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ABSTRACT

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

This Paper studies the theoretical and econometric implications of agents' uncertainty about their future consumption when a monopolist offers them either a unique, mandatory non-linear tariff, or a choice in advance among a menu of optional two-part tariffs. In this model, agents' uncertainty is resolved through individual and privately known shocks on 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 non-linear schedule. The theoretical implications of the model are used to evaluate the tariff experiment run by South Central Bell in two cities of 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. Both expected welfare and profits fail to increase with the introduction of optional tariffs for the estimated value of the parameters.

JEL Classification: D42, D82 and L96
Keywords: asymmetric information, optional tariffs and type shocks

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NON-TECHNICAL SUMMARY

Over the last decade, the use of optional calling plans for pricing telephone services has grown in popularity in the United States. 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 unmeasured local service, which was available at zero marginal charge despite a low but generally positive marginal cost of service. Today, optional calling plans are not restricted to local service, and long distance carriers have intensively used them as marketing strategies in a competitive environment since the divestiture of AT&T.

The critical feature that distinguishes optional calling plans from standard non-linear 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 in use. At the beginning of the billing period, consumers choose the tariff plan that the local monopolist will apply for pricing their future consumption. Later, given their tariff choice, they decide on telephone usage. The standard theory of non-linear 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. This makes the marginal tariff an endogenous variable in the econometric estimation but without explicitly modelling the choice of that variable.

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, in many other cases, the total amount paid under any of the options is so low relative to consumers' income that risk aversion is unlikely to be important. 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 *à la carte* menu. In all these cases, the option chosen *ex ante* need not 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 if they had 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 a 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.

In this Paper I first 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, I show that under this additional source of uncertainty for consumers, the monopolist may discriminate among them by offering a menu of optional calling plans. Finally, I 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. The stochastic structure of the empirical model differs from standard non-linear pricing because of the two-stage decision process of optional calling plans. This feature of the empirical specification enables me to account for the effect of different sources of asymmetric information and the importance of consumers' uncertainty for the optimal design of optional calling plans.

I analyse the demand for residential local telephone service in two cities of Kentucky in the fall of 1986. The data is suitable for testing the two different stochastic structures of the model. In Bowling Green all customers were placed on mandatory measured service and a standard non-linear pricing model is estimated. The choice of tariff and usage are simultaneous decisions and estimation of monthly expenditures on local telephone service incorporates the distribution of the *ex post* asymmetric information parameter. In Louisville, consumers chose between a flat tariff and local measured service. In this case 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 distributions of the *ex ante* type and the type shock are also characterized.

Results show that income has a significant negative effect while the size of the household has a positive effect on usage. Local telephone demand is also significantly different between optional and mandatory measured service and between optional measured and optional flat tariffs. Differences in local telephone demand exceed what differences in demographic characteristics of each local exchange could explain. There is also some evidence that commonly observed strong effects of teenagers on demand might be due to misspecification of the econometric model and lack of treatment of asymmetry of information.

I also find that the hypothesis of symmetric information is always rejected, both under mandatory measured and optional two-part tariffs. Finally, the model provides us with an appropriate framework to evaluate the welfare effects of the introduction of optional calling plans. Using the estimates of the structural parameters of the model I conclude that the optional non-linear tariff offered in Louisville failed to increase expected profits and welfare so that consumers also suffered from its application.

1 Introduction

Over the last decade, the use of optional calling plans for pricing telephone services has grown in popularity in the United States. 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 despite a low but generally positive marginal cost of service [Martins–Filho and Mayo (1993), Shin and Ying (1992)]. Today, optional calling plans are not restricted to local service, and interexchange carriers have intensively used them 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 in use. At the beginning of the billing period, consumers choose the tariff plan that the local monopolist will apply for 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. This makes the marginal tariff an endogenous variable in the econometric estimation but without explicitly modeling the choice of that variable.

The objective of this paper is threefold. First, I 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, I show that under this additional source of uncertainty for consumers, the monopolist may discriminate among them by offering a menu of optional calling plans. Finally, I 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. The stochastic structure of the empirical model differs from standard nonlinear pricing because of the two–stage decision process of optional calling plans. This feature of the empirical specification enables me to account for the effect of different sources of asymmetric information and the importance of consumers' uncertainty for the optimal design of optional calling plans.

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, in many other cases, the total amount paid under any of the options is so low relative to consumers' income that risk aversion is unlikely to be important. 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 *à la carte* menu. 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 if they had 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]

This paper assumes risk neutrality towards monthly bill variation, and suggests 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. The consumer only knows her *ex–ante* type when she chooses among tariff options. The privately known *ex–ante* type therefore determines the choice among optional calling plans. 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 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 where 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).[2] Another result of this paper is to show that both nonlinear tariffs coincide if there is not significant difference between the *ex–ante* and the *ex–post* type. The empirical evidence presented in this paper shows that the introduction of optional tariffs increases the local carrier's expected profits. This result is consistent with the proliferation of optional tariffs across telecommunications services since deregulation.

I analyze the demand for residential local telephone service in two cities of Kentucky in the fall of 1986. The data is suitable to test the two different stochastic structures of the model. In Bowling Green all customers were placed on mandatory measured service

[1] Miravete (2000) uses the same data of the present paper to show that this common belief is not supporter by the data. The reported evidence also rules out risk aversion as a sensible explanation of the observed tariff choices.

[2] The monopolist decision on the information structure does not constrain consumers to use that particular information set to decide on consumption [Lewis and Sappington (1994)], but it affects consumers' purchases through the *ex–ante* choice of tariff since that defines their *ex–post* nonlinear budget sets.

and a standard nonlinear pricing model is estimated. The choice of tariff and usage are simultaneous decisions and estimation of monthly expenditures on local telephone service incorporates the distribution of the *ex–post* asymmetric information parameter. In Louisville consumers chose between a flat tariff and local measured service. In this case 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 distributions of the *ex–ante* type and the type shock are also characterized.

Regarding the estimation of local telephone demand, my empirical results show that income has a significant negative effect while the size of the household has a positive effect on usage. Local telephone demand is also significantly different between optional and mandatory measured service and between optional measured and optional flat tariffs. Differences in local telephone demand exceeds what differences in demographic characteristics of each local exchange could explain. There is also some evidence that commonly observed strong effects of teenagers on demand might be due to misspecification of the econometric model and lack of treatment of asymmetry of information.

The model is carefully constructed to provide closed form solutions for both pricing problems. Closed form solution of the theoretical model identifies 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, *i.e.*, the asymmetry of information is present both *ex–ante* and *ex–post*. This result calls into question empirical work on telecommunications demand where the asymmetry of information between consumers and the monopolist has been neglected. Finally, the model provides with an appropriate framework to evaluate the welfare effects of the introduction of optional calling plans. Using the estimates of the structural parameters of the model I conclude that the optional nonlinear tariff offered in Louisville failed to increase expected profits and welfare so that consumers also suffered from its application.

The theoretical literature on optional tariffs is not extensive. 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. They assume that regularity conditions hold so that the options of a menu of linear *ex–ante* tariffs are self–selecting. This result is proved by Miravete (1996) working 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 (1992a, 1992b) 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. As for the empirical issues addressed by this paper, the empirical auction literature [Donald and Paarsch (1996); Laffont, Ossard, and Vuong (1995); Paarsch (1997)] is most relevant as it explicitly addresses the identification of asymmetric information effects through functional form assumptions on the distribution of types. Very few empirical papers have explicitly incorporated asymmetries of information in estimation of demand [Feinstein and Wolak (1991), Ivaldi and Martimort (1994), and Wolak (1996)], and none deal with final consumer goods. The closest to my model, among the many empirical studies on telecommunications demand, is the work of Hobson and Spady (1988). They obtain an estimate of the distribution of an asymmetric information parameter, but this is done without an 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. Section 2 characterizes the optimal, *ex–post* nonlinear tariff for measured telephone service and optional calling plans. Section 3 describes the tariff experiment undertaken by South Central Bell in two cities of Kentucky in 1986 and points out the links between the features of the experiment and the theoretical approach of this paper. Section 4 analyzes statistical differences between the two cities and presents reduced form estimates for monthly payments and tariff choice. In section 5, I present the estimates of the determinants of monthly payments under symmetric and asymmetric information. Section 6 evaluates the welfare effects of introducing optional two–part tariffs in Louisville after estimating the structural parameters of the model. Finally, section 7 concludes.

2 Ex–Ante and Ex–Post Tariffs

Consider the following game. A monopolist without capacity constraints offers customers either the choice among *ex–ante* tariff plans or an *ex–post* tariff. 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 that she chose. In the case of the *ex–post* tariff (Bowling Green), consumers choose simultaneously the usage and the marginal tariff when they know their *ex–post* type. The *ex–ante* type θ_1 is always private information for consumers, while the monopolist only knows its distribution. However, the monopolist and consumers share the same prior on the distribution of the shock θ_2. In this section I show that 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. The particular choices of functional form of demand and distribution of types are made to obtain a close form solution to both problems that later guide the empirical analysis of this paper.

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. For the *ex–ante* pricing problem the argument is less straightforward. 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] I thus 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 of the nonlinear pricing model, I specify a linear demand for telephone service as follows:

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

where θ represents the consumer's *ex–post* type, and θ_0 is assumed to be a large enough parameter 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:

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

I consider a two stage model to address the decision process with optional tariffs. In the first stage consumers are asked to choose among a set of tariffs, and in the second they decide how much to consume of the subscribed service while their payments are billed according to the previously chosen calling plan. As consumers' information sets increase over time, 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. The shock, which is

[3] See Miravete (1999, §3). There are some other minor justifications to follow this approach. Two–part tariffs are commonly used in many industries because of their simplicity and low monitoring costs. Furthermore, Miravete (1999, §6) also estimates that the increase in expected profits from implementing a menu of fully nonlinear options instead of 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 case the average share of telephone expenditure ranges from 1.6% in Bowling Green to 2.8% in Louisville. See Miravete (2000, §3.3) for an explicit test of this hypothesis.

privately known by consumers at the time of purchase, defines the *ex–post* type "around" the *ex–ante* type, meaning that the realized state of the world may induce higher or lower consumption than the conjectured state. For analytical convenience I assume that the *ex–ante* type, *ex–post* type, and the shock are related as follows:[6]

$$\theta = \theta_1 \theta_2. \tag{4}$$

There are some technical issues that I should address before characterizing the optimal *ex–ante* and *ex–post* nonlinear schedules. 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, and without loss of generality, the distribution of the *ex–ante* type is assumed to be a particular specification of the beta distribution:

$$\theta_1 \sim \beta\left[1, \frac{1}{\lambda_1}\right] \quad \text{on } \Theta_1 = [0, 1]; \quad \lambda_1 > 0. \tag{5}$$

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

$$\varepsilon \sim \beta\left[1 + \frac{1}{\lambda_1}, \frac{1}{\lambda} - \frac{1}{\lambda_1}\right] \quad \text{on } 0 \leq \varepsilon \leq 1; \quad 1 + \frac{1}{\lambda_1} > 0, \quad \frac{1}{\lambda} - \frac{1}{\lambda_1} > 0, \tag{6}$$

so that in order to ensure that the shock moves the *ex–post* type around the corresponding *ex–ante* type, I can define the normalized shock as:

$$\theta_2 = 1 + \varepsilon - \mu_\varepsilon, \tag{7}$$

which has the same distribution than ε, but over the unit interval defined by:

$$\Theta_2 = [1 - \mu_\varepsilon, 2 - \mu_\varepsilon], \tag{8}$$

where μ_ε accounts for the mean of ε:

$$\mu_\varepsilon = \frac{\lambda(1 + \lambda_1)}{\lambda_1(1 + \lambda)}. \tag{9}$$

[6] This one–dimensional characterization of the *ex–post* type avoids the complex issues associated with multidimensional types. If *ex–post* types are ordered along one single dimension there is no qualitative difference between the multidimensional and the one–dimensional formulation. See Armstrong (1996, §1), Rochet and Choné (1998), and Wilson (1993, §8.4). However, given the present specification of demand (2), the inverse demand function $P(x, \theta)$ is linear in θ. For this case, Rochet (1985) has shown that the generalized single crossing property [McAfee and McMillan (1988)] holds. The present formulation enables me to characterize the optimal nonlinear *ex–ante* and *ex–post* tariffs because it ensures that the single crossing property (SCP) holds both *ex–ante*, with respect to θ_1, and *ex–post*, with respect to θ.

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 beta distributed with distribution function defined by:[7]

$$\theta \sim \beta\left[1, \frac{1}{\lambda}\right] \quad \text{on } \Theta = [0, \overline{\theta}] = [0, 2 - \mu_\varepsilon]; \quad \lambda > 0. \tag{10}$$

The beta is a very flexible distribution that, under the present parameterization for θ and θ_1, allows me to find a closed form solution for the optimal nonlinear tariff problem. The beta distribution of the first kind with parameters 1 and $1/\lambda$, also known as Burr distribution of type XII, can be integrated analytically and is characterized by an inverse hazard rate proportional to λ. Thus, the pricing problems are well defined as long as λ and λ_1 are positive. Observe that (6) 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. 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.[8]

The independence of the distribution of the *ex–ante* type and the shock may also appear restrictive at first sight. It is reasonable to consider that shocks are related to consumers' characteristics, as for example their known *ex–ante* type. However, as shown in Miravete (1999, §4.3), such approach does not easily produce analytical solutions for the present type of models. Thus, I avoid working with conditional distributions for the shock. But the multiplicative structure of equation (4) captures the observed fact that higher telephone demand consumers are also those with higher variability of calling patterns.[9]

An *ex–post* two–part tariff (usually called tapers in telecommunications), consists of a pair of functions $\{\hat{p}(\theta), \hat{A}(\theta)\}$ that assigns a fee and a marginal tariff to each *ex–post* consumer type. The mechanism is incentive compatible if it induces consumers to truthfully reveal their *ex–post* type in the corresponding communication game. The mechanism is called feasible if it is globally incentive compatible. The *ex–ante* tariff $\{\tilde{p}(\theta_1), \tilde{A}(\theta_1)\}$ is defined accordingly with respect to the *ex–ante* type. In the next two subsections I will explicitly solve the optimal *ex–ante* and *ex–post* mechanisms.

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[7] See Jambunathan (1954), Kotlarski (1962), and the particular derivation of Appendix 1.

[8] 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. If $\lambda \to 0$, the distribution of θ becomes degenerate at 0 and $\mu_\varepsilon \to 0$. If $\lambda_1 \to 0$, 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 if $\lambda \to \infty$, $\mu_\varepsilon \to 1$. Finally, if $\lambda \to \lambda_1$, $\mu_\varepsilon \to 1$. In all cases, the support of the distributions of θ_1 and θ_2 is 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 $\theta_2 = 1$.

[9] This observation may cause the heteroscedastic errors of the usage equations that I find at the estimation stage. The theoretical model is thus well suited to deal with this empirical feature of the data.

2.1 Ex–Post Nonlinear Tariff: Bowling Green

In this section the optimal *ex–post* nonlinear tariff is derived, as are some of its properties. 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, this is not any "choice of tariff plan" in the sense explained before. The choice of marginal tariff is made when consumers' uncertainty has been resolved. The problem is therefore standard. The monopolist's optimal design of the *ex–post* tariff may be solved using the Revelation Principle. 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 minimizing consumers' informational rents. The usage/tariff choice decision may be written as the following communication game:

$$\theta \in \arg\max_{\theta'} \left[\frac{(\theta_0 + \theta - p(\theta'))^2}{2b} - A(\theta') \right]. \tag{11}$$

The necessary and sufficient conditions for this mechanism to be IC are:[10]

$$A'(\theta) = -\frac{\theta_0 + \theta - p(\theta)}{b} p'(\theta), \tag{12}$$

$$p'(\theta) \leq 0. \tag{13}$$

Next, define the consumer surplus for *ex–post* type θ as:

$$V(\theta) = \frac{(\theta_0 + \theta - p(\theta))^2}{2b} - A(\theta), \tag{14}$$

then, using the IC condition, the Envelope Theorem ensures that:

$$V'(\theta) = \frac{\partial V(p(\theta), A(\theta), \theta)}{\partial \theta} = \frac{\theta_0 + \theta - p(\theta)}{b}. \tag{15}$$

The monopolist uses a constant return to scale technology. His cost function is given by:

$$c(x) = K + cx, \tag{16}$$

which is intended to capture all costs of a telephone firm and seems adequate for an industry characterized by very high fixed costs and very low marginal costs. The cost

[10] Observe that the necessary condition may be written as $-(\theta_0 + \theta - p(\theta'))p'(\theta')/b - A'(\theta') = 0$ at $\theta' = \theta$. Totally differentiating this expression with respect to θ and θ' leads to the condition that $-p'(\theta)/b \geq 0$ ensures the consumer's problem to be concave. Since $b > 0$, it follows that $p'(\theta) \leq 0$ is sufficient for equation (12) to characterize an IC tariff schedule.

function is defined on usage which is an aggregate of duration, distance, peak periods, and setup charges.

Given the consumer's optimal tariff/usage choice, the regulated monopolist's problem is to maximize total expected surplus subject to the individual rationality and incentive compatibility constraints. The monopolist's *ex–post* problem may be written as follows:

$$\max_{p(\theta),V(\theta)} \int_{\Theta} \left[A(\theta) + (p(\theta) - c)\frac{\theta_0 + \theta - p(\theta)}{b} - K\right] dF(\theta), \tag{17a}$$

$$\text{s.t.} \quad V(\theta) = \frac{(\theta_0 + \theta - p(\theta))^2}{2b} - A(\theta), \tag{17b}$$

$$V'(\theta) = \frac{\theta_0 + \theta - p(\theta)}{b}, \tag{17c}$$

$$V(0) \geq 0. \tag{17d}$$

The solution for this *ex–post* tariff is:

$$\hat{p}(\theta) = c + \lambda[\overline{\theta} - \theta], \tag{18a}$$

$$\hat{A}(\theta) = \frac{\lambda(1+\lambda)\theta^2}{2b}. \tag{18b}$$

It is assumed that all consumes 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 (18a):

$$\theta_0 = c + \lambda\overline{\theta}, \tag{19}$$

which has already been used in deriving (18b). Thus, $\hat{A}(0) = 0$ and $V(0) = 0$.

As shown in Appendix 2, the outcome functions of the *ex–post* mechanism $\{\hat{p}(\theta), \hat{A}(\theta)\}$ are directly obtained solving the variational problem (17). The conditions embodied in (18) suffice to characterize the optimal nonlinear tariff. This result is presented here as a proposition.

PROPOSITION 1: *Necessary conditions to solve problem* (17) *suffice to characterize the monopolist's maximum expected profit as long as* $b > 0$.

PROOF: For the *ex–post* necessary conditions to be sufficient, it must be the case that the objective and constraint functions are concave in p and V, given that $V(\theta)$ and the Lagrange multiplier $\delta(\theta)$ are continuous, and $\delta(\theta) \geq 0$ [Kamien and Schwartz (1991, §II.3)]. It is straightforward to show that the Hessians of these functions are always singular matrices. Therefore, it suffices that the second derivative of the objective and constraint functions with respect to p be negative. The concavity of the constraint is ensured by the linearity of the demand in p. Also because of linearity of demand, the concavity of the objective function only requires that the *ex–post* SCP holds: $V_{p\theta} = 1/b > 0$. ∎

Given the present formulation I can also write in closed form the optimal purchase for a consumer of type θ who is offered the *ex–post* tariff (18):

$$\hat{x}(\theta) = \frac{\theta_0 + \theta - \hat{p}(\theta)}{b} = \frac{(1+\lambda)\theta}{b}. \tag{20}$$

Then, the following results may be proved (see Appendix 2):

Proposition 2: *The solution to the ex–post pricing problem has the following properties:*

(a) 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,

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

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

(d) The ex–post tariff may be implemented by a continuum of self–selecting two–part tariffs if $\lambda > 0$.

The fact that the inverse hazard rate of the suggested beta distribution is linear in λ is a key feature that allows me to solve the model in closed form. As equations (18) – (20) show, the marginal tariff and optimal purchase are linear in θ while the fixed fee and the optimal *ex–post* tariff, $\hat{T}(\theta) = \hat{A}(\theta) + \hat{p}(\theta)\hat{x}(\theta)$, 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 constitute a separating equilibria (with respect to *ex–post* types).

Since the value of λ is closely related to the spread of the distribution of θ, I can use this statistic to measure the magnitude of the asymmetry of information. A very low value of λ implies that the distribution of θ is almost degenerate, so that consumers' unobservable characteristics are not very different, and therefore not significant for pricing decisions. A two–part tariff will be optimal if $K > 0$ in this case. However, an increase in the value of λ means that the monopolist is more uncertain about 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. As a consequence, the *ex–post* tariff will include more important price distortions for higher values of λ. This result is formally proved in the following proposition:[11]

[11] If $\lambda > \lambda'$, and $T(x, \lambda)$ denotes the optimal outlay schedule as derived in (13) – (15) then, $T(x, \lambda)$ is more concave than $T(x, \lambda')$ because $T(x, \lambda) - T(x, \lambda')$ is a concave function of x. This follows directly from part b) of Proposition 3, and is an application of the pricing solutions obtained with two distributions that can be ordered with respect to their hazard rates. See Maskin and Riley (1984, §4).

PROPOSITION 3: *When* $\theta \sim \beta[1, \lambda^{-1}]$ *on* $[0, \overline{\theta}]$*, the ex–post tariff is such that* $\forall \theta < \overline{\theta}$*:*
(a) The price margin is increasing in λ*,*
(b) The price per unit is increasing in λ*.*

PROOF: *(a)* Using equation (18), differentiate $(\hat{p}(\theta) - c)/\hat{p}(\theta)$ with respect to λ to obtain $[\overline{\theta} - \theta]c/\hat{p}(\theta)^2 > 0$. *(b)* Using equations (18) – (20) and $\hat{T}(\theta) = \hat{A}(\theta) + \hat{p}(\theta)\hat{x}(\theta)$, differentiate $\hat{T}(\theta)/\hat{x}(\theta)$ with respect to λ to obtain $[\overline{\theta} - \theta]\hat{A}(\theta)/b\hat{x}^2(\theta) > 0$. ∎

2.2 Ex–Ante Nonlinear Tariff: Louisville

In this section the optimal *ex–ante* nonlinear tariff is characterized. I show that the monopolist is still able to screen consumers when they are partially uncertain about their future *ex–post* type. In Louisville consumers are offered a choice of tariff plans at the beginning of the billing period. 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 which implicitly embodies their expected future consumption. 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 the previous section, both the choice of tariff plan and the usage decision may be written as communication games. The consumer's choice of plan solves:

$$\theta_1 \in \arg\max_{\theta_1'} \int_{\Theta_2} \left[\frac{(\theta_0 + \theta_1\theta_2 - p(\theta_1'))^2}{2b} - A(\theta_1') \right] dF_2(\theta_2). \tag{21}$$

The necessary and sufficient conditions for this mechanism to be *ex–ante* IC are:

$$A'(\theta_1) = -\int_{\Theta_2} \left[\frac{\theta_0 + \theta_1\theta_2 - p(\theta_1)}{b} p'(\theta_1) \right] dF_2(\theta_2) = -\frac{\theta_0 + \theta_1 - p(\theta_1)}{b} p'(\theta_1), \tag{22}$$

$$p'(\theta_1) \leq 0. \tag{23}$$

Now denote by $\vartheta(\theta_1)$ the expected rent for a consumer of *ex–ante* type θ_1:

$$\vartheta(\theta_1) = E_2[V(p(\theta_1), A(\theta_1), \theta)] = \int_{\Theta_2} \left[\frac{(\theta_0 + \theta_1\theta_2 - p(\theta_1))^2}{2b} - A(\theta_1) \right] dF_2(\theta_2). \tag{24}$$

In order to incorporate the IC constraint into the monopolist's ex-ante problem I apply the Envelope Theorem to equation (24) to get:

$$E_2 \left[\frac{\theta_0 + \theta_1\theta_2 - p(\theta_1)}{b} \cdot \theta_2 \right] = \frac{\theta_0 + \theta_1(1 + \sigma_2^2) - p(\theta_1)}{b}, \tag{25}$$

where σ_2^2 accounts for the variance of the shock:

$$\sigma_2^2 = \sigma_\varepsilon^2 = \frac{\left(\frac{1}{\lambda} - \frac{1}{\lambda_1}\right)\left(1 + \frac{1}{\lambda_1}\right)}{\left(1 + \frac{1}{\lambda}\right)^2 \left(2 + \frac{1}{\lambda}\right)}. \tag{26}$$

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. For the set of optional tariff plans to be *ex–post* I.C., the consumers optimal purchase solves:

$$\theta_2 \in \arg\max_{\theta_2'} \left[[\theta_0 + \theta_1\theta_2 - p(\theta_1)]x(\theta_1\theta_2') - \frac{1}{2}bx^2(\theta_1\theta_2') - A(\theta_1) \right], \tag{27}$$

which implies:

$$\theta_0 + \theta_1\theta_2 - bx(\theta_1\theta_2) = p(\theta_1), \tag{28}$$

that is, *ex–post*, consumers buy the amount of good that equates their marginal utility to the marginal tariff of the chosen plan. Although for particular realizations of the shock consumers could have been better off *ex–post* if they had chosen a different tariff plan *ex–ante*, they do not violate any IC constraint [Clay, Sibley, and Srinagesh (1992)]. 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, the monopolist's *ex–ante* problem may be written as:

$$\max_{p(\theta_1),\vartheta(\theta_1)} \int_{\Theta_1} \left[A(\theta_1) + (p(\theta_1) - c)\frac{\theta_0 + \theta_1 - p(\theta_1)}{b} - K \right] dF_1(\theta_1), \tag{29a}$$

$$\text{s.t.} \quad \vartheta(\theta_1) = E_2\left[\frac{(\theta_0 + \theta_1\theta_2 - p(\theta_1))^2}{2b}\right] - A(\theta_1), \tag{29b}$$

$$\vartheta'(\theta_1) = \frac{\theta_0 + \theta_1(1 + \sigma_2^2) - p(\theta_1)}{b}, \tag{29c}$$

$$\vartheta(0) \geq 0, \tag{29d}$$

whose solution gives the *ex–ante* menu of two–part tariffs:

$$\tilde{p}(\theta_1) = c + \lambda_1[1 - \theta_1], \tag{30a}$$

$$\tilde{A}(\theta_1) = \frac{\lambda_1(1 + \lambda_1)\theta_1^2}{2b}. \tag{30b}$$

The formal similarity with the *ex–post* case allows me to show straightforwardly that problem (29) is globally concave.[12]

PROPOSITION 4: *Necessary conditions to solve problem* (29) *suffice to characterize the monopolist's maximum expected profit as long as* $b > 0$.

Because of the definition of the *ex–post* type in equation (4), and the assumed relationship among the distributions of θ_1, θ_2, and θ, the *ex–ante* mechanism is a standard nonlinear pricing problem defined over consumers' *ex–ante* type θ_1 but it has different implications. In particular for any demand specification $x(p, \theta(\theta_1, \theta_2))$ such that $\partial\theta/\partial\theta_i > 0$ and $x_{\theta_1}(p, \theta) > 0$ (*ex–ante* SCP), the monopolist can screen consumers according to the *ex–ante* types by offering a continuum of self–selecting two–part tariffs. Proposition 5 proves this result within the framework of the present model:

PROPOSITION 5: *The solution to the ex–ante pricing problem has the following properties:*

(a) 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,*

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

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

(d) The mathematical lower envelope of the ex–ante tariff is concave if $\lambda_1 > 0$.

Proposition 5 shows that the monopolist can screen consumers according to their *ex–ante* types by offering a set of exclusive tariff plans. When the monopolist offers a discrete number of plans, the support of the *ex–ante* type distribution $[0, 1]$ may be partitioned into a set of S intervals defined by $S_i = [\tilde{\theta}_{1i-1}, \tilde{\theta}_{1i}]$. All those consumers whose *ex–ante* type lies in a particular interval S_i will choose the corresponding $\tilde{T}_i$ tariff plan.

Proposition 5 also shows that the mathematical lower envelope of the *ex–ante* tariff is concave. However, the meaning of concavity differs from the usual interpretation given in the nonlinear pricing literature because individual demands are stochastic. It means that the monopolist can design a continuum of two–part tariffs from which consumers will self–select according to their *ex–ante* type. But this lower envelope does not represent the *ex–post* optimal combinations of tariff–purchases for each *ex–post* type. The explanation is that each consumer chooses a particular tariff plan at the beginning of the billing period, and when the shock is realized, its effect is to move the consumer's combination of total payments and purchases along the chosen linear tariff instead of the lower envelope. But the concavity of the lower envelope enables the monopolist to discriminate among consumers when they are partially uncertain about their future type by offering a set of *ex–ante* optional, self–selecting, two–part tariff plans. Furthermore, since by definition both the *ex–ante* type and the shock are directly related to the *ex–post* type, countervailing incentives

[12] Note that the non-negativity constraint on demand for this problem, $\theta_0 = c + \lambda_1$, has been imposed to obtain (30*b*).

[Lewis and Sappington (1989)] are not present and pooling is not an equilibrium feature within each option: payments are different for all *ex–post* consumer types with the same *ex–ante* type.

The last proposition has also proved that on average consumers with higher *ex–ante* valuations tend to buy larger quantities of the monopolized good, but the final result is ambiguous because shocks move consumers along their chosen plan and not along the lower envelope. Hence, there is no longer a one–to–one relationship between optimal purchase, marginal tariff and fixed fee. With an *ex–ante* nonlinear tariff schedule it is possible that the same quantity be billed at different prices to different consumers since they may have chosen different tariff plans *ex–ante*. Therefore, consumers who chose different plans because their *ex–ante* type were different, may end up purchasing the same amount if they receive the appropriate shocks. In this sense, the model leads to bunching since consumers with different *ex–ante* types may end up buying the same amount of good. This result is also true even for consumers with different *ex–post* types if they differ also in the *ex–ante* type because they choose different options.

Finally, 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.[13]

As before, I can 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 $\tilde{T}_i$ where $i = 1, \ldots, S$, and where each tariff plan consists of a two–part tariff defined as $\tilde{T}_i(\theta) = \tilde{A}_i(\theta_1) + \tilde{p}_i(\theta_1)\tilde{x}(\theta)$. The optimal consumption of an *ex–post* type θ with *ex–ante* type θ_1 under *ex–ante* tariff billing is:

$$\tilde{x}(\theta) = \tilde{x}(\theta_1, \theta_2) = \frac{(\theta_2 + \lambda_1)\theta_1}{b}. \tag{31}$$

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 $\tilde{T}_i(\theta)$ is quadratic in θ_1. In addition, the concavity of the *ex–ante* outlay schedule is directly related to the degree of asymmetric information relative to the distribution of consumers' *ex–ante* type:

Proposition 6: *When* $\theta_1 \sim \beta[1, \lambda_1^{-1}]$ *on* $[0, 1]$*, the ex–ante tariff is such that* $\forall \theta_1 < \overline{\theta}_1$*:*

[13] It should also be mentioned 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 (1999, §6) shows that the increase of profits from screening consumers' type shocks is very small. But optional two–part tariffs are a deliberate choice to obtain solutions in closed form. The empirical analysis will be able to estimate λ_1 that determines the shape of the mathematical lower envelope of the *ex–ante* tariff, which happens to be common for a menu of optional two–part tariffs and also for fully nonlinear options.

(a) The price margin is increasing in λ_1*,*

(b) The price per unit is increasing in λ_1*.*

Therefore, 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 3 and 6 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. But for the distribution of the shock to be properly defined it must be the case that $\sigma_2^2 \geq 0$, which according to equation (26) requires that $\lambda \leq \lambda_1$. Therefore, the mark–up for any unit sold is larger in the case of the *ex–ante* tariff since the monopolist screens consumers only with respect to the *ex–ante* type.[14]

2.3 Testable Hypotheses

In this section I present testable hypotheses based on structural relationships of the theoretical model. The estimation of monthly payments for local telephone service will include the effect of several demographic and economic characteristics, and it will be possible to evaluate the importance of the asymmetry of information between consumers and the monopolist, and relevance of consumers' uncertainty for the design of optional tariffs.

Many empirical studies of telephone demand concentrate on a few tariff dimensions (calls, peak *vs.* off–peak,...). The present approach avoids this problem. Equation (2) suggests a simple demand function where the effect of several individual characteristics on the demand of an abstract measure of telephone usage may be included. The solutions to problems (17) and (29) provide the link for the estimation of some of these structural parameters.

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 the individuals, 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. The intercepts of consumers' demands are partially explained by their observable characteristics. Let $W = (w_1, \ldots, w_m)$ denote the set of consumers' observable characteristics both for the monopolist and the econometrician. Let $\psi = (\psi_1, \ldots, \psi_m)'$ denote the vector of associated parameters. In general the intercept may be any function

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[14] This result is obtained whenever θ_1 dominates in hazard rate to θ. Appendix 1 shows that this is the case when $\lambda \leq \lambda_1$.

of individual characteristics, but for simplicity, I will use a linear specification.[15] For the case of an *ex–post* tariff this is:

$$\breve{\theta} = \theta + W\psi, \tag{32}$$

so that the distributions of $\breve{\theta}$ and θ are beta distributions but with different supports:

$$\breve{\theta} \sim \beta\left[1, \frac{1}{\lambda}\right] \quad \text{on } [W\psi, \overline{\theta} + W\psi]. \tag{33}$$

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 that is significant for the optimal tariff design may also be tested. If unobserved heterogeneity is not relevant for pricing decisions, the monopolist then maximizes the sum of consumer surplus and profits:

$$\arg\max_x \left[(\theta_0 + \theta)x - \frac{b}{2}x^2 - cx - K\right] = \frac{\theta_0 + \theta - c}{b}, \tag{34}$$

but this is the solution for $\hat{x}(\theta)$ in equation (20) for the incomplete information case when $\lambda = 0$. This is equivalent to the beta distribution being degenerate at $\underline{\theta} = 0$. Following the implications of Propositions 2 and 3, a degenerate distribution of θ implies that consumers do not differ significantly in anything else than demographics that is relevant for the monopolist pricing decision. Therefore, 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 consumers' uncertainty is relevant for the design of optional tariffs. The measure of consumers' uncertainty is the variance of the shock. If the variance of the shock is zero, there is no difference between the tariffs of the *ex–ante* and *ex–post* case. This result is here stated as a theorem:

PROPOSITION 7: *If $\lambda = \lambda_1$ the ex–ante and ex–post tariffs are the same.*

PROOF: If $\lambda = \lambda_1$, the variance of the shock vanishes, $\sigma_2^2 = \sigma_\varepsilon^2 = 0$, and the distribution of θ_2 is degenerate with shock's mean $\mu_2 = 1$. It follows that all differences

[15] 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 5.3 I show that the data is consistent with this redefinition of the type.

in consumers' tastes are completely captured by the distribution of the *ex–ante* type since $F(\theta) = F(\theta_1) = F_1(\theta_1)$. Therefore $V(\theta) = \vartheta(\theta_1)$ and $\{\hat{p}(\theta), \hat{A}(\theta)\} = \{\tilde{p}(\theta_1), \tilde{A}(\theta_1)\}$. ■

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 outlay. 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 7, it suffices to test whether $\lambda = \lambda_1$ (or equivalently $\sigma_2^2 = 0$). If the test rejects this hypothesis, testing $\lambda < \lambda_1$ will confirm whether the hazard rate dominance prediction of the model holds, and consequently the relative ordering of markups, would be supported by the data.

3 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" according to the commission criteria. As a result SCB carried out an extensive experiment in the second half of 1986 in two cities of Kentucky to provide the commission with evidence in favor of extending optional local measured service.

In spring, when all households in Kentucky were on flat rates, SCB collected demographic and economic information for about 5000 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), and payments for a period of three months. The experiment defined two different scenarios. In the Bowling Green exchange all customers were placed on a mandatory local measured service trial from July 1st through December 31st, 1986. In Louisville, over the same period, local telephone customers were offered a choice between local measured or flat rate service. After a three month adjustment period, SCB collected another three months of billing and usage data for residential customers. This is the data used here.

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. 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 measured service with a monthly fee of \$14.02. The measured service tariff in Louisville was more complex than in Bowling Green because it 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 Sunday; 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 had a 100% surcharge at any period of time relative to the corresponding tariff for band A.

Because of KPSC's desire to minimize the impact of the experiment on individual customers, a \$21.50 tariff cap was established in Bowling Green. In contrast, Louisville customers who chose the measured option (with a higher fixed fee than in Bowling Green) were granted an allowance so that the first \$5 of measured service were not billed. The establishment of fixed fees, allowances, and caps means that many customers may have faced zero marginal tariffs.

All these features lead to a multidimensional nonlinear tariffs, so that the present model is suitable for the data from this experiment. The situation in Bowling Green mimics the theoretical model of Section 2.1. Consumers were not given a choice of tariffs, and because of the nonlinearity of the tariff schedule, the marginal tariff is determined by the level of consumption. That is, marginal tariffs depend on telephone usage, *i.e.,* number of calls, duration, and temporal calling patterns according to the measured tariff in Bowling Green. This scenario corresponds to my *ex–post* tariff formulation, and estimates will provide information regarding the importance of asymmetries of information between the local telephone monopolist and consumers. 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 price cap forces me to control for a censored endogenous variable (switching regression in Section 4 and selection model in Section 5).

The Louisville scenario resembles the case of *ex–ante* tariffs in Section 2.2. 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.[16]

Tariff choice in Louisville was not among two two–part tariffs, but rather among 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

[16] For an exhaustive analysis of the reasons driving "mistakes" in the choice of tariff, see Miravete (2000).

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. The generality of the model allows me to account for the role of consumers' uncertainty in the optimal tariff design by a monopolist in addition to measuring the importance of information asymmetry and the effectiveness of the regulatory constraint.

4 Tariff Choice and Telephone Usage

In this section I examine the determinants of local telephone demand and tariff choice without imposing any restrictions of the structural model developed in Section 2. These reduced form estimates do not account for the effect of the asymmetry of information between the local regulated monopolist and telephone users.

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. There are dummy variables that identify the age group of the head of the household; whether the head of the household has a college degree, or if the head of the household is married, retired, black, or single and male; whether the household moved in the last five years, receives any type of benefits, or if the telephone is used for charity or church activities. In addition, the number of individuals and the number of teenagers living in the household, total annual income, and the standard deviation of the number of monthly calls for each household over the months of the experiment are also reported. For Louisville, a dummy variable indicates whether a household was on measured service in a particular month. Two additional dummies control for different calling patterns over time with October as the reference category. Monthly payments and income are measured in 1986 dollars.

Only active users are considered. Households that did not make any calls over the three months when the data was 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. Missing observations are recoded at the average expected income, $19,851. 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).[17]

Table 1 presents descriptive statistics and compare household characteristics in Bowling Green and Louisville, as well as households on measured and flat rate in Louisville.

[17] See Appendix 3 for the construction of a continuous income variable from the reported categorical data.

The last two columns of this table present values of t–statistics to test the null hypothesis that the population means are the same in the corresponding samples. 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 the following sections are consistent with this result.

Table 2 presents several demand estimates. For each city I include a selection equation (whether the monthly bill is above the $21.50 level in Bowling Green, or whether customers prefer the flat tariff option in Louisville), as well as two usage equations, one for each regime. In these usage equations, the endogenous variable is the share of telephone bills over total monthly income. The estimated equations fit reasonably well with adjusted R^2 ranging from 0.78 to 0.99.[18]

Starting with the selection equations, observe that consumers with lower income, larger households and more teenagers reach more frequently the tariff cap of $21.50 in Bowling Green. These variables have exactly the same effect on the probability of choosing the optional flat service in Louisville. Customers aged fifty five or older, black households, or those who receive any kind of social benefits tend to reach the tariff cap, while married couples, retired citizens, those who have recently moved, or single and male households are more likely to consume below the tariff cap in Bowling Green. In Louisville, those with a college degree clearly prefer the measured service, while young households, blacks, and those who receive benefits or use the telephone for charity prefer to remain in the flat tariff option.

[18] The estimation applies the two–stage method devised by Heckman (1979) to correct for selectivity bias. Heteroscedastic–consistent errors are always computed. Choice equations use the covariance matrix of Manski and McFadden (1981, §1.6). For the case of Louisville I also have to 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. In this case 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)]. The same procedure is applied to estimates presented later in the paper.

There are some interesting common patterns to every usage regime in both cities, once I correct for sample selection. For instance, the effect of income on telephone expenditure shares is always negative. The squared income regressor accounts for possible nonlinearities on this variable. The effect of this variable is always positive. Together it means that for most households the share of local telephone services decreases with income, but at a decreasing rate.[19] Because of this nonlinear effect of income on demand, and because of the estimate used to recode missing values, the positive coefficient of DINCOME indicates that most households who did not report their annual income belong to income categories above the sample average. The size of the household has a positive effect only in Bowling Green, but it is remarkable that the presence of teenagers is never significant across cities and usage regimes.

The estimates also provide information on how differences in demographics affect telephone demand. Some common patterns include the age and racial composition of the households. Age appears to have a positive effect on demand. Senior (non–retired) households in Bowling Green have significantly higher demands than the reference group while young households have lower demands than average. Furthermore, this interpretation is consistent with the fact that young customers prefer the flat tariff choice in Louisville. As for racial composition, black households appear to have larger demands than non–black households.[20] The effect is clear in Bowling Green, but in Louisville, the selection effect dominates: black households do not have the largest demand among those customers who select the flat tariff, but most of them prefer this option.

In addition to these common patterns, Table 2 also show that for the case of Bowling Green, households with a college degree, married couples, households that have moved during the last five years, and single and males have lower demand than average for local telephone service independently of whether they are above or below the tariff cap.

5 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.[21] 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)]

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[19] Observe that the effect of income in the selection and usage equations in Bowling Green confirms that local telephone demand (and not only the share of its expenses) decreases with income, or at least for a range of income. This is not a troubling results as we are studying the demand for local telephone service and not that of long distance and/or wireless communications, which are in general positively associated to higher income levels.

[20] This empirical regularity, as well as the negative effect of income on demand for local telephone service has already been documented in another local telephone demand study by Kling and Van Der Ploeg (1990).

[21] See Kling and Van Der Ploeg (1990), Mitchell and Vogelsang (1991, §12.2.2), and Wolak (1993) for overviews.

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) shares some similarities with the present paper, but they account 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.

While dealing with asymmetry of information empirically, I am concerned with the existence of characteristics of consumers that are not observable neither for the monopolist or the econometrician, and which are relevant for consumption decisions, and therefore for second degree price discrimination. Symmetric information does not mean that the monopolist observes all characteristics of consumers so that he can apply first degree price discrimination (personalized pricing), which is ruled out by regulation, or third degree price discrimination (provided that arbitrage among consumers is not allowed, and that the cost of offering several different options is low enough), but that he knows all consumer differences that are relevant for pricing decisions. This is however an unlikely event since there may be many sources of individual heterogeneity. The solution of the theoretical model allows for quantification of the magnitude of this asymmetry of information, through the estimation of λ and λ_1 in a cross–section estimation.

In this section I explicitly account for asymmetries of information between telephone users and the local monopolist *i.e.*, non–observed consumer heterogeneity. I first present the econometric implications of the existence of asymmetric information as modeled in Section 2. Because of the closed form of the solution of the theoretical model and specific properties of the beta and truncated beta distributions, I obtain an econometric specification that, under the symmetric information hypothesis, is a nested version of the model under asymmmetric 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 it is not reasonable, from an econometric point of view, to assume that the choice of tariff and usage decision are simultaneous.

5.1 Econometric Approach

The theoretical model provides me with a closed form solution for the optimal tariff functions under two different regimes. Conditioning on the observable demographics as in (32), these solutions become:

$$\hat{T}(\breve{\theta}) = \hat{A}(\breve{\theta}) + \hat{p}(\breve{\theta})\hat{x}(\breve{\theta}), \tag{35a}$$

$$\tilde{T}(\breve{\theta}) = \tilde{A}(\breve{\theta}_1) + \tilde{p}(\breve{\theta}_1)\tilde{x}(\breve{\theta}). \tag{35b}$$

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 present formulation explains the total consumption pattern conditional on some demographics, whose distribution on the population can be used by the monopolist to design the tariff. The following econometric specification explains the total bill payment depending on two factors: the average consumer type conditional on demographics and the unobservable type:

$$T = E_{\Theta}[\hat{T}(\breve{\theta})] + \theta, \tag{36a}$$

$$T = E_{\Theta_1}[\tilde{T}(\breve{\theta})] + \theta_1. \tag{36b}$$

First consider Bowling Green. Using the properties of the beta distribution and the identification condition $\overline{\theta} = 1$ (in the case of Bowling Green there is no distinction between *ex–ante* and *ex–post* types), it can be shown that the expected monthly payment for those customers below the cap, conditional on individual characteristics W is:

$$\begin{aligned} E_{\Theta}[\hat{T}(\breve{\theta}) \mid \theta \leq \theta^{\star}] = {} & \frac{(1+\lambda)}{b}\mu^{\star}\left[c + \lambda - \frac{\lambda^2}{1+2\lambda}\right] \\ & + \frac{1+\lambda}{b}\left[c+\lambda-\lambda\mu^{\star}\right]\sum_{i=1}^{m}\psi_i w_i \\ & - \frac{\lambda(1+\lambda)}{2b}\sum_{i=1}^{m}\sum_{j=1}^{m}\psi_i w_i \psi_j w_j, \end{aligned} \tag{37}$$

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

$$T = \gamma_0 + \sum_{i=1}^{m}\gamma_i w_i + \sum_{i=1}^{m}\sum_{j=1}^{m}\gamma_{ij} w_i w_j + \theta. \tag{38}$$

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 the theoretical approach. The existence of individual heterogeneity that makes that some households with similar characteristics are above and some below the tariff cap identifies λ:

$$T = \gamma_0 + \sum_{i=1}^{m}\gamma_i w_i + \sum_{i=1}^{m}\sum_{j=1}^{m}\gamma_{ij} w_i w_j + \theta \qquad \text{if } \theta \leq \theta^{\star}, \tag{39a}$$

$$T = 21.50 \qquad \text{if } \theta > \theta^{\star}, \tag{39b}$$

$$I^{\star}(\theta) = \theta^{\star} - \theta \quad ; \quad \theta \sim \beta\left(1, \frac{1}{\lambda}\right) \quad \text{on } [0,1]. \tag{39c}$$

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 W is:

$$
\begin{aligned}
E_\Theta[\hat{T}(\breve{\theta}) \mid \theta_1 \le \theta_1^\star] = \frac{1}{b} &\left[(c+\lambda_1)\mu^\circ - \left(\lambda_1(c+\lambda_1) - \frac{\lambda_1^2(1+\lambda_1)}{1+2\lambda_1}\right)\mu_1^\star\right] \\
&+ \frac{1+\lambda_1}{b}\left[c+\lambda_1-\lambda_1\mu_1^\star\right]\sum_{i=1}^{m}\psi_i w_i \\
&- \frac{\lambda_1(1+\lambda_1)}{2b}\sum_{i=1}^{m}\sum_{j=1}^{m}\psi_i w_i \psi_j w_j,
\end{aligned}
\tag{40}
$$

where $\mu_1^\star = E(\theta_1 \mid \theta_1 \le \theta_1^\star)$, and $\mu^\circ = E(\theta \mid \theta_1 \le \theta_1^\star) = E(\theta \mid \theta \le \theta_1^\star\overline{\theta}_2)$. 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:

$$
T = \gamma_0 + \sum_{i=1}^{m}\gamma_i w_i + \sum_{i=1}^{m}\sum_{j=1}^{m}\gamma_{ij}w_i w_j + \theta \qquad \text{if } \theta_1 \le \theta_1^\star, \tag{41a}
$$

$$
T = 18.70 \qquad \text{if } \theta_1 > \theta_1^\star, \tag{41b}
$$

$$
I^\star(\theta_1) = \theta_1^\star - \theta_1 \quad ; \quad \theta_1 \sim \beta\left(1, \frac{1}{\lambda_1}\right) \quad \text{on } [0,1]. \tag{41c}
$$

Equation (39) is the econometric specification of the model for Bowling Green under asymmetric information and mandatory measured service. A structural implication of the econometric specification (37) 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 (41) for Louisville.[22] The set 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 (39c) and (41c) should be defined

[22] In this latter case, the estimation is slightly more involved because of the existence of a monthly allowance and a choice of tariffs. To control for the \$5 allowance I use the "virtual income" approach developed by Hausman (1985) to estimate demand in the presence of nonlinear budget sets. Monthly income of consumers on optional measured service is increased by \$5 and monthly payment neglects the allowance.

on $\Re$. 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, without loss of generality in the case of Louisville. For Bowling Green just delete the subindices of the θ's and λ's.

Whenever $I^\star(\theta_1) > 0$ in the selection equation ($41c$), 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 the following index function:

$$\theta_1^\star = \Im(X\zeta_1) : \Re^n \to [0,1], \tag{42}$$

where $X \subseteq W$ 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:

$$\begin{aligned} \Pr[I^\star(\theta_1) \geq 0 \mid X] &= \Pr[\theta_1 \leq \Im(X\zeta_1)] \\ &= \Pr\left[\ln\left(z_1 = \frac{\theta_1}{1-\theta_1}\right) \leq \ln\left(\frac{\Im(X\zeta_1)}{1-\Im(X\zeta_1)}\right)\right]. \end{aligned} \tag{43}$$

The advantage of using the monotone transformation z_1 instead of θ_1 is that z_1 is distributed on $\Re$ instead of on a bounded support. Furthermore, 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 close form representation and depends exclusively on the same parameter λ_1 of the theoretical model. Thus, to simplify, and without loss of generality, assume the following specific functional form for the index function:

$$\Im(X\zeta_1) = \frac{\exp(X\zeta_1)}{1+\exp(X\zeta_1)}, \tag{44}$$

so that the selection rule becomes:

$$I^\star(z_1) = X\zeta_1 - z_1 \qquad ; \qquad \frac{\exp(z_1)}{1+\exp(z_1)} \sim \beta\left(1, \frac{1}{\lambda_1}\right). \tag{45}$$

Equations (39) and (41) 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:

$$T = \gamma_0 + \sum_{i=1}^{m} \gamma_i w_i + \sum_{i=1}^{m}\sum_{j=1}^{m} \gamma_{ij} w_i w_j - \sigma_1 \rho \frac{\phi[J(X\zeta_1)]}{F(X\zeta_1)} + \eta, \tag{46}$$

where ρ accounts for the correlation coefficient of disturbance θ_1 and $J(X\zeta)$ [Maddala (1983, §9.4)], so that $E[\eta \mid I = 1, W, Z] = 0$. The probability that $z_1 \leq X\zeta_1$ is denoted by $F(Z\zeta_1)$ which represents the exponential generalized beta probability distribution function with parameters 1 and λ_1^{-1}:

$$F(X\zeta_1) = 1 - [1 + \exp(X\zeta_1)]^{\frac{1}{\lambda_1}} . \tag{47}$$

Finally, $J(\cdot)$ is defined as:

$$J(X\zeta_1) = \Phi^{-1}[F(X\zeta_1)], \tag{48}$$

where $\phi(\cdot)$ in equation (46) and $\Phi(\cdot)$ in equation (48) account respectively for the standard normal density and distribution functions.

5.2 Estimates

To estimate the model 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 only in some econometric specifications of the model.

Tables 3 and 4 present estimates of equations (39) and (41). 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.[23] 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 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. The estimate of λ_1 is not significantly different from 1 and thus, the unobserved heterogeneity parameter can be assumed to

[23] The estimates of the effect of demographics and those of λ or λ_1 are obtained simultaneously by maximum–likelihood. See Appendix 1 where the the likelihood function of this model is presented.

be uniformly distributed. However, the distribution of θ_1 is skewed to the right. This result could be interpreted as the 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 4.

I regress monthly payments on several combination of variables. I estimate four versions of the usage equations. Only two of them include demographics while the other 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.[24] This last result supports the specification of the model in equations (37) and (40) 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 who 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 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. Married and/or retired households, and with college degree have lower demands. Similarly happens for single and male households and 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 5. 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 3 and 4 is non–significant in Bowling Green and positive only under a specification consistent with symmetric information. However, accounting for direct and indirect effects of the specification under asymmetric information, the effect of teenagers is

[24] 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.

not only positive and significant, but also generally much larger than the estimation under the symmetric information assumption. A similar conclusion is obtained for the case of the size of the household. Direct and indirect effects of income remain non–significant.

5.3 Specification Tests

In the previous analysis, results were consistent with the structural implications of the model with asymmetric information (including cross–products of demographics). However, it could 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 in 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 (the uniform distribution as a special case $\lambda = 1$), 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 asymmetric information.

The second issue is that of the redefinition of types in equation (32). Economic theory does not impose any restriction on this matter. I assumed that those household 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 (32) as the first order Taylor's expansion of any general function $\breve{\theta} = \theta(W)$. 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. To see why, suppose that the alternative hypothesis is a quadratic approximation to $\theta(W)$. Then, monthly payments become a fourth degree polynomial in W under the hypothesis of asymmetric information and quadratic redefinition of types. The regression model is only linear in W if the information is symmetric and the 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 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. Therefore, there is no way of distinguishing whether the quadratic definition of types leads to a better specified model than the linear definition.

This result is not very restrictive either. Equation (32) 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 falls is just a particular case of a broader discussion on the right specification of 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. Thus there is no obvious choice. To evaluate the actual linear specification, let consider the following non–nested hypotheses:

$$H_1 : \ \breve{\theta} = \theta + \sum_{i=1}^{m} \psi_i^{(1)} w_i, \tag{49a}$$

$$H_2 : \ \breve{\theta} = \theta + \sum_{i=1}^{m} \psi_i^{(2)} g(w_i). \tag{49b}$$

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 nonnested, the econometric analysis is based on the construction of J–tests [Davidson and MacKinnon (1993, §11.3) and Greene (1993, §7.7)]. The way to proceed is to artificially nest two competing type definition hypotheses:

$$H_C' : \ \breve{\theta} = \theta + (1-\alpha) \sum_{i=1}^{m} \psi_i^{(1)} w_i + \alpha \sum_{i=1}^{k} \psi_i^{(2)} g(w_i). \tag{50}$$

Since this model is not estimable because parameters α, $\psi^{(1)}$, and $\psi^{(2)}$, are not all separately identifiable, the idea is to replace the second summation with a consistent estimate under H_2. Therefore, to test the linear definition *vs.* one of the non–linear definitions, H_1 *vs.* H_C', I first estimate equation (49b) and compute the predicted values of the regression that are then included in the estimation of (50) instead of the second summation to test whether $\alpha = 0$. It is evident that I could also test H_2 against H_C'. In this case the first stage is to obtain a consistent estimate of the first summation under H_1, and then test whether $(1-\alpha) = 0$. Results are presented in Table 6.

Because of the additively separable form of the type redefinition, it is still possible to identify the specifications of the model under symmetric or asymmetric information.

Results are not conclusive when the alternative model is the logarithmic redefinition of types. The linear model is rejected in favor of the logarithmic one in Bowling Green while the logarithmic redefinition is never rejected in favor of the linear one. In Louisville this result is not so obvious. The linear model is the preferred specification whenever cross–product of demographics are included (specification under asymmetric information) while the logarithmic redefinition fits better the data if cross–product of demographics are excluded (specification under symmetric information). 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 of $g(w_i)$ analyzed in Table 6 are intended to shed some light on the robustness of this result. For the case of Bowling Green, the square root of demographics are preferred to the linear redefinition of types under asymmetric information. In all other cases, the linear and the alternative model are rejected. The case of Louisville is again more dependent of the informational framework and the functional form assumed for the redefinition of types. Thus, the linear model is always preferred if the alternative is the exponential; both the linear and the inverse exponential redefinitions are rejected in all cases; and the linear and square root of demographics perform equally well under asymmetric information, or equally wrong under symmetric information. I thus conclude that while there is evidence that in some cases other functional forms may perform better than the linear model,[25] there is no clear indication of which alternative functional redefinition of types will produce better results.

6 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. I 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 (40).

From equation (40) it is obvious that structural parameters are only identifiable up to a scale factor. Therefore, I normalize the slope of demand in the theoretical model to $b = 1$ at this stage.[26] Marginal cost cannot be identified from demand data exclusively and I therefore assume $c = 0$. This assumption is not unreasonable. By middle of the eighties, marginal costs were positive but very small. They were mostly generated at the local switch. After the introduction of digital technology they became more related to

[25] This conclusion should however account for the fact that the J–test is a very low power test.

[26] Parameter b only affects the magnitude of the welfare, but does not change the nature of the results. See Miravete (1996, §4.1). The effect of the shock therefore keeps a one to one relationship with unforeseen changes in monthly usage as defined by bill payments.

the connection than to the duration of the calls. More importantly, all these costs were insignificant compared with the fixed cost of maintaining the local network.

I estimate equation (40) 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 imposes 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. Besides the assumed $b = 1$ and $c = 0$, I 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 (44) and the estimates of the selection equation, ζ_1 (first column of Table 4).

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 $\lambda_1 = 1.0353$ obtained in the previous section.[27] This result has several implications that I now present in the remaining of this section.[28]

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 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 (26)– 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

[27] 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 4.

[28] 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 (1999), 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.

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 $F_1(\theta_1)$ 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.[29] 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 than if they had chosen the flat tariff option instead. However, we should not expect that this event involves too many households because both the distribution of θ is more skewed to the right than that of θ_1. This means that the smaller θ_1 the larger the expected type shock has to be to exceed the threshold value that justifies the *ex–ante* choice of the flat tariff option. As there is a concentration of mass of probability around the higher values of θ, the previous argument implies that the customers who belong to the higher fractiles of $F(\theta)$ are likely to be also 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 3 and Proposition 6, 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. Intuitively, since there is an important concentration of probability for high values of θ, the monopolist has to introduce important distortions to separate 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 the KPSC was to decide *ex–ante* whether they should introduce and/or approve the introduction of optional calling plans. Thus, welfare, and 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.

[29] This systematic underestimation of future usage is documented and analyzed in Miravete (2000) making use of the actual and expected number of weekly calls for each household.

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 of λ and 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, consumers are significantly more diverse *ex–post* than *ex–ante*. Thus, an *ex–ante* tariff introduces 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. Now type shocks are such that consumers become relatively more homogeneous *ex–post* than *ex–ante*. Observe again in Figure 1 that while θ_1 is uniformly distributed, an important mass of probability is concentrated around the highest values of θ under the *ex–post* type distribution. Therefore, the more homogeneous the *ex–post* consumers become relative to their *ex–ante* distribution, the higher is the expected profit difference in favor of $\tilde{T}$ over $\hat{T}$. The argument, as before, is based on the need of a more powerful mechanism to induce self–selection when consumer types are heavily concentrated. 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, 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. For the scenario of the estimated parameters, their expected surplus difference under $\hat{T}$ is only 50% the maximum of their informational rents difference under $\tilde{T}$, which happens at $\lambda = 0$. Consumers seem to prefer $\tilde{T}$ to $\hat{T}$ as long has type shocks do not make them "excessively" similar *ex–post*. When they are relatively heterogeneous, there is no need of a very powerful mechanism to extract their informational rents. For values of the ratio $r_\lambda < 1.946$, high *ex–post* markups leave little or no rent to many consumers. For $r_\lambda > 1.946$ they are considerably more diverse *ex–ante* than *ex–post*. Subscribing an optional calling plan in such scenario only increases the probability of choosing a tariff option that proves to be inadequate for their *ex–post* usage.

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 (1999) rejects using individual usage information for this same data set.

A final remark is needed. It can be argued whether the monopolist could 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 the 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 should evaluate offering fully nonlinear tariff options. The monopolist must introduce further incentives for consumers to self–select accordingly also after the realization of the shock, so that those that 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 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 tariffs involving quantity premia, this conclusion could be reversed. If the optimal fully nonlinear options is characterized by higher rates per unit as consumers depart from their expected usage, then the expected difference in profits depicted in Figure 2 should be considered a lower bound.

Similarly, the expected difference in consumer surplus would be only 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 reductions in expected consumer surplus which is already negative, *i.e.*, in favor of the measured service option.

7 Conclusions

The present work has developed a theoretical model that explicitly accounts for the role of asymmetry of information on 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.

Empirical results suggest that demand for local telephone service depends on the nature of the tariff (mandatory or optional), consumers' choices (optional measured or optional flat rate), and consumption network externalities associated to the size of local interexchanges. The econometric estimation of the model in the case of Louisville does not consider the choice of marginal tariff as exogenous to the analysis. Instead, the theoretical model explicitly address the tariff choice stage providing a beta distributed selection rule for a sample selection model. This econometric procedure characterizes the distribution of consumers' *ex–ante* types as well as the effect of several variables on the demand for local telephone service.

The theoretical model also provides a method to test the significance of asymmetry of information both under mandatory measured and with optional tariffs. The closed form solution of the model leads to different econometric specifications of demand for local telephone service assuming either symmetric or asymmetric information. Estimates reject the null hypothesis of symmetric information. Therefore, failure to explicitly model the asymmetry of information involved in the design of optimal telephone tariffs leads to misspecified econometric estimations of demand for local telephone service.

Finally, 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 degenerate. I show that the introduction of optional calling plans in Louisville was not welfare increasing, and that neither consumers or the local telephone monopolist will be better of relative to the scenario where usage is measured.

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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)]:

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

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:

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

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:

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

and next observe that the hazard rate is:

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

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, to show that:

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

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

$$\lambda_1(1 - z) > k\lambda_1 b(1 - z), \tag{A.6}$$

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 $\Re$. 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 (δ,σ,c,p,q) defined on $-\infty<(z-\delta)/\sigma<\ln(1/(1-c)$ is as follows:

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

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)]:

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

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

Define $y_i=1$ when consumers subscribe the measured option and $y_i=0$ otherwise. Therefore $P[y_i=1]=P[z_1<X\zeta]=F(X\zeta)$ 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:

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

Appendix 2

• **Derivation of the Ex–Post Tariff**

The Hamiltonian of the *ex–post* pricing problem is:

$$H[V,p,\theta]=\left[\frac{(\theta_0+\theta-p(\theta))^2}{2b}+(p(\theta)-c)\frac{\theta_0+\theta-p(\theta)}{b}-K-V(\theta)\right]f(\theta)$$
$$+\delta(\theta)\frac{\theta_0+\theta-p(\theta)}{b}, \tag{A.19}$$

with first order necessary conditions:

$$H_p : (p(\theta) - c)f(\theta) + \delta(\theta) = 0, \tag{A.20a}$$

$$H_V : f(\theta) = \delta'(\theta) \; ; \; \delta(\overline{\theta}) = 0. \tag{A.20b}$$

The transversality condition vanishes at $\overline{\theta}$ since $V'(\theta) = (\theta_0 + \theta - p(\theta))/b > 0 \; \forall \theta$, so that the participation constraint is only binding at $\underline{\theta} = 0$ [Kamien and Schwartz (1991, §II.7)]. Therefore:

$$\delta(\theta) = \int_{\overline{\theta}}^{\theta} f(z)dz = F(\theta) - 1, \tag{A.21}$$

which leads to equation (18a) in the text. The fixed fee is computed from the definition of $V(\theta)$ in equation (14), so that:

$$\hat{A}(\theta) = \frac{(\theta_0 + \theta - \hat{p}(\theta))^2}{2b} - V(0) - \int_0^{\theta} \frac{\theta_0 + z - p(z)}{b} dz. \tag{A.22}$$

• **Proposition 2**

(a) Let $\theta > \theta'$. Incentive compatibility implies:

$$\frac{(\theta_0 + \theta - \hat{p}(\theta))^2}{2b} - \hat{A}(\theta) \geq \frac{(\theta_0 + \theta - \hat{p}(\theta'))^2}{2b} - \hat{A}(\theta'),$$
$$\frac{(\theta_0 + \theta' - \hat{p}(\theta'))^2}{2b} - \hat{A}(\theta') \geq \frac{(\theta_0 + \theta' - \hat{p}(\theta))^2}{2b} - \hat{A}(\theta). \tag{A.23}$$

Adding these two inequalities yields:

$$\begin{aligned}\int_{\theta'}^{\theta} \frac{\theta_0 + z - \hat{p}(\theta)}{b} dz =& \frac{(\theta_0 + \theta - \hat{p}(\theta))^2}{2b} - \frac{(\theta_0 + \theta' - \hat{p}(\theta))^2}{2b} \\ \geq& \left[\frac{(\theta_0 + \theta - \hat{p}(\theta))^2}{2b} - \hat{A}(\theta)\right] - \left[\frac{(\theta_0 + \theta' - \hat{p}(\theta'))^2}{2b} - \hat{A}(\theta')\right] \\ \geq& \frac{(\theta_0 + \theta - \hat{p}(\theta'))^2}{2b} - \frac{(\theta_0 + \theta' - \hat{p}(\theta'))^2}{2b} = \int_{\theta'}^{\theta} \frac{\theta_0 + z - \hat{p}(\theta')}{2b} dz.\end{aligned} \tag{A.24}$$

This inequality together with the single crossing property (SCP), $b > 0$, implies that $\hat{p}(\theta) \leq \hat{p}(\theta')$. Therefore, since $\hat{p}(\theta)$ is monotone, it is almost everywhere continuous and differentiable. Observe that $\hat{p}'(\theta) < 0$, *i.e.,* higher consumer types pay lower marginal tariffs. This result holds globally because of the SCP, and ensures that local maximum of the consumer tariff choice is also a global maximum. For the mechanism to be almost everywhere differentiable, it remains to be proved that the other outcome function, $\hat{A}(\theta)$, is also almost everywhere differentiable. Observe that IC also implies:

$$\frac{(\theta_0+\theta-\hat{p}(\theta))^2-(\theta_0+\theta-\hat{p}(\theta'))^2}{2b} \geq \hat{A}(\theta)–\hat{A}(\theta') \geq \frac{(\theta_0+\theta'-\hat{p}(\theta))^2-(\theta_0+\theta'-\hat{p}(\theta'))^2}{2b}. \quad (A.25)$$

Then, taking limits and using equation (4), it follows:

$$\lim_{\theta'\to\theta} \frac{(\theta_0+\theta-\hat{p}(\theta))^2-(\theta_0+\theta-\hat{p}(\theta'))^2}{2b(\theta-\theta')} = \lim_{\theta'\to\theta} \frac{(\theta_0+\theta'-\hat{p}(\theta))^2-(\theta_0+\theta'-\hat{p}(\theta'))^2}{2b(\theta-\theta')} =$$
$$\lim_{\theta'\to\theta} \frac{(\theta_0+\theta-\hat{p}(\theta))^2-(\theta_0+\theta-\hat{p}(\theta'))^2}{2b(\hat{p}(\theta)-\hat{p}(\theta'))} \frac{\hat{p}(\theta)-\hat{p}(\theta')}{\theta-\theta'} = -\frac{\theta_0+\theta-\hat{p}(\theta)}{b}\hat{p}'(\theta) = \hat{A}'(\theta). \quad (A.26)$$

Therefore higher consumer types pay higher fixed fees.

(b) For an *ex–post* consumer type θ, the optimal purchase is $\hat{x}(\theta) = x(\hat{p}(\theta), \theta)$. Consumers with higher valuations purchase larger amounts of good because if $\lambda > 0$:

$$\hat{x}'(\theta) = \frac{1+\lambda}{b} > 0, \quad (A.27)$$

(c) At $\theta = \overline{\theta}$, $F(\overline{\theta}) = 1$ and the second term of the marginal tariff equation (18a) equals zero, so that $\hat{p}(\overline{\theta}) = c$.

(d) A continuum of self–selecting two–part tariffs is such that each *ex–post* consumer type chooses the one which maximizes her utility. Each linear tariff is the optimal solution for only one *ex–post* consumer type, and therefore equilibrium is completely separating. A sufficient condition for a continuum of two–part tariffs to be self–selecting is that its lower envelope be concave in consumption [Faulhaber and Panzar (1977)]. Provided that $\hat{x}'(\theta) > 0$, it suffices to prove that $\hat{p}'(\theta) < 0$. But $\hat{p}'(\theta) = -\lambda < 0$ when $\lambda > 0$. ∎

• Proof of Proposition 5

Part *(a)* is proved similarly to Proposition 2 by substituting $(\theta_0 + \theta - \hat{p}(\theta'))^2/2b - \hat{A}(\theta')$ for $E_2[(\theta_0 + \theta_1\theta_2 - \tilde{p}(\theta_1'))^2/2b - \tilde{A}(\theta_1')]$ in order to show that $\tilde{p}'(\theta_1) < 0$ and $\tilde{A}'(\theta_1) > 0$. To prove part *b)* observe that the optimal purchase of an *ex–post* consumer type θ with an *ex–ante* type θ_1 defined in (31) is such that:

$$\frac{\partial\tilde{x}}{\partial\theta_1} = \frac{\theta_2+\lambda_1}{b} > 0, \quad (A.28)$$

as long as $\lambda_1 > 0$ because θ_2 is always positive. Part *c)* makes use of the distribution of θ_1 so that $F_1(1) = 1$ and $\tilde{p}(1) = c$. Finally, part *d)* the concavity of the lower envelope is ensured by $\tilde{p}'(\theta_1) = -\lambda_1 < 0$ when $\lambda_1 > 0$. ∎

Appendix 3

- **Description of Variables**

The data includes the following set of variables. Most of them are dummy variables that take value equal to 1 for the indicated case:

AGE1	The age of the household head is between 15 and 34 years.
AGE3	The age of the household head is above 54 years.
BENEFITS	The household receives some benefits such as Food Stamps, Social Security, Federal Rent Assistance, Aid to Families with Dependent Children,...
BILL	Total monthly expenditure in local telephone service.
BLACK	Ethnic group of the household head is black.
CHURCH	Household uses the telephone for charity or church work.
CITY	Household with residence in Louisville.
COLLEGE	The household head is at least a college graduate.
DINCOME	The household did not answer questions about total annual income.
HHSIZE	Number of people who regularly live in the household.
INCOME	Estimated total monthly income of the household.
MARRIED	The household head is married.
MEASURED	The household is on local measured service.
MOVED	The household moved at least once in the past five years.
ONLYMALE	The household head is male and single.
RETIRED	The household head is retired.
SHARE	Percentage of monthly income accounted for telephone expenditure.
TEENS	Number of teenagers (between 13 and 19 years old) living in the household.
NOV	Observation for the month of November 1986.
DEC	Observation for the month of December 1986.

• **Estimation of Income**

Income was reported as a categorical variable. Categories were defined by known income ranges. In order to define a continuous variable that represents the income of households, I assume that income is distributed according to a displaced gamma distribution [Johnson, Kotz, and Balakrishnan (1995, §17)], so that the probability density function is:

$$g(x, \alpha, \beta, \gamma) = \frac{(x-\gamma)^{\alpha-1} \exp\left[-\dfrac{x-\gamma}{\beta}\right]}{\beta^{\alpha}\Gamma(\alpha)} \tag{A.29}$$

for $\alpha > 0$, $\beta > 0$, and $x > \gamma$. As the number of cases in each category (once missing values have been deleted) is known, these parameters may be estimated by maximum likelihood. The estimates of these estimates are presented in Table A1.

Table A1

Parameters	ML Estimates	t–Statistic
α	1.4224	232.805
β	2.6672	134.917
γ	-10.2914	-506.386

The number of observations is 24,132.

Then the estimated income of a household with a particular category defined by known thresholds, is computed as the expected conditional income for that category using the estimated gamma distribution. This is particularly important for the highest open ended category. Table A2 presents the annual income estimates for each reported income category.

Table A2

Category	Cases	Estimate
0–4999	2,616	1,821.97
5,000–7,499	1,998	6,070.36
7,500–9,999	1,800	8,568.68
10,000–14,999	3,366	11,794.58
15,000–19,999	3,180	16,788.45
20,000–24,999	3,120	21,784.23
25,000–34,999	3,966	27,477.29
35,000–49,999	2,586	37,667.15
$\geq$ 50,000	1,500	52,714.91

Table 1. Descriptive Statistics

	Bowling Green		Louisville (All)		LV: Measured		LV: Flat		TEST	
Variable	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	T1	T2
BILL	14.1510	(5.4334)	19.3023	(4.4871)	20.7795	(8.1531)	18.7000	(0.0000)	-56.92	10.25
SHARE	0.0154	(0.0239)	0.0268	(0.0354)	0.0248	(0.0360)	0.0275	(0.0351)	-20.20	-2.57
INCOME	1.8282	(1.0840)	1.4892	(0.9217)	1.5362	(0.9298)	1.4701	(0.9178)	18.53	2.42
DINCOME	0.1592	(0.3659)	0.1858	(0.3890)	0.1276	(0.3337)	0.2095	(0.4070)	-3.84	-7.79
HHSIZE	2.7674	(1.2956)	2.4932	(1.4621)	2.1003	(1.2784)	2.6534	(1.5016)	10.81	-13.91
TEENS	0.3755	(0.7207)	0.2437	(0.6241)	0.1207	(0.4490)	0.2939	(0.6764)	10.74	-11.17
AGE1	0.0628	(0.2427)	0.0829	(0.2757)	0.0755	(0.2643)	0.0858	(0.2802)	-4.19	-1.30
AGE2	0.2310	(0.4215)	0.2500	(0.4331)	0.2861	(0.4521)	0.2353	(0.4242)	-2.42	3.87
AGE3	0.7061	(0.4556)	0.6671	(0.4713)	0.6384	(0.4806)	0.6789	(0.4670)	4.59	-2.88
COLLEGE	0.2709	(0.4445)	0.2192	(0.4137)	0.3108	(0.4630)	0.1818	(0.3857)	6.61	9.89
MARRIED	0.6721	(0.4695)	0.4864	(0.4999)	0.4563	(0.4982)	0.4986	(0.5001)	20.90	-2.87
RETIRED	0.1722	(0.3776)	0.2582	(0.4377)	0.2601	(0.4388)	0.2575	(0.4373)	-11.45	0.20
BLACK	0.0670	(0.2501)	0.1275	(0.3336)	0.0904	(0.2868)	0.1426	(0.3498)	-11.11	-5.77
CHURCH	0.2116	(0.4085)	0.1661	(0.3722)	0.1393	(0.3464)	0.1770	(0.3817)	6.40	-3.57
BENEFITS	0.2285	(0.4199)	0.3187	(0.4660)	0.2780	(0.4482)	0.3353	(0.4721)	-11.07	-4.26
MOVED	0.4684	(0.4990)	0.4333	(0.4956)	0.4638	(0.4988)	0.4209	(0.4938)	3.86	2.92
ONLYMALE	0.0430	(0.2028)	0.1044	(0.3058)	0.1406	(0.3477)	0.0896	(0.2857)	-12.76	5.21
NOV	0.3348	(0.4720)	0.3336	(0.4715)	0.3375	(0.4730)	0.3320	(0.4710)	0.15	0.39
DEC	0.3324	(0.4711)	0.3343	(0.4718)	0.3337	(0.4717)	0.3345	(0.4719)	-0.22	-0.05
MEASURED	1.0000	(0.0000)	0.2896	(0.4536)	1.0000	(0.0000)	0.0000	(0.0000)	116.93	
Observations	6445		5576		1615		3961			

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 dollars.

Table 2. Telephone Demand. Switching Regression

	Bowling Green			Louisville		
	CAP=1	Below Cap	Above Cap	FLAT=1	Flat Option	Measured Service
Constant	-1.0607 (4.10)	0.6003 (62.18)	1.0120 (97.69)	1.3903 (8.09)	0.9623 (5.71)	0.9145 (303.70)
INCOME	-0.1403 (3.88)	-0.1561 (56.37)	-0.2613 (101.29)	-0.0776 (3.43)	-0.2901 (13.91)	-0.2328 (240.59)
HHSIZE	0.1962 (8.75)	0.0150 (8.19)	0.0028 (2.41)	0.1624 (6.97)	0.0281 (1.38)	0.0004 (1.36)
TEENS	0.2740 (7.02)	0.0030 (0.90)	-0.0017 (1.12)	0.1073 (2.66)	-0.0126 (0.39)	0.0000 (0.03)
DINCOME	0.2012 (3.12)	0.0022 (10.20)	0.0049 (6.24)	0.3718 (7.96)	-0.0199 (0.79)	0.0022 (8.83)
INCOME2		0.0102 (50.07)	0.0165 (82.19)		0.0198 (13.30)	0.0150 (191.79)
HHSIZE2		-0.0001 (1.76)	-0.0000 (0.49)		-0.0004 (1.69)	0.0000 (1.94)
TEENS2		0.0001 (1.25)	0.0002 (1.36)		-0.0033 (1.65)	-0.0000 (1.18)
INCHHS		-0.0018 (7.46)	-0.0000 (0.25)		-0.0047 (2.79)	-0.0001 (2.95)
INCTNG		-0.0004 (0.90)	0.0006 (2.69)		0.0008 (0.19)	-0.0000 (0.56)
HHSTNG		0.0003 (3.05)	-0.0002 (1.02)		0.0019 (1.53)	0.0000 (0.91)
AGE1	0.0073 (0.06)	-0.0005 (0.98)	-0.0000 (0.01)	0.1660 (2.45)	-0.0116 (1.05)	-0.0005 (2.25)
AGE3	0.1821 (2.62)	0.0003 (1.29)	0.0014 (1.97)	0.0371 (0.85)	-0.0017 (0.58)	-0.0002 (1.90)
COLLEGE	-0.0979 (1.66)	-0.0008 (5.10)	-0.0014 (2.99)	-0.2657 (6.72)	0.0155 (0.86)	0.0002 (0.66)
MARRIED	-0.2275 (3.49)	-0.0018 (5.99)	-0.0024 (2.81)	-0.0469 (1.08)	-0.0003 (0.07)	0.0006 (6.00)
RETIRED	-0.3794 (4.06)	-0.0008 (1.91)	-0.0023 (1.65)	-0.0816 (1.68)	0.0048 (0.86)	-0.0002 (1.58)
BLACK	0.5735 (7.23)	0.0021 (2.95)	0.0050 (2.63)	0.1831 (3.50)	-0.0063 (0.51)	-0.0007 (3.94)
CHURCH	0.0667 (1.15)	-0.0001 (0.31)	0.0011 (2.93)	0.1134 (2.50)	-0.0080 (1.07)	-0.0002 (1.49)
BENEFITS	0.1654 (2.10)	0.0005 (1.53)	-0.0000 (0.06)	0.1268 (2.90)	-0.0104 (1.22)	-0.0006 (4.95)
MOVED	-0.1059 (1.98)	-0.0005 (2.99)	-0.0011 (2.25)	-0.0144 (0.37)	0.0028 (2.03)	0.0001 (0.52)
ONLYMALE	-0.6385 (3.23)	-0.0012 (2.07)	-0.0056 (2.21)	-0.0190 (0.32)	-0.0030 (1.65)	0.0001 (0.66)
NOV	-0.0949 (1.59)	-0.0003 (1.53)	-0.0010 (2.05)	-0.0169 (0.44)	-0.0004 (0.24)	-0.0000 (0.00)
DEC	0.0095 (0.16)	0.0003 (1.55)	0.0002 (0.48)	-0.0066 (0.17)	0.0002 (0.20)	-0.0000 (0.00)
IMR		0.0088 (2.21)	0.0111 (2.45)		-0.0772 (0.93)	-0.0067 (2.52)
Observations	6445	5892	553	5576	1615	3961
R^2/Log-likelihood	-1628.974	0.918	0.995	-3145.701	0.783	0.995

The dependent variable in the first column equals one when total bill payments exceed $21.50, and when consumers choose the flat service option in the fourth, but it is the share of telephone expenditure in monthly income in all the others. Income is measured in logarithm. IMR denotes the Inverse Mill's Ratio for correction of selectivity bias. Selection equations are estimated as probits and usage equations by OLS. Absolute heteroscedastic–consistent t–statistics are shown between parentheses.

Table 3. 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)
INCOME2		0.0880 (1.79)		0.1194 (2.53)	
HHSIZE2		-0.0689 (2.56)		-0.0115 (0.48)	
TEENS2		-0.1236 (4.55)		-0.1150 (5.08)	
INCHHS		-0.1397 (2.85)		-0.1736 (3.71)	
INCTNG		0.0238 (0.21)		-0.0045 (0.04)	
HHSTNG		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.	6445	5892	5892	5892	5892
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 4. 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)
INCOME2		0.4730 (1.61)		0.6184 (2.14)	
HHSIZE2		-0.3039 (2.95)		-0.0916 (1.36)	
TEENS2		-0.5704 (0.58)		-0.0143 (0.01)	
INCHHS		0.0015 (0.00)		-0.4733 (1.48)	
INCTNG		2.9018 (2.38)		2.8020 (2.16)	
HHSTNG		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.	5576	1615	1615	1615	1615
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 5. 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 6. 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	0.05	0.18	6.90	4.62	1.50	0.79
	1.64	0.79	3.03	2.07	4.83	2.39	1.21	0.19
$\exp(w_i)$	3.19	4.01	8.03	6.96	5.37	5.82	6.23	5.84
	1.88	0.77	43.37	0.00	1.04	0.95	4.14	5.48
$\exp(-w_i)$	5.51	5.04	3.26	4.45	6.05	4.39	5.07	2.49
	4.58	4.36	3.17	3.09	4.10	2.47	3.29	2.77
$\sqrt{(w_i)}$	3.03	2.88	1.25	1.32	6.80	4.33	3.92	1.97
	1.94	0.89	1.13	0.79	4.72	2.36	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.

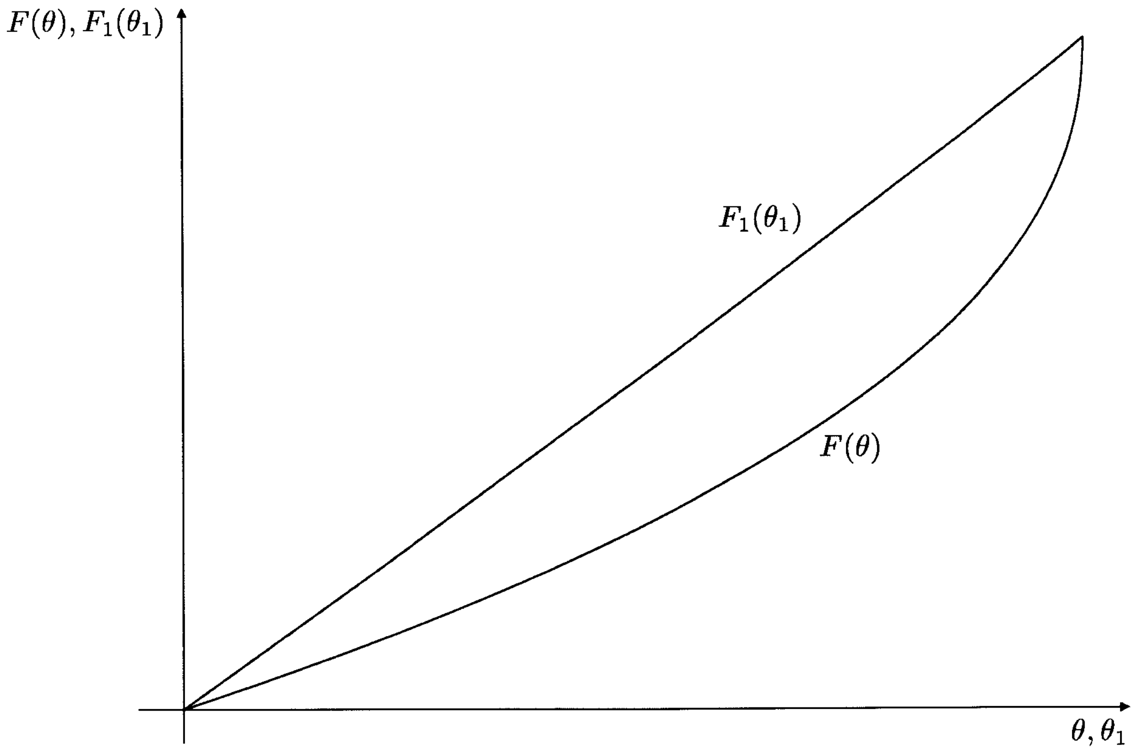

Figure 1: Distribution Functions

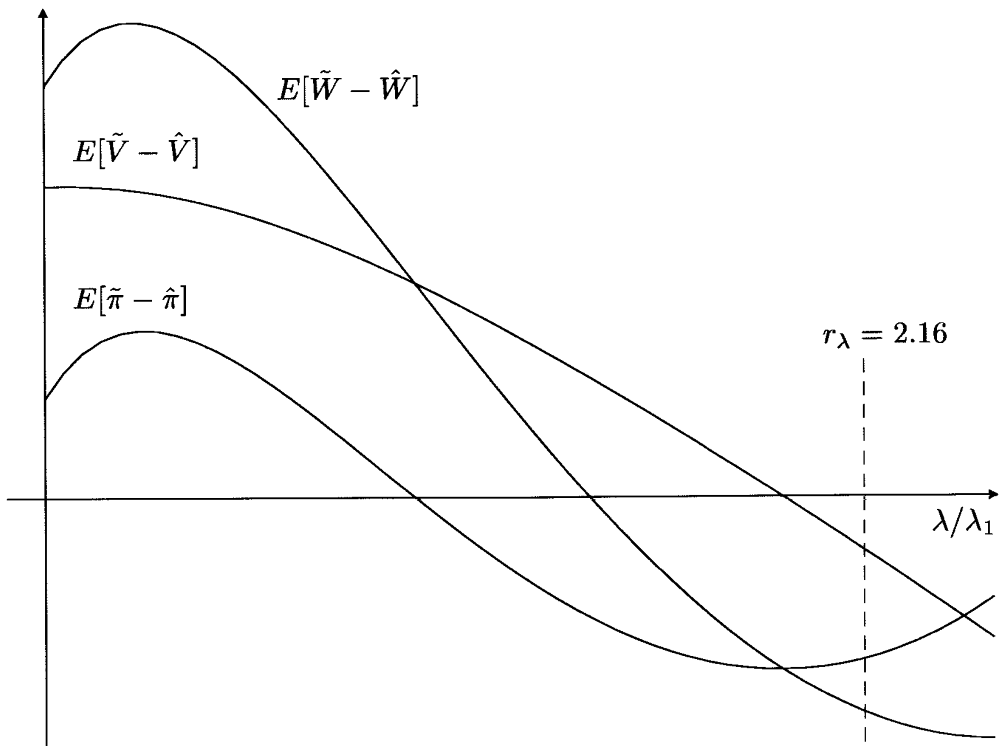

Figure 2: Welfare Analysis