ELSEVIER Information Economics and Policy 8 (1996) 163–203

INFORMATION ECONOMICS AND POLICY

# Can universal service survive in a competitive telecommunications environment? Evidence from the United States consumer expenditure survey

Frank A. Wolak*

*Department of Economics, Stanford University, Stanford, CA 94305-6072, USA*

Received 15 February 1996; accepted 20 March 1996

**Abstract**

The purpose of this paper is to assess the impact of reducing the magnitude of the cross-subsidy in telephone services pricing that flows from long-distance service to local service on the consumption of both of these telephone services and consumer welfare at the household level. Our econometric modeling framework specifies a complete system of household-level demand functions that are derived from the assumption of utility maximization. Our primary data source is the Bureau of Labor Statistics' *Survey of Consumer Expenditures*. In addition to local and long-distance phone service, food, clothing, and other non-durable expenditures are included in our five-good demand system. All of our price change scenarios point to the conclusion that there appears to be little loss in household-level welfare and little, if any, reduction in the number of households connected to the local telephone network, due to the projected reductions in this cross-subsidy brought about by an increasing amount of competition in all telecommunications markets. In addition, we find that if these local service price increases are coupled with reductions in the price of long-distance service, the net effect can actually be a gain in consumer welfare for a large fraction of households which in the aggregate yields a net benefit to the population of United States households.

*Key words:* Telecommunications demand; Universal service; Separability testing; Household-level demand analysis

*JEL Classification:* C50; D12; L51

* Corresponding author: E-mail: wolak@zia.stanford.edu.


*PII* S0167-6245(96)00009-1

## 1. Introduction

Increasingly vigorous competition in all aspects of telecommunications supply make it difficult to maintain substantial cross-subsidies in the pricing of products in these markets. In the United States, one cross-subsidy that is currently under increasing scrutiny flows from long-distance telephone service to local exchange telephone service. Between 45 to 49 percent of the total cost of a inter-LATA long-distance call is paid to the local exchange carriers at the originating and terminating points of the call. According to Sievers (1994), approximately half of these payments to local exchange carriers are in excess of the costs of local access. A major rationale for these subsidies is to maintain a low price for local residential service in order to achieve the goals of universal service. Although there is considerable debate over the exact fraction of local access charges that are subsidies, there is little debate over the existence of these subsidies. Competition in the inter-LATA long-distance market and the growing number of competitive providers of long-distance access service, particularly for large business customers, traditionally the largest source of revenues for local exchange carriers, has led to consideration of large rate increases for the price of local residential phone service. For example, the California Public Utilities Commission recently approved a more than 35 percent increase in the price of local residential service provided by Pacific Bell, the Regional Bell Operating Company serving California (Banks, 1994).

In light of these recent increases in the price of local residential phone service, the purpose of this paper is to assess the impacts of these price changes on the consumption of both local and long-distance phone service and consumer welfare at the household-level. Because our unit of observation is the household, we can determine how these changes in telephone service consumption and consumer welfare vary with the attributes of the household. Quantifying how these impacts vary with household characteristics is necessary to understand how the aggregate burden of these price changes is shared across different types of households in our sample. Because the dataset we use is a probability sample of the population of U.S. households, we can also compute estimates of the welfare impacts of these price changes for the population of U.S. households.

In the process of answering these questions we also investigate the validity of several hypotheses about the structure of demand for both local and long-distance telephone service. One issue of particular interest is the validity of assuming that both local and long-distance telephone service are separable from all other goods in the household's utility function. This assumption is implicit in all studies of telecommunications demand which focus on the household's allocation of total telephone expenditures between local and long-distance service independent of the level of consumption or the prices of all others goods. Our econometric modeling framework and household-level database provides an opportunity to test this often maintained assumption.

Our econometric modeling framework specifies a complete system of consumer demand functions that are derived from the assumption of household-level static utility maximization. We specify a parametric functional form for the household-level indirect utility function and derive an expression for the demand system through an application of duality theory. We estimate demand systems derived from both the translog indirect utility function similar to Jorgenson et al. (1982) and Integrable Quadratic Almost Ideal Demand System (IQUAIDS) introduced by Banks et al. (1994). Although both functional forms are second-order flexible, neither can be nested within the other. Consequently, we present our consumption change calculations and household-level welfare change calculations for both models as a check of the robustness of our conclusions to the choice of the functional form for the household's indirect utility function.

Our primary data source is the Bureau of Labor Statistics' (BLS) *Survey of Consumer Expenditures*. For this survey, the BLS collects information on household consumption, broken down by many classes of goods, income and various demographic characteristics on a quarterly basis. From 1988 to the end of 1991, this survey collected data on household-level telephone consumption broken down by local and long-distance service. The survey also collects information on the consumption of several categories of non-durable expenditures which we include in our demand system. In addition to local and long-distance phone service, food, clothing, and other non-durable expenditures are included in our five-good demand system. We also include demographic variables interacted with prices in the both the translog and IQUAIDS indirect utility functions, which implies that the demand functions associated with each indirect utility function differ across households according to these demographic characteristics. Because of our focus on modeling the demand for both local and long-distance service within the context of a complete system of consumer demand functions, we utilize data from January 1988 (the first month the local/long-distance consumption split was collected) through February of 1991 (the last month in which this data was collected).

There is a substantial amount of agreement in the structure of consumer preferences recovered from the translog and IQUAIDS models. Both estimated indirect utility functions provide substantial evidence against homothetic preferences. The distributions of price elasticity estimates across our sample of households obtained from the translog and IQUAIDS model are very similar. We also find that for all but a small fraction of observations in the sample, the curvature restrictions on the indirect utility function implied by optimizing behavior hold for the translog model. For the IQUAIDS model this fraction is noticeably higher, but the vast majority of observations still satisfy these curvature restrictions, so that we feel confident in using both models to compute welfare changes. We consider a set of price change scenarios and compute the expenditures on both local and long-distance services expected under these new prices and the compensating variation relative to the current prices facing that household for each price change scenario. We find that for price change scenarios which balance the percent increase in local service with a corresponding reduction in long-distance service – for instance, a twenty percent increase in the price of local service coupled with a twenty percent decrease in the price of long-distance service – the mean of the compensating variations over our sample of households is negative. Using the sampling weights associated with each household in the dataset, we compute an estimate of the population (of U.S. households) mean compensating variation associated with each price change scenario. We find that for the combination local and long-distance price changes scenarios the mean compensating variation for the U.S. population is also negative. This result implies that if all households in the U.S. were remunerated positively or negatively according to their compensating variation for this pair of price changes, the process would generate a net increase in revenue to society. This result follows from the estimated price-inelastic demand for local service and very price-elastic demand for long-distance service for the vast majority of households in the sample.

For local price increases unaccompanied by decreases in the price of long-distance service, although the mean compensating variation over all households in our sample is positive (implying that consumers must be compensated a positive amount to be indifferent to this type of price change), it is still a very small fraction of the household's total non-durable expenditure. For example, a forty percent increase in the price of local service only results in an average quarterly compensating variation of $17.59 in January 1988 dollars for the translog model. This figure is approximately 0.6 percent of the sample mean of household total nondurable expenditure, which is $3,017.52 in January 1998 dollars. The IQUAIDS model produces an estimate of $17.89 for this same value for our sample of households. The 5th percentile to 95th percentile range of compensating variations for this price change for the translog model is $13.01 to $22.60, ($13.10 and $25.52 for the IQUAIDS model), so that even for the extremes of the sample, the welfare losses associated with these price changes seem relatively minor. Even for a price increase of this magnitude (40%), for both models there are no households in the sample expected to consume a positive amount of local service before the price change that are predicted to reduce their consumption of telephone service to zero, which we equate with disconnecting from the local exchange network. All of these results point to the conclusion that there appears to be little loss in consumer welfare and little, if any, reduction in the fraction of households connected to the local telephone network, due to the projected increases in the price of local service brought about by an increasing amount of competition in all telecommunications markets. In addition, we find that if these local price increases are coupled with long-distance price decreases, the net effect can actually be a gain in consumer welfare for a large fraction of households.

We implement our separability test only on the translog model, because this functional form allows us to impose separability globally by restricting only the parameters of the model. For this functional form, we find substantial evidence against the null hypothesis of separability of local and long-distance service from all other goods. This result implies that, at least for our dataset, to consistently estimate the structure of household-level demand for local and long-distance service, a two-stage budgeting approach, which first estimates the demand for telephone service expenditure (the sum of local and long-distance consumption) and then conditional on this total expenditure level allocates it between local and long-distance service according to the relative price of local versus long-distance service, cannot be utilized. The household-level demand for local and long-distance service must specified to depend on total expenditure and the prices of all other goods consumed by the household because of the non-separability between local and long-distance telephone consumption and the consumption of other goods.

If the burden of these price changes can be measured by the ratio of household-level compensating variation (due to the price changes) to total expenditure, we find that the burden of these price changes is more than proportionately borne by the lower income (or total expenditure) segment of the sample. In addition, this burden is borne to a greater extent by the older-headed households, those in urban areas, those with the spouse employed, those with more children in the age range from 2 through 15 years old. This relative burden is also increasing in the number of hours the head works and the number of hours the spouse works annually.

The paper reaches these conclusions in the following fashion. Section 2 describes our data set and presents some descriptive statistics on the distribution of total phone expenditures, local-service expenditures and long-distance expenditures across our sample of households. This helps to motivate the results from our demand system estimation procedure. Section 3 presents the translog and IQUAIDS indirect utility functions and derives their respective demand systems. This section also motivates the stochastic specifications chosen for each demand system. Section 4 describes our estimation procedure and gives the results of our model estimation and several model specification tests, including our test for separability of local and long-distance phone service from all other goods. Section 5 first describes the price change scenarios that we consider for our welfare analyses. This section then characterizes how these welfare impacts vary across households in the sample. In light of the welfare calculations and telephone consumption changes presented in Section 5, the paper concludes with a discussion of the answer to the question posed in the title of the paper concerning the viability of universal service in a competitive telecommunications environment.

## 2. Local and long-distance telephone expenditure distribution

In this section we describe the dataset used in our analysis. We then examine the distribution of quarterly total telephone expenditure, local telephone expenditure and long-distance telephone expenditure for each of the four quartiles of the total quarterly non-durable expenditure distribution for our sample of households.

Our dataset consists of quarterly observations on consumption expenditures on local telephone service, long-distance service, food, clothing and other non-durable good consumption. We focus our demand analysis on total non-durable consumption to avoid the issues associated with the distinction between the price of single quarter's service flow from a good versus the purchase price of the good. This distinction arises whenever the good purchased provides services for a longer time than the period in which the purchase is observed (in our case a quarter). By definition, this type of good is a durable good, hence our focus on total nondurable consumption, which we henceforth refer to as total expenditure for expositional ease. We now describe each of the goods classes in more detail.

Local telephone service, as defined in the *Survey of Consumer Expenditures* (CES), is all expenditures for local telephone service for that household. It includes the cost of local phone service for all phones in all dwellings the household may own and as well as the installation charges associated with these phones, if an installation charge occurs within one of the quarters in our sample period. Long distance telephone service consumption is defined by the CES as the total of all long-distance calling charges where the cost of a single call is broken out in detail on the phone bill. Food consumption is defined as all expenditures on food consumed both within the household and outside of the household (in, for example, restaurants). Clothing consumption is the total of all clothing purchases made by the household during that quarter. Other non-durable consumption is the residual category of non-durable consumption. It includes commodities such as fuel for automobiles and household heating, electricity, transportation services, and other non-durable consumption services. For the purposes of our analysis we convert all nominal magnitudes to January 1988 dollars using the BLS *Consumer Price Index Detailed Report* total nondurable goods price index normalized to have January 1988 as the base period. Consequently, all dollar magnitudes discussed in the paper are real January 1988 dollars deflated in this manner.

To provide intuition for the demand system estimation results in Section 3, we decompose the distributions of local, long-distance, and total telephone expenditures across our sample of households according the quartiles of the total expenditure distribution. We perform this decomposition with respect to the total expenditure distribution rather than the distribution of household income, because for a large fraction of households, income in a given time period is a very poor predictor of the household's total consumption expenditures during that period. For the usual life-cycle, permanent-income considerations, we would expect total expenditure in any period to be more highly correlated with permanent income than would be actual income for that period. This point is discussed in Blundell et al. (1993) for the *Family Expenditure Survey* (the United Kingdom's analogue to the CES) and by Lusardi (1993) for the CES.

Our decomposition procedure first computes the three breakpoints of the quartiles of the distribution of quarterly total expenditure ($QTE$) for our sample of $N = 11{,}467$ households. These quartiles in January 1988 dollars are: 1st – ($QTE < 1{,}688.00$), 2nd – ($1{,}688.00 \leq QTE < 2{,}594.37$), 3rd – ($2{,}594.37 \leq QTE < 3{,}787.27$), 4th – ($QTE \geq 3{,}687.27$). The mean level of total expenditure in January 1988 dollars for the four quartiles are: 1,192.43, 2,126.83, 3,127.46, and 5,624.48, respectively. For each of these total expenditure quartile subsamples, we compute a kernel estimate of the density of expenditure on local telephone service, long-distance service, and the total telephone bill. The general form for this density estimate is

$$\hat{f}(x) = \frac{1}{Th} \sum_{i=1}^{T} K\left(\frac{x - X_i}{h}\right)$$

where $X_i$ is the value of any of the three telephone expenditure values for the $i$th observation, $T$ is the total number of observations in the quartile, and $h$ is the smoothing or bandwidth parameter. The kernel $K(t)$ is a symmetric function satisfying

$$\int K(t)\mathrm{d}t = 1, \quad \int tK(t)\mathrm{d}t = 0, \text{ and } \int t^2K(t)\mathrm{d}t < \infty.$$

We use the Gaussian kernel

$$K_G(t) = \frac{1}{\sqrt{2\pi}}\exp(-\frac{1}{2}t^2).$$

We use the automatic bandwidth selection procedure recommended by Silverman (1986), $h = 0.9AT^{-1/5}$, where

$$A = \min(\text{sample standard deviation of the } X_i,$$

$$(\text{inter-quartile range of the } X_i / 1.34)).$$

Before describing the density estimates, we note one small-sample irregularity. Because of the local smoothing property of the kernel estimation process, these densities can take on values slightly greater than zero for negative values of telephone expenditure despite the fact that the actual data contains no negative values for expenditure. In the limit as the number observations gets large, this estimated probability mass on negative values would tend to zero.

Fig. 1 plots estimates of the density of total quarterly local telephone expenditure for all of the quartiles of the total expenditure distribution. The striking aspect of this figure is the invariance of the density of total local telephone expenditures across the quartiles of the total expenditure distribution. Even comparing the first to the fourth quartile, there is very little difference between the two local telephone expenditure density estimates. Fig. 2 plots estimates of the density of total quarterly long-distance service expenditures. A very different story

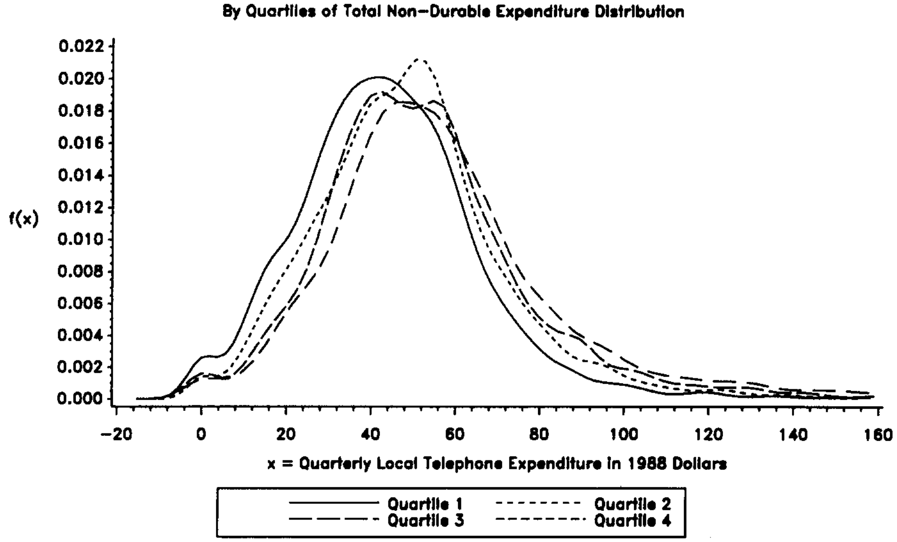


Fig. 1. Kernel density estimate of local telephone expenditure.

emerges from these density estimates. We find that each expenditure quartile is generally a rightward shift and more positively skewed version of the quartile below it. This pattern indicates that for any value of quarterly long-distance expenditures, the proportion of households in each total expenditure quartile

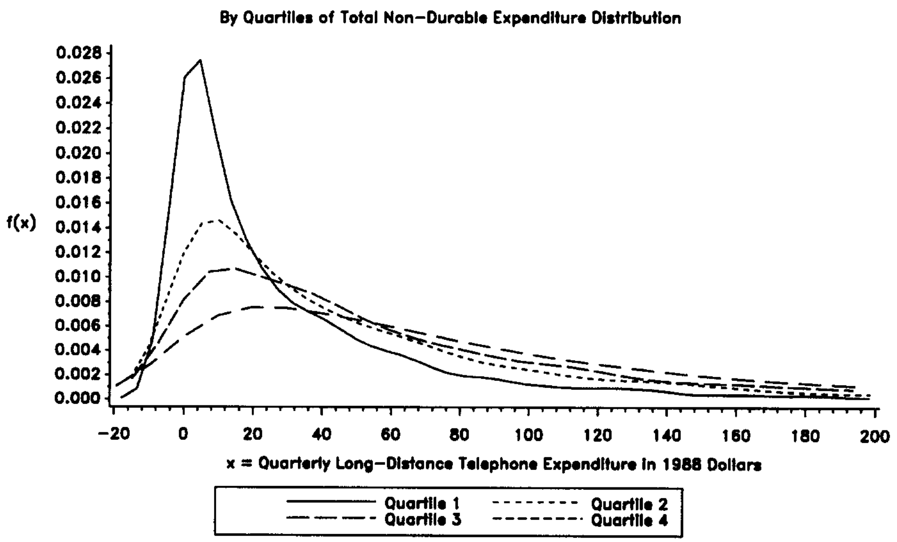


Fig. 2. Kernel density estimate of long-distance telephone expenditure.

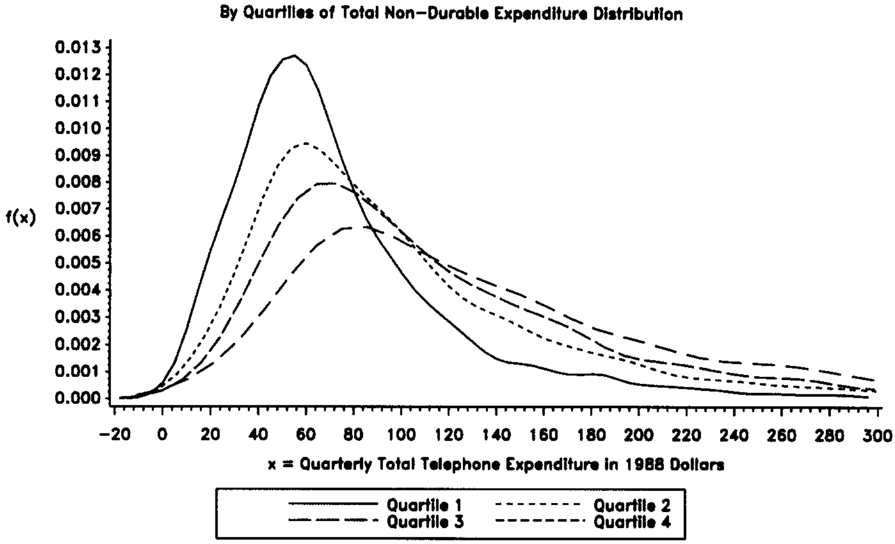


Fig. 3. Kernel density estimate of total telephone expenditure.

whose long-distance expenditures exceed this value is larger for each successively higher total expenditure quartile.

Fig. 3 investigates which of the above two patterns of changes in density shapes dominates in characterizing the changes in the shape of the density of total quarterly telephone service expenditures. The pattern of density shifts across expenditure quartiles is similar to the pattern for long-distance service, although the increase in the positive skewness associated with moving to higher total expenditure quartiles is much less pronounced. The sample correlation between total local service expenditure and total long-distance expenditure at the household level is 0.17, which implies a surprisingly small degree of positive linear dependence between these two components of the total telephone bill. This small correlation partially explains why the pattern of density shifts across total expenditure quartiles for total phone expenditures is less pronounced than for long-distance expenditures.

We close this section with a discussion of the price data used in our econometric analysis. A major problem faced by all cross-section demand system studies using either U.S. or U.K. household-level data is the fact that there is no cross-section dataset of commodity prices which can be linked to the sample of households, so that all price series used in these analyses vary only over time. Even if there were commodity-specific price data available at some degree of cross-sectional disaggregation, say at the state level, our analysis (and other analyses using household-level data such as the CES) are complicated by the fact that confidentiality considerations preclude the BLS from releasing any state identifiers for a substantial fraction of our household-level observations. Until very recently, the BLS did not release information on state identifies for any households. The degree of geographic detail available for all observations is whether or not that household lives in one of the four Census regions of the U.S. or any rural area in the U.S. Although the BLS does compute price indexes for food, clothing and other non-durable goods on a monthly basis for each of these four Census regions, it only computes price-indexes for local and long-distance service on a monthly basis for the entire nation. It is plausible to assume that all households face the same price for long-distance service, given the nature of the good being purchased. However, for the same reason, this assumption is less tenable for local service. Nevertheless, an argument in favor of local prices moving together is the fact that they are all regulated at the state level and there is a considerable amount of across-state communication between regulatory bodies in setting these prices. Regardless of the validity of this argument, for data availability reasons and confidentiality considerations we are forced into making this unsatisfying assumption.

Nevertheless, we attempt to make maximum use of the available regional and time series variation in the price data in our demand system estimation. For food, clothing and other non-durable expenditures we use the Census Region level of cross-sectional variation in prices contained in the BLS *Consumer Price Index Detailed Report* to assign prices for these commodities to households based on the census region in which that household resides. For households in rural areas we use the national price index for these three goods. For local and long-distance phone service prices we assign the aggregate price index for both of these commodities to all observations in our sample. To account for the fact the initial level of relative prices for food, clothing and other non-durable goods across Census Regions is unobserved (because all regional prices indexes are only defined relative to a base year), we include regional dummy variables in the household-level demand functions for all goods.

Because of the rolling panel nature of the CES data collection process we are able to introduce some additional meaningful cross-sectional variation in prices for a given consumer within a given quarter. Data for the CES is collected on a quarterly basis for each household for the consumption amounts in the previous three months. This means that in any given month a different set of consumers is being retrospectively interviewed for their consumption in the previous quarter. Households usually remain in the survey for four quarters and then exit. Because of the retrospective nature of the questionnaire and the fact that all of the BLS price index series are collected on a monthly basis, we use the price index for each good for the most recent month of three months covered by the retrospective survey.

Fig. 4 plots the monthly nominal and real price indexes for local and long-distance phone service from January 1988 to June 1994. Recall that the real and nominal series for both local and long-distance phone service are normalized

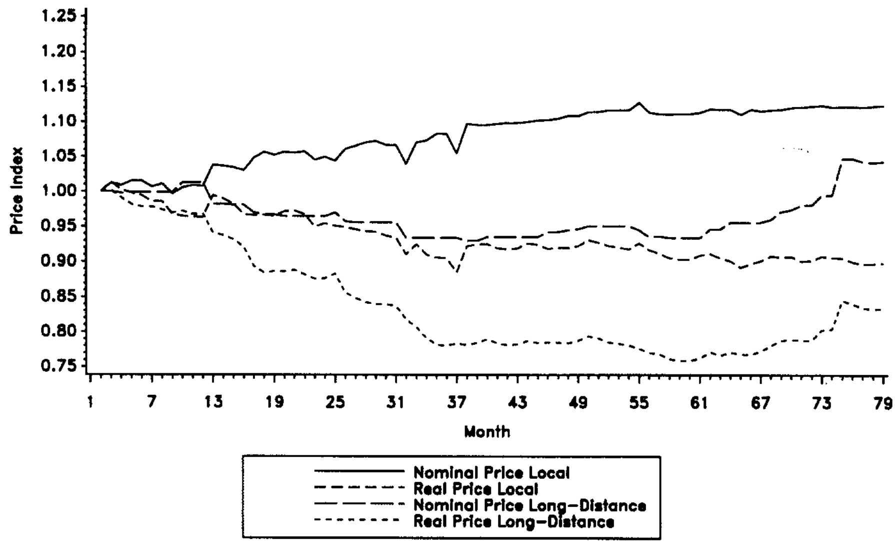


Fig. 4. Monthly price series (month 1: January 1988).

to one in January 1988. The figure shows that the nominal price of phone service increased over this time period, whereas the real price (deflated by the U.S. total nondurable goods price deflator) remains almost constant throughout this four-year period. The long-distance price index showed modest nominal declines and substantial real declines over this sample period. These price trends have continued up through the present for local service. For long-distance service, since early 1993 there has been a steady increase in the nominal and real price of long-distance phone service.

These figures summarize the characteristics of the data that will be used to estimate our household-level demand system for local phone service, long-distance phone service, food, clothing and other nondurable expenditures. The major shortcoming of our analysis is certainly the price series available. Optimally, we would like household-level price indexes for the five goods. A less ambitious goal would be to have state-level price indexes. Even if these state-level prices were available, as mentioned above, for confidentiality reasons we cannot assign a substantial fraction of our households to any state. Nevertheless, given the price information we do have, as discussed above, we attempt to make use of it in a way that introduces as much across-household price heterogeneity as is possible.

## 3. Econometric modeling framework

Our econometric modeling framework must be sufficiently flexible to encompass several empirical and theoretical considerations. The first empirical regularity is the long history of work indicating the existence of non-homothetic preferences. The most well-known result is Engel's Law, which states that the budget share devoted to food is declining in total expenditure. Deaton and Muellbauer (1980b) survey the evidence against homothetic preferences. Recently there has been research providing evidence against traditional Engel curve representations with budget shares linear in the log of total expenditure, first specified by Working (1943) and Leser (1963). Bierens and Pott-Butter (1990) and Hausman et al. (1995) are representative of this line of research. Given the strong empirical evidence against homotheticity, we must select an underlying household-level utility function which is non-homothetic and allows for budget shares that are nonlinear functions of the log of total expenditure.

From the theoretical perspective, there are several requirements for our underlying utility function. The first is the ability to impose the restrictions implied by utility maximizing behavior on the demand functions estimated in a data-independent fashion. Our second requirement is second-order flexibility of the underlying utility function, which means that for any point in the data space, the functional form can exactly reproduce any theoretically possible value of the function, its gradient, and matrix of second-partial derivatives through appropriate choice of the parameters of the functional form. The final theoretical requirement is that our modeling framework allows the imposition of the restrictions implied by separability of local and long-distance phone service from all other commodities in a data-independent manner using restrictions on the parameters of the demand system. In this way, the null hypothesis of separability can be tested using conventional parametric hypothesis testing techniques.

The translog indirect utility function satisfies these theoretical and empirical criteria. A second functional form which satisfies all but the last of these requirements is the IQUAIDS demand system. Because neither of these parametric families of demand systems can be nested within the other (see Lewbel, 1989), and each has its advantages relative to the other, we estimate both models and perform our welfare calculations for both, as a check of the robustness of the magnitudes calculated. Both of these models are derived by an application of duality theory to an indirect utility function. We utilize duality theory to recover the parameters of the indirect utility function from the Marshallian demand functions because we have observations on prices indexes associated with the five goods consumed rather than quantity indexes, which are required to recover estimates of the parameters of the direct utility function for the cases in which the underlying utility functions are not self-dual in the sense discussed by Houthakker (1965).

We now present each functional form and discuss how to impose the restrictions implied by utility maximizing behavior on the demand functions estimated. These restrictions are: (1) homogeneity of degree zero of the demand functions in prices and total expenditure, (2) symmetry of the Slutsky matrix (the matrix of compensated own- and cross-price effects) and (3) quasi-convexity of the indirect utility function in prices, which is equivalent to negative semidefiniteness of the Slutsky matrix.

Before presenting each functional form we define the following notation. Let $p_i$ denote the price of good $i$, $x_i$ the quantity of good $i$ consumed, and $M$ the total expenditure. In terms of this notation we have

$$M = \sum_{i=1}^{N} p_i x_i \text{ and } w_i = \frac{p_i x_i}{M}$$

where $w_i$ is the share of total expenditure spent on the $i$th good and $N$ is the total number of goods consumed. The translog indirect utility function is

$$\ln V(P, M) = \alpha_0 + \sum_{i=1}^{N} \alpha_i \ln\left(\frac{p_i}{M}\right) + \frac{1}{2} \sum_{i=1}^{N}\sum_{j=1}^{N} \beta_{ij} \ln\left(\frac{p_i}{M}\right) \ln\left(\frac{p_j}{M}\right) \tag{3.1}$$

where $P = (p_1, p_2, \ldots, p_N)'$ is the vector of prices for the N goods. Applying the logarithmic version of Roy's Identity to this indirect utility function yields share equations

$$w_i(P, M) = \frac{\alpha_i + \sum_{j=1}^{N} \beta_{ij} \ln(p_j/M)}{\sum_{k=1}^{N} \alpha_k + \sum_{k=1}^{N} \sum_{j=1}^{N} \beta_{kj} \ln(p_j/M)}. \tag{3.2}$$

Because the share equation (3.2) is homogeneous of degree zero in the parameters $\alpha_i$ and $\beta_{ij}$, a single normalization restriction must be imposed to identify the remaining parameters. The usual restriction is to impose $\sum_{i=1}^{N} \alpha_i = -1$. By inspection, the share equation is homogeneous of degree zero in the vector of prices $P$ and total expenditure $M$, so that homogeneity imposes no restrictions on the parameters of the model. As discussed in Jorgenson et al. (1982), the Slutsky matrix is

$$S = \Pi^{-1}\left(\frac{1}{D(P, M)}(I - \iota w')\Delta_{pp}(I - \iota w') + ww' - W\right)\Pi^{-1}, \tag{3.3}$$

where $\Pi$ is the $(N \times N)$ diagonal matrix with $(p_i/M)$ as the $i$th diagonal element, $w = (w_1, w_2, \ldots, w_N)'$, $\Delta_{pp}$ is an $(N \times N)$ matrix with $\beta_{ij}$ as the $(i,j)$th element, $W$ is the $(N \times N)$ diagonal matrix with $w_i$ as the $i$th diagonal element and $\iota$ is an $N$-dimensional vector of 1's. The function $D(P, M)$ is the denominator of the fraction on the right-hand side of Eq. (3.2). This expression implies that symmetry of $S$ is equivalent to symmetry of $\Delta_{pp}$, which holds if $\beta_{ij} = \beta_{ji}$ for all $i$ and $j$. Eq. (3.3) also shows that quasi-convexity of the indirect utility function in prices, which implies $S$ is negative semi-definite, is a data dependent restriction. The Slutsky matrix, $S$, explicitly depends on the prices, shares and total expenditure. Consequently, our strategy is first to estimate the model without imposing this restriction. Given the parameter estimates obtained, we check to see whether this constraint holds for each of the points in our dataset before performing our welfare calculations on that observation, because it makes little economic sense to perform welfare calculations for observations failing the conditions for integrability.

The IQUAIDS model is an extension of the Almost Ideal Demand System of

Deaton and Muellbauer (1980a) which includes an additional term in the square of the log of total expenditure in each share equation to allow greater flexibility in modeling expenditure effects in a demand system derived from utility maximizing behavior. Banks et al. (1994) show that the IQUAIDS demand system has rank 3, which Gorman (1981) showed is the maximal rank that can be attained by an integrable demand systems with expenditure shares that are linear in functions of total expenditure. The rank of a demand system is essentially a measure of the variety of shapes of Engel curves that can be generated from it. Lewbel (1991) gives a discussion of the rank of a demand system and its implications for empirical analyses of demand. Banks et al. (1994) also show that among demand systems that can be written as

$$w_i = A_i(P) + B_i(P)\ln(x) + C_i(P)g(x) \tag{3.4}$$

where $x = M/a(P)$ and $A_i(P)$, $B_i(P)$, $C_i(P)$, and $g(x)$ are any differential functions, integrability of the demand system requires that $g(x) = (\ln(x))^2$, or the imposition of a restriction on the function $C_i(P)$ which reduces the rank of the demand system. The function $a(P)$ is a homogeneous of degree 1 price index function. Given recent evidence against Engel curve models of the form given in (3.4) with $C_i(P) = 0$, the IQUAIDS model provides a demand system derived from utility maximizing behavior which can capture these empirically important nonlinear relationships the between expenditure share and $\ln(x)$.

The indirect utility function underlying the IQUAIDS model is

$$\ln(V(P, M)) = \left(\left\{\frac{(\ln(M) - \ln(a(P)))}{b(P)}\right\}^{-1} - \lambda(P)\right)^{-1}, \tag{3.5}$$

with the following notational definitions:

$$\ln(a(P)) = \delta_0 + \sum_{i=1}^{N}\delta_i \ln(p_i) + \frac{1}{2}\sum_{i=1}^{N}\sum_{j=1}^{N}\gamma_{ij}\ln(p_i)\ln(p_j), \tag{3.6}$$

$$b(P) = \prod_{i=1}^{N} p_i^{\beta_i}, \text{ and } \lambda(P) = \sum_{i=1}^{N}\lambda_i \ln(p_i), \text{ where } \sum_{i=1}^{N}\lambda_i = 0. \tag{3.7}$$

In terms of this notation the expenditure share equations for the IQUAIDS models take the form

$$w_i = \delta_i + \sum_{j=1}^{N}\gamma_{ij}\ln(p_j) + \beta_i \ln(M/a(P)) - \frac{\lambda_i}{b(P)}(\ln(M/a(P)))^2. \tag{3.8}$$

The expenditure and uncompensated price elasticities can be computed in terms of

$$\mu_M = \frac{\partial w_i}{\partial \ln(M)} = \beta_i + \frac{2\lambda_i}{b(P)}\left[\ln\left(\frac{M}{a(P)}\right)\right], \tag{3.9}$$

$$\mu_{ij} = \frac{\partial w_i}{\partial \ln(p_j)}$$
$$= \gamma_{ij} - \mu_M\left(\delta_i + \sum_{k=1}^{N} \gamma_{jk} \ln(p_j)\right) - \frac{\lambda_i \beta_j}{b(P)}\left[\ln\left(\frac{M}{a(P)}\right)\right]^2. \quad (3.10)$$

In terms of this notation the uncompensated price elasticities are given by

$$e^{\mathrm{u}}_{ij} = \frac{\mu_{ij}}{w_i} - \delta_{ij}, \quad (3.11)$$

where $\delta_{ij}$ is the Kronecker delta. The elasticity version of the Slutsky equation is

$$e^{\mathrm{c}}_{ij} = e^{\mathrm{u}}_{ij} + e_M w_j. \quad (3.12)$$

Imposing symmetry and negative semi-definiteness on the matrix with $(i,j)$ element $[w_i \times e^{\mathrm{c}}_{ij}]$ is equivalent to those two properties holding for the Slutsky matrix. By inspection of the terms comprising this matrix, we can see that symmetry requires $\gamma_{ij} = \gamma_{ji}$ for all $i$ and $j$. However, similar to the translog, negative semi-definiteness of the Slutsky matrix depends on the prices and expenditure shares at the point of evaluation of the demand system. Consequently, we follow the same strategy as for the translog-checking negative semi-definiteness at all data points after we have estimated the model. There are two sets of additional restrictions required to make the IQUAIDS model consistent with utility maximization. The first is implied by the fact that budget shares must sum to one. The summability restrictions are

$$\sum_{i=1}^{N} \delta_i = 1, \quad \sum_{i=1}^{N} \beta_i = 0, \text{ and } \sum_{i=1}^{N} \gamma_{ik} = 0 \text{ for all } k. \quad (3.13)$$

The second set of restrictions are homogeneity of degree zero of the demand system in prices and total expenditure. These restrictions are

$$\sum_{j=1}^{N} \gamma_{ij} = 0 \text{ for all } i, \quad (3.14)$$

in addition to the restriction on the $\lambda_i$ given in Eq. (3.7).

Because we are using across-household differences in consumption patterns to identify the parameters of the demand systems, we would like to distinguish between consumption differences due to differences in prices and total expenditure and those that are due to differences in household preferences. For this reason we include household demographic characteristics in both the translog and IQUAIDS indirect utility functions. In order for these differences to be econometrically identified, they must enter interacted with functions of prices and total expenditure. If we define $A_k$ as the $k$th demographic characteristic and $A$ as the $K$-dimensional vector of these characteristics, the translog indirect utility function with demographic characteristics becomes

$$\ln V(P, M, A) = \alpha_0 + \sum_{i=1}^{N} \alpha_i \ln\left(\frac{p_i}{M}\right) + \frac{1}{2} \sum_{i=1}^{N}\sum_{j=1}^{N} \beta_{ij} \ln\left(\frac{p_i}{M}\right) \ln\left(\frac{p_j}{M}\right) + \sum_{i=1}^{N}\sum_{k=1}^{K} \eta_{ik} \ln\left(\frac{p_i}{M}\right) A_k. \quad (3.15)$$

Applying the logarithmic version of Roy's identity, the translog share equations become

$$w_i(P, M, A) = \frac{\alpha_i + \sum_{j=1}^{N} \beta_{ij} \ln(p_j/M) + \sum_{k=1}^{K} \eta_{ik} A_k}{\sum_{k=1}^{N} \alpha_k + \sum_{k=1}^{N} \sum_{j=1}^{N} \beta_{kj} \ln(p_j/M) + \sum_{i=1}^{N} \sum_{k=1}^{K} \eta_{ik} A_k}. \quad (3.16)$$

The assumption of utility maximizing behavior does not impose any restrictions on the $\eta_{ik}$. To compute the Slutsky matrix for this demand system, the variable $D(P, M)$ in Eq. (3.3) now becomes $D(P, M, A)$ and is given by the denominator of Eq. (3.16).

For the IQUAIDS there is no straightforward way to include demographics in the demand system. Because it results in additive demographics in the share equations, we chose the following re-formulation of the $a(P)$ function to include demographics:

$$\ln(a(P, A)) = \delta_0 + \sum_{i=1}^{N}\sum_{k=1}^{K} \phi_{ik} A_k \ln(p_i) + \sum_{i=1}^{N} \delta_i \ln(p_i) + \frac{1}{2}\sum_{i=1}^{N}\sum_{j=1}^{N} \gamma_{ij} \ln(p_i) \ln(p_j). \quad (3.17)$$

Allowing the household demographic characteristics to enter in this manner requires the following restrictions on the $\phi_{ik}$ in order for the summability restriction to hold:

$$\sum_{i=1}^{N} \phi_{ik} = 0 \text{ for all } k = 1,2,\ldots,K.$$

The equations for the Slutsky matrix defined above continue to hold with $a(P, A)$ in place of $a(P)$.

To estimate our econometric demand systems, we must specify a stochastic structure which accounts for differences between the observed expenditure shares and those predicted by our two functional forms for the household's indirect utility function. To allow for these differences, we append to the share equation for either model an additive mean zero error, $\epsilon_i$, which can be contemporaneously correlated with the errors from the other share equations for a given household, yet are independently distributed across households. If $\epsilon = (\epsilon_1, \epsilon_2, \ldots, \epsilon_N)$ is the vector of additive share equation disturbances for a given household, our assumption is that $E(\epsilon)=0$ and $E(\epsilon\epsilon')=\Omega(P,M)$, where $\Omega(P,M)$ is a matrix which depends on the prices and total expenditure associated with that observation. We interpret $\epsilon$ as the unobservable (to the econometrician) portion of the household's indirect utility function, so that the general form for the household-level indirect utility function is $V(P, M, A, \epsilon)$. Brown and Walker (1989) show that, in general, this interpretation of the stochastic structure of the demand system and associated indirect utility function implies that the vector of additive share equation disturbances, $\epsilon$, must be heteroscedastic conditional on prices and total expenditure.

Choosing a specific functional form for the way in which $\epsilon$ enters the indirect utility function would imply a specific functional form for the dependence of $\Omega$ on $P$ and $M$. This dependence could be exploited to yield more efficient estimates of the parameters of the model. However, this approach has the drawback that if our functional form for the $\Omega(P,M)$ matrix is incorrect, this fact can adversely effect the consistency of all of the estimated parameters of the indirect utility function. Consequently, our strategy is to acknowledge the results of Brown and Walker (1989) and that there is an unobservable portion of the household's indirect utility function so that it takes the form $V(P, M, A, \epsilon)$. However, we do not explicitly model how $\epsilon$ enters the indirect utility function. Instead, we only require that it enters in a way that yields additive disturbances to the share equations which satisfy the moment restrictions given above for $\epsilon$. To investigate the empirical importance of these considerations, we will test for the existence of heteroscedastic disturbances to the share equations conditional on $P$ and $M$. If there is evidence for the dependence of $\Omega$ on these variables, rather than selecting a parametric model for this dependence and re-estimating the model, our strategy is instead to construct standard error estimates which are consistent in the presence of this form of heteroscedasticity. As a result, although our parameter estimates will be less efficient, all of our inferences will be based on asymptotically valid standard error estimates. Given our relatively large sample size, this seems to be the best research strategy to balance our competing goals of parametric flexibility and consistent estimation of the parameters of the model.

The final issue associated with making the stochastic structure consistent with utility maximization is that summability of the observed budget shares implies that the sum of the $\epsilon_i$ over all of the goods is identically zero. In terms of our earlier notation, this restriction is $\iota'\epsilon=0$, which implies $\iota'\Omega(P,M)\iota=0$, so that $\Omega(P,M)$ is a matrix of rank $N-1$. To estimate this model we drop one of the share equations and estimate the parameters of the demand system using the remaining $N-1$ share equations. So long as we utilize the Gaussian quasi-maximum likelihood approach to estimate the model, the parameter estimates will be invariant to the share equation that we drop from our model. In addition, so long as certain higher order moments of $\epsilon$ exist, this estimation procedure will yield consistent estimates of the parameters of the demand system even if the distribution of $\epsilon$ is non-Gaussian.

Gourieroux et al. (1984) prove the consistency of these quasi- (or pseudo) maximum likelihood estimators for the parameters of our demand system, and White (1982) provides consistent standard error estimates for case in which $\Omega$ depends on $P$ and $M$ under general distributional assumptions for $\epsilon$. Using these results, we have a general estimation strategy which is consistent with utility maximizing behavior yet does not require a specific distributional assumption for $\epsilon$ or a parametric form for the dependence of $\Omega$ on $P$ and $M$, yet still allows asymptotically valid inferences to be made about the parameters of the demand systems.

## 4. Estimation results

This section describes the estimation of both the translog and IQUAIDS models. We first describe the specific household characteristics entering into our model and then proceed with a discussion of our estimation procedure and its output. We then perform our separability test for the translog specification, because this restriction can be examined in a data-independent manner for the translog demand system only.

In our choice of household characteristics to include in the vector $A = (A_1, A_2, \ldots, A_K)$ we attempt to control for across-household differences in the preferences for goods that do not depend on the prices faced by the household or its total expenditure. We include the age and age-squared of the head of household, because we expect cohort differences in the demand for telephone service. We also include dummies for whether the household contains a non-working head and a dummy for whether or not the household contains a non-working spouse, because there is strong evidence, for example in Browning and Meghir (1991), that employment status affects consumption patterns. We also include variables measuring the annual hours of work for both the head and spouse, because we expect household preferences to depend on the extent of attachment both the head and spouse have to the labor market.

To account for both the differences in the geographic location of the household to the extent possible given confidentiality constraints and the fact mentioned earlier that the base period relative prices of food, clothing and other non-durable goods across the regions is unobserved, we include a dummy for each of the four Census regions, with the excluded type of household being those in rural areas of the U.S. Because we believe that a household's demographic composition will influence its preferences, we include variables measuring the total number of persons in the household, the number of persons 65 years old and over, the number of males between 2 and 15 years old, and the number of females 2 to 15 years old. Finally, we include dummy variables for race and educational status. Specifically, we include dummies for whether the head of the household is white, whether the head is female, whether the head is a high school graduate, whether the head is a college graduate, and whether the head is single or married. We also include a dummy variable for whether the head works in a professional occupation. While there are other household characteristics we could include in our model, based on our preliminary model estimations, the variables we selected appeared sufficient, relative to models with more household characteristics included, to explain much of the across-household differences in consumption not due to differences in prices and total expenditure.

As discussed in the previous section, the estimation procedure utilizes the multivariate normal quasi-likelihood function. Define $y_j$ to be the $(N-1)$-dimensional vector of expenditure shares for the $j$th household and $f_j(P^j, M^j, A^j, \theta)$ to be the $(N-1)$ dimensional vector of fitted expenditure shares which are functions of prices, expenditure, and household characteristics for the $j$th household for the translog or IQUAIDS model. Let $J$ denote the total number of households in our sample. In this notation $\theta$ denotes the vector of parameters for the demand system to be estimated. To compute the quasi-maximum likelihood (QML) estimate of $\theta$ we maximize

$$L(\theta, \Sigma) = -J(N-1)\ln(2\pi) - \frac{J}{2}\ln\det(\Sigma)$$
$$-\sum_{j=1}^{J}\frac{1}{2}(y_j - f_j(P^j, M^j, A^j))'\Sigma^{-1}(y_j - f_j(P^j, M^j, A^j)).$$

with respect to $\theta$ and the elements of the matrix $\Sigma$.

To test the null hypothesis of homoscedastic disturbances to the share equations, we perform the Breusch and Pagan (1979) Lagrange Multiplier test for homoscedasticity against the alternative that the variance of the disturbances to each share equation depends on prices and total expenditure as is implied by the results of Brown and Walker (1989) discussed earlier. This test is implemented by taking the residuals from the QML estimation of each of the share equations and regressing these residuals squared on a constant and the log of prices and log of total expenditure and all of the unique cross products of these log prices and log of total expenditures. Taking $J$ times the $R^2$ from this regression yields the LM statistic, which is asymptotically distributed as $\chi^2$ with degrees of freedom equal to $\frac{1}{2}((N+1)^2 - (N+1)) + 2(N+1)$, where $N$ is the number of goods in our model. For our $N=5$ good model, this number is 27.

Table 1 presents the QML estimates of the parameters of translog demand system in terms of the notation for the translog indirect utility function given in Eq. (3.15). Table 2 contains the QML parameters estimates for the IQUAIDS model in terms of the notation for the IQUAIDS indirect utility function given in Eqs. (3.5), (3.7) and (3.17). Both the translog and IQUAIDS models are estimated with the summability, homogeneity and symmetry restrictions imposed so that the resulting demand systems can be used to perform welfare calculations. Despite the price data used, both models yield fairly precisely estimated own-price effect

Table 1
Coefficient estimates for translog model
(L = local, D = long-distance, F = Food, C = clothing, O = other, and M = total expenditure)
$\beta_{Mi} = \sum_{j=1}^{N} \beta_{ij}$, assuming $\beta_{ij} = \beta_{ji}$

| Coefficient | Estimate | Standard error |
|---|---|---|
| $\alpha_L$ | −1.86e−02 | 1.54e−03 |
| $\alpha_D$ | −3.49e−02 | 3.91e−03 |
| $\alpha_F$ | −2.77e−01 | 1.88e−02 |
| $\alpha_C$ | −1.52e−01 | 1.20e−02 |
| $\beta_{LL}$ | −2.07e−03 | 1.12e−02 |
| $\beta_{LD}$ | −1.02e−03 | 4.71e−03 |
| $\beta_{LF}$ | −6.46e−03 | 1.01e−02 |
| $\beta_{LC}$ | −5.27e−03 | 5.38e−03 |
| $\beta_{DD}$ | 1.90e−02 | 5.92e−03 |
| $\beta_{DF}$ | −3.48e−02 | 1.41e−02 |
| $\beta_{DC}$ | −1.26e−02 | 8.49e−03 |
| $\beta_{FF}$ | 1.22e−01 | 6.43e−02 |
| $\beta_{FC}$ | −6.71e−02 | 3.13e−02 |
| $\beta_{CC}$ | 1.07e−01 | 2.77e−02 |
| $\beta_{ML}$ | −9.92e−03 | 1.54e−03 |
| $\beta_{MD}$ | 4.07e−04 | 7.62e−04 |
| $\beta_{MF}$ | 3.42e−02 | 8.62e−03 |
| $\beta_{MC}$ | 4.94e−02 | 7.14e−03 |
| $\beta_{MO}$ | 1.59e−01 | 2.31e−02 |
| $\eta_{Lk}$, $k = 1,2,\ldots,K$ (Demographics for local share equation) | | |
| Age of Head | −8.11e−04 | 1.02e−02 |
| (Age of Head)$^2$ | −6.61e−04 | 1.04e−02 |
| Number of Family Members | −1.12e−03 | 3.02e−04 |
| Members ≥65 years old | −3.48e−04 | 7.26e−04 |
| Head White (Dummy) | 7.32e−04 | 6.96e−04 |
| Head Female (Dummy) | 4.42e−06 | 7.80e−04 |
| Head College Graduate (Dummy) | 1.32e−03 | 1.11e−03 |
| Head HS Graduate (Dummy) | 3.77e−04 | 7.24e−04 |
| Head Single (Dummy) | 7.91e−05 | 1.25e−03 |
| Head Professional (Dummy) | 8.59e−05 | 7.79e−04 |
| Head Hours Worked per Year | −4.19e−03 | 4.81e−03 |
| Spouse Hours Worked per Year | −1.13e−04 | 4.78e−03 |
| Head Non-Worker (Dummy) | 1.98e−03 | 1.10e−03 |
| Spouse Non-Worker (Dummy) | 1.99e−04 | 1.15e−03 |
| Males Age 2 through 15 | 1.16e−03 | 4.23e−04 |
| Females Age 2 through 15 | 6.17e−04 | 3.74e−04 |
| North East (Dummy) | 3.20e−03 | 1.46e−03 |
| North Central (Dummy) | 1.85e−03 | 1.07e−03 |
| South (Dummy) | 2.22e−03 | 1.10e−03 |
| West (Dummy) | 3.18e−03 | 1.06e−03 |
| $\eta_{Dk}$, $k = 1,2,\ldots,K$ (Demographics for long−distance share equation) | | |
| Age of Head | 3.85e−02 | 1.25e−02 |
| (Age of Head)$^2$ | −1.57e−02 | 1.23e−02 |

Table 1 (Contd.)

| Coefficient | Estimate | Standard error |
|---|---|---|
| Number of Family Members | −3.11e−03 | 7.19e−04 |
| Members ≥65 years old | −3.13e−04 | 8.39e−04 |
| Head White (Dummy) | −1.93e−04 | 1.09e−03 |
| Head Female (Dummy) | −2.92e−03 | 1.20e−03 |
| Head College Graduate (Dummy) | −3.29e−03 | 1.74e−03 |
| Head HS Graduate (Dummy) | −3.93e−04 | 8.93e−04 |
| Head Single (Dummy) | −1.41e−03 | 2.00e−03 |
| Head Professional (Dummy) | 6.57e−04 | 1.16e−03 |
| Head Hours Worked per Year | 9.62e−03 | 7.08e−03 |
| Spouse Hours Worked per Year | 9.92e−03 | 7.18e−03 |
| Head Non-Worker (Dummy) | 5.55e−03 | 1.82e−03 |
| Spouse Non-Worker (Dummy) | 6.57e−05 | 1.71e−03 |
| Males Age 2 through 15 | 4.94e−03 | 1.04e−03 |
| Females Age 2 through 15 | 3.88e−03 | 8.84e−04 |
| North East (Dummy) | 1.06e−02 | 2.16e−03 |
| North Central (Dummy) | 9.38e−03 | 1.71e−03 |
| South (Dummy) | 7.68e−03 | 1.69e−03 |
| West (Dummy) | 2.96e−03 | 1.79e−03 |
| $\eta_{Fk}$, 1,2, . . . ,$K$ (Demographics for food share equation) | | |
| Age of Head | −1.21e−03 | 1.49e−01 |
| (Age of Head)$^2$ | −8.22e−03 | 1.44e−01 |
| Number of Family Members | −1.94e−02 | 5.40e−03 |
| Members ≥65 years old | −1.35e−02 | 1.22e−02 |
| Head White (Dummy) | −3.88e−02 | 1.18e−02 |
| Head Female (Dummy) | 3.56e−02 | 1.31e−02 |
| Head College Graduate (Dummy) | 3.56e−02 | 1.79e−02 |
| Head HS Graduate (Dummy) | 1.47e−02 | 1.07e−02 |
| Head Single (Dummy) | −5.43e−03 | 2.45e−02 |
| Head Professional (Dummy) | 6.18e−03 | 1.36e−02 |
| Head Hours Worked per Year | 7.60e−03 | 7.79e−02 |
| Spouse Hours Worked per Year | −1.13e−02 | 9.74e−02 |
| Head Non-Worker (Dummy) | 3.07e−02 | 1.73e−02 |
| Spouse Non-Worker (Dummy) | −1.34e−02 | 2.40e−02 |
| Males Age 2 through 15 | 1.04e−02 | 8.60e−03 |
| Females Age 2 through 15 | 7.64e−04 | 6.32e−03 |
| North East (Dummy) | −1.81e−02 | 2.93e−02 |
| North Central (Dummy) | 1.58e−02 | 1.82e−02 |
| South (Dummy) | 2.95e−02 | 1.77e−02 |
| West (Dummy) | −2.46e−02 | 2.20e−02 |
| $\eta_{Ck}$, $k=1,2,\ldots,K$ (Demographics for clothing share equation) | | |
| Age of Head | 2.16e−01 | 3.44e−02 |
| (Age of Head)$^2$ | −1.64e−01 | 3.16e−02 |
| Number of Family Members | 1.94e−03 | 1.25e−03 |
| Members ≥65 years old | −6.07e−04 | 2.63e−03 |
| Head White (Dummy) | 5.46e−03 | 2.78e−03 |
| Head Female (Dummy) | −2.05e−02 | 4.40e−03 |
| Head College Graduate (Dummy) | −4.65e−03 | 5.27e−03 |

Table 1 (Contd.)

| Coefficient | Estimate | Standard error |
|---|---|---|
| Head HS Graduate (Dummy) | 4.72e−05 | 2.52e−03 |
| Head Single (Dummy) | 1.48e−02 | 6.39e−03 |
| Head Professional (Dummy) | −1.37e−02 | 4.33e−03 |
| Head Hours Worked per Year | −2.74e−02 | 2.45e−02 |
| Spouse Hours Worked per Year | 3.42e−02 | 2.80e−02 |
| Head Non-Worker (Dummy) | 2.64e−03 | 4.80e−03 |
| Spouse Non-Worker (Dummy) | 9.10e−03 | 6.26e−03 |
| Males Age 2 through 15 | 4.68e−03 | 2.35e−03 |
| Females Age 2 through 15 | −1.80e−03 | 2.11e−03 |
| North East (Dummy) | −9.77e−03 | 7.15e−03 |
| North Central (Dummy) | −7.19e−03 | 4.71e−03 |
| South (Dummy) | 3.60e−04 | 4.74e−03 |
| West (Dummy) | 3.60e−04 | 4.74e−03 |
| $\eta_{Ok}$, $k = 1,2,\ldots,K$ (Demographics for other share equation) | | |
| Age of Head | 1.16e−01 | 1.73e−01 |
| (Age of Head)$^2$ | −2.12e−01 | 1.65e−01 |
| Number of Family Members | 3.23e−03 | 4.85e−03 |
| Members ≥65 years old | −1.89e−02 | 1.59e−02 |
| Head White (Dummy) | −5.26e−02 | 1.47e−02 |
| Head Female (Dummy) | −2.98e−03 | 1.58e−02 |
| Head College Graduate (Dummy) | 1.38e−02 | 2.43e−02 |
| Head HS Graduate (Dummy) | −9.38e−03 | 1.33e−02 |
| Head Single (Dummy) | 4.45e−02 | 3.04e−02 |
| Head Professional (Dummy) | 7.79e−03 | 1.78e−02 |
| Head Hours Worked per Year | 4.22e−02 | 1.01e−01 |
| Spouse Hours Worked per Year | 6.30e−02 | 1.33e−01 |
| Head Non-Worker (Dummy) | 7.12e−02 | 2.23e−02 |
| Spouse Non-Worker (Dummy) | 2.53e−02 | 3.01e−02 |
| Males Age 2 through 15 | 2.19e−02 | 9.68e−03 |
| Females Age 2 through 15 | 5.90e−03 | 7.61e−03 |
| North East (Dummy) | 3.81e−02 | 3.68e−02 |
| North Central (Dummy) | 5.77e−02 | 2.44e−02 |
| South (Dummy) | 8.40e−02 | 2.46e−02 |
| West (Dummy) | 5.11e−02 | 2.60e−02 |

coefficients and some precisely estimated cross-price effects. The small standard errors relative to coefficient estimates for the majority of demographic variables implies that these variables substantially improve the explanatory power of both demand systems. A Wald test of the null hypothesis that all demographic variables do not enter any of the share equations is overwhelmingly rejected for both models. In Table 3 we report the LM statistics against heteroscedasticity of the elements of $\epsilon$ conditional on the log of prices and log total expenditure for all five share equations for both models. For all share equations in both models, the test statistic is substantially larger than the critical value for any conventional size

Table 2
Coefficient estimates for IQUAIDS model
(L = local, D=long-distance, F=Food, C=clothing, O=other, and M = total expenditure)

| Coefficient | Estimate | Standard error |
|---|---|---|
| $\alpha_L$ | 7.25e−02 | 2.71e−03 |
| $\alpha_D$ | 4.43e−02 | 4.50e−03 |
| $\alpha_F$ | 3.85e−01 | 1.68e−02 |
| $\alpha_C$ | 1.09e−01 | 1.14e−02 |
| $\gamma_{LL}$ | −1.66e−03 | 1.10e−02 |
| $\gamma_{LD}$ | 1.68e−03 | 4.64e−03 |
| $\gamma_{LF}$ | −1.58e−03 | 1.02e−02 |
| $\gamma_{LC}$ | 6.94e−03 | 5.24e−03 |
| $\gamma_{DD}$ | −2.31e−02 | 5.93e−03 |
| $\gamma_{DF}$ | 4.07e−02 | 1.48e−02 |
| $\gamma_{DC}$ | 1.79e−02 | 9.47e−03 |
| $\gamma_{FF}$ | −1.66e−01 | 7.04e−02 |
| $\gamma_{FC}$ | 1.11e−01 | 3.45e−02 |
| $\gamma_{CC}$ | −1.40e−01 | 2.77e−02 |
| $\beta_L$ | −4.37e−02 | 1.77e−03 |
| $\beta_D$ | −2.81e−03 | 2.49e−03 |
| $\beta_F$ | −8.41e−02 | 1.06e−02 |
| $\beta_C$ | 2.20e−02 | 6.54e−03 |
| $\lambda_L$ | 7.17e−03 | 4.29e−04 |
| $\lambda_D$ | −1.20e−03 | 6.01e−04 |
| $\lambda_F$ | 5.72e−03 | 2.85e−03 |
| $\lambda_C$ | 3.97e−03 | 1.90e−03 |
| $\eta_{Lk}$, $k=1,2,\ldots,K$ (Demographics for local share equation) | | |
| Age of Head | 2.05e−02 | 6.77e−03 |
| (Age of Head)$^2$ | −2.11e−02 | 6.94e−03 |
| Number of Family Members | 6.05e−04 | 1.58e−04 |
| Members ≥65 years old | −7.01e−04 | 3.36e−04 |
| Head White (Dummy) | −5.04e−03 | 5.05e−04 |
| Head Female (Dummy) | 3.30e−04 | 3.76e−04 |
| Head College Graduate (Dummy) | −3.77e−04 | 4.36e−04 |
| Head HS Graduate (Dummy) | −1.31e−04 | 3.86e−04 |
| Head Single (Dummy) | 5.84e−04 | 5.13e−04 |
| Head Professional (Dummy) | 8.66e−06 | 3.27e−04 |
| Head Hours Worked per Year | 6.49e−03 | 2.34e−03 |
| Spouse Hours Worked per Year | 3.34e−04 | 1.80e−03 |
| Head Non-Worker (Dummy) | 2.16e−03 | 6.39e−04 |
| Spouse Non-Worker (Dummy) | −3.07e−04 | 4.47e−04 |
| Males Age 2 through 15 | −6.40e−05 | 2.32e−04 |
| Females Age 2 through 15 | −3.21e−04 | 2.35e−04 |
| North East (Dummy) | −2.77e−03 | 6.32e−04 |
| North Central (Dummy) | 2.29e−04 | 4.13e−04 |
| South (Dummy) | 1.53e−03 | 4.15e−04 |
| West (Dummy) | −2.28e−03 | 4.75e−04 |

Table 2 (Contd.)

| Coefficient | Estimate | Standard error |
|---|---|---|
| $\eta_{Dk}$, $k=1,2,\ldots,K$ (Demographics for long-distance share equation) | | |
| Age of Head | −2.22e−02 | 1.28e−02 |
| (Age of Head)$^2$ | −6.08e−03 | 1.28e−02 |
| Number of Family Members | 2.85e−03 | 5.09e−04 |
| Members ≥65 years old | −8.45e−04 | 7.05e−04 |
| Head White (Dummy) | −4.18e−03 | 1.21e−03 |
| Head Female (Dummy) | 3.61e−03 | 8.60e−04 |
| Head College Graduate (Dummy) | 5.08e−03 | 1.03e−03 |
| Head HS Graduate (Dummy) | 8.16e−04 | 7.87e−04 |
| Head Single (Dummy) | 3.23e−03 | 1.33e−03 |
| Head Professional (Dummy) | −2.60e−04 | 7.97e−04 |
| Head Hours Worked per Year | −6.72e−03 | 5.50e−03 |
| Spouse Hours Worked per Year | −9.54e−03 | 4.81e−03 |
| Head Non-Worker (Dummy) | −1.38e−03 | 1.57e−03 |
| Spouse Non-Worker (Dummy) | 7.08e−04 | 1.12e−03 |
| Males Age 2 through 15 | −3.96e−03 | 7.15e−04 |
| Females Age 2 through 15 | −4.22e−03 | 6.80e−04 |
| North East (Dummy) | −1.06e−02 | 1.12e−03 |
| North Central (Dummy) | −8.23e−03 | 1.01e−03 |
| South (Dummy) | −4.25e−03 | 1.10e−03 |
| West (Dummy) | −2.29e−03 | 1.16e−03 |
| $\eta_{Fk}$, $k=1,2,\ldots,$K (Demographics for food share equation) | | |
| Age of Head | 2.53e−01 | 5.53e−02 |
| (Age of Head)$^2$ | −2.62e−01 | 5.96e−02 |
| Number of Family Members | 1.44e−02 | 1.63e−03 |
| Members ≥65 years old | 7.24e−04 | 3.31e−03 |
| Head White (Dummy) | 6.70e−03 | 4.02e−03 |
| Head Female (Dummy) | −3.84e−02 | 3.41e−03 |
| Head College Graduate (Dummy) | −2.21e−02 | 4.14e−03 |
| Head HS Graduate (Dummy) | −1.57e−02 | 3.51e−03 |
| Head Single (Dummy) | 2.54e−02 | 5.27e−03 |
| Head Professional (Dummy) | −5.64e−03 | 2.95e−03 |
| Head Hours Worked per Year | −6.22e−03 | 1.87e−02 |
| Spouse Hours Worked per Year | 5.04e−02 | 2.07e−02 |
| Head Non-Worker (Dummy) | 1.63e−02 | 5.57e−03 |
| Spouse Non-Worker (Dummy) | 2.08e−02 | 4.58e−03 |
| Males Age 2 through 15 | 5.38e−03 | 2.48e−03 |
| Females Age 2 through 15 | 2.09e−03 | 2.40e−03 |
| North East (Dummy) | 2.72e−02 | 5.10e−03 |
| North Central (Dummy) | 1.29e−02 | 3.85e−03 |
| South (Dummy) | 1.91e−02 | 3.96e−03 |
| West (Dummy) | 4.16e−02 | 4.39e−03 |
| $\eta_{Ck}$, $k=1,2,\ldots,K$ (Demographics for clothing share equation) | | |
| Age of Head | −1.83e−01 | 3.26e−02 |
| (Age of Head)$^2$ | 1.04e−01 | 3.28e−02 |
| Number of Family Members | −4.97e−03 | 1.01e−03 |

Table 2 (Contd.)

| Coefficient | Estimate | Standard error |
|---|---|---|
| Members ≥65 years old | 5.69e−04 | 1.97e−03 |
| Head White (Dummy) | −1.80e−02 | 2.74e−03 |
| Head Female (Dummy) | 2.60e−02 | 2.20e−03 |
| Head College Graduate (Dummy) | 1.06e−02 | 2.71e−03 |
| Head HS Graduate (Dummy) | 2.70e−03 | 2.08e−03 |
| Head Single (Dummy) | −1.20e−02 | 3.54e−03 |
| Head Professional (Dummy) | 1.47e−02 | 1.96e−03 |
| Head Hours Worked per Year | 3.44e−02 | 1.34e−02 |
| Spouse Hours Worked per Year | −3.12e−02 | 1.45e−02 |
| Head Non-Worker (Dummy) | 3.86e−03 | 3.46e−03 |
| Spouse Non-Worker (Dummy) | −5.85e−03 | 3.09e−03 |
| Males Age 2 through 15 | −9.21e−05 | 1.55e−03 |
| Females Age 2 through 15 | 3.12e−03 | 1.59e−03 |
| North East (Dummy) | 1.15e−02 | 3.28e−03 |
| North Central (Dummy) | 1.42e−02 | 2.33e−03 |
| South (Dummy) | 6.37e−03 | 2.47e−03 |
| West (Dummy) | −1.98e−03 | 2.82e−03 |
| $\eta_{Ok}$, $k = 1,2,\ldots,K$ (Demographics for other share equation) | | |
| Age of Head | −6.83e−02 | 5.74e−02 |
| (Age of Head)$^2$ | 1.85e−01 | 6.17e−02 |
| Number of Family Members | −1.29e−02 | 1.64e−03 |
| Members ≥65 years old | 2.52e−04 | 3.51e−03 |
| Head White (Dummy) | 2.05e−02 | 4.26e−03 |
| Head Female (Dummy) | 8.48e−03 | 3.67e−03 |
| Head College Graduate (Dummy) | 6.82e−03 | 4.46e−03 |
| Head HS Graduate (Dummy) | 1.23e−02 | 3.67e−03 |
| Head Single (Dummy) | −1.72e−02 | 5.64e−03 |
| Head Professional (Dummy) | −8.78e−03 | 3.20e−03 |
| Head Hours Worked per Year | −2.80e−02 | 2.03e−02 |
| Spouse Hours Worked per Year | −1.00e−02 | 2.24e−02 |
| Head Non-Worker (Dummy) | −2.10e−02 | 5.90e−03 |
| Spouse Non-Worker (Dummy) | −1.54e−02 | 4.93e−03 |
| Males Age 2 through 15 | −1.26e−03 | 2.60e−03 |
| Females Age 2 through 15 | −6.63e−04 | 2.57e−03 |
| North East (Dummy) | −2.53e−02 | 5.51e−03 |
| North Central (Dummy) | −1.91e−02 | 4.14e−03 |
| South (Dummy) | −2.27e−02 | 4.25e−03 |
| West (Dummy) | −3.51e−02 | 4.58e−03 |

hypothesis test, providing strong evidence against the null hypothesis of homoscedasticity versus the alternative hypothesis of heteroscedasticity conditional on the log of prices and log of total expenditure as is implied by including $\epsilon$ as an unobservable vector of household-level variables in the indirect utility function in manner consistent with utility maximizing behavior. For this reason, the standard errors given in Table 1 and Table 2 are computed from the heteroscedasticity-

Table 3
Heteroscedasticity LM test statistics, distributed as $\chi^2_{25}$ under homoscedasticity null hypothesis[a]

| Share equation | Translog model | IQUAIDS model |
|---|---|---|
| Local expenditure | 268.38 | 307.39 |
| Long-distance expenditure | 101.94 | 101.46 |
| Food expenditure | 288.16 | 305.89 |
| Clothing expenditure | 164.80 | 161.78 |
| Other expenditure | 288.08 | 314.13 |

[a] $\chi^2_{27,\alpha=0.05} = 40.11$, $\chi^2_{27,\alpha=0.01} = 46.96$.

consistent QML covariance matrix estimates given in White (1982). Because of the presence of heteroscedasticity, comparing values of the QML objective function cannot be used to compute valid test statistics, so that all hypotheses will be examined using Wald statistics based on this covariance matrix estimate.

We now discuss our test for homothetic separability of local and long-distance phone service consumption from food, clothing, and all other nondurable goods in the household-level direct utility function for the translog model. Our motivation for testing this restriction is to examine whether or not the household-level direct utility function, $U(x_1, x_2, \ldots, x_N)$, can be written as $U(g(x_1, x_2), x_3, \ldots, x_N)$, where $x_i$ is the quantity consumed of good $i$ and $g(.,.)$ is a homothetic aggregator function. Assuming $x_1$ and $x_2$ are the amount of local and long-distance service consumed, $g(x_1, x_2)$ represents the telephone service aggregate good. If this restriction on the household-level direct utility function holds, then the consumer's demand problem can be thought of in the two-stage budgeting context, where the household first determines its demands for the telephone aggregate and the other three goods using a price index for this telephone aggregate and the prices of the other goods; then conditional of the total amount of telephone expenditures determined from this first-stage problem, the household solves for the optimal local versus long-distance split by maximizing $g(x_1, x_2)$ subject to $p_1x_1 + p_2x_2 = M_T$, where $M_T$ is total telephone expenditures determined from the first-stage optimization problem. Homothetic separability of household-level direct utility function is implicit in any analysis of the demand for local and long-distance service which ignores the household's demand (or the prices) for all other goods consumed. Without this restriction on its utility function, the household cannot solve for the optimal local versus long-distance split without regard to its optimal demand for all other goods. Most existing analyses of telecommunications demand make this two-stage budgeting assumption. Our household-level dataset with many goods provides an ideal opportunity to investigate its empirical validity. See Taylor (1994) for more evidence in favor of the importance of examining the validity of this restriction on household preferences.

Although the restriction of homothetic separability given above is specified in terms of the household's direct utility function, Theorem 4.4 of Blackorby et al. (1978) shows that homothetic separability of the direct utility function is equivalent to homothetic separability of the indirect utility function. Therefore, we perform our test of homothetic separability on the translog indirect utility function. In terms of the parameters of the translog function, global imposition (for all values of prices and total expenditure) of this restriction implies

$$\frac{\alpha_1}{\alpha_2} = \frac{\gamma_{1i}}{\gamma_{2i}} \text{ for } i = 1,2,\ldots,N.$$

This test involves $N=5$ nonlinear restrictions on the parameters of the translog demand system. We investigate it validity using a Wald test. The test statistic for this null hypothesis is 25.07, which is larger than the $\chi^2_5$ critical value for all conventional test sizes. We should caution that this rejection of homothetic separability, and therefore the validity of the two-stage budgeting assumption, is conditional on our maintained hypothesis of a translog indirect utility function. In addition, as discussed in Blackorby et al. (1977) imposing homothetic separability on the translog function destroys its second-order flexibility property. Unfortunately, this property – global imposition of separability destroys second-order flexibility – is shared by all other existing flexible functional forms that have been used to test for separability. Although we can reject homothetic separability conditional on the assumed translog indirect utility function, we do not know if this is due to a misspecification of the true underlying utility function or because the homothetic separability null hypothesis is false. We are unable to rigorously test for homothetic separability in the IQUAIDS model because there does not exist a way to impose this restriction for all prices and total expenditures. An informal test in terms of the household-level income and price elasticities performed for all points in the sample found substantial deviations from this null hypothesis for the IQUAIDS model as well.

Given these separability results, for both the translog and IQUAIDS models we use the demand system estimates which do not impose these separability restrictions to perform the welfare calculations and consumption simulations given in the next section.

## 5. Assessing the welfare impacts of price increases for local telephone service

In this section, we utilize the two integrable demand systems estimated to assess the impact of various price change scenarios for local and long-distance service on household consumption patterns and consumer welfare. We consider four price change scenarios. Two involve increases in the price of local service alone. The second two involve price increases in local service accompanied by equivalent decreases in the price of long-distance service. The first two scenarios attempt to capture the range of impacts of the current proposals for increases in the price of local service. The second two attempt to assess the likely impacts of balancing this local service price increase with a corresponding decrease in the price of long-distance service as might be expected if long-distance access charges are reduced in response to local price increases, as was done in connection with the recent increase in the price of local service in California discussed earlier. Our framework also allows us to assess the regressivity of these proposed price increases as well as determine which types of households, as measured by their observable characteristics, bear a greater portion of the burden of these price changes.

Before presenting the results of these simulations, we discuss the sample average own- and cross-price and total expenditure elasticity estimates given in Table 4 for the translog and IQUAIDS model. These mean elasticity estimates are surprisingly similar across the two indirect utility functional forms. Both the household-level mean own-price and income elasticities for local service are small in absolute value. For both models, the mean income elasticity is slightly positive, whereas the 5th to 95th percentile ranges contains small positive and negative values. For long-distance service we find, at the mean values for our sample, a very price-elastic demand for long-distance service and a higher income elasticity than for local service. Particularly, for long-distance telephone service and clothing, there is substantial heterogeneity across households in these elasticities as is indicated by the 5th percentile to 95th percentile ranges given in Table 4. This heterogeneity is indicative of the potential for substantial differences in the welfare and consumption impacts of these price changes across households.

The four price change scenarios we consider are: (1) a 20 percent increase in the price of local service alone, (2) a 40 percent increase in the price of local service alone, (3) a 20 percent increase in the price of local service accompanied by a 20 percent decrease in the price of long-distance service, and (4) a 40 percent increase in the price of local service accompanied by a 40 percent decrease in the

Table 4
Own price and expenditure elasticity estimates
(L = local, D = long-distance, F = food, C = clothing, O = other, M = expenditure)
Mean, 5th percentile, and 95th percentile values

| Elasticity | Translog | | | IQUAIDS | | |
|---|---|---|---|---|---|---|
| | Mean | 5th% | 95th% | Mean | 5th% | 95th% |
| $e_{LL}$ | −0.88 | −0.94 | −0.79 | −0.95 | −0.98 | −0.87 |
| $e_{DD}$ | −2.07 | −3.02 | −1.62 | −2.17 | −3.06 | −1.61 |
| $e_{FF}$ | −1.35 | −1.49 | −1.27 | −1.39 | −1.48 | −1.32 |
| $e_{CC}$ | −2.85 | −4.99 | −1.70 | −2.52 | −3.54 | −1.96 |
| $e_{OO}$ | −1.00 | −1.03 | −0.98 | −0.98 | −1.02 | −0.95 |
| $e_{ML}$ | 0.11 | −0.28 | 0.33 | 0.16 | −0.18 | 0.56 |
| $e_{MD}$ | 0.75 | 0.63 | 0.82 | 0.60 | 0.23 | 0.85 |
| $e_{MF}$ | 0.84 | 0.74 | 0.89 | 0.82 | 0.79 | 0.85 |
| $e_{MC}$ | 1.61 | 1.12 | 2.51 | 1.38 | 1.26 | 1.59 |
| $e_{MO}$ | 1.11 | 1.06 | 1.22 | 1.11 | 1.04 | 1.19 |

price of long-distance service. For each of these price change scenarios we compute both the local and long-distance expenditures expected as a result of these price changes. For comparison, we also compute the expected expenditures on both goods at the current prices faced by the household.

For our welfare analysis we also compute the compensating variation associated with each of these price changes. Assuming $P^0$ is the initial vector of prices, $M^0$ is initial total expenditure, $P^1$ is the proposed vector of prices, and $V(P, M, A)$ is the indirect utility for a household facing prices $P$, and having total expenditure $M$ and characteristics $A$, the compensating variation is the value of $CV$ which solves the following nonlinear equation:

$$V(P^1, M^0 + CV, A) = V(P^0, M^0, A).$$

In words, $CV$ is the amount of additional total expenditure which must be paid to a household with characteristics $A$ in order for it to be indifference between total expenditure $M^0 + CV$ and prices $P^1$ and total expenditure $M^0$ and prices $P^0$. To get an idea of the magnitude of the welfare burden of these price changes, we compute $CV/M^0$ for each observation in our sample. Examining how this ratio changes with $M^0$ allows us to determine the extent of the regressivity or progressivity of these price change scenarios. Finally, in order to assess how this burden relates to the characteristics of the household, we estimate best linear predictor functions for $CV/M^0$ as a function of our household characteristics and total expenditure.

For all of these price change scenarios to yield theoretically valid household-level consumption changes and compensating variations, the estimated demand system must satisfy all the restrictions implied by utility maximization behavior at the prices, total expenditures, and household characteristics for that observation. Because we estimate both the translog and IQUAIDS models with summability, homogeneity, and symmetry imposed, we only need to check that negative semidefiniteness of the Slutsky matrix holds at each observation. For the translog model, this process involves computing the Slutsky matrix given in Eq. (3.3) for each observation and then computing the five eigenvalues of this matrix and verifying whether they are all nonpositive. For the IQUAIDS model we must compute the matrix described following (3.12) for each observation and repeat this process. As discussed earlier, because there are no necessary and sufficient conditions on the parameters of either the translog or IQUAIDS model which guarantee the Slutsky matrix is negative semidefinite for all prices and expenditures, we must follow this procedure to select those observations that can be used to compute the telephone consumption changes and household-level welfare changes that result from our price-change scenarios. For the translog model we lose only a small percentage of observations due to failure of the negative semidefiniteness of the Slutsky matrix. Even for those observations that fail this restriction there is only one positive eigenvalue and this eigenvalue is a small fraction, less than 1/5, of the value of the smallest, in absolute value, negative eigenvalue. There are 11,346 observations out of the full sample of 11,467 observations which satisfy the necessary curvature restrictions. All of the calculations for the translog model in this section are based on this regular sample. For the IQUAIDS model we lose substantially more observations due to the failure these curvature restrictions. In this case, 8,245 of the observations satisfy the negative semidefiniteness of the Slutsky matrix. For this case as well, those observations that fail the restriction have just a single marginally positive eigenvalue that is a small fraction of the smallest, in absolute value, negative eigenvalue. All calculations for the IQUAIDS model presented in this section are for this restricted sample of regular observations.

Table 5 computes the mean, 5th percentile, and 95th percentile of the predicted local and long-distance expenditures as a result of each of the four price change scenarios for the translog and IQUAIDS models. The predicted expenditures at current prices for both models are included in the first row. This table quantifies several of the qualitative conclusions made based on the elasticity estimates. First, for both of the local price increase scenarios there is an increase in the mean expenditure on local phone service, with the larger increase for the 40 percent local service price increase. However, the other impact of this price increase is a

Table 5
Predicted telephone expenditure for observed prices and four price change scenarios (January 1988 dollars)[a]

| | Translog | | IQUAIDS | |
|---|---|---|---|---|
| Scenario | Local | Long-distance | Local | Long-distance |
| Panel A – Means | | | | |
| Observed prices | 51.23 | 64.97 | 52.66 | 70.70 |
| $P_L$ +20%, $P_D$ −0% | 52.31 | 65.43 | 53.17 | 71.92 |
| $P_L$ +20%, $P_D$ −20% | 51.72 | 76.31 | 51.83 | 85.65 |
| $P_L$ +40%, $P_D$ −0% | 53.22 | 65.82 | 53.60 | 72.94 |
| $P_L$ +40%, $P_D$ −40% | 52.13 | 85.87 | 51.10 | 98.26 |
| Table 5B – 5th% | | | | |
| Observed prices | 37.91 | 18.28 | 38.53 | 25.18 |
| $P_L$ +20%, $P_D$ −0% | 38.60 | 18.57 | 38.92 | 25.65 |
| $P_L$ +20%, $P_D$ −20% | 38.22 | 24.69 | 37.71 | 31.85 |
| $P_L$ +40%, $P_D$ −0% | 39.21 | 18.81 | 39.21 | 26.05 |
| $P_L$ +40%, $P_D$ −40% | 38.47 | 29.76 | 36.95 | 37.00 |
| Panel C – 95th% | | | | |
| Observed prices | 65.65 | 134.28 | 74.41 | 133.96 |
| $P_L$ +20%, $P_D$ −0% | 67.18 | 135.03 | 75.77 | 136.25 |
| $P_L$ +20%, $P_D$ −20% | 66.36 | 154.15 | 73.40 | 160.67 |
| $P_L$ +40%, $P_D$ −0%11 | 68.47 | 135.67 | 76.47 | 138.20 |
| $P_L$ +40%, $P_D$ −40% | 66.96 | 170.97 | 72.50 | 185.39 |

[a] $P_L$ = price of local service, $P_D$ = price of long-distance service.

slight increase in the consumption of long-distance service. For the local price increase and equal long-distance price decrease scenarios, there is an increase in the mean amount of local service consumed. The major difference from the local price-increase only scenario is the large relative increase in long-distance expenditures as a result of the corresponding percentage decrease in the price of long-distance service. For example, for the translog model and the 40 percent two-price change scenario, the mean long-distance expenditure increases by more than 30 percent, from \$64.97 to \$85.87. For the IQUAIDS model this effect is noticeably larger, from \$70.70 to \$98.26. This difference between the two models is due in part to the fact that most of the observations lost due to the failure of the integrability conditions of the IQUAIDS model are at the low end of the total expenditure distribution. The relative and absolute values of the 5th percentile base price versus price change scenario expenditure differences tend to be smaller than the corresponding values at the mean, and the analogous 95th percentile values tend to be larger.

Fig. 5 plots the kernel estimate of the density of the ratio of household-level compensating variation to total expenditure (*CV/M*) for the translog model estimates. Table 6 shows that the sample means of the quarterly compensating variation (in January 1988 dollars) for the four price change scenarios for the translog model are: Two-price 20%, −\$3.49; Single price 20%, \$9.45; Two-price 40%, −\$7.96, and Single price 40%, \$17.59. The IQUAIDS model yields slightly larger absolute magnitudes. They are: −\$4.72, \$9.65, −\$10.94 and \$17.89, respectively. These mean values indicate that although both single price increase

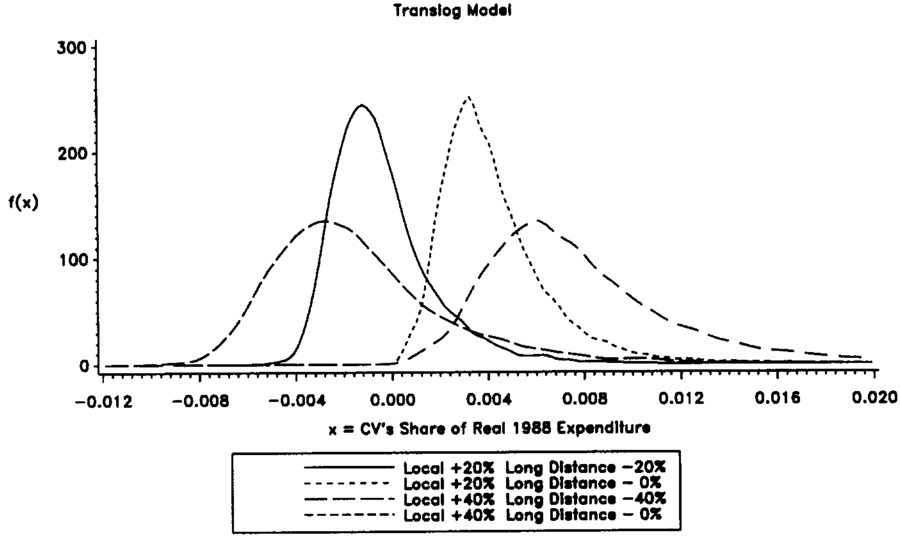


Fig. 5. Kernel density estimate of CV's share of real expenditure (translog model).

Table 6
Compensating variations for price change scenarios (January 1988 dollars) (translog and IQUAIDS models)[a]

| Scenario | Translog | | | IQUAIDS | | |
|---|---|---|---|---|---|---|
| | 5th% | Mean | 95th% | 5th% | Mean | 95th% |
| $P_L+20\%$, $P_D-0\%$ | 6.98 | 9.45 | 12.12 | 7.07 | 9.65 | 13.73 |
| $P_L+20\%$, $P_D-20\%$ | −15.95 | −3.49 | 4.04 | −16.14 | −4.72 | 2.86 |
| $P_L+40\%$, $P_D-0\%$ | 13.01 | 17.59 | 22.60 | 13.10 | 17.89 | 25.52 |
| $P_L+40\%$, $P_D-40\%$ | −32.17 | −7.96 | 6.52 | −33.43 | −10.94 | 4.26 |

[a] $P_L$ = price of local service, $P_D$ = price of long-distance service.

scenarios are on average welfare reducing, the two equal price change scenarios are on average welfare improving. The sample mean values of *CV/M* for these four scenarios for the translog model are: −0.00044, 0.0042, −0.0013, and 0.0078, respectively. These figures indicate that even for the single 40 percent price increase scenario, the mean quarterly relative burden is less than one percent. Balancing this 40 percent price increase with an equivalent 40 percent price decrease for long-distance service yields a negative mean value of *CV/M*, indicating an average predicted relative benefit to households from this set of price changes.

Because the CES is a probability sample of the population of U.S. households, associated with each observation is a weight giving the number of households in the U.S. that it represents. Consequently, we can apply these weights to compute an estimate of the population (of U.S. households) mean compensating variation associated with each of the price change scenarios and demand system estimates. For both models, the population mean results are very similar to the sample mean results discussed above in terms of signs and relative magnitudes. For example, for the translog model, the four estimated population mean values of household-level compensating variation are: Two-price 20%, −$2.95; Single price 20%, $9.43; Two-price 40%, −$6.92, and Single price 40%, $17.55. The estimated population mean values for the IQUAIDS model are: −$4.20, $9.61, −$9.89, and $17.83, respectively. Consequently, the earlier conclusion that paying or charging all households in our sample their compensating variation generates net revenue, carries over to the population of U.S. households. However, the estimated U.S. population mean compensating variation is smaller in absolute value for all four price change scenarios relative to the sample mean values, and this difference is most noticeable for the two-price-changes scenarios.

The IQUAIDS model yields similar values for the mean of *CV* and *CV/M* for the same price change scenarios despite losing a larger fraction of observations due to the local failure of the curvature restrictions on the demand system. The estimates of the density of *CV/M* for the four price change scenarios are given in Fig. 6. These density estimates, for the same price change scenario, are similar to

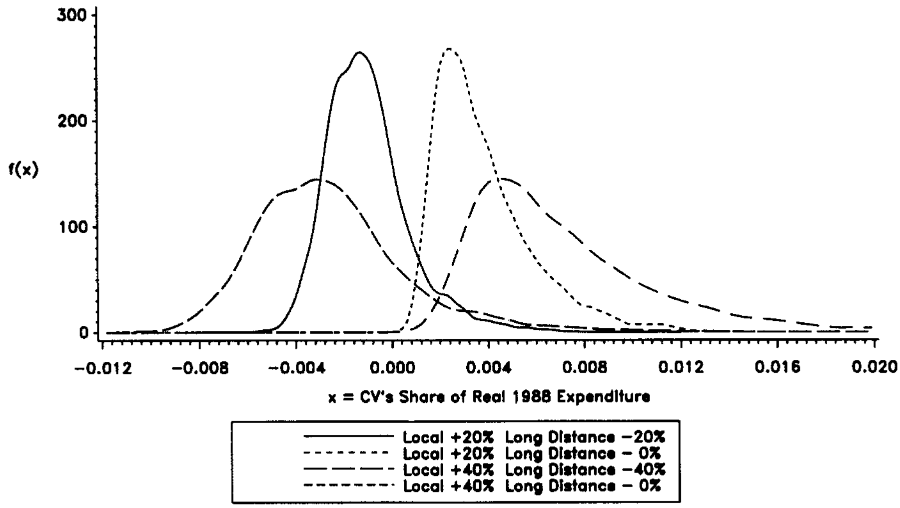


Fig. 6. Kernel density estimate of CV's share of real expenditure (IQUAIDS model).

those in Fig. 5, which at least allows us to conclude that the distribution of our household-level welfare calculations are largely invariant this change in specification of the demand system. An important point to note about the results of Table 6 and Fig. 5 and Fig. 6 is that even for the two-price-changes scenarios, there is a still a substantial fraction of households who view the combination of the local service price increase with a long-distance service price decrease as welfare reducing. This is indicated by the large amount of the support of the estimated density of *CV*/*M* for the two-price-changes scenarios that is greater than zero in both Fig. 5 and Fig. 6. As will be shown below, households experiencing welfare losses from these two-price changes scenarios also tend to be those from the bottom of the total quarterly expenditure distribution, the kinds of households that the traditionally low local service prices are designed to benefit.

An argument often raised against increasing the price of local phone service is that it causes households to disconnect from the telephone network. Unfortunately, the current version of our model is poorly suited to addressing the problem of zero consumption because both of our underlying demand systems implicitly assume that all observed demands are the result of the satisfaction of the first-order conditions for utility maximization without the imposition of the Kuhn–Tucker conditions for non-negative consumption. Nevertheless, our demand system estimates can contribute some useful input to this debate by way of the following thought experiment. Suppose a household faces a new vector of prices. If the predicted expenditure for either local or long-distance is negative for this vector of prices, we know the household would not be predicted to have positive consumption of this good arising from the household-level utility maximization process which included the Kuhn–Tucker conditions for non-negative consumption.

Consequently, we first compute the predicted local telephone expenditures for every regular observation at existing prices. Then for each of our price change scenarios, we compute the number of negative predicted local or long-distance demand shares that result. We interpret a negative predicted share for local service as a disconnection from the telephone network. All four of the price change scenarios yield no negative predicted shares for either local or long-distance service for either the translog or IQUAIDS models for all households having expected nonzero consumption at the initial prices; indicating that at least for our sample of households, disconnection from the network appears to be an unlikely response to increases in the price of local service of the magnitudes we consider.

In order to assess how the burden of these price changes are shared across households we compute best linear predictor functions for *CV/M* as a function of the household characteristics included in our demand system and the log of total expenditure. The coefficients from these regressions provide an estimate of how the best linear predictor of *CV/M* changes as a result of changes in any of the household characteristics or log of total expenditure. Table 7 and Table 8 present these best linear predictor functions for both the translog and IQUAIDS models for the local +20% and long-distance −20% price changes scenario. The best linear predictor functions are very similar across the two tables. The negative coefficient

Table 7
Best linear predictor of *CV/M* for translog model ($P_L$ +20% and $P_D$ −20% scenario)[a]

| Variable | Estimated coefficient | Standard error | *t*-statistic |
|---|---|---|---|
| Constant | 1.34e−02 | 1.37e−04 | 97.21 |
| Age of Head | 6.51e−03 | 3.38e−04 | 19.28 |
| (Age of Head)$^2$ | −2.60e−04 | 3.85e−04 | −0.68 |
| Number of Family Members | −3.47e−04 | 7.56e−06 | −45.83 |
| Members ≥65 years old | −2.38e−04 | 1.84e−05 | −12.96 |
| Head White (Dummy) | −3.30e−04 | 2.64e−05 | −12.52 |
| Head Female (Dummy) | −5.32e−04 | 1.47e−05 | −36.25 |
| Head College Graduate (Dummy) | −1.03e−03 | 2.09e−05 | −49.26 |
| Head HS Graduate (Dummy) | −2.73e−04 | 1.72e−05 | −15.84 |
| Head Single (Dummy) | −3.55e−04 | 2.15e−05 | −16.51 |
| Head Professional (Dummy) | 1.57e−04 | 1.10e−05 | 14.27 |
| Head Hours Worked per Year | 3.42e−03 | 7.51e−05 | 45.57 |
| Spouse Hours Worked per Year | 1.96e−03 | 8.18e−05 | 23.99 |
| Head Non-Worker (Dummy) | 1.05e−03 | 2.75e−05 | 38.13 |
| Spouse Non-Worker (Dummy) | −1.76e−04 | 1.86e−05 | −9.50 |
| Males Age 2 through 15 | 7.05e−04 | 1.13e−05 | 62.18 |
| Females Age 2 through 15 | 6.44e−04 | 1.23e−05 | 52.38 |
| North East (Dummy) | 1.46e−03 | 1.96e−05 | 74.56 |
| North Central (Dummy) | 1.63e−03 | 1.59e−05 | 102.77 |
| South (Dummy) | 1.02e−03 | 1.66e−05 | 61.40 |
| West (Dummy) | 1.30e−04 | 1.63e−05 | 8.00 |
| Log of NonDurable Expenditure | −2.18e−03 | 1.74e−05 | −125.28 |

[a] White (1980) heteroscedasticity-consistent standard error estimates.

Table 8
Best linear predictor of *CV/M* for IQUAIDS model ($P_L + 20\%$ and $P_D - 20\%$ scenario)[a]

| Variable | Estimated coefficient | Standard error | *t*-statistic |
|---|---|---|---|
| Constant | 1.04e−02 | 2.71e−04 | 38.55 |
| Age of Head | 5.62e−03 | 4.98e−04 | 11.29 |
| (Age of Head)$^2$ | −1.30e−04 | 5.74e−04 | −0.23 |
| Number of Family Members | −3.69e−04 | 1.12e−05 | −32.85 |
| Members ≥65 years old | −1.63e−04 | 2.86e−05 | −5.72 |
| Head White (Dummy) | −1.97e−04 | 2.71e−05 | −7.25 |
| Head Female (Dummy) | −4.97e−04 | 2.27e−05 | −21.87 |
| Head College Graduate (Dummy) | −9.72e−04 | 2.90e−05 | −33.56 |
| Head HS Graduate (Dummy) | −2.55e−04 | 2.54e−05 | −10.05 |
| Head Single (Dummy) | −3.14e−04 | 3.45e−05 | −9.11 |
| Head Professional (Dummy) | 1.23e−04 | 1.84e−05 | 6.68 |
| Head Hours Worked per Year | 2.28e−03 | 1.40e−04 | 16.30 |
| Spouse Hours Worked per Year | 1.96e−03 | 1.47e−04 | 13.28 |
| Head Non-Worker (Dummy) | 8.70e−04 | 4.38e−05 | 19.84 |
| Spouse Non-Worker (Dummy) | −1.63e−04 | 3.28e−05 | −4.96 |
| Males Age 2 through 15 | 6.40e−04 | 1.54e−05 | 41.58 |
| Females Age 2 through 15 | 6.71e−04 | 1.72e−05 | 38.97 |
| North East (Dummy) | 1.50e−03 | 3.66e−05 | 41.07 |
| North Central (Dummy) | 1.72e−03 | 2.76e−05 | 62.46 |
| South (Dummy) | 1.03e−03 | 2.97e−05 | 34.89 |
| West (Dummy) | 2.58e−04 | 2.79e−05 | 9.24 |
| Log of NonDurable Expenditure | −1.77e−03 | 3.54e−05 | −50.02 |

[a] White (1980) heteroscedasticity-consistent standard error estimates.

on the log of total expenditure indicates that households with higher total expenditures are predicted to have lower values of *CV/M*, which implies that the relative burden is more heavily borne by the lower total expenditure portion of the population. We estimate a steep regressivity to the relative burden of these price changes. For the translog model, this coefficient estimate implies that the best linear predictor of *CV/M* increases by 0.001 for a 50 percent decrease in total expenditures. (Recall that the interquartile range of total expenditures is $1688.00 to $3787.27, so this is not a large percentage change in total expenditures.) Comparing this magnitude to the range for *CV/M* given in Fig. 5 for the local +20% and long-distance price −20% price scenario, approximately to −0.004 and 0.01, quantifies the large regressivity of the relative burden of these price changes.

Several other conclusions emerge from these regressions. The best predictor of the relative burden is increasing the age of the head of household, indicating that older households bear an increasing (at a decreasing rate) relative burden of these price changes. Urban households are predicted to bear a greater relative burden. (Recall that the excluded type of household lives in rural areas.) Households with males age 2 through 15 and females 2 through 15 are both predicted to bear a greater relative burden. White headed-households are predicted to bear less relative burdens, as do college graduate-headed households, although professional-headed households bear a greater relative burden. Households with the head or spouse working a greater number of hours annually also bear a greater relative burden.

These differences in relative burden are driven by the fact that we allow household characteristics to shift the household-level indirect utility function and expenditure share demand equations, so that the price and income elasticities differ across households according to these characteristics. Within in the context of our modeling framework, variability in these relative burdens can only be explained by differences in price and income elasticities due to household characteristics. Because our model takes the household as the unit of analysis, we are unable to distinguish between the many within-household explanations for these differences in price and income responsiveness using our data and modeling framework.

As a final check of the models, we investigate the extent to which the presumed non-homotheticity of the demand for the five goods is captured by the two models. Figs. 7–11 plot the Engel curves giving the expenditure share on each good as a function of the log of total expenditure evaluated at the sample mean of the prices and demographics for both models. All goods indicate substantial non-homotheticity of demand for both the translog and IQUAIDS models. For local phone service, food, clothing and other non-durable expenditure there is close agreement between the translog and IQUAIDS Engel curves, particularly for the range of values of the log of total expenditure that contain the bulk of our sample of households. For food, the linear in the log of total expenditure Engel curve provides an adequate fit. This is clear from Fig. 9, as well as from the fact that estimate of the $\lambda_F$ in Table 2 is not large relative to its standard error. The two

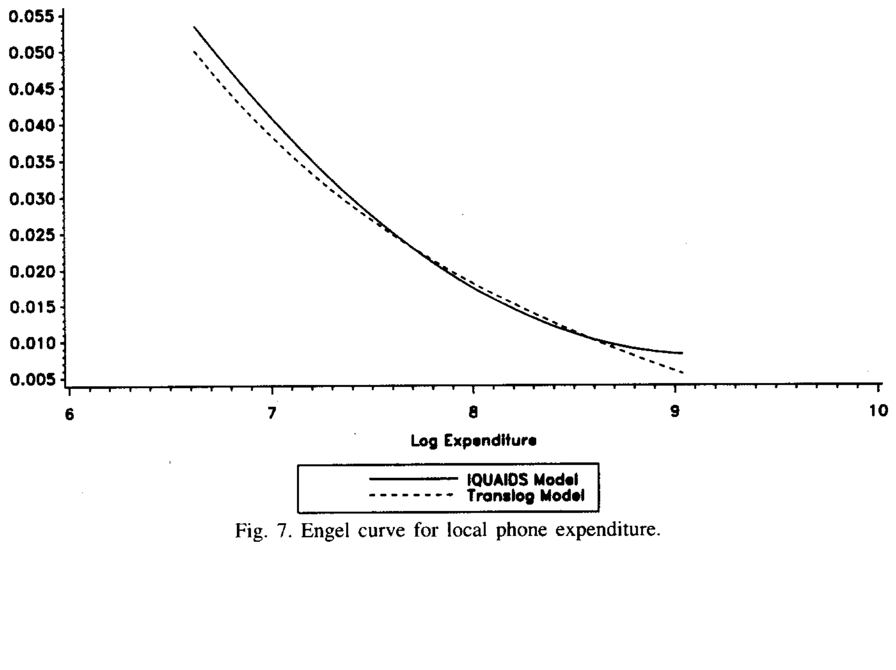


Fig. 7. Engel curve for local phone expenditure.

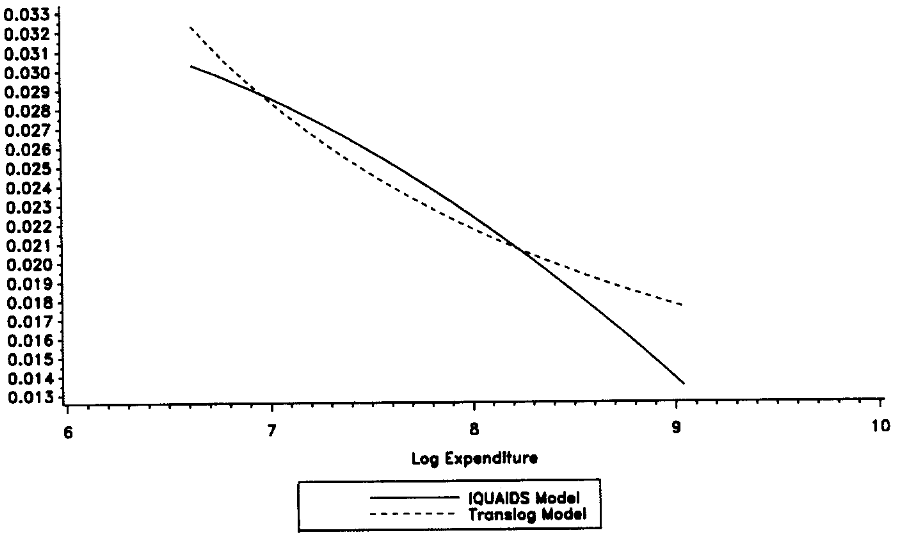


Fig. 8. Engel curve for long-distancel phone expenditure.

models differ most in the estimated Engel curve for long-distance service. The IQUAIDS estimates show a more rapid decline in the expenditure share with increases in total expenditure, but there is still a considerable amount of agreement over the interquartile range of total expenditure. The discrepancy is less than 5% of the sample mean of the long-distance expenditure share for all but very large

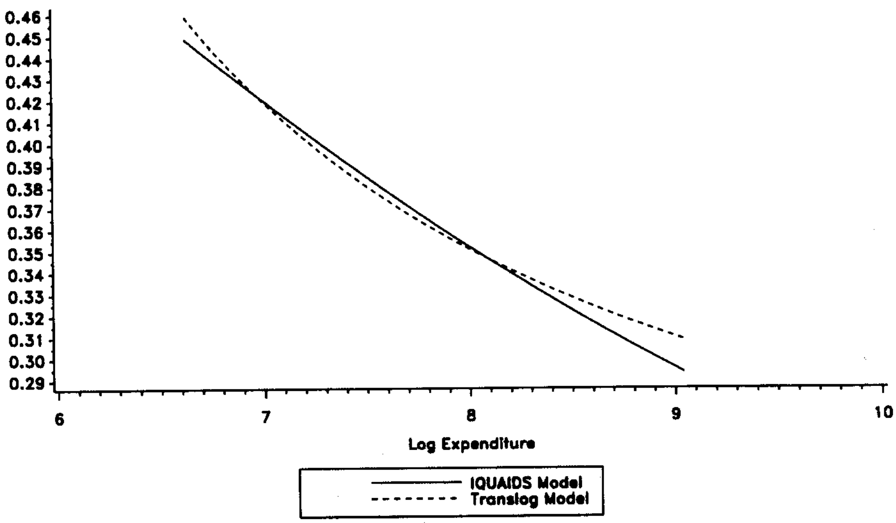


Fig. 9. Engel curve for food expenditure.

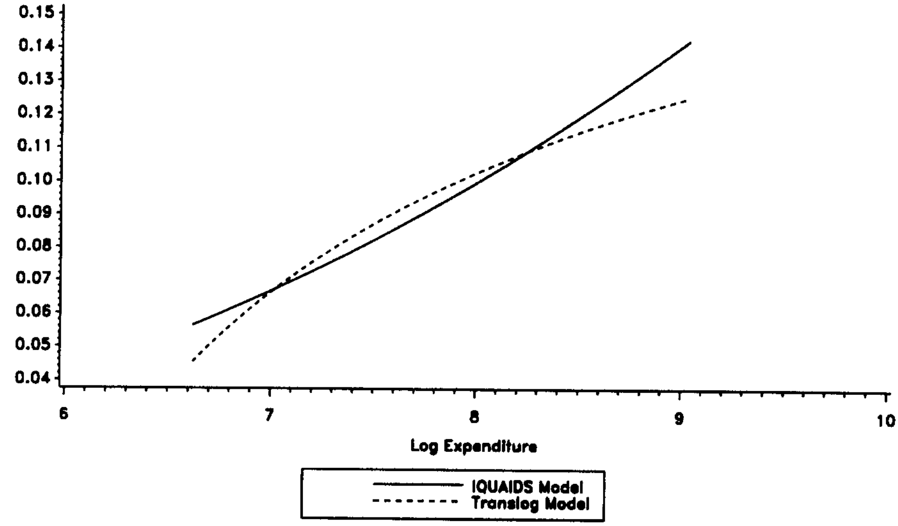


Fig. 10. Engel curve for clothing expenditure.

levels of total expenditure. Consequently, what difference there is in the Engel curves tends to be at the extremes of total expenditure. This fact could explain the large difference in the fraction of regular observations in the translog versus IQUAIDS model.

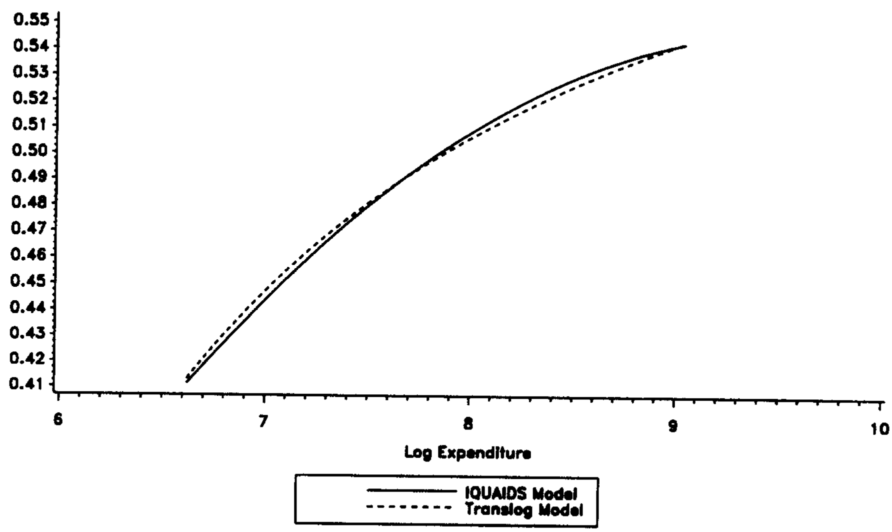


Fig. 11. Engel curve for other expenditure.

## 6. The viability of universal service in competitive telecommunications markets

Assuming our estimated demand systems are valid descriptions of the observed pattern of household-level preferences, the calculations of the previous section provide the following answer to the question posed in the title of the paper. All of the evidence seems consistent with the view that the substantial increases in the price of local service proposed as a result of increasing competition in telecommunications markets will result in small losses in consumer welfare to the vast majority of U.S. households. Balancing these local service price increases with reductions in long-distance access charges as would result if the cross-subsidies from long-distance service to local service were eliminated, appears to result in net consumer welfare gains to the majority of households in our sample.

Our results, however, do not overturn the conventional belief that local service price increases more than proportionately burden low-income (or in our case low total-expenditure households), and older-headed households. In particular, even for the balanced price changes scenarios, there are many households who experience reductions in welfare. Nevertheless, our estimation results suggest that for the price changes we consider, this burden is not sufficiently large to merit disconnection from the local telecommunications network.

Our separability test results signal the importance of modeling telephone demand jointly with the demand for all other goods in order to accurately measure price and income elasticities and to perform theoretically valid welfare calculations. Further work on the impact of making this very convenient assumption seems called for given the caveats associated with our rejection of separability between telephone expenditures and other goods.

In conclusion, so long as increases in the price of local service brought about by competition are accompanied by close to equivalent long-distance price decreases, the welfare cost to the vast majority of households should be minor, and for the majority of households, the net benefits are in fact estimated to be positive. Consequently, it seems plausible that appropriate compensation schemes can be designed to sufficiently benefit those households experiencing welfare losses from these prices changes, so that we can answer the question posed in the title in the affirmative, qualified, of course, by the many modeling assumptions and data shortcomings.

## Acknowledgments

I would like to thank Greg Crawford and Matt Shum for excellent research assistance. Partial financial support for this research was provided by the Markle Foundation and National Science Foundation Grant SBR-90-57511. A preliminary version of this paper was presented at the 'Networks and Competition' conference at the Institut D'Economié Industrielle in Toulouse, France, October 20–22, 1994.

## References

Banks, D., 1994, MCI to build local phone networks, San Jose Mercury News, January 5, p. C-1.

Banks, J., R.W. Blundell and A. Lewbel, 1994, Quadratic Engel curves, indirect tax reform and welfare measurement, Working paper 94-04 (Institute for Fiscal Studies, University College, London).

Bierens, H.J. and H.A. Pott-Butter, 1990, Specification of household Engel curves by nonparametric regression, Econometric Reviews 9, 123–184.

Blackorby, C., D. Primont and R.R. Russell, 1977, On testing separability restrictions with flexible functional forms, Journal of Econometrics 5, 195–209.

Blackorby, C., D. Primont and R.R. Russell, 1978, Duality, separability, and functional structure (Elsevier, New York).

Blundell, R.W., P. Pashardes and G. Weber, 1993, What do we learn about consumer demand patterns from micro-data?, American Economic Review 83, no. 3, 570–597.

Breusch, T. and A. Pagan, 1979, A simple test for heteroscedasticity and random coefficient variation, Econometrica 47, 1287–1294.

Brown, B.W. and M.B. Walker, 1989, The random utility hypothesis and inference in demand systems, Econometrica 57, no. 4, 815–829.

Browning, M.J. and C. Meghir, 1991, The effects of male and female labour supply on commodity demands, Econometrica 59, 925–951.

Deaton, A.S. and J. Muellbauer, 1980a, An almost ideal demand system, American Economic Review 70, 312–336.

Deaton, A.S. and J. Muellbauer, 1980b, Economics and consumer behavior (Cambridge University Press, Cambridge).

Gourieroux, C., A. Monfort and A. Trognon, 1984, Pseudo maximum likelihood methods: Theory, Econometrica 52, no. 3, 681–700.

Gorman, W.M., 1981, Some Engel curves, In: Angus Deaton, ed., The theory and measurement of consumer behaviour (Cambridge University Press, Cambridge).

Hausman, J.A., E.K. Newey and J. Powell, 1995, Nonlinear errors in variables: Estimation of some Engel curves, Journal of Econometrics 65, no. 1, 205–233.

Houthakker, H.S., 1965, A note on self-dual preferences, Econometrica 33, 797–801.

Jorgenson, D.W., L.J. Lau and T.M. Stoker, 1982, The transcendental logarithmic model of aggregate consumer behavior, In: R. Basmann and G. Rhodes, eds., Advances in econometrics, Vol. 1 (JAI Press, Greenwich, CT) 97–238.

Leser, C.E.V., 1963, Forms of Engel functions, Econometrica 31, 694–703.

Lewbel, A., 1989, Nesting the AIDS and translog demand systems, International Economic Review 30, 349–356.

Lewbel, A., 1991, The rank of demand systems: Theory and nonparametric estimation, Econometrica 59, 711–730.

Lusardi, A., 1993, Euler equations in micro data: Merging data from two samples, CentER discussion paper 9304 (Tilburg University, Tilburg).

Sievers, M., 1994, Should the InterLATA restrictions be lifted? Analysis of the significant issues, Paper presented at Rutgers University Advanced Workshop in Regulation and Public Utility Economics, July 6–8, 1994.

Silverman, B.W., 1986, Density estimation for statistics and data analysis (Chapman and Hall, London).

Taylor, L.D., 1994, Telecommunications demand in theory and practice (Kluwer Academic Publishers, Boston, MA).

White, H., 1980, A heteroscedasticity-consistent covariance matrix estimator and direct test for heteroscedasticity, Econometrica 48, 817–838.

White, H., 1982, Maximum likelihood estimation of misspecified models, Econometrica 50, no. 1, 1–25.

Working, H., 1943, Statistical laws of family expenditure, Journal of the American Statistical Association 38, 43–56.