below the tariff cap in Bowling Green and the second to those customers on optional measured service in Louisville. The first column of these tables shows the estimate of the selection rule, as well as the estimate of λ in Bowling Green and λ_1 in Louisville.[9] For the selection rule I have included all available demographic variables. Observe that the effects of demographics are similar in both selection equations, as those who reach the tariff cap in Bowling Green and those that self–select into the flat tariff option are more likely to be intensive users of local telephone service. Thus, intensive users in Bowling Green include low income households, large households and/or with teenagers, blacks, and those who receive social benefits. Among the non–intensive customers we find those who are single, retired, have moved in the past five years, or are single and males. In Louisville few variables are significant in the selection rule. The choice of tariffs appears to be explained mostly by the size of the household and by whether the head of the household has a college degree or not.

But in addition to these effects, the estimation of the selection rules identifies the unobserved heterogeneity as playing an important role, both in determining the intensity of the usage in Bowling Green, and in the choice of tariff plans in Louisville. The estimates of λ and λ_1 are significant, which means that there is important asymmetric information between consumers and the monopolist in both exchanges.[10] The estimate of

[9] The estimates of the effect of demographics on the selection rule and those of λ or λ_1 are obtained simultaneously by maximum–likelihood. Appendix 1 presents the corresponding likelihood function.

[10] Observe that this test of hypothesis occurs at the boundary of the parameter space. In accordance with the results of Gouriéroux, Holly, and Monfort (1982) for linear models, McDonald and Xu (1992) found that that the size of the likelihood ratio test should be expected to be smaller than the suggested by a standard $\chi^2_{0.95}(1)$ when evaluating limiting cases of parameters of beta distributions. This presents difficulties when the null is not rejected, which is not the present case. The Wald test (asymptotically equivalent to the Likelihood Ratio test) for the significance of λ and λ_1 is 549.43 and 1321.32 respectively.

λ_1 in Louisville is not significantly different from 1 and thus, θ_1 can be considered to be uniformly distributed. However, the distribution of θ in Bowling Green is skewed to the right. This result could be interpreted as unobservable characteristics having a particularly strong positive effect on usage in Bowling Green, *i.e.*, the available information may explain most consumption decisions for low and medium intensity consumers, but unobservable heterogeneity is an important issue to predict the usage of intensive consumers. The uniform distribution of θ_1 in Louisville implies that unobservable heterogeneity has a balanced effect across users of different types, which is consistent with the low power of the available information in predicting the choice of tariff plan as reported in Table 3.

Next, I estimate four versions of the usage equations.[11]. Only two of them include demographics and two equations include the cross–products of the three key variables. The goodness of fit always improves with the inclusion of demographics (dummies) and with the cross–products. The hypothesis of exclusion of any of these groups of variables is always rejected in both cities.[12] This last result supports the specification of the model in equations (23) and (26) in which, consistent with the structural restrictions, cross–products should be jointly non–significant only when $\lambda = 0$ or $\lambda_1 = 0$.

Usage is nonlinear in income, size of the household, and number of teenagers both in Bowling Green and Louisville. In Bowling Green the most important variable is the size of the household. Demand for usage among those whoever do not reach the tariff cap is higher in December, for senior households, those who receive benefits, and who use the telephone for charity. Demand is significantly lower for married couples or those

Both tests have p–values much smaller than 0.01, under the standard distribution theory, which represents an upper bound of the actual p–value of these tests at the boundary of the parameter space.

[11] Heteroscedastic–consistent standard errors are always computed. Choice equations use the covariance matrix of Manski and McFadden (1981, §6). For the case of Louisville I also correct the maximization procedure to account for a choice–biased sample, as during this period only 10% of the population actually chose the measured service, while they amount to almost 30% of the observations in the sample. I use the sampling correction method of Manski and Lerman (1977). Standard errors in the usage equations are corrected for sample selection [Lee, Maddala, and Trost (1980)], as well as for heteroscedasticity [MacKinnon and White (1985)].

[12] The critical value of $\chi^2_{0.95}(6)$ is 12.59. For Bowling Green the likelihood ratio tests are 63.22 and 60.22 for the model with and without demographic dummies. In the case of Louisville, these statistics are 44.99 and 27.74 respectively. As for excluding the demographics, this hypothesis is rejected whenever the corresponding likelihood ratio tests exceeds $\chi^2_{0.95}(12) = 21.03$. In Bowling Green this test reaches the value of 94.00 while in Louisville is 36.52.

who moved recently. Income has a stronger negative effect on usage than on the selection equation in Louisville. The same happens with the size of the household. The direct effect of teenagers is positive or negative depending on whether cross–products are included. The effect of demographics in Louisville is as follows. Black households and those who use the telephone for charity are among the intensive users even if they subscribe the optional measured service. This empirical regularity regarding race, as well as the negative effect of income on demand for local telephone service has already been documented by Kling and Van Der Ploeg (1990). The effect of income on usage is negative but it increases with income. Together with the positive sign of the DINCOME variable on the selection rule, it implies that most households who did not report their annual income belong to income categories well above the sample average. Married and/or retired households, and with college degree have lower demands. Similar results hold for single and male households and for local telephone usage during November.

Taking the estimation with cross–products and demographics as the correctly specified model, I can compare estimates of alternative models to analyze the implications of neglecting an explicit treatment of asymmetric information. The following analysis makes use of the marginal effects evaluated a the sample means presented in Table 4. For instance, most empirical studies on telephone demand claim that the existence of teenagers in a household leads to higher demand for telephone services. The direct effect of this variable reported in Tables 2 and 3 is non–significant in Bowling Green and positive only for misspecified versions of the model (not consistent with symmetric information according to the present model). However, accounting for direct and indirect effects of demographics, the effect of teenagers is not only positive and significant, but also generally much larger than if estimation does not include those cross–products. A similar conclusion is obtained for the case of the size of the household. Direct and indirect effects of income remain non–significant.

4.3 Specification Tests

In the previous analysis, results were consistent with the structural implications of the model with asymmetric information. However, it could also be argued that the approach followed allows for such interpretation only because it relies heavily on simplifying assumptions that are not strongly justified from the theoretical perspective. In this subsection I briefly discuss some of these arguments to conclude that in general the econometric specification is flexible enough to accommodate the data.

A first issue is the assumption on the distribution of errors. The approach followed here does not make the results dependent on the normality of residuals that imposes normal kurtosis and symmetric distribution of disturbances. The Burr type II distribution used for the estimation is quite flexible, includes many distributions with very varied shapes and allows for skewed distributions (both right and left) as well as for distribution of errors with different degrees of kurtosis. The estimation shows that the errors are particularly biased, leading to distributions with a thick right tail. Thus, I do not find that the statistical assumptions are restrictive since it is actually the analysis of the distribution which is estimated jointly with the parameters what reveals the nature of the asymmetry of information.

The second issue is that of the redefinition of types in equation (18). Economic theory does not impose any restriction on this matter. I assumed that those characteristics that are observed by, or whose distribution is commonly available to the monopolist, enter linearly in the redefinition of types. The linearity assumption simplifies the analysis considerably because it leads to an econometric specification where monthly payments are regressed against a linear function of the individual characteristics and their cross–products. I will discuss two arguments in favor of the approach followed in this paper.

I could interpret equation (18) as the first–order Taylor–series expansion of any general function $\breve{\theta} = \theta(\mathbf{W})$. The immediate question is whether I can ignore higher order expansion terms. This is most relevant because the inclusion of higher order terms conditions how I test for asymmetric information. However, there is no way to determine if higher order expansion terms should be considered. The regression model is only linear in $\mathbf{W}$ if the information is symmetric and the type redefinition is linear. But it is impossible to distinguish between asymmetry of information and linearity of type redefinition because both, a linear redefinition under asymmetric information and a quadratic redefinition under symmetric information, leads to a regression of monthly payments against a second degree polynomial in $\mathbf{W}$. Considering even higher order expansion terms for the type redefinition only repeats the problem at higher degree polynomials of the regression equation. Thus, different structural assumptions could lead to observationally equivalent testable implications of the model.

This result is not very restrictive either. Equation (18) distinguishes two components behind the different intercept of the individual demand function: one explained by demographics, and one due to unobserved heterogeneity. Thus, whether the redefinition of the types is linear or quadratic is just a particular case of a broader discussion on the right specification of the functional form that relates demographics and types. In principle, different functional forms will lead to parameter estimates with different values but equivalent economic interpretation as long as the functional forms are monotone transformations of the demographics. Since there is no obvious choice, the approach followed here is to study whether the linear type redefinition is particularly wrong given the goodness of fit of the model to the data. I study the performance of the linear redefinition against the logarithmic, exponential, inverse exponential, and square root of demographics. Since these hypotheses are non–nested, the econometric analysis is based on the construction of J–tests [Davidson and MacKinnon (1993, §11.3)]. I test the linear definition *vs.* one of several non–linear definitions, H_1 *vs.* H'_C, as well as the opposite, H_2 against H'_C. Results are presented in Table 5.

Results are ambiguous when the alternative model is the logarithmic redefinition of types. While in the case of Bowling Green there appears to exist some evidence in favor of the linear model, results are inconclusive for Louisville. The few alternative functional forms analyzed in Table 5 are intended to shed some light on the robustness of this result. In general, I conclude that while there is evidence that in some cases other functional forms may perform better than the linear model, there is no clear indication of which alternative functional redefinition of types will be more appropriate.

5 Were Optional Calling Plans "Proper"?

The ultimate objective of this paper is to develop a framework to evaluate whether the introduction of optional calling plans in Louisville was welfare enhancing. In order to do so, I still have to estimate λ for this local exchange. While the selection equation identifies λ_1, the observed usage decision of those who choose the measured service option allows me to identify λ. In order to do so, I have to make use of the cross–product restrictions among the ψ_i's and the relationships between the structural parameters of the model and the estimates embodied in equation (26).

Since marginal cost cannot be identified from demand data exclusively I also assume $c = 0$. This assumption is not unreasonable. By middle of the eighties, marginal cost of local telephony was positive but very small. They were mostly generated at the local switch. After the introduction of digital technology they became more related to the connection than to the duration of the calls. More importantly, all these costs were quite small compared with the fixed cost of maintaining the local network.

I estimate equation (26) by Nonlinear Least Squares using the sample of Louisville customers that chose the optional measured service (later correcting for sample selection). This is in fact a linear model with nonlinear restrictions among structural parameters. In particular the model requires that the parameters of products involving INCOME, HHSIZE and TEENS be proportional to the product of the ψ_i's of the corresponding variables. Once the parameters have been estimated imposing these constraints, the estimate of the intercept identifies λ. Thus in accordance with the theoretical model, the effect of the shock just shifts the demand around the expected usage, which is partially explained by the demographics and the distribution of θ_1. I also make use of the consistent estimate of λ_1 obtained in the previous section, and the threshold $\theta_1^\star$, which is also consistently estimated at the average of the sample values according to (28) and the estimates of the selection equation, $\boldsymbol{\zeta_1}$ (first column of Table 3). The estimate of the parameter of the distribution of *ex post* types is $\lambda = 2.2357$. This estimate is significantly larger than the value of $\lambda_1 = 1.0353$ previously obtained.[13] This result has several implications that I now present in the remaining of this section.[14]

First, there is important *ex ante* and *ex post* asymmetry of information in the local exchange of Louisville. Both parameters are significantly positive, which rules out that the distributions $F(\theta)$ and $F_1(\theta_1)$ are degenerate. While the distribution of θ_1 is close to the uniform, the distribution of θ is skewed to the right, similarly to the distribution of θ in Bowling Green, but with a smaller degree of skewness. The implication is that usage related unobserved heterogeneity affects mostly to those customers that end up making an intensive use of the telephone, even though they had previously self–selected into the measured service option that was intended for low usage customers.

Second, estimates of λ and λ_1 are significantly different from each other. This result confirms that there is an additional source of asymmetric information due to the fact that consumers are uncertain about their future usage levels when they subscribe an optional tariff. Thus, the use of sequential screening is justified.

Third, the value of λ exceeds that of λ_1. The estimates violate a internal consistency condition –equation (7)– of the theoretical model. But violation of this condition does not lead to any bias in the estimation as any of the estimates have been obtained under such constraint. This result is however very informative of the nature and role of the asymmetry of information, and ultimately it is responsible for whether the introduction of optional tariffs can be considered welfare enhancing.

Figure 1 shows the distribution functions of θ and θ_1 (the support is normalized in both cases to the unit interval). If the shocks were balanced, we would expect that $F(\theta)$ and

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[13] The consistent t–statistic of the estimate of λ is 112.31. Estimates of INCOME, HHSIZE, and TEENS are not significant. The rest of estimates are the same of those in the third column of Table 3.

[14] The estimates of λ and λ_1, and their relative magnitude characterize the distribution of the *ex ante* and *ex post* types in a similar manner than Miravete (2001b, §5), who makes use of nonparametric methods and direct observations of individual consumer types to analyze this same data set. The implications on expected profitability and welfare are also similar, which supports the robustness of the present estimates.

$F_1(\theta_1)$ were also very similar. But for the estimated parameters, θ first order stochastically dominates θ_1, leading to the conclusion that consumers systematically underestimate their future usage. As tariff choice is made on the basis of expected usage, some customers choose the measured service although their *ex post* usage amounts to a lower bill had they chosen the flat tariff option instead. However, we should not expect that this event involves too many households because the distribution of θ is more skewed to the right than that of θ_1, *i.e.*, $F(\cdot)$ is more favorable than $F_1(\cdot)$. As there is a concentration of mass of probability around the higher values of θ, customers who belong to higher fractiles of $F(\theta)$ are also more likely to be those who had sufficiently high expectations about their future usage to justify their subscription to the flat tariff option. Thus, important type shocks are associated to high usage volumes, but since high usage consumers are also more likely to have high usage expectations, forecast errors have little effect on consumers' utility because most consumers have previously chosen the flat tariff option.

The relative magnitude of λ and λ_1 also implies, according to the results of Proposition 2 and Proposition 4, that markups are uniformly higher under the mandatory measured service than under optional calling plans. The result follows from the hazard rate dominance of θ_1 over θ (given the parameter estimates), which is briefly discussed in Appendix 1.[15] Intuitively, since there is an important concentration of probability for high values of θ, the monopolist has to introduce important distortions to distinguish among these very similar customers. Unless price distortions are not sufficiently important, high consumer types will find profitable to imitate lower types and keep a larger fraction of their informational rent.

Another implication that follows from the uniform dominance of the *ex post* over the *ex ante* markups is that the monopolist should always prefer an *ex post* based tariff. This is always true if we evaluate the problem *ex post* since θ captures all the actual differences among consumers. But the problem faced by SCB and the KPSC was to decide *ex ante* whether they should introduce and/or approve the introduction of optional calling

[15] First order stochastic dominance is a consequence of the preservation of the increasing hazard rate property under the composition low defined by equation (5). See Miravete (2001a, §5) and Miravete (2002).

plans. Thus, welfare components should be evaluated in expectation. The rest of the analysis confirms that λ is sufficiently larger than λ_1 as to make the *ex post* tariff socially and privately preferable to optional calling plans. This evidence questions that the actual tariff options introduced in Louisville were optimal.

Figure 2 presents the differences of expected profits (*ex ante* minus *ex post*), expected consumer surplus, and expected welfare. Since all these expressions depend on λ and λ_1, they have been computed for the estimated value of $\lambda_1 = 1.0353$ and several different values of λ. Thus, the horizontal axis represents the value of the ratio λ/λ_1. Given the estimate of $\lambda = 2.2357$, the value of the ratio that represents the situation in Louisville according to my estimates is $r_\lambda = \lambda/\lambda_1 = 2.1595$. Magnitudes are conveniently scaled so that when $\lambda = \lambda_1$, $E[\tilde{\pi} - \hat{\pi}] = 0$, and when $\lambda = 0$, $E[\tilde{\pi} - \hat{\pi}] = 1$, *i.e.*, when the distribution of the shock is degenerate, the distribution of *ex ante* types captures all existing asymmetry of information, and thus the expected profits of using each tariff should be the same.

Observe that for values of the ratio $r_\lambda < 1$, *i.e.*, according to the assumptions of the theoretical model, the monopolist prefer $\tilde{T}$ over $\hat{T}$. As long as the condition $\lambda < \lambda_1$ is met, the distribution of consumers would be less concentrated around the highest type *ex post* than *ex ante*. Thus, an *ex ante* tariff would introduce higher distortions at every consumption level in order to effectively separate consumers types. These higher markups suffice to ensure that expected revenues under optional calling plans exceed those of the measured service. However, the estimated value of λ exceeds that of λ_1, and in the region where $r_\lambda > 1$ the situation is reversed. The more concentrated *ex post* consumers become close to $\overline{\theta}$ relative to their *ex ante* distribution, the higher is the expected profit difference in favor of $\hat{T}$ over $\tilde{T}$. The argument, as before, is based on the need of a more powerful mechanism to induce self–selection when consumer types are heavily concentrated around the highest type. In this case, the hazard rate dominance of θ_1 over θ induces uniformly higher markups for $\hat{T}$ at every usage level. At the estimated parameter values, the expected profit difference in favor of $\hat{T}$ amounts to 88% of the maximum expected profit difference in favor of $\tilde{T}$, which happens at $r_\lambda = 0.263$.

Consumers' expected payoff difference decreases monotonically with r_λ. Their expected surplus difference under $\hat{T}$ is only 50% of the maximum of their informational rents difference under $\tilde{T}$, which happens at $\lambda = 0$. For values of the ratio $r_\lambda < 1.946$, high *ex post* markups leave little or no rent to many consumers, thus making *ex ante* pricing the preferred tariff. For $r_\lambda > 1.946$ they are considerably more concentrated around $\overline{\theta}$ *ex post* than *ex ante* from the consumer surplus perspective, therefore making *ex post* pricing the preferred alternative.

Given the positive magnitude of $E[\tilde{\pi} - \hat{\pi}]$ for small values of r_λ, and even larger $E[\tilde{V} - \hat{V}]$ for a wider range of r_λ, the analysis concludes that the introduction of optional calling plans cannot be considered welfare enhancing at the estimated parameter values. The ratio r_λ implied by the estimates is significantly higher (in an statistical sense) than 1.946, the minimum necessary so that at least a group of agents, consumers, prefer the optional calling plans. At the estimated parameter values, the increase in expected welfare by using $\hat{T}$ amounts to 43% of the maximum expected welfare increase of using $\tilde{T}$, which happens at $r_\lambda = 0.230$. Therefore, measured service is not only the optimal tariff *ex post* but also *ex ante*. This second result would be reversed if $\lambda < \lambda_1$, but this means that consumers systematically overestimate their future usage, *i.e.*, the traditional argument in favor of the profitability of optional callings plans that Miravete (2000a) rejects using individual usage information for this same data set.

6 Conclusions

I have developed a theoretical model that explicitly accounts for the role of information asymmetries in the design of optional nonlinear tariffs. The model is solved for two different scenarios: when all consumers are placed on mandatory measured service (tapers or *ex post* tariffs), and when consumers have a choice between tariffs (optional calling plans or *ex ante* tariff). The closed–form solutions of the model provide the theoretical background to test whether asymmetry of information is relevant for the design of telephone tariffs in two cities of Kentucky. Using the structural estimates of the parameters of the distribution of types in Louisville, I show that there is important evidence of both *ex ante* and *ex* *post* asymmetry of information. Furthermore, the distribution of type shocks cannot be considered welfare increasing, and neither consumers or the local telephone monopolist will be better off relative to the scenario where usage is measured. Analyzing the empirical frequency functions of the present usage data, Miravete (2001b, §5) also concludes that there is sufficient evidence to support the opinion that the options offered by SCB were far from optimal.

Alternatively, as mentioned at the beginning of Section 3, it could also be argued that the monopolist would do better by offering a menu of nonlinear optional tariffs. If λ were very similar to λ_1, *i.e.*, if the distribution of type shocks were close to degenerate, the monopolist will effectively screen consumers according to their *ex ante* type by means of a menu of two–part tariffs. The optimal design of these options will reduce the informational rent of consumers, almost exclusively due to their *ex ante* type (expected usage). The linearity of the options will suffice to capture the effect of minor deviations so that the monopolist could profit from the lock–in effect as tariff choices cannot be renegotiated once usage is realized.

But type shocks appear not to have a degenerate distribution, and a monopolist seeking to offer an optimal menu of optional tariffs could evaluate offering fully nonlinear tariff options. The monopolist could introduce further incentives for consumers to self–select accordingly also after the realization of the shock, so that those who receive different shocks while sharing a common *ex ante* tariff do not keep all the informational rents associated to θ_2. However, the lower envelopes of a menu of optional two–part tariffs and a menu of fully nonlinear options are both determined by λ_1. Since the estimate of λ exceeds that of λ_1 markups from standard nonlinear pricing will always be higher at every usage level if options are two–part tariffs or if they lead to further quantity discounts. For fully nonlinear tariff options involving quantity premia, this conclusion could be reversed. If the optimal fully nonlinear options are characterized by higher rates per unit as consumers depart from their expected usage, then the expected difference in profits depicted in Figure 2 should only be considered a lower bound. Similarly, the expected difference in consumer surplus would be an upper bound as the monopolist successfully extracts a larger share of consumers' informational rents. However, given the magnitude of the welfare estimates, it is not very likely that in any event, more complicated optional tariffs will make them socially optimal because additional profit gains will be compensated by further reductions in expected consumer surplus which is already negative, *i.e.*, in favor of the measured service option.

References

Bousquet, A., and M. Ivaldi (1997): "Optimal Pricing of Telephone Usage: An Econometric Implementation." *Information Economics and Policy*, 9, 219–239.

Clay, K., D.S. Sibley, and P. Srinagesh (1992): "Ex Post *vs.* Ex Ante Pricing: Optional Calling Plans and Tapered Tariff." *Journal of Regulatory Economics*, 4, 115–138.

Courty, P. and H. Li (2000): "Sequential Screening." *Review of Economic Studies*, 67, 697-717.

Davidson, R., and J.G. MacKinnon (1993): *Estimation and Inference in Econometrics.* (New York, NY: Oxford University Press).

Donald, S.G., and H.J. Paarsch (1996): "Identification, Estimation, and Testing in Parametric Empirical Models of Auctions within the Independent Private Values Paradigm." *Econometric Theory*, 12, 517-567.

Feinstein, J.S., and F.A. Wolak (1991): "The Econometric Implications of Incentive Compatible Regulation," in G.F. Rhodes (ed.): *Advances in Econometrics*, Vol 9. (Greenwich, CT: JAI Press).

Gouriéroux, C., A. Holly, and A. Monfort (1982): "Likelihood Ratio Test, Wald Test, and Kuhn–Tucker Test in Linear Models with Inequality Constraints on the Regressions Parameters." *Econometrica*, 50, 63–80.

Hausman, J.A. (1985): "The Econometrics on Nonlinear Budget Sets." *Econometrica*, 53, 1255–1282.

Heckman, J. (1979): "Sample Selection Bias as a Specification Error." *Econometrica*, 47, 153–161.

Hobson, M. and R.H. Spady (1988): "The Demand for Local Telephone Service Under Optional Local Measured Service." Bellcore Economics Discussion Paper No. 50.

Ivaldi, M. and D. Martimort (1994): "Competition under Nonlinear Pricing." *Annales d'Economie et de Statistique*, 34, 71–114.

Johnson, N.L., S. Kotz, and N. Balakrishnan (1995): *Continuous Univariate Distributions*, 2nd edition. (New York, NY: John Wiley & Sons).

Kling, J.P. and S.S. van der Ploeg (1990): "Estimating Local Call Elasticities with a Model of Stochastic Class of Service and Usage Choice," in A. de Fontenay, M.H. Shugard, and D.S. Sibley (eds.): *Telecommunications Demand Modelling.* (Amsterdam: North–Holland).

KOTLARSKI, I. (1962): "On Groups of n Independent Random Variables Whose Product Follows the Beta Distribution." *Colloquium Mathematicum*, 9, 325–332.

LAFFONT, J.J., H. OSSARD, AND Q. VUONG (1995): "Econometrics of First Price Auctions." *Econometrica*, 63, 953–980.

LEE, L.F. (1983): "Generalized Econometric Models with Selectivity." *Econometrica*, 51, 507–513.

LEE, L.F., G.S. MADDALA, AND R.P. TROST (1980): "Asymptotic Covariance Matrices of Two–Stage Probit and Two–Stage Tobit Methods for Simultaneous Equations Models with Selectivity." *Econometrica*, 48, 491–503.

LEWIS, T.R., AND D.E.M. SAPPINGTON (1994): "Supplying Information to Facilitate Price Discrimination." *International Economic Review*, 35, 309–327.

MACKIE–MASON, J.K. AND D. LAWSON (1993): "Local Telephone Calling Demand when Customers Face Optimal and Nonlinear Price Schedules." Working Paper. Department of Economics. University of Michigan.

MACKINNON, J. AND H. WHITE (1985): "Some Heteroskedasticity Consistent Covariance Matrix Estimators with Improved Finite Sample Properties." *Journal of Econometrics*, 29, 305–325.

MADDALA, G.S. (1983): *Limited–Dependent and Qualitative Variables in Econometrics.* (Cambridge: Cambridge University Press).

MASKIN, E. AND J. RILEY (1984): "Monopoly with Incomplete Information." *Rand Journal of Economics*, 15, 171–196.

MANSKI, C.F., AND S.R. LERMAN (1977): "The Estimation of Choice Probabilities from Choice Based Samples." *Econometrica*, 45, 1977-1988.

MANSKI, C.F., AND D. MCFADDEN (1981): "Alternative Estimators and Sample Designs for Discrete Choice Analysis," in C.F. Manski and D. McFadden (eds.): *Structural Analysis of Discrete Data with Econometric Applications*. (Cambridge, MA: MIT Press.)

MCDONALD, J.B. AND Y.J. XU (1992) "An Empirical Investigation of the Likelihood Ratio Test when the Boundary Condition is Violated ." *Communications in Statistics*, 21, 879–892.

MCDONALD, J.B. AND Y.J. XU (1995) "A Generalization of the Beta Distribution with Applications." *Journal of Econometrics*, 66, 133–152; plus "Errata," 69, 427–428.

MIRAVETE, E.J. (1996): "Screening Consumers Through Alternative Pricing Mechanisms." *Journal of Regulatory Economics*, 9, 111–132.

MIRAVETE, E.J. (2000): "Choosing the Wrong Calling Plan? Ignorance, Learning, and Risk Aversion." CEPR Discussion Paper No. 2562.

MIRAVETE, E.J. (2001a): "Screening Through Bundling." CARESS Working Paper #01–01. University of Pennsylvania.

MIRAVETE, E.J. (2001b): "Quantity Discounts for Taste–Varying Consumers." CEPR Discussion Paper No. 2699.

Miravete, E.J. (2002): "Preserving Log–Concavity Under Convolution: Comment." *Econometrica*, 70, 1253–1254.

Mitchell, B.M. and I. Vogelsang (1991): *Telecommunications Pricing. Theory and Practice*. (Cambridge, UK: Cambridge University Press).

Paarsch, H.J. (1992): "Deciding between the Common and Private Value Paradigms in Empirical Models of Auctions." *Journal of Econometrics*, 51, 191–215.

Paarsch, H.J. (1997): "Deriving an Estimate of the Optimal Reserve Price: An Application to British Columbian Timber Sales." *Journal of Econometrics*, 78, 333–357.

Park, R.E., B.M. Wetzel, and B.M. Mitchell (1983): "Price Elasticities for Local Telephone Calls." *Econometrica*, 51, 1699–1730.

Panzar, J.C. and D.S. Sibley (1978): "Public Utility Pricing under Risk: The Case of Self–Rationing." *American Economic Review*, 68, 887–895.

Spulber, D.F. (1992): "Optimal Nonlinear Pricing and Contingent Contracts." *International Economic Review*, 33, 747–772.

Taylor, L.D. (1994): *Telecommunications Demand in Theory and Practice*, 2nd edition. (Dordrecht: Kluwer Academic Publishers).

Train, K.E., M. Ben–Akiva, and T. Atherton (1989): "Consumption Patterns and Self–Selecting Tariffs." *The Review of Economics and Statistics*, 50, 62–73.

Train, K.E., D.L. McFadden, and M. Ben–Akiva (1987): "The Demand for Local Telephone Service: A Fully Discrete Model of Residential Calling Patterns and Service Choices." *Rand Journal of Economics*, 18, 109–123.

Wolak, F.A. (1993): "Telecommunications Demand Modeling." *Information Economics and Policy* 5, 179–196.

Wolak, F.A. (1994): "Estimating Regulated Firm Production Functions with Private Information: An Application to California Water Utilities." *Annales d'Economie et de Statistique*, 34, 13–69.

Appendix 1

• Generalized Beta Distribution

A random variable x has a generalized beta distribution with parameters a, $b > 0$, $0 \leq c \leq 1$, $p > 0$, and $q > 0$ if its probability density function can be written as [McDonald and Xu (1995, §2)]:

$$G\beta[x;a,b,c,p,q] = \frac{|\,a\,|\; x^{ap-1}\left[1-(1-c)\left(\frac{x}{b}\right)^{a}\right]^{q-1}}{b^{ap}B(p,q)\left[1+c\left(\frac{x}{b}\right)^{a}\right]^{p+q}}, \quad \text{for} \quad 0 < x^a < \frac{b^a}{1-c}, \qquad (A.1)$$

where $B(p,q) = \Gamma(p)\Gamma(q)/\Gamma(p+q)$ is the Beta function.

• Standard Beta with General Closed Support

The theoretical model makes use of the standard beta distribution $\theta_1 \sim \beta(1, \lambda_1^{-1})$ on $\Theta_1 = [0,1]$, and $\theta \sim \beta(1, \lambda^{-1})$ on $\Theta = [0, \overline{\theta}]$. The probability density functions of these distributions are easily derived from the generalized beta distribution given above. In both cases $a = 1$ and $p = 1$. When $c = 0$ we obtain the family of beta distributions of the first kind. Thus:

$$\beta[x;a=1,b,p=1,q] = f(x) = \frac{q}{b}\left[1-\left(\frac{x}{b}\right)\right]^{q-1}, \quad \text{for} \quad 0 < x < b. \qquad (A.2)$$

Hence, $q = \lambda_1^{-1}$ and $b = 1$ for θ_1 while $q = \lambda^{-1}$ and $b = \overline{\theta} = 2 - \mu_\varepsilon$ for θ. For this family of Burr type XII distributions the cumulative distribution function can be written analytically:

$$F(x) = 1 - \left[1-\left(\frac{x}{b}\right)\right]^{q}. \qquad (A.3)$$

Next observe that the hazard rate is:

$$r(x) = \frac{f(x)}{1-F(x)} = \frac{q}{b-x}, \qquad (A.4)$$

so that it is bounded from below at $(b\lambda)^{-1}$. Since $\lambda < \lambda_1$ from the theory model, it is straightforward to prove that $r(z) > r_1(z)$, $\forall z$. Normalizing the support, we have

$$\frac{1}{\lambda b(1-x/b)} > \frac{1}{\lambda_1(1-x_1)}. \qquad (A.5)$$

Let define $\lambda = k\lambda_1$ for $0 \leq k \leq 1$ and $0 \leq z \leq 1$. Thus, the previous inequality implies:

$$\lambda_1(1-z) > k\lambda_1 b(1-z), \qquad (A.6)$$

which requires that $kb - 1 < 0$. Substituting $b = \overline{\theta} = 2 - \mu_\varepsilon$ into this last inequality, and after some algebra we get:

$$\lambda_1 > \frac{(k-1)^2}{-k}, \tag{A.7}$$

which always holds since $\lambda_1 > 0$. Thus, the internal consistency of the product of the two beta distributed type components of the theory model requires that θ dominates in hazard rate to θ_1.

Finally, some moments of this distribution are used to define the nonlinear regression. I present here these moments for the case of $q = \lambda^{-1}$. Remember that if $q = \lambda_1^{-1}$, then $b = 1$:

$$\mu = \frac{bp}{p+q} = \frac{b\lambda}{1+\lambda}, \tag{A.8}$$

$$\sigma^2 = b^2 pq(p+q)^{-2}(p+q+1)^{-1} = \frac{b^2\lambda^2}{(1+\lambda)^2(1+2\lambda)} = \frac{\mu^2}{(1+2\lambda)}, \tag{A.9}$$

$$\alpha = \mu^2 + \sigma^2 = 2\mu^2 \frac{1+\lambda}{1+2\lambda} = 2b\mu \frac{\lambda}{1+2\lambda}, \tag{A.10}$$

$$\mu^\star = E(\theta \mid \theta < \theta^\star) = \mu - \frac{\lambda + \theta^\star}{1+\lambda} \left[1 - \theta^\star\right]^{\frac{1}{\lambda}}, \tag{A.11}$$

$$\alpha^\star = E(\theta^2 \mid \theta < \theta^\star) = \frac{2b\lambda}{1+2\lambda} E(\theta \mid \theta < \theta^\star). \tag{A.12}$$

• Product of Beta Distributed Variables

To show this relationship I will assume without loss of generality that the variables are distributed on the unit interval. Thus, the central moment of order r of a random variable $\theta_i \sim \beta(p_i, q_i)$, $i = 1, 2$, is:

$$\mu'_r(\theta_i) = \frac{B(p_i + r, q_i)}{B(p_i, q_i)} = \frac{\Gamma(p_i + r)\Gamma(p_i + q_i)}{\Gamma(p_i)\Gamma(p_i + q_i + r)}. \tag{A.13}$$

Then, $\theta = \theta_1\theta_2 \sim \beta(p, q)$ only if its central moments can be written as the product of the central moments of θ_1 and θ_2 [Johnson, Kotz, and Balakrishnan (1995, §25.8)]:

$$\mu'_r(\theta) = \frac{\Gamma(p_1 + r)\Gamma(p_1 + q_1)}{\Gamma(p_1)\Gamma(p_1 + q_1 + r)} \cdot \frac{\Gamma(p_2 + r)\Gamma(p_2 + q_2)}{\Gamma(p_2)\Gamma(p_2 + q_2 + r)} = \frac{\Gamma(p + r)\Gamma(p + q)}{\Gamma(p)\Gamma(p + q + r)}. \tag{A.14}$$

For the requirements of canceling of terms, either $p_1 = p_2 + q_2$ or $p_2 = p_1 + q_1$. That always implies $q = q_1 + q_2$ and $p = p_2$ in the first case or $p = p_1$, in the second. The second condition is used in the theoretical model where $p = p_1 = 1$, $q_1 = \lambda_1^{-1}$, $q = \lambda^{-1}$, and therefore $p_2 = 1 + \lambda_1^{-1}$ and $q_2 = \lambda^{-1} - \lambda_1^{-1}$.

• Likelihood Function

The beta distribution of the theoretical model is not suitable for the empirical analysis since the support of the error term does not necessarily fall in $[0, 1]$. The way to proceed is to transform the index function in such a way that the distribution is a known transformation of the beta and the error term can take any value in $\mathbb{R}$. Since the variable θ_1 has a beta distribution of the first kind on $0 < \theta_1 < b$, the ratio $y = \theta_1/(1-\theta_1)$ has a beta distribution of the second kind defined on $0 < y < \infty$ [Johnson, Kotz, and Balakrishan (1995, §25.7)]. Furthermore, $z = \ln(y)$ has an exponential beta distribution defined on $-\infty < z < \infty$. The probability density function of the exponential generalized beta distribution with parameters $(\delta, \sigma, c, p, q)$ defined on $-\infty < (z-\delta)/\sigma < \ln(1/(1-c)$ is as follows:

$$EG\beta[z; \delta, \sigma, c, p, q] = \frac{\exp\left[p\left(\frac{z-\delta}{\sigma}\right)\right]\left\{1-(1-c)\exp\left[\frac{z-\delta}{\sigma}\right]\right\}^{q-1}}{\mid \sigma \mid B(p,q)\left\{1+c\cdot\exp\left[\frac{z-\delta}{\sigma}\right]\right\}^{p+q}}. \qquad (A.15)$$

Thus, for θ_1, $\delta = \ln(b) = 0$, $\sigma = a^{-1} = 1$, $c = 1$ (that characterizes the family of beta distributions of the second kind), $p=1$, and $q=\lambda_1^{-1}$, so that $B(1, \lambda_1^{-1}) = \lambda_1$, which leads to the following Burr type II distribution [Johnson, Kotz, and Balakrishnan (1995, §12.4.5)]:

$$f(z) = \frac{\exp(z)}{\lambda[1+\exp(z)]^{1+\frac{1}{\lambda}}}, \qquad (A.16)$$

$$F(z) = 1 - [1+\exp(z)]^{-\frac{1}{\lambda}}. \qquad (A.17)$$

Define $y_i = 1$ when consumers subscribe the measured option and $y_i = 0$ otherwise. Therefore $P[y_i = 1] = P[z_1 < \mathbf{X}\boldsymbol{\zeta_1}] = F(\mathbf{X}\boldsymbol{\zeta_1})$ from the above specification of the distribution and the definition of the index function made in the text. Thus, the log–likelihood function to estimate parameters b of the selection rule is given by:

$$\ln L(\boldsymbol{\zeta_1}, \lambda_1; \mathbf{X}) = \sum_{i=1}^{n}\left[y_i \ln\left[1-[1+\exp(\mathbf{X}\boldsymbol{\zeta_1})]^{-\frac{1}{\lambda_1}}\right] - (1-y_i)\frac{1}{\lambda_1}\ln\left[1+\exp(\mathbf{X}\boldsymbol{\zeta_1})\right]\right]. \quad (A.18)$$

Table 1. Descriptive Statistics

		Bowling Green		Louisville (All)		LV: Measured		LV: Flat		TEST	
Variable	Description	Mean	S.D.	Mean	S.D.	Mean	S.D.	Mean	S.D.	T1	T2
BILL	Monthly expenditure in local telephone service	14.1510	(5.433)	19.3023	(4.487)	20.7795	(8.153)	18.7000	(0.000)	-56.92	10.25
INCOME	Monthly income of the household	1.8282	(1.084)	1.4892	(0.922)	1.5362	(0.930)	1.4701	(0.918)	18.53	2.42
DINCOME	Household did not provide income information	0.1592	(0.366)	0.1858	(0.389)	0.1276	(0.334)	0.2095	(0.407)	-3.84	-7.79
HHSIZE	Number of people who live in the household	2.7674	(1.296)	2.4932	(1.462)	2.1003	(1.278)	2.6534	(1.502)	10.81	-13.91
TEENS	Number of teenagers (13–19 years)	0.3755	(0.721)	0.2437	(0.624)	0.1207	(0.449)	0.2939	(0.676)	10.74	-11.17
AGE1	Head of household is between 15 and 34 years old	0.0628	(0.243)	0.0829	(0.276)	0.0755	(0.264)	0.0858	(0.280)	-4.19	-1.30
AGE2	Head of household is between 35 and 54 years old	0.2310	(0.421)	0.2500	(0.433)	0.2861	(0.452)	0.2353	(0.424)	-2.42	3.87
AGE3	Head of household is above 54 years old	0.7061	(0.456)	0.6671	(0.471)	0.6384	(0.481)	0.6789	(0.467)	4.59	-2.88
COLLEGE	Head of household is at least a college graduate	0.2709	(0.445)	0.2192	(0.414)	0.3108	(0.463)	0.1818	(0.386)	6.61	9.89
MARRIED	Head of household is married	0.6721	(0.470)	0.4864	(0.500)	0.4563	(0.498)	0.4986	(0.500)	20.90	-2.87
RETIRED	Head of household is retired	0.1722	(0.378)	0.2582	(0.438)	0.2601	(0.439)	0.2575	(0.437)	-11.45	0.20
BLACK	Head of household is black	0.0670	(0.250)	0.1275	(0.334)	0.0904	(0.287)	0.1426	(0.350)	-11.11	-5.77
CHURCH	Telephone is used for charity and church purposes	0.2116	(0.409)	0.1661	(0.372)	0.1393	(0.346)	0.1770	(0.382)	6.40	-3.57
BENEFITS	Household receives some federal or local benefits	0.2285	(0.420)	0.3187	(0.466)	0.2780	(0.448)	0.3353	(0.472)	-11.07	-4.26
MOVED	Head of household moved in the past five years	0.4684	(0.499)	0.4333	(0.496)	0.4638	(0.499)	0.4209	(0.494)	3.86	2.92
ONLYMALE	Head of household is single and male	0.0430	(0.203)	0.1044	(0.306)	0.1406	(0.348)	0.0896	(0.286)	-12.76	5.21
NOV	Dummy variable for November observations	0.3348	(0.472)	0.3336	(0.472)	0.3375	(0.473)	0.3320	(0.471)	0.15	0.39
DEC	Dummy variable for December observations	0.3324	(0.471)	0.3343	(0.472)	0.3337	(0.472)	0.3345	(0.472)	-0.22	-0.05
MEASURED	Household on measured service this month	1.0000	(0.000)	0.2896	(0.454)	1.0000	(0.000)	0.0000	(0.000)	116.93	
Observations		6,445		5,576		1,615		3,961			

Mean and standard deviations of demographics for the fall of 1986. The column "TEST" shows the test of differences of means for each variable. "T1" compares the samples of Bowling Green and Louisville while "T2" compares the measured and flat options in Louisville. Income is measured in thousand of 1986 dollars.

Table 2. Bowling Green: Mandatory Measured Service (below tariff cap)

Variable	CAP=1	Asymmetric Info.	Symmetric Info.	Asymmetric Info.	Symmetric Info.
Constant	4.8610 (2.76)	14.3044 (6.70)	13.9547 (22.46)	17.8128 (8.96)	15.4512 (40.21)
INCOME	-0.8251 (3.65)	-0.8762 (1.39)	0.0779 (1.03)	-1.1680 (1.91)	0.0421 (0.71)
HHSIZE	1.1719 (6.58)	1.9502 (5.48)	0.4367 (5.83)	1.5228 (4.34)	0.2703 (5.80)
TEENS	2.7083 (8.83)	0.4397 (0.56)	0.1365 (1.04)	0.3498 (0.44)	0.1103 (1.07)
DINCOME	1.3137 (3.32)	0.1039 (0.64)	-0.0663 (0.48)	-0.0193 (0.16)	-0.0585 (0.48)
INCOME*INCOME		0.0880 (1.79)		0.1194 (2.53)	
HHSIZE*HHSIZE		-0.0689 (2.56)		-0.0115 (0.48)	
TEENS*TEENS		-0.1236 (4.55)		-0.1150 (5.08)	
INCOME*HHSIZE		-0.1397 (2.85)		-0.1736 (3.71)	
INCOME*TEENS		0.0238 (0.21)		-0.0045 (0.04)	
HHSIZE*TEENS		0.0496 (0.86)		0.0711 (1.31)	
AGE1	0.1149 (0.15)	0.2341 (1.39)	0.2715 (1.61)		
AGE3	0.6048 (1.57)	0.3673 (3.09)	0.3594 (3.18)		
COLLEGE	-0.7580 (2.03)	-0.1927 (1.63)	-0.0697 (0.67)		
MARRIED	-1.4694 (3.73)	-0.6414 (3.70)	-0.3744 (2.74)		
RETIRED	-1.6107 (3.46)	-0.2939 (1.51)	-0.1912 (1.13)		
BLACK	4.0078 (5.47)	0.2431 (0.81)	-0.0625 (0.25)		
CHURCH	0.5696 (1.66)	0.2228 (1.97)	0.1339 (1.26)		
BENEFITS	0.9084 (2.14)	0.3430 (2.28)	0.2519 (1.82)		
MOVED	-0.7420 (2.46)	-0.5062 (4.76)	-0.4117 (4.19)		
ONLYMALE	-3.3906 (2.88)	-0.0498 (0.12)	0.4956 (1.55)		
NOV	-0.5160 (1.58)	0.0419 (0.41)	0.1049 (1.07)		
DEC	0.0197 (0.06)	0.3981 (4.23)	0.3938 (4.17)		
IMR		0.7683 (1.87)	1.3782 (4.73)	1.6753 (3.78)	1.7339 (14.70)
λ	29.2777 (23.44)				
Obs.	6,445	5,892	5,892	5,892	5,892
R^2		0.158	0.149	0.144	0.149
log-L	-1611.662	-14663.652	-14695.265	-14710.653	-14740.762

The dependent variable in the first equation equals one when total bill payments exceed \$21.50, and it is the total monthly telephone bill for local service in the rest of columns. Income is measured in logarithm. IMR denotes the Inverse Mill's Ratio for correction of selectivity bias. The selection equation is estimated by maximum likelihood asuming that the error term follows an exponential generalized beta distribution while usage equations are estimated by OLS. Absolute heteroscedastic–consistent t–statistics are shown between parentheses.

Table 3. Louisville: Optional Measured Service

Variable	FLAT=1	Asymmetric Info.	Symmetric Info.	Asymmetric Info.	Symmetric Info.
Constant	2.3358 (2.87)	37.6718 (3.06)	23.4476 (8.40)	45.2317 (3.82)	22.7048 (11.34)
INCOME	-0.1506 (1.37)	-7.6125 (2.07)	-0.1451 (0.36)	-7.8496 (2.13)	-0.0488 (0.14)
HHSIZE	0.3811 (4.92)	4.8252 (2.09)	0.8913 (1.99)	4.9178 (2.16)	0.7704 (3.04)
TEENS	0.2466 (1.32)	-16.4190 (1.96)	3.2388 (4.17)	-16.5483 (1.85)	3.4583 (4.54)
DINCOME	0.7738 (3.38)	3.4302 (1.76)	-0.4554 (0.36)	-0.7591 (0.82)	-1.0021 (1.19)
INCOME*INCOME		0.4730 (1.61)		0.6184 (2.14)	
HHSIZE*HHSIZE		-0.3039 (2.95)		-0.0916 (1.36)	
TEENS*TEENS		-0.5704 (0.58)		-0.0143 (0.01)	
INCOME*HHSIZE		0.0015 (0.00)		-0.4733 (1.48)	
INCOME*TEENS		2.9018 (2.38)		2.8020 (2.16)	
HHSIZE*TEENS		0.0749 (0.11)		-0.0424 (0.06)	
AGE1	0.3429 (1.13)	1.2727 (1.20)	-0.9234 (1.06)		
AGE3	0.1038 (0.53)	0.7373 (1.21)	0.0669 (0.12)		
COLLEGE	-0.5114 (3.11)	-3.8068 (2.83)	-0.2364 (0.27)		
MARRIED	-0.1402 (0.74)	-1.9219 (3.14)	-0.6983 (1.38)		
RETIRED	-0.1518 (0.68)	-1.7582 (2.32)	-1.2923 (1.88)		
BLACK	0.3768 (1.46)	3.8362 (3.64)	1.9671 (2.22)		
CHURCH	0.2754 (1.32)	1.5169 (2.12)	0.0734 (0.12)		
BENEFITS	0.2670 (1.20)	1.3433 (1.56)	0.1476 (0.20)		
MOVED	-0.0077 (0.05)	0.3832 (0.84)	0.3521 (0.76)		
ONLYMALE	0.0053 (0.02)	-1.4660 (3.10)	-1.4656 (3.08)		
NOV	-0.0423 (0.24)	-1.4973 (3.13)	-1.2546 (2.66)		
DEC	-0.0247 (0.14)	-0.1974 (0.41)	-0.0413 (0.09)		
IMR		-22.9902 (1.75)	11.9618 (1.60)	13.4512 (4.27)	14.7086 (5.22)
λ_1	1.0353 (36.35)				
Obs.	5,576	1,615	1,615	1,615	1,615
R^2		0.191	0.168	0.164	0.149
log-L	-1671.010	-5508.791	-5531.264	-5535.656	-5549.525

The dependent variable in the first column equals one when consumers subscribe the flat rate option, and it is the total monthly telephone bill for local service in all the other columns. Income is measured in logarithm. IMR denotes the Inverse Mill's Ratio for correction of selectivity bias. The selection equation is estimated by weighted maximum likelihood assuming that the error term follows an exponential generalized beta distribution while usage equations are estimated by OLS. Absolute heteroscedastic–consistent t–statistics are shown between parentheses.

Table 4. Marginal Effects

	Asymmetric Info.		Symmetric Info.	
	Dummies	No Dumm.	Dummies	No Dumm.
Bowling Green				
INCOME	0.0363	0.1014	0.0779	0.0421
	(0.36)	(1.34)	(1.03)	(0.71)
HHSIZE	0.5789	0.2210	0.4367	0.2703
	(5.26)	(4.35)	(5.83)	(5.80)
TEENS	0.6664	0.4339	0.1365	0.1103
	(3.11)	(3.12)	(1.04)	(1.07)
Louisville				
INCOME	-0.5306	0.2902	-0.1451	-0.0488
	(1.05)	(0.74)	(0.36)	(0.14)
HHSIZE	3.5679	1.1622	0.8913	0.7704
	(3.87)	(3.45)	(1.99)	(3.04)
TEENS	4.2378	3.2871	3.2388	3.4583
	(3.04)	(2.55)	(4.17)	(4.54)

Absolute heteroscedastic–consistent t–statistics are displayed between parentheses.

Table 5. J–Tests

	Asymmetric Info.		Symmetric Info.	
Alternative	H_1 *vs.* H'_C	H_2 *vs.* H'_C	H_1 *vs.* H'_C	H_2 *vs.* H'_C
$\log(w_i)$	3.44 3.11 1.64 0.79	0.05 0.18 3.03 2.07	6.90 4.62 4.83 2.39	1.50 0.79 1.21 0.19
$\exp(w_i)$	3.19 4.01 1.88 0.77	8.03 6.96 43.37 0.00	5.37 5.82 1.04 0.95	6.23 5.84 4.14 5.48
$\exp(-w_i)$	5.51 5.04 4.58 4.36	3.26 4.45 3.17 3.09	6.05 4.39 4.10 2.47	5.07 2.49 3.29 2.77
$\sqrt{(w_i)}$	3.03 2.88 1.94 0.89	1.25 1.32 1.13 0.79	6.80 4.33 4.72 2.36	3.92 1.97 2.83 0.96

Absolute t–ratios. First columns of each alternative corresponds to the model with demographic dummy variables, and the second to the model without dummy variables. Similarly, each first row reports the results of the tests for Bowling Green and the second for Louisville.

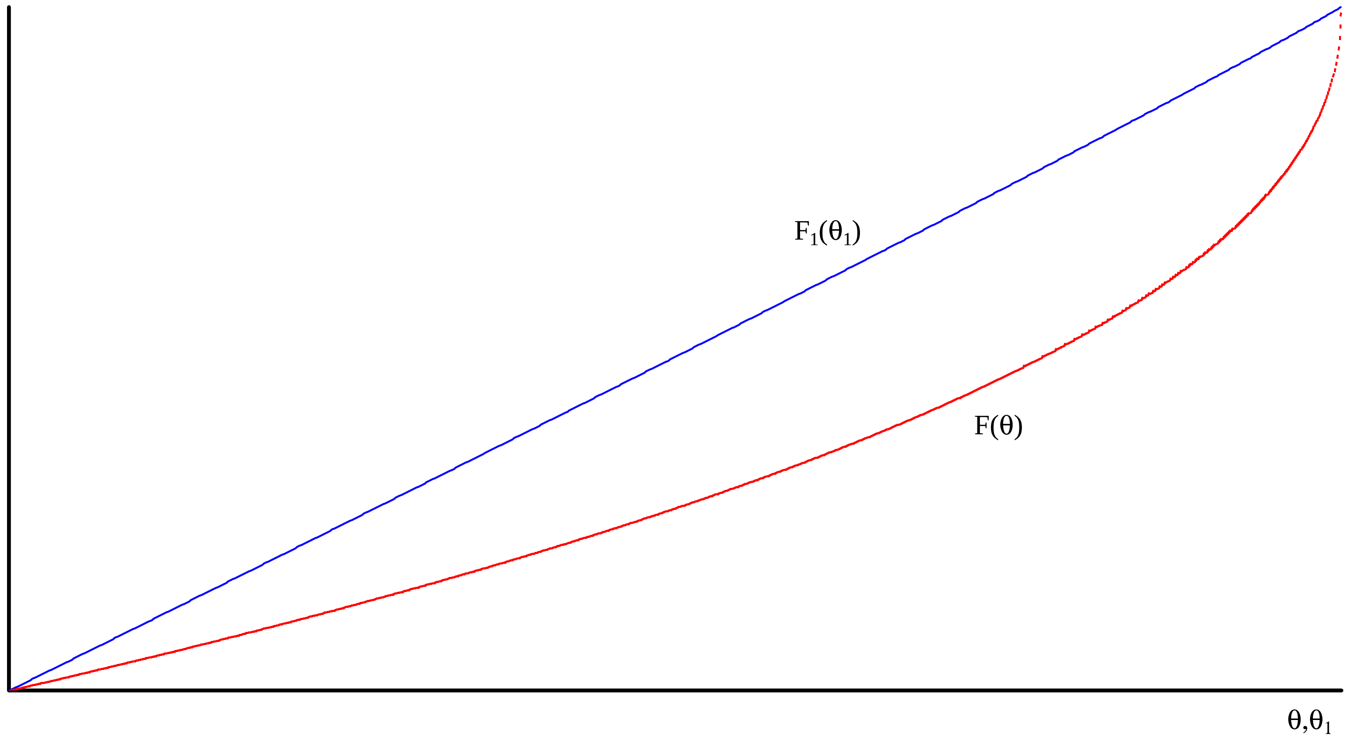

Figure 1: Distribution Functions (Louisville)

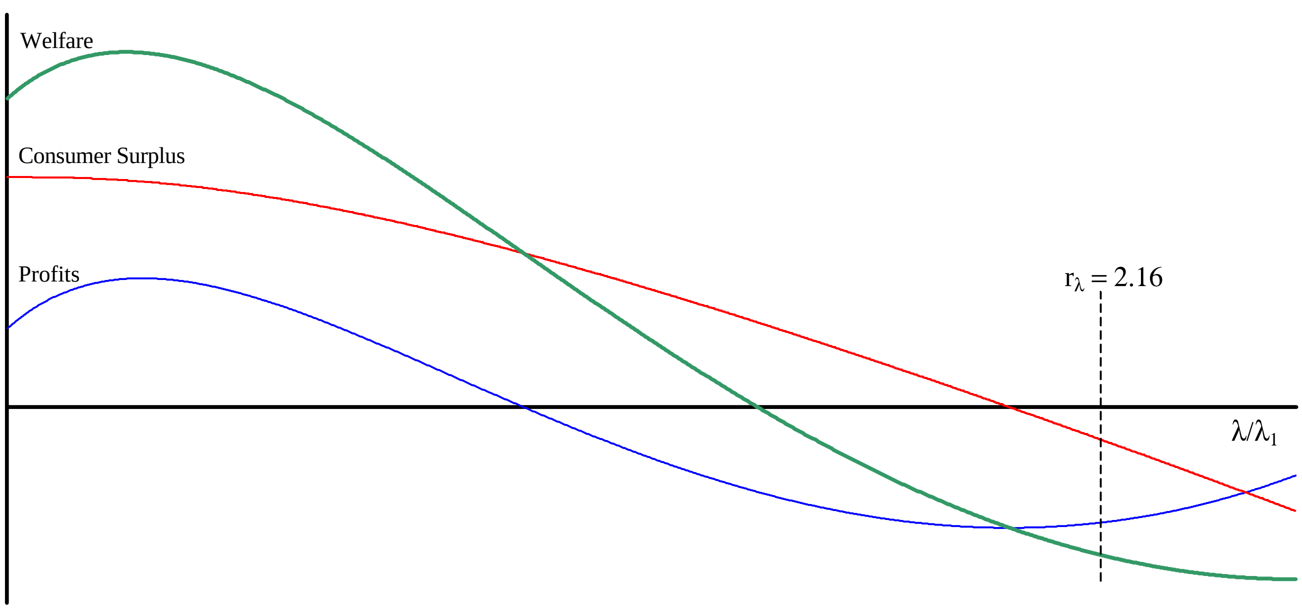

Figure 2: *Ex Ante* Minus *Ex Post* Difference of Welfare Components (Louisville)