The Consumer Gains from Direct Broadcast Satellites and the Competition with Cable TV*

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Abstract

This paper examines Direct Broadcast Satellites (DBS) as a competitor to cable. We first estimate a structural consumer level demand system for satellite, basic cable, premium cable and local antenna using micro data on more than 30,000 households in 317 markets, including extensive controls for unobserved product quality and allowing the distribution of unobserved tastes to follow a fully flexible multivariate normal distribution. The estimated elasticity of expanded basic is about -1.5, with the demand for premium cable and DBS more elastic. The results identify strong correlations in the taste for different products not captured in conventional logit models. Estimates of the supply response of cable suggest that without DBS entry cable prices would be about 15 percent higher and cable quality would fall. We find a welfare gain of between $125 and $190 per year (aggregate $2.5 billion) for satellite buyers, and about $50 (aggregate $3 billion) for cable subscribers.

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1 Introduction

Although cable television began in the 1950s with the modest goal of improving network broadcast signals to rural households, it's rise to prominence in the last 25 years has been extraordinary. Today, almost 70 percent of U.S. households subscribe to cable. Its success is largely due to Americans strong taste for watching television, by far their most popular leisure time activity.

Historically, cable systems have not faced much competition. Until the Telecommunications Act of 1996, they were primarily viewed as natural monopolies, given exclusive franchises, and directly regulated. The Act phased out almost all price regulation of cable. Instead, it argued that competition arising from the entry of telephone companies and new cable start-ups (known as "overbuilds") would keep prices down. In reality, though, few companies attempted entry and those that did have had major financial problems. Indeed, were it not for the growth of Direct Broadcast Satellites (DBS) as an alternative source of programming (and one that was barely considered at the time of the Telecommunication Act), most markets in the U.S. would be classified as unregulated monopolies.

The role of DBS as the only competitor to cable makes understanding the nature of this competition fundamental for telecommunications policy. It is central to the debate over the extent of cable market power, where the rapid increase in the real price of cable since deregulation has many consumer groups calling for re-regulation.[1] It has also been a key issue in discussions of antitrust policy toward mergers in the cable and satellite industries.

In this paper we provide direct estimates of the nature of competition between cable and DBS.[2] We do this in three ways. First, we estimate own- and cross-price

[1] See the discussions in Consumers Federation of America (2001), Kimmelman (1998), and Gregory, Brenner, Schooler, and Nicoll (2000).

[2] The paper is related to an older literature that sought to examine the demand for cable television

elasticities for cable and satellite to get the degree of demand substitutability between products. Second, we look at the supply side response of cable systems to the rise of DBS by examining how cable prices and characteristics respond to DBS. Third, we compute the implied change in consumer welfare caused by entry of DBS for adopters and non-adopters.

To estimate the demand system, we use detailed micro data on television choices - expanded basic cable, premium cable, DBS, and local antenna only - and the cable system characteristics of more than 30,000 households living in 317 cable system areas. There is substantial cross-sectional variation in prices of expanded basic and premium cable across cable franchises. For DBS, however, there is no cross-sectional price variation. To identify this price elasticity we rely on Slutsky symmetry and the fact that the market shares sum to one. Together, they imply that the impact on household demand of raising the satellite price by one dollar is the same as the impact of lowering the price of all substitute products by one dollar.

Since, at the micro level of our data, households choose only one alternative, we use a discrete choice model to characterize demands (e.g., see McFadden (1986)). In particular, we use a multinomial probit model, avoiding the restrictions on taste heterogeneity implied by conventional logit and nested logit models (and the well known problems associated with such restrictions). By estimating all of the parameters of an unconstrained multivariate normal (MVN) variance-covariance matrix we allow unobserved taste heterogeneity to vary product by product and to be correlated across products (e.g., a strong taste for DBS can imply strong taste for premium ca-

(then newly available) and the impact it had on the demand for network television (see Ellickson (1979) and Park (1971)), as well as the large literature testing for market power among cable companies (such as Wildman and Dertouzos (1990), Rubinovitz (1993), Jaffe and Kanter (1990), Prager (1990), Zupan (1989), Mayo and Otsuka (1991), or Hazlett and Spitzer (1997)). Within the literature on cable, Crawford (1997), Crawford (2000) and Chipty (2001) are the first papers to apply modern industrial organization/econometric methods to the issue.

ble). The results show that this flexibility is important. All of the parameters in the variance-covariance matrix are significant, tastes are strongly correlated across products, and the implications of these correlations for elasticities and welfare are economically meaningful and cannot be captured in a conventional logit model.

We also control for unobserved product quality differences by including separate product fixed effects for each market. This approach requires a two-step estimator to recover the price elasticity (as in, for example, Berry, Levinsohn, and Pakes (1998)). Without such controls for unobserved quality, the demand elasticities are strongly biased toward zero. Our corrected estimates indicate that the demand for DBS and for premium cable are more elastic than the demand for expanded basic and that consumers view DBS to be a closer substitute for premium cable than it is for expanded basic. We find the implied average annual surplus for DBS subscribers between $130 and $190.

After estimating the demand system, we turn to the supply side response of cable systems to the rise of DBS. We model cable prices as a function of the observed and unobserved factors that affect the quality of DBS, cable and premium television in the market and other exogenous factors such average consumer demographics. Here our estimates from the demand side play an important role, providing us with a set of controls that reflect the unobserved quality and tastes for each product in each market. The results suggest that higher DBS quality in a market is correlated with lower cable prices. We also look at changes in cable characteristics and find evidence of modest improvements in quality in response to DBS entry. Overall, the annual average per capita surplus ranges between $50 and $60 a year for consumers that remain with cable.

The paper proceeds in nine sections. In section 2, we give background on the industry. In section 3 we describe our data and identification strategy. In section 4 we outline our demand model and in section 5 we describe the two-step estimation method. In section 6 we discuss the basic results and price elasticities. In section 7 we examine the cable response to DBS growth. In section 8 we compute the consumer gains from the existence of DBS. In section 9 we conclude.

2 The Market for Television Services

In 2001, virtually every household in the U.S. had a television.[3] The three principle ways to receive television programming in the U.S. are via local antenna reception (i.e., over-the-air), via cable, or via DBS. Local antenna reception is free but offers only the local broadcast stations (channels 2-13 and the local UHF channels), and the reception quality tends to be low.

For an average price of about \$28 per month, consumers can instead choose expanded basic cable, which typically delivers about 60 channels (including the local broadcast channels). For an additional fee of about \$10 per channel, they can become premium cable subscribers, adding channels like Home Box Office (HBO). In most locations, households have no choice in who provides them cable.[4]

The other alternative households can choose is DBS. Although it delivers many of the same channels as cable, DBS systems are quite different from local cable. DBS systems are national companies that broadcast directly from satellites in geosynchronous orbit to individual home satellite dish receivers around 18 inches in diameter.[5] The two leading providers of DBS in 2001 were DirecTV and the DISH Network (Echostar). Both of them charged the same price in every cable system area across

[3]The fraction in 2001 was over 99% (see FCC (2001)).

[4]FCC (2002a) indicates that in 2002 overbuilds accounted for only about 2 percent of systems. Their market share is low even in places where they exist.

[5]Being in geosynchronous orbit means that the satellite remains at a fixed point in the sky, rotating exactly with the earth's surface. This prevents the receiver from having to track the satellite's movement in the sky. To remain in the orbit, the satellite must be over the equator at approximately 23,000 miles above the earth. Owen (1999) gives a history of DBS systems.

the country (one charged $30 per month for the basic package, the other $32). In the 30 largest television markets, users could also get their local broadcast stations for an additional $5 per month. Households subscribing to satellite must pay more than just the monthly fees, however. They must also purchase the satellite dish and a tuner, typically at a cost of $100 to $200. Our calculation of the DBS price will treat this as an ongoing annual expense of $50 per year. This is more appropriate than viewing it as a fixed cost since used equipment trades frequently in an active resale market (such as the Satellite section of eBay where there are thousands of listings at any given time).

While on average slightly more expensive than cable, DBS has a number of appealing features, including more than 200 available channels. It also has more extensive pay per view options than most cable systems, some exclusive sports and international programming, and better quality audio and video than typical cable.

The main drawback of DBS, other than the higher price, comes from the potential for signal interference. Households need an unobstructed line-of-sight toward the satellite in order to receive the signal.[6] Anything that might block the view, such as mountains, buildings, and trees can affect the quality of the reception. As a result things like terrain, elevation and latitude have a direct impact on the quality of satellite in an area.[7] Further, people living in a single unit dwelling are more likely to have a clear line-of-sight in the direction of the satellite than are people living in a multi-unit dwelling. Similarly, renters are not allowed to put a dish on the roof of their building without the landlord's permission, so they typically have more difficultly getting a clear line-of-sight.

[6]The exact direction depends on the location of the customer's house relative to the satellite location, but the satellite is above the equator so from North America it is always to the south.

[7]Regarding latitude, a person in Seattle, for example, needs a clear line of site at 31.5 degrees above the horizon. In Houston, they need a clear line at only 55 degrees. This problem of greater interference at higher latitudes is well known in the industry (see, for example, Owen (1999)).

Many consumers have embraced DBS since its introduction at the end of 1994. As table 1 shows, from a base of only 400,000 users in 1994, the number of households subscribing to satellite rose to more than 5 million by 1997 (FCC (2001)). Since then, two important factors have helped the number of DBS subscribers continue to climb. First, the relative price of DBS has fallen substantially over time, both because cable prices have been rising steadily since deregulation and also because DBS equipment prices have fallen. Second, in 1999, Congress repealed the rule preventing DBS from providing local networks. With the change, almost 75% of DBS subscribers can get this local network content for an additional $5 per month.

3 Data and Identification

3.1 Household Level Data on Television Choices

Our analysis starts with household level information on television choices. This comes from a survey sponsored by Forrester Research as part of their Technographics 2001 program.[8] Forrester is a leading market research company focusing on the information economy and its annual survey of many thousands is designed to be nationally representative.[9] The survey provides various demographic information about households in addition to their zip code and their television market.[10] The survey also asks people several questions about their television choices. Households report whether they have cable. Cable subscribers report yes or no to whether they receive any

[8]The survey was conducted by mail from December 2000 to February 2001 and the field work was done by NFO Worldwide using their ongoing consumer panel.

[9]More details on the survey itself can be found in McQuivey, et. al. (1998) or Goolsbee (2000).

[10]The television market is known as the Designated Market Area or DMA and it typically contains several cable franchise areas.

premium channels (e.g., HBO). Respondents also report whether they have DBS.[11]

From these questions we construct four mutually exclusive television choices. Cable users are classified as expanded basic customers unless they report subscribing to any premium service, in which case they are defined as premium subscribers. Anyone reporting having DBS is classified as a satellite subscriber.[12] Households reporting no cable and no DBS are classified as local-antenna only.[13]

We restrict our sample to people living in cable franchise areas with at least 30 respondents in the Forrester survey. In doing so, we keep the demand system computationally manageable (with 951 fixed effects) and we reduce the sampling error in the estimated market shares, each of which is important for the estimation procedure. This yields a sample of approximately 30,000 households in 317 cable franchise areas spread across 118 television markets. Table 2 provides summary statistics for these 30,000 households.[14] Approximately 20% of them choose antenna reception only, 45%

[11] We drop from the sample households reporting DBS service provided from someone other than DirecTV or DISH Network since they are likely to be owners of C-band dishes (the older, 9-foot type). Such services are notably inferior to the new generation of DBS and are not considered competitive with cable. Their use is in serious decline and almost entirely restricted to rural areas.

[12] Satellite users can, of course, simultaneously subscribe to cable and about 2% of the sample does so. Since the higher channel offerings on satellite fully dominate the cable offerings in almost all markets during our sample, we assume that anyone reporting that they subscribe to both satellite and cable are subscribing to the minimum cable package only to get access to the local broadcast networks. Thus we classify these households as DBS subscribers.

[13] Because virtually all U.S. households have a television, local-antenna reception is the equivalent of the outside good in our framework.

[14] Restricting the sample in this way means, of course, that it does not represent the entire nation. Instead, we are estimating the degree of competition between products in large cable franchises. These large franchise areas tend to be more concentrated, as one might expect, in the larger markets. About half our sample respondents are in the top 15 DMAs versus 37 percent of households in the U.S. population. More than 75 percent of our sample lives in the top 35 DMAs versus 57 percent nationally (see Nielsen Media Research (2002) for data on the U.S. population).

expanded basic cable, 23% premium cable, and 10% DBS.

For demographic information, we include household income, size, an indicator for male and single and an indicator for female and single, and average years of schooling among the household's adults, all of which may affect tastes for television. We also include one indicator for renter status and one for single unit dwelling status since they may influence the ability to receive satellite.

3.2 Cable Franchise Characteristics

To these household data on choices we match information on the prices and cable franchise characteristics each household would have if they chose cable. These data come from Warren Publishing's 2002 *Television and Cable Factbook*, the standard source for cable system characteristics in the industry. We match households to the cable company in their area using Warren's ICA system identification number, which is based on zip code information. This allows us to match the survey households to a local cable franchise even if they do not subscribe to cable. The cable characteristics we include in the demand system are the channel capacity of the cable system, whether pay per view is available, the price of basic plus expanded basic service, the price of premium (where the regular cable price plus the price of HBO serves as the proxy), the year the franchise system began (as a proxy for technology), dummy variables for the number of major premium movie channels available (there are six major ones), and ownership dummies for the seven largest multiple system operators (MSOs). We also use the city franchise tax/fee on the cable system, and the number of over-the-air channels in the local franchise area (as measured by the number of local channels carried on the cable system). Table 3 reports summary statistics on the franchises for 2001.[15]

[15] When we look at the supply response of cable companies to DBS entry, we will also use information on the same systems in 1994 using the 1995 Television and Cable Factbook.

As table 3 indicates, there is substantial variation in expanded basic and premium prices across cable franchise areas. For example, expanded basic prices range from \$15.50 dollars per month (\$186 per year) to almost \$45 per month (\$540 per year). This compares to an annual cost of DBS that is \$422 per year in every franchise area (i.e., an average DBS price of \$31 per month plus the annualized \$50 cost of the equipment). While our results will control for many other factors, in the raw data we do see that high relative prices appear to be correlated with low market shares. Cable systems with prices in the top one percent, for example, have average expanded basic prices of about \$44.50 per month and the DBS share in those system areas exceeds 25 percent (some two and a half times greater than the average DBS share of 10 percent in systems with average prices of \$28 per month).

3.3 Identifying the DBS Price Elasticity

Since DBS prices are the same across markets, our demand model will rely on Slutsky symmetry to identify the DBS price elasticities, using variation in price differences to estimate demands. Slutsky symmetry holds for optimizing consumers in standard settings, and Anderson, de Palma, and Thisse (1992) show that it also holds in discrete choice settings if the marginal utility of income is constant over the relevant price changes. By using household level data we avoid imposing a constant marginal utility of income across people, which is often rejected by data. The specification allows for income effects to enter flexibly, although we find only a very small income effect for these products, as the compensated and uncompensated demand elasticities are equal frequently to the second decimal place.[16]

[16] Symmetry will not hold if the products have large one-way switching costs (say, from learning), or if the market is not in equilibrium. Although DBS shares were growing before 2001, as we mentioned before, this growth can be explained by the steep fall in relative prices and the regulatory changes that were enhancing the quality of satellite.

With symmetry, identifying the own- and cross-price elasticities in a demand system of J products requires variation in only $J-1$ prices. In our case, to identify the satellite own-price elasticity, we know (because the shares sum to one in this market) that

$$\frac{\partial s_{Sat}}{\partial p_{Sat}} = -\frac{\partial s_{Base}}{\partial p_{Sat}} - \frac{\partial s_{Prem}}{\partial p_{Sat}} - \frac{\partial s_{Ant}}{\partial p_{Sat}}, \tag{1}$$

where s_{Sat}, s_{Base}, s_{Prem}, and s_{Ant} are market shares for satellite, expanded basic, premium, and antenna respectively. Slutsky symmetry says that

$$\frac{\partial s_{Base}}{\partial p_{Sat}} = \frac{\partial s_{Sat}}{\partial p_{Base}}, \tag{2}$$

and similarly for premium and antenna. Thus, observing the response of satellite share to price changes of expanded basic cable, premium cable, and antenna is enough to infer the satellite own-price elasticity. Intuitively, increasing the price of a good by one dollar has the same effect on demands as decreasing the price of all competing goods by one dollar.

We face one additional problem, though; antenna is free in all markets. In terms of our example, we do not directly "observe" $\frac{\partial s_{Sat}}{\partial p_{Ant}}$ because the price of antenna does not vary (although its quality does). If the substitutability between satellite and antenna is negligible, we might simply assume that this derivative is zero. In fact, the stark differences in these products' characteristics are suggestive of a very low cross-price elasticity, as DBS systems offer more than 200 channels delivered with digital audio/video quality at a cost of more than $400 per year, while local antenna reception costs nothing and offers on average 13 channels of mediocre quality.[17]

[17]We tried regressing the local antenna share in a franchise market on factors correlated with satellite quality, such as dish angle, weather, and elevation, to get a sense of a quality adjusted cross-price effect (holding price constant and improving quality is similar to holding quality fixed and lowering price). We found no evidence that antenna share is greater in markets where satellite quality is lower.

Instead of ruling out such substitution, we estimate a characteristics-based demand model, in which preferences over product characteristics are parameterized as a function of observed and unobserved household characteristics (e.g., see Lancaster (1971), McFadden (1973), and McFadden (1986)). Following the literature, given the tastes for product characteristics implied by a household's demographics, we assume that the marginal effect on utility for a given household from a small change in a product characteristic (like price) is common across products.[18] Our approach captures any substitution effects between antenna and satellite arising because of differences in the price or number of channels, and the results suggest that the substitution between antenna and satellite is quite small.

4 Demand Specification

Given the discrete nature of household-level demand, we choose a discrete choice demand specification. We index our specification by household (n), product (j), and franchise (market) area (m).[19] The price and attributes of the products vary over M market areas. The product price and product attributes we observe in the data are denoted p_{mj} and x_{mj}. For example, $x_{m,Base}$ contains the characteristics of expanded basic for the local cable franchise such as channel capacity, whether pay-per-view is available, and so on. Of course, some of the attributes of the products and of the consumers are not observable in the data. The utility that customer n living in market area m gets from product j is decomposed into observed and unobserved parts

$$U_{nj} = V(p_{mj}, x_{mj}, z_n) + e_{nj}, \tag{3}$$

[18]Note that assuming an equal effect on utility from raising price by one dollar across products is not the same thing as assuming an equal effect of raising price on demand across products since this is a discrete choice setting.

[19]The more correct but more cumbersome notation would be to write $m(n)$ as the market in which n resides.

where z_n denotes the observed characteristics of the customer, $V(\cdot)$ is a calculable function up to parameters of the observable data and e_{nj} is defined as the difference that makes the equation an identity.

Our specification for $V(\cdot)$ and e_{nj} is given as:

$$U_{nj} = \alpha_0 p_{mj} + \sum_{g=2}^{5} \alpha_g p_{mj} d_{gn} + \beta x_{mj} + \sum_{l=1}^{L} \beta_{jl} z_{nl} + (\xi_{mj} + \epsilon_{nj}). \tag{4}$$

The price effect common to all households is α_0. Price sensitivity is allowed to vary by 5 household income levels, with the lowest income group taken as the base, and the dummy variable d_{gn} equal to one if household n is in income group g and zero otherwise. Thus, the price coefficient for a household in the lowest income group is α_0 while that for a household in a higher income group, is $\alpha_0 + \alpha_g$.[20] We also include product-specific constants for each of the alternatives in each of the markets (in the x_{mj}). In addition, for each product j, these constants are interacted with L demographic variables - given by the term $\sum_{l=1}^{L} \beta_{jl} z_{nl}$ - which permits demographics to affect purchase probabilities for each of the four products differently.[21]

The error term e_{nj} is given in parentheses in terms of its two components ξ_{mj} and ϵ_{nj}. ξ_{mj} represents the market area-wide average value of omitted attributes and other unobserved factors (it is constant across people in the same cable franchise area). ϵ_{nj} is the unobserved idiosyncratic taste household n has for product j. Household n in market m chooses good j (conditional on z_n, x_{mj}, p_{mj}) with probability

$$s_{nj} = \int_{A_{nj}} dF(e_n)$$

[20]We tried many alternative specifications, including product specific income coefficients, income group dummy variables combined with the product specific coefficients, and also leaving the income effects out entirely. All of these approaches yielded almost identical demand results.

[21]Since local antenna reception is the outside good, we make all of the necessary normalizations to it. For example, since choice data only identifies relative rankings, we normalize $\beta_{Ant} = 0$ and $\xi_{m,Ant} = 0$ for all markets m.

where $A_{nj} = \{e_n \mid U_{nj} > U_{nk} \; \forall k \neq j\}$ and $dF(e_n)$ is the density of this composite error.

If the omitted attributes of the products, ξ_{mj}, are correlated with prices p_{mj}, then e_{nj} will not be independent of the regressors. Ignoring this correlation (i.e., assuming e_{nj} and p_{mj} are independent), will make consumers look less price sensitive than they actually are. For example, we do not observe customer service. Local cable franchises with good service will typically charge a higher price and have higher demand than their high price indicates they should (suggesting to the econometrician that consumers are not responsive to price).

We deal with this potential bias by following Berry, Levinsohn, and Pakes (1998), including alternative-specific constants (fixed-effects) for each product in each area.[22] Denote the fixed effects as δ_{mj}. When they are included in the specification, they incorporate the value of all observed and unobserved attributes for product j that do not vary within the market area, that is, from (4),

$$\delta_{mj} = \alpha_0 p_{mj} + \beta x_{mj} + \xi_{mj}. \tag{5}$$

Rewriting (4) with these fixed effects,

$$U_{nj} = \delta_{mj} + \sum_{g=2}^{5} \alpha_g p_{mj} d_{gn} + \sum_{l=1}^{L} \beta_{jl} z_{nl} + \epsilon_{nj}. \tag{6}$$

Since we now condition on the part of the error correlated with price (ξ_{mj}), the remaining household error term, ϵ_{nj}, is uncorrelated with price. For a very general class of choice models Berry (1994) shows that such fixed effects exist and are unique.

This approach places few restrictions on either the mean or the variance of the unobserved quality/taste error ξ_{mj} for each product, or on the covariance in the errors between products. The trade-off for this flexibility is the need to estimate a fixed effect for every product (other than the outside good) in every market area.

[22] See Petrin and Train (2003) for an alternative approach based on control functions.

In our case this adds 951 parameters to the estimation. In principle, these could be estimated by maximum likelihood with the other parameters, but in practice the high dimensional parameter space coupled with the non-linear environment makes locating the maximum a difficult computational problem.

The typical solution to the problem is to assume that the error term ϵ_{nj} is distributed as extreme value (so when normalized to the outside good the error term is distributed logit). The logit distributional assumption reduces computational burden because a closed form solution exists for the integration over the unobserved idiosyncratic term ϵ_{nj}. In addition, for the logit case, Berry, Levinsohn, and Pakes (1995) have developed an algorithm for locating the fixed effects conditional on the other parameters in the model, so they can be concentrated out during estimation.

The cost of the logit assumption, however, is well known. Logit models impose restrictions on the substitution patterns that are often unreasonable.[23] Although computationally more demanding, a probit-type model with an unrestricted (multivariate normal) covariance matrix will not suffer from these problems. In the case of television, we believe that allowing unobserved tastes to vary product by product and to covary across products is crucial for correctly estimating consumer substitution patterns. The variance-covariance matrix adds a vector of five parameters which

[23] See the discussions in Hausman and Wise (1978), Berry, Levinsohn, and Pakes (1995), Berry and Pakes (1999), Petrin (2002), Train (1986) or Bajari and Benkard (2001).

we call σ.[24]

5 Estimation

Our estimation consists of two steps. First, we maximize the likelihood function using the household level data and including the separate product specific dummies in every market. This identifies all of the parameters except the ones absorbed in the fixed effects (i.e., the market level parameters α_0 and β). To estimate the α_0 and β, we use a second stage regression of the estimated fixed effects on price and product characteristics (i.e., we estimate equation (5)), and instrument the p_{mj} price with a cost shifter (the franchise tax rate) to control for correlation with the error ξ_{mj}.

In the first stage of estimation, for any candidate values of $\theta = (\alpha, \beta_{Base}, \beta_{Prem}, \beta_{Sat}, \sigma)$ and vector of fixed effects δ, the probability that household n purchases good j is $s_{nj} = s_j(\theta, \delta; z_n)$, given by

$$s_{nj} = \int_{B_{nj}} dP(\epsilon_n) \tag{9}$$

where $B_{nj} = \{\epsilon_n \mid U_{nj} > U_{nk}\ \forall k \neq j\}$ is the set of ϵ_n such that product j provides maximal utility. We compute it by conditioning on z_n, θ, and δ, and numerically integrating out the multivariate normal errors.[25] The likelihood function is then

[24] Since choice data can only identify relative rankings, we again normalize to local antenna reception, i.e.,

$$\epsilon_n^* = (\epsilon_{Base} - \epsilon_{Ant}, \epsilon_{Prem} - \epsilon_{Ant}, \epsilon_{Sat} - \epsilon_{Ant}) \sim \text{MVN}(0, \Omega^*) \tag{7}$$

which also is distributed multivariate normal with a variance covariance matrix given by

$$\Omega^* = \begin{bmatrix} 1 & \sigma_{B,P} & \sigma_{B,S} \\ \dots & \sigma_P^2 & \sigma_{P,S} \\ \dots & \dots & \sigma_S^2 \end{bmatrix}. \tag{8}$$

$\sigma_{B,P}$ is (for example) the covariance between the Base minus Antenna only term and the Premium minus Antenna only term. σ_B^2 is normalized to one.

[25] We use a frequency simulator with 4000 draws per household.

given by

$$L = \Pi_{n=1}^{N} \Pi_{j=1}^{J} s_{nj}^{j(n)} \tag{10}$$

where $j(n)$ is the indicator function

$$\jmath(n) = \begin{cases} 1 & \text{if household } n \text{ chose } j \\ 0 & \text{otherwise.} \end{cases}$$

We do not maximize the likelihood over the entire space of (θ, δ) directly. Instead, in the spirit of Berry, Levinsohn, and Pakes (1998), we concentrate out the likelihood function and only search over the space of θ. To do this we condition on θ and, market by market, solve for the vector $\delta_m(\theta)$ that matches observed market shares to those predicted by the model.[26] The likelihood function value for θ is then computed at $(\theta, \delta(\theta))$. Thus, parameter values θ are chosen to maximize a likelihood function that, conditional on θ, includes additional fixed effect parameters that exactly match observed to predicted market shares for each product in each market.[27]

6 Demand Results

In our first stage, we include the variables described in section 3: household income, size, and education, and indicators for single male and single female. Each demographic enters each product's utility equation. In the satellite equation we also add

[26] At each θ and for each market m we use a non-linear least squares search routine to solve for

$$\delta_m(\theta) = argmin_{\delta_m} \sum_{j=1}^{3} (s_{mj}(\theta, \delta_m) - \hat{s}_{mj})^2,$$

where $s_{mj}(\theta, \delta_m)$ is j's predicted share in market m at δ_m (and θ), and $\hat{s}_{mj}$ is the observed market share from our data. Because the fixed effects exist and are unique, the $\delta_m(\theta)$ that sets this objective function to zero exists and, once located, is known to be the unique minimum.

[27] The usual formulas for the standard errors of θ will not apply in this case since there is sampling error in the observed market shares $\hat{s}_{mj}$ across locations in the Forrester data. The appendix describes our methodology for deriving the appropriate standard errors.

controls for renter and single unit dwelling status. Finally, we include fixed effects for each product-market pair. These control for observed and unobserved factors that vary at the market level, including the price and characteristics of the local cable franchise, the quality differences across markets in satellite and the local antenna only option, and unobserved product-market specific tastes.[28]

As with any non-linear framework, the coefficients from table 4 do not give the marginal effects on purchase probabilities. Instead, they indicate the change in the value of the latent variable (utility) relative to the local antenna only option. Almost every variable enters the specification significantly. The four price coefficients imply that higher income households are less price sensitive than lower income households.[29] Living in a single unit dwelling and not renting is associated with higher utility from satellite. Single males get greater utility from premium and satellite while single females get less. Increases in the average level of household education are associated with increases in the utility of antenna only (relative to multichannel video).

We translate the parameters from table 4 into marginal effects in table 5 (holding the fixed effects constant at their estimated levels). For any marginal effect, we compute the average percentage change in the four purchase probabilities (holding the households' other characteristics fixed). For example, in column 1 we report the average percentage change in purchase probability for multi-unit (MU) dwellers that move to single-unit dwellings. DBS adoption responds strongly, increasing (for multi-unit dwellers) by almost 100% (from 4.8% to 9.5%), largely at the expense of

[28] While satellite programming characteristics are the same across markets (up to the availability of local channels), geographic differences like dish angle, the weather, elevation and its variance, lead to differences in the quality of reception. Similarly, the number of over-the-air channels available and the quality of this reception varies across markets.

[29] Since the price sensitivity for the lowest income group is α_0 and for any other group g is $\alpha_0 + \alpha_g$, the positive and increasing parameter estimates for α_g indicate that as income increases, consumers become less price sensitive.

expanded basic and premium cable (which have smaller percentage changes because their base probability for multi-unit dwellers is much larger). Similarly, from column 2, switching from renter to non-renter also substantially increases the satellite adoption probability. Thus, even after conditioning on demographics and product-market effects, factors influencing the household's ease of line-of-sight play an important role in determining the likelihood of dish adoption.

Column 3 reports the effect of increasing income 10%. This leads to a 0.6% increase in satellite adoption, implying a very modest income elasticity of less than 0.1. When coupled with the fact that total television expenditures are, for almost all households, less than one-half of one percent of their budget, it follows that the compensated and uncompensated demand curves are almost identical. This finding is a common one in the demand literature for goods with small budget shares (e.g., see Willig (1976)).

The parameters of the multivariate normal are reported at the bottom of table 4. They are all highly significant, rejecting the restrictions typically imposed by logit and nested logit approaches. More than statistical significance, however, these parameters substantially affect estimated consumer substitution patterns, as shown in table 6. In our setting, when DBS is removed from the choice set, 46.7% of DBS subscribers switch to premium cable, 45.6% switch to expanded basic, and only 7.7% switch to local-antenna reception. These substitution patterns are very different from what comes out of reestimating our model with a logit error.[30] In that case, the results indicate substitution along the lines of aggregate market shares of the products: 25.9% to premium cable, 50.9% to expanded basic and 23% to local antenna. The logit model does not capture the high correlation in tastes between DBS and premium cable, understating their substitutability by almost one-half. Similarly, it overstates

[30]We do not report the parameter estimates on all the demographics from the logit model but they are similar to those in table 4.

the substitutability of DBS and local antenna by a factor of almost three.

Estimates of the price elasticities require estimates of α_0 and the β's, which necessitates a second step of estimation. In it we regress the estimated fixed effects on product characteristics and prices (that is, we estimate the system of equations implied by (5)). Since satellite prices and characteristics do not vary across markets, we use only two of the three equations for the $\delta' s$: expanded basic and premium cable. We estimate this system in two ways, first without instruments, using seemingly unrelated regressions (denoted SUR), and then with instruments, using three stage least squares (denoted 3SLS). The tax on local cable franchise revenue is the price instrument in the 3SLS setup. This tax is typically negotiated and then fixed for long periods of time (i.e., decades), and is reported in Warren Publishing (2002). The other characteristics entering the system are those that vary only at the market level (and are thus absorbed in the fixed effects). These include the cable company characteristics, like channel capacity, pay-per-view availability, the franchise age, and dummy variables for the MSOs and for the number of the six most popular premium channels available. To account for differences in the amount of sampling error arising from the estimated market shares we weight the market-level regressions by the estimated variances of the δ's from the first stage. The standard errors also account for the fact that the unobserved $\xi_{Base,m}$ and $\xi_{Prem,m}$ in each market may be correlated. Except for the coefficients on the 7 MSO indicators and the premium channel dummy variables, we test and cannot reject that the marginal effects of characteristics on utility are the same for both the expanded basic and premium equations (this is the standard characteristics-based restriction).[31] We report the results with the restrictions imposed (for efficiency purposes) in table 7.

Columns 1 and 2 of table 7 report the results from the SUR regression of the

[31] We could not reject the restrictions for the five characteristics where we imposed them either jointly or individually.

estimated fixed effects on the observed characteristics and prices, which is equivalent to assuming no correlation between price and the unobservable component of utility (which reflects, e.g., unobserved quality). We have argued that this price coefficient will be biased toward zero. Columns 3 and 4 contain the estimates from 3SLS. The SUR coefficient on price is only one-sixth the magnitude of that from the 3SLS system. This large difference suggests that the unobserved quality of cable in an area is positively correlated with the price of cable. The coefficients we obtain on the other variables are consistent with our priors, as the characteristics we associate with higher quality increase δ (and thus the purchase probability). Higher quality systems are those systems that have pay-per-view, more channel capacity, and more pay channels. Most of the coefficients in this regression are significant. Over the air channels enter significantly and negative, meaning improved local antenna options reduce the market share of both cable packages.

With the estimates of all of the model parameters, we can compute the own and cross-price elasticities for each television choice. In column 1 of table 8 we present results using the SUR estimates from the fixed effects regression. Assuming no endogeneity problems leads to estimates of the own price elasticities of expanded basic, premium, and satellite that are slightly positive, implying that television demand is almost completely insensitive to price.[32]

Column 2 uses the 3SLS estimates and yields results that are economically sensible. Consistent with economic theory, all elasticities now exceed 1 (in absolute value). Expanded basic cable has the lowest own-price elasticity, at about -1.53, suggesting reasonably inelastic demand for the basic cable package.[33] The higher quality offerings of premium cable and DBS are substantially more price elastic. The own-price

[32]For a discussion of the importance of instruments in the context of estimating welfare gains, see the debate in Hausman (1997b), Hausman (1997a), Bresnahan (1997), and Petrin (2002).

[33]This cable price elasticity is similar to the results found in the literature such as Crawford (1997), Hazlett and Spitzer (1997), U.S. General Accounting Office (2000), or Crawford (2000).

elasticity of premium is about -3.18 and the DBS elasticity is -2.45. A DBS elasticity higher than expanded basic's does suggest that the marginal costs of DBS are higher than for cable (since the prices are similar).[34] Finally, as mentioned above, the compensated and uncompensated demand curves (columns 2 and 3) are almost identical (often equal to the second decimal place).

In terms of substitutability, the demand elasticities reflect the correlation of observed and unobserved tastes. DBS's closest substitute appears to be premium cable. Raising the price of DBS leads to the largest percentage increase in premium followed then by expanded basic. Similarly, for premium price increases, DBS has the largest cross-price, followed again by expanded basic. Consistent with the differences in observed characteristics, local antenna is not a close substitute to DBS.

7 Cable's Response to DBS Entry

The incumbent firm's response to entry is often an important part of the product-market competition. In most empirical work this effect is ignored, even though it can substantially impact consumer welfare. In the cases where it is recognized (such as Petrin (2002) and Hausman and Leonard (2002)), incumbents are assumed to respond only by changing prices. Because of the rapid rise of DBS and the fact that it is a higher quality alternative on many dimensions, we examine the response of both cable prices and cable characteristics to entry. To do so we use our 2001 cross-section of cable systems. As a complement to these results, we also match the 2001 cable systems to their prices and characteristics in 1994 to ask the same questions.[35]

The only previous analyses of the impact of satellite competition on cable prices come from two government reports: U.S. General Accounting Office (2000) and FCC

[34] According to satellite industry sources, this may arise from higher programming costs for DBS than for cable MSOs (Willig (2001)).

[35] This information comes from the 1995 *Television and Cable Factbook*.

(2002b). These studies use cross-sectional data and regress cable prices on satellite share, controlling for product characteristics and demographic averages. These studies have found that increasing satellite penetration is associated with *higher* cable prices. Our data set (which comes from a different source) yields a similar finding in such regressions. However, since satellite share is neither exogenous nor likely to enter the equilibrium pricing function linearly, it is problematic to conclude that the positive correlation of prices is, in fact, causal. For example, the positive correlation may simply reflect the fact that within markets unobserved tastes for multichannel video are correlated.[36]

These difficulties are symptomatic of the fact that supply responses are hard to estimate. This is especially true when a more structural approach is desired. In addition to the demand side estimation, a structural approach requires cost measures for each product. It also requires the type of competition to be completely specified (e.g. Cournot or Bertrand-Nash, static or dynamic).[37] We - like most researchers - neither observe marginal costs nor know exactly what the mode of conduct is for the supply side. Given these difficulties, we introduce a reduced-form approach that exploits our demand side estimates to construct an estimator for the value to consumers of the entry-induced changes in cable price and characteristics.

Our approach is to estimate an equilibrium pricing function directly and then use it to ask whether cable prices vary systematically with the quality of satellite, holding the other market-level factors entering the equation constant. The demand side model coupled with micro/econometric theory implies that the price for cable is potentially a non-linear function of all exogenous factors that affect demand and

[36]Some of the previous work has tried instrumenting for the satellite share. Unfortunately, this does not correct the inconsistency problem if the pricing equation is a non-linear function of the DBS share (or exogenous instruments for satellite entry are not available).

[37]Often, the assumption regarding the type of competition is combined with estimates of demand side elasticities to infer markups and (thus) marginal costs.

supply. These include demographics, the observed characteristics of all the products in the market (including the identity of the sellers), the franchise fee, and proxies for the unobserved factors of each product (which, for cable, table 7 shows are strongly correlated with prices).[38]

To provide a measure of these unobserved factors, we turn to the estimates from our demand model. They provide us with consistent estimates of the magnitude of unobserved characteristics and tastes (together) in the units of utility for every product-market pair. These estimates are given by the residuals in (5), and are computed as a function of the estimated $\delta's$, $\beta's$, and observed characteristics:

$$\xi_{mj} = \delta_{mj} - \beta x_{mj} - \alpha_0 p_{mj}. \tag{11}$$

Estimates for δ_{mj}, βx_{mj}, and $\alpha_0 p_{mj}$ for expanded basic and premium cable are available directly from the second stage regressions in table 7. For satellite, programming characteristics and prices do not vary across U.S. markets, so an equivalent decomposition yields a residual that only differs from $\delta_{m,Sat}$ by a constant; we just use $\delta_{m,Sat}$ as the satellite regressor (and provide a robustness check below).

Our basic cable pricing equation is given in column 1 of table 9. It treats cable prices as a linear function of many market-specific factors: the mean demographics, the observed characteristics of both expanded basic and premium cable, including indicator variables for those systems belonging to the largest seven multiple system operators and indicators for the number of leading premium channels equal to two, three, four, or five (six is the excluded group), the city franchise fee, residuals for expanded basic and premium cable from (11) and, similarly, the satellite fixed effect $\delta_{m,Sat}$.

Many variables enter significantly and with the anticipated sign, including pay-per-view, over-the-air channels, and some of the premium and MSO ownership indicators.

[38] We present only the regressions using the prices of expanded basic for simplicity but the results for premium were very similar in every case.

In particular, entering most significantly are $\delta_{m,Sat}$ and both of the cable residuals.[39] Overall, the r-squared of 0.62 suggests these factors have reasonable explanatory power for the cable price. We estimate several specifications that included higher-order terms to check for the importance of non-linearities, testing for the significance of the sets of additional coefficients. This included second order approximations using all squared and interaction terms for all programming characteristics (136 total extra regressors), using just demographics (an additional 28 regressors), and using just $\delta_{m,Sat}$ and the cable residuals jointly (an additional 6 regressors). None of these specifications rejected the null hypothesis that all additional higher order coefficients equal zero. Finally, the results do suggest that, holding other factors constant, where satellite quality is higher, prices for cable are lower.[40]

Our preferred approach to evaluating the magnitude of the price effect would be to assume that eliminating satellite is equivalent to reducing $\delta_{m,Sat}$ until demand is zero (by, for example, raising satellite's price to its reservation level). In our data, the lowest satellite share we observe is 2%, so instead we use the pricing function to ask how cable prices would change if we reduced satellite quality in every market

[39] The standard errors for this and future regressions are all corrected for the fact that many regressors are estimated (our approach follows Murphy and Topel (1985)).

[40] As an alternative specification, we also tried using OLS to decompose $\delta_{m,Sat}$ into a component related to geographic factors that may affect satellite reception such as the the dish angle in the franchise area (computed using the DirecTV Dishpointer, DirecTV (2000) for the primary zip code in the DMA), the average elevation in the market and the variance in elevation in the market (computed for us by David Rowley of the Geophysical Sciences department at the University of Chicago using the one degree U.S. Geologic Survey Digital Elevation Model for a 30 pixel by 30 pixel area centered at the geographic center of the DMA), and the climate mildness and brightness score (as measured by the Places Rated Almanac Savageau and Lotus (1997)). The R^2 of 0.05 yields estimates of satellite residuals that differ little from the estimates of $\delta_{m,Sat}$. Including both the predicted value and the residual separately in the column one regression (instead of $\delta_{m,Sat}$) results in an insignificant parameter estimate on the predicted value and a very significant coefficient on the residual that is virtually identical to the coefficient on $\delta_{m,Sat}$ from column 1.

to the estimated $\delta_{m,Sat}$ from this market, holding all other observed and unobserved characteristics and tastes for the products constant.[41] The price regression suggests that doing so would raise the average cable price by $4.17 per month, an increase of about 15%, holding other factors constant.

In general, the specification in column 1 can be very demanding on the data. Even with just four goods, a linear approximation has twenty-five regressors, and a complete second order approximation has 357 regressors (which exceeds the number of market level observations in our sample). A natural, more parsimonious alternative, and one that is perhaps easier to interpret, is to include just the price-adjusted quality index given by

$$\delta^*_{mj} = \beta x_{mj} + \xi_{mj}, \tag{12}$$

which adjusts each δ_{mj} by the price effect $\alpha_0 p_{mj}$ (see (5)). For each good j, including the δ^*_{mj} index in place of each of the x_{mj} and the ξ_{mj} as regressors in the pricing function imposes k_j restrictions, where k_j is the number of observed characteristics for good j.[42] This set of restrictions is equivalent to assuming that the δ's from the demand side are sufficient proxies for both the demand and supply factors entering the equilibrium pricing function. Column 2 reports this regression, imposing these restrictions for both expanded basic and premium cable. The r-squared falls from 0.62 to 0.58, an insignificant amount for the number of restrictions, so the more parsimonious approach is not rejected. The coefficient on $\delta^*_{m,Sat}$ is almost identical to column 1, leading to a similar predicted price change.[43]

We also use our matched sample of 1994-2001 cable systems to ask whether systems appear to have changed their prices and characteristics in response to DBS entry

[41] This is Greenfield, Wisconsin near Milwaukee.

[42] The $k_j + 1$ coefficients on each of the $\beta_k x_{mjk}$ and the ξ_{mj} are assumed to be identical.

[43] We also consider a restriction which specified prices as a function of relative quality indices, which leads to $\delta^*_{m,Sat} - \delta^*_{m,Base}$ and $\delta^*_{m,Prem} - \delta^*_{m,Base}$ replacing $\delta^*_{m,Sat}, \delta^*_{m,Base}$, and $\delta^*_{m,Prem}$ as regressors. The r-squared falls from 0.62 to 0.32, which rejects this restriction.

(since DBS essentially did not exist in 1994). The added complication, however, is that the differences in prices over time are potentially a non-linear function of all exogenous factors from both 1994 and 2001. Since the Forrester data was not collected in 1994, we cannot estimate the demand system for that year. Instead, we assume that the tastes β for observed characteristics are constant over time and construct an equivalent 1994 βx_{mj} index for expanded basic, up to controlling for MSO effects (the MSOs in 1994 are different from 2001). Because of the data limitations, we cannot correct for either the unobserved factors for 1994 (derived from the δ's) or the 1994 demographic averages.

Column 3 reports this estimated specification. The r-squared for the price differences is still reasonably high at 0.38.[44] The coefficients on the δ^* terms are similar to the cross-sectional regression and all enter significantly. In particular, the coefficient on $\delta_{m,Sat}$ is -2.22 (vs. -2.39 in the cross-section), suggesting a relative price increase of about \$3.86 per month for the average system without DBS entry.[45] Thus, the matched sample and cross-sectional predictions are very similar.

These estimated price changes hold cable characteristics constant. Although harder to measure than changing prices, cable systems may have responded to DBS competition by innovating on characteristics like channel capacity and pay per view availability, especially given the seven year time frame. Indeed, some industry sources have argued that changing quality has been one of the most important ways that cable responded to DBS growth (see Watts (2003)). To investigate this question we ask whether, for the characteristics we observe, higher measured observed characteristics

[44]We again found that all of the specifications including higher-order terms and interaction terms to test for non-linearities as described above lead to no rejections of the null hypothesis of linearity.

[45]Note that this does not imply average prices were 12% lower in 2001 than they were in 1994, only that prices were 12% lower than they would have been without DBS entry. Indeed, average cable prices rose by more than 25% over this time period, much of it (according to the cable franchises) due to the rising costs of programming (see the discussion in FCC (2002b)).

are positively correlated with higher DBS quality.

We look at both the cross-sectional data from 2001 (column 4) and the matched sample (column 5). To construct the dependent variable, we again use the weights from the utility function to construct a characteristics' index defined as $\beta x_{m,Base}$. Since we are trying to explain the characteristics' index (and the change in it), we only use as regressors the market demographics, the franchise fee, and $\delta^*_{m,Sat}$. The r-squared is 0.19 and $\delta^*_{m,Sat}$ has a positive and significant coefficient. The average predicted change for this index if every satellite system were moved to its lowest observed quality level is between 0.075 and 0.1 in the two specifications. If we use the estimate of the marginal utility of income from table 7 to translate this into a "price-equivalent", the regressions suggest that, moving to the characteristics prior to satellite results in a welfare loss equal to between \$1.04 to \$1.38 per month.[46] We emphasize that this calculation provides only a lower bound on the improvements in quality that systems undertook, as this index does not include all characteristics (e.g whether digital cable is available), just the ones we observe.

8 The Change in Consumer Welfare

The results from sections 6 and 7 suggest that the introduction of DBS has impacted the welfare of both DBS and cable subscribers.[47] In this section we use the demand and supply estimates to compute the overall changes to each group, and the aggregate effect (weighting each household equally). Our base level of welfare is that achieved

[46]The price equivalent is calculated by solving for the price increase that exactly offsets the increase in utility from the improved characteristics, or the Δp such that (say) $0.075 - 0.312\Delta p = 0$, where 0.312 is the estimated marginal utility of income from table 7. Since the original demand specification is estimated using weekly prices, this Δp must then be multiplied by (52/12) to get the monthly price equivalent.

[47]Local antenna subscribers in 2001 experience no change in welfare; without DBS, local antenna remains free and the cable alternatives get worse.

by households in 2001 with DBS available as an alternative. We ask how much income would have to change for a household to achieve that same utility level in 2001 with DBS not available (i.e., the compensating variation).[48]

We write compensating variation for household n as Δy_n, Formally defined, it is the difference in the value of the expenditure function between the two economic environments under consideration. Let $\tilde{V}_n = \tilde{V}(p_0, y_n, n)$ be the welfare level in the base environment, that is, the highest utility household n with income y_n can achieve when facing prices/characteristics p_0 available in 2001 with DBS in the market. The expenditure necessary for household n to achieve $\tilde{V}_n$ is given by $e_n(p_0, \tilde{V}_n)$. Defining p_1 as the prices/characteristics faced in the counterfactual of no DBS entry, the expenditure necessary for household n to achieve $\tilde{V}_n$ is $e_n(p_1, \tilde{V}_n)$. Our measure of household level compensating variation is then

$$\Delta y_n = e_n(p_1, \tilde{V}_n) - e_n(p_0, \tilde{V}_n) = e_n(p_1, \tilde{V}_n) - y_n. \tag{13}$$

Welfare increases for satellite consumers for two reasons. First, at the observed 2001 prices and characteristics, DBS consumers' willingness to pay for satellite exceeds the price. Second, the results from sections 6 and 7 suggest that, without DBS, almost all of these satellite consumers would subscribe to cable at both a higher price and a lower quality than that observed in 2001. Similarly for cable consumers, surplus increases because they pay both lower prices and get higher quality cable than they would have without DBS entry.

A standard empirical concern when calculating welfare gains from new goods is that much of the estimated gain may come from extrapolating an estimated functional form to areas outside the region of the observed price and quantity variation (i.e., that the true maximum willingness to pay is not observed in the actual price variation). To determine how important this is in our case, we provide a lower bound estimate

[48]See Diamond and McFadden (1974) or Hicks (1946). Hause (1975) and Mishan (1977) also provide helpful discussions.

on welfare by raising the price of satellite to the largest observed difference in the actual data (as opposed to raising the satellite price to infinity as with compensating variation). Similarly, as described in the previous section, for cable subscribers our lower bound estimate reduces the quality of satellite only to the lowest observed level of $\delta^*_{m,Sat}$ observed in our data.

We begin by reiterating the importance of the controls for unobserved quality. Generally, when the price sensitivity is biased toward zero, welfare is biased upward. Without the controls for unobserved quality our estimates (in column 1 of table 7) are so biased that the demand curve is actually slightly upward sloping, a case for which welfare estimation makes little economic sense.

If cable prices and characteristics are held constant at their observed 2001 levels, the corrected (i.e. 3SLS) estimate for the welfare gain to DBS users averages about \$127 per year, or \$10.58 per month above what they pay for the service. When satellite prices are raised to only the maximum observed price difference in the sample, the per capita average is about \$100. This confirms that most of our estimated welfare gain for DBS subscribers (almost 80%) comes from a part of the demand curve that is not extrapolated. If we calculate welfare assuming annual cable prices would be \$4 higher per month without DBS (from the supply regressions), the average welfare gain increases to \$175 per year rather than \$127. Finally, adding the characteristics' effect yields a gain of almost \$190 per year, the average additional amount a DBS consumer would willingly pay every year (over what they currently pay), once differences in cable franchises induced by DBS entry are taken into account.

Although our sample is not representative of the entire country, if we extrapolate our results to all 16 million DBS adopters at the time of our sample (rather than just to the urban customers), it would imply an aggregate welfare gain of between \$2.5 billion and \$3 billion for DBS subscribers.

For cable subscribers, our results suggest that cable prices are at least \$4 per month lower than they would have been. In the aggregate, given the 70 million cable subscribers, the price effect yields a total welfare gain of close to $3.3 billion for the consumers that stay with cable. The quality improvements to cable characteristics are worth approximately another $1 per month of surplus, which adds another $800-900 million to the welfare change. In the end, while these supply-side calculations are less structural than the demand side estimates, they point to a substantial aggregate welfare gain from DBS entry of as much as $4 billion per year for cable consumers.

9 Conclusion

Because DBS is the only direct competitor to cable in most markets, the nature of its competition with cable television is fundamentally important for developing telecommunications policy. This paper examines the introduction of Direct Broadcast Satellites (DBS), the nature of that competition, and the welfare gains satellites generate for consumers. We estimate a household-level discrete choice demand system for satellite, basic cable, premium cable and local antenna using micro data on the television choices of more than 30,000 households, as well as price and characteristics data on cable companies throughout the nation. Our structural demand framework has extensive controls for unobserved product quality and permits the distribution of unobserved tastes to follow a fully flexible multivariate normal distribution.

The results indicate that the own-price elasticity of expanded basic is at about -1.5 while the demands for premium cable and DBS are substantially more elastic (-3.2 and -2.4). The cross-price elasticities suggest that DBS and premium are the closest substitutes. The flexibility of the multivariate normal distribution is crucial for understanding consumers' true substitution patterns as the correlation of unobserved tastes for DBS and premium cable are particularly high and are not captured in a conventional logit model.

Our approach to inferring the supply side response is more reduced form in nature, using an estimated equilibrium pricing equation to ask whether cable prices vary systematically with the level of competition provided by satellite. The supply side results exploit the estimated controls from the structural demand side model and suggest that more competition from DBS is correlated with lower cable prices and somewhat higher quality cable. Overall there is a significant welfare gain to the 16 million satellite buyers between $2-3 billion dollars, depending upon whether changes in cable prices and characteristics are added back into the calculation. The aggregate gains to the 70 million cable users amount to between roughly $3-4 billion. In the end, our results suggest large gains from DBS entry, some of which are not captured if the price and characteristics' response is ignored. The overall gains from this product introduction may be as large as $7 billion, illustrating once again the importance of understanding the impact of new goods on consumer welfare.

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Appendix

If there were no sampling error in the product-market shares, the usual formulas for the standard errors of the likelihood function estimates are valid for $\hat{\theta}$ (with the fixed effect constraints imposed). When sampling error enters the share estimates, it must be accounted for when computing the standard errors for $\hat{\theta}$ (and $\hat{\delta}$).[49] This appendix describes one method that accounts for this source of error.

Before describing the method we follow, we begin at a more natural starting point for thinking about this correction, the first order conditions of the likelihood function written with δ as a function of θ, or

$$\frac{\partial L(\hat{\theta}_k, \hat{\delta}(\hat{\theta}))}{\partial \theta_k} = 0 \qquad k = 1, ..., K.$$

These moment conditions include a term that accounts for the derivative of the likelihood function with respect to θ (holding δ constant), the derivative of the likelihood function with respect to each δ, and the derivative of each δ with respect to each θ. Obtaining estimates of the standard errors of $\hat{\theta}$ and $\hat{\delta}$ requires one to differentiate these first order conditions, a non-trivial task requiring the evaluation of some computationally difficult terms.

Instead of following this approach, we make use of the fact that we can characterize the asymptotic behavior of the parameter estimates using a generalized method of moments (GMM) representation (similar, e.g., to Petrin (2002) or Berry, Levinsohn, and Pakes (1998)). This characterization provides a set of moments that have derivatives which are much easier to evaluate numerically.

There are two different types of moments that are zeroed at the estimates $(\hat{\theta}, \hat{\delta})$. First, the difference between the observed and the fitted shares for each product-

[49] Our estimates for market shares are constructed using between thirty and several hundred observations per market.

market pair is zeroed, or

$$s_{mj}(\hat{\theta}, \hat{\delta}) - \hat{s}_{mj} = 0 \quad \forall j \; \forall m, \tag{14}$$

where $s_{mj}(\hat{\theta}, \hat{\delta})$ is the predicted market share for product j in market m and $\hat{s}_{mj}$ is the observed market share. $J * (M - 1)$ of these moments are independent (one moment in each market is redundant because the shares in each market sum to one). In each market m there are N_m observations on each moment (where N_m is the number of households observed in market m).

The second set of moments derive from the likelihood function itself. At the maximum, the first order conditions for the likelihood function are zeroed for each of the K parameters in θ, holding $\hat{\delta}$ constant. Thus, an additional K moments are given by

$$\frac{\partial L(\hat{\theta}_k, \hat{\delta})}{\partial \theta_k} = 0 \qquad k = 1, ..., K.$$

There are N observations on each of these moments, where $N = \sum_{m=1}^{M} N_m$.

To compute the standard errors, all of the moments must be written in terms of the *overall* N observations. To do so, the first set of moments in (14) are reexpressed to yield the share of adopters of good j in market m in the *entire* sample of N observations. For good j in market m this moment is given by the

$$\frac{\sum_{n=1}^{N} I_{m(n)} I_{j(n)}}{N}$$

where $I_{m(n)}$ is an indicator that is on if individual n resides in market m, and $I_{j(n)}$ is an indicator that is on if individual n purchases good j.[50] The transformed moments - which are also zeroed by the parameter estimates - provide for a common index $n = 1, ..., N$ for all of the observations entering the estimator, and is equivalent to multiplying each of the moments from (14) by N_m/N, a weight that accounts for the differences in market sizes.

[50]Predicted market shares are computed in a similar manner using the model.

Having transformed the moments so that each observation contributes to every moment, the asymptotic behavior of all parameter estimates can be represented as a function of N. If the standard regularity conditions for GMM estimators hold in this data, the parameter estimates $(\hat{\theta}, \hat{\delta})$ are asymptotically normally distributed around the true values (θ_0, δ_0) and can be written in terms of N, the variance of the (stacked) moments V evaluated at (θ_0, δ_0), and the matrix of moment gradients Γ with respect to θ and δ (also evaluated at (θ_0, δ_0)):

$$\sqrt{N} \begin{pmatrix} \hat{\theta} - \theta_0 \\ \hat{\delta} - \delta_0 \end{pmatrix} \overset{A}{\sim} N(0, (\Gamma'\Gamma)^{-1} \Gamma' V \Gamma (\Gamma'\Gamma)^{-1}).$$

Because Γ is a square invertible matrix, we can rewrite this as

$$\sqrt{N} \begin{pmatrix} \hat{\theta} - \theta_0 \\ \hat{\delta} - \delta_0 \end{pmatrix} \overset{A}{\sim} N(0, \Gamma^{-1} V \Gamma'^{-1}).$$

Reported standard errors evaluate Γ and V at $(\hat{\theta}, \hat{\delta})$ using numerical approximations when necessary.

Table 1

Cable and DBS Subscribers (millions of households)

Year	Cable	DBS
1994	59.4	0.4
1995	62.1	2.2
1996	63.5	4.3
1997	64.9	5.0
1998	66.1	7.2
1999	66.7	10.1
2000	67.7	13.0
2001	69.0	16.1

Source: FCC, 2001 CS 01-001.

Table 2

Summary Statistics: Consumer Data

Variable	Mean	Std. Dev.
Indicators		
Local Antenna Only	0.218	0.413
Expanded Basic Cable	0.446	0.497
Premium Cable	0.233	0.423
Satellite	0.100	0.300
Rent	0.171	0.376
Single Unit Dwelling	0.834	0.371
Single Male	0.093	0.290
Single Female	0.097	0.297
Average Education (yrs)	14.417	2.151
Household Size	2.562	1.225
Household Income	$63,605	$47,185
Observations	29,484	

Source: Authors' calculations using Forrester Technographics, 2001.

Table 3

Summary Statistics: Television Markets

Variable	Mean	Std. Dev.
Monthly Cable Price	\$27.11	\$5.69
Premium Cable Price	\$38.60	\$6.06
Channel Capacity	66.14	21.87
Pay-Per-View Available	.785	.411
Year Franchise Began	1973.44	8.86
No. Top 6 Premium Channels	4.26	0.96
Number of Over-The-Air Channels	10.82	4.26
City Fixed Fee (Tax)	4.1%	1.4%
Indicator: ATT	.277	.448
Indicator: Adelphia	.078	.269
Indicator: Cablevision	.037	.191
Indicator: Charter Comm	.135	.342
Indicator: Comcast	.126	.332
Indicator: Cox Comm	.034	.183
Indicator: Time-Warner	.173	.379
Observations	317	

Source: Authors' calculations using Warren Publishing 2002 Television and Cable Factbook.

Table 4

First Stage Parameter Estimates

Explanatory Variable	Coefficient	Standard Error
Price for income group 2	0.046	0.002
Price for income group 3	0.071	0.002
Price for income group 4	0.087	0.003
Price for income group 5	0.098	0.003
Interactions with Expanded Basic Dummy		
Education	-0.030	0.001
Male*Single	-0.100	0.022
Female*Single	0.071	0.021
HHSize	-0.037	0.004
Interactions with Premium Dummy		
Education	-0.060	0.002
Male*Single	0.022	0.016
Female*Single	-0.072	0.025
HHSize	-0.003	0.002
Interactions with Satellite Dummy		
Education	-0.066	0.003
Male*Single	0.116	0.042
Female*Single	-0.479	0.059
HHSize	-0.002	0.004
Single Unit Dwelling	0.490	0.032
Rent	-0.174	0.036
Multivariate Normal		
$\sigma_{B,P}$	0.596	0.036
$\sigma_{B,S}$	1.042	0.069
$\sigma_{P,S}$	1.425	0.080
σ_P^2	1.186	0.065
σ_S^2	3.126	0.135
Log Likelihood	-34768	
Observations	29454	

Note: Specification is estimated using 951 product-market fixed effects for the 317 markets. See footnote 25 for a description of the multivariate normal parameters. The Appendix describes the approach to estimating the reported asymptotic standard errors.

Table 5

Marginal Effects on Purchase Probabilities (Estimated Percentage Changes)

For	MU Dweller	Renter	Household
Changing to	SU Dweller	Non Renter	Income
Change in probability (in %):			Increases 10%
Antenna Only	-1.81	-0.72	-4.32
Expanded Basic	-4.33	-1.67	0.42
Premium	-8.95	-3.43	2.61
Satellite	95.83	25.57	0.61

For	Not Male Single	Not Female Single	High School Educ.
Changing to	Male Single	Female Single	College Educ.
Change in probability (in %):			
Antenna Only	6.84	-0.99	22.79
Expanded Basic	-11.85	15.72	1.45
Premium	8.11	-5.56	-17.52
Satellite	19.34	-46.10	-12.08

Notes: The table reports the average percentage change in purchase probabilities arising from changing all people with the characteristic in the top row to having the characteristic listed in the bottom row. Because they are percentage changes, they do not sum to one. MU/SU Dwelling is Multi-Unit/Single Unit Dwelling, and Educ. is an index of average household education.

Table 6

Multivariate Normal vs. Logit: Substitution Patterns for DBS Subscribers (with DBS unavailable)

% Substituting to:	Multivariate Normal	Logit
Antenna Only	7.7%	23.0%
Expanded Basic	45.6%	50.9%
Premium	46.7%	25.9%

Source: Authors' calculation as described in the text.

Table 7

Stage Two Estimates from
Fixed Effects Regressions

Standard errors in parentheses

Explanatory Variable	SUR Exp Basic	SUR Premium	3SLS Exp Basic	3SLS Premium
Price	-.056 (.014)	-.056 (.014)	-.312 (.095)	-.312 (.095)
Channel capacity	-.001 (.001)	-.001 (.001)	.001 (.001)	.001 (.001)
Pay per view available	.095 (.052)	.095 (.052)	.323 (.105)	.323 (.105)
Year Franchise Began	-.010 (.002)	-.010 (.002)	-.010 (.002)	-.010 (.002)
No. Over-the-air channels	-.002 (.004)	-.002 (.004)	-.010 (.006)	-.010 (.006)
Two Premium Channels		.189 (.182)		.219 (.199)
Three Premium Channels		.078 (.063)		.024 (.069)
Four Premium Channels		-.048 (.053)		-.115 (.061)
Five Premium Channels		-.033 (.058)		-.060 (.062)
ATT is MSO	-.295 (.076)	-.107 (.080)	-.495 (.115)	-.343 (.130)
Adelphia Comm is MSO	-.001 (.110)	.263 (.115)	.032 (.128)	.266 (.141)
Cablevision is MSO	-.313 (.101)	.455 (.107)	-.295 (.118)	.402 (.132)
Charter Comm is MSO	-.227 (.093)	-.082 (.097)	-.324 (.113)	-.193 (.125)
Comcast is MSO	-.211 (.087)	.133 (.091)	-.366 (.117)	-.039 (.129)
Cox Comm is MSO	-.093 (.096)	.076 (.101)	-.145 (.113)	-.010 (.127)
Time-Warner is MSO	-.103 (.080)	.175 (.083)	-.031 (.096)	.222 (.103)
Constant	21.220 (4.160)	21.009 (4.164)	23.579 (5.148)	24.113 (5.159)
Observations	254	254	254	254

Note: Specification is estimated using the 254 markets for which the tax on franchise revenues is reported in Warren Publishing. SUR is seemingly unrelated regressions (not instrumented). 3SLS is three stage least squares using the tax to instrument price. Over the air channels appears here because local antenna only is the good to which the system is normalized. SUR regressions with all 317 observations yielded a very similar estimate of the price coefficient.

Table 8

Estimated Demand Elasticities

Marshallian and Hicksian

Method:	SUR	3SLS	3SLS
		Marshallian	Hicksian
Price of Expanded Basic			
Antenna only share	0.020	1.301	1.323
Expanded Basic share	0.014	-1.538	-1.516
Premium share	-0.040	1.263	1.284
Satellite share	-0.014	0.929	0.951
Price of Premium			
Antenna only share	-0.000	0.917	0.932
Expanded Basic share	-0.030	0.924	0.938
Premium share	0.074	-3.175	-3.161
Satellite share	-0.035	1.173	1.187
Price of Satellite			
Antenna only share	0.001	0.123	0.129
Expanded Basic share	-0.005	0.286	0.292
Premium share	-0.015	0.492	0.498
Satellite share	0.050	-2.448	-2.442

Note: Specification is estimated using the 254 markets for which the tax on franchise revenues is reported in Warren Publishing. SUR is seemingly unrelated regressions (not instrumented). 3SLS is three stage least squares using the tax to instrument price.

Table 9

Supply Side Response of Cable Systems

Explanatory Variable	Price Exp Basic	Price Exp Basic	Δ Price Exp Basic	$\beta x_{m,Base}$ (Chars Ind)	$\Delta\beta x_{m,Base}$ (Δ Chars Ind)
$\delta^*_{m,Sat}$	-2.390	-2.447	-2.215	0.043	0.056
	(0.552)	(0.527)	(0.804)	(0.021)	(0.022)
$\delta^*_{m,Base}$		5.038	4.083		
		(0.817)	(1.255)		
$\delta^*_{m,Prem}$		3.990	3.980		
		(0.709)	(1.096)		
$\xi_{m,Base}$	5.175				
	(0.896)				
$\xi_{m,Prem}$	3.761				
	(0.755)				
Channel Capacity	0.019				
	(0.012)				
Pay per view	3.340				
	(0.692)				
Year Fran Began	-0.004				
	(0.033)				
Over the Air Channels	-0.122				
	(0.066)				
City Fee (Tax)	0.244	0.441	0.038	-0.004	-0.013
	(0.243)	(0.226)	(0.357)	(0.010)	(0.065)
$\beta x_{m,Base}$ (for 1994)			-8.861		
			(3.377)		
HHsize	0.868	1.588	4.156	-0.037	-0.135
	(1.415)	(1.419)	(2.198)	(0.060)	(0.065)
MaleSingle	2.746	5.173	11.994	-0.737	-0.624
	(6.287)	(6.134)	(9.575)	(0.262)	(0.282)
FemaleSingle	4.308	4.096	1.715	-0.353	-0.550
	(4.625)	(4.610)	(7.006)	(0.193)	(0.205)
Rent	2.062	0.525	-5.365	0.371	0.400
	(4.943)	(4.857)	(7.700)	(0.207)	(0.229)
SingleUnit	-1.377	-2.121	-9.855	-0.281	-0.095
	(4.434)	(4.430)	(6.762)	(0.187)	(0.201)
Income	-0.813	-0.887	-1.201	-0.031	-0.013
($10,000)	(0.369)	(0.348)	(0.519)	(0.015)	(0.016)
Education	2.816	2.955	1.993	0.051	0.006
	(1.059)	(1.051)	(1.559)	(0.044)	(0.046)
94/01 Owner Ind	No/Yes	No/No	Yes/No	No/Yes	Yes/Yes
94/01 Prem Ind	No/Yes	No/No	Yes/No	No/Yes	No/No
R^2	0.627	0.582	0.377	0.187	0.704
Observations	250	250	243	250	247

Notes: The levels/differences regressions use 2001/1994-2001 data. The δ^*, $\beta x_{m,Base}$, and ξ's are demand side estimates of observed and unobserved quality (described in text). Demographics are franchise area averages. "94/01 Owner Ind" is Yes if separate indicators are included for the 1994 or 2001 multiple system operators (all regressions include some kind of proxy for ownership variables), and similarly for the top six premium channel indicators with "94/01 Prem Ind". Standard errors are corrected for the estimated nature of some regressors.