Quick-Service Expenditure in Ireland: Parametric vs. Semiparametric Analysis

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1. Introduction

Ireland has seen record levels of economic growth in recent times coupled with rising incomes and increased employment. These prosperous times have seen corresponding growth in the level of weekly household expenditure and labour force participation. As a result, food consumed away from home (FAFH) constitutes an increasingly important part of Irish food expenditure.[1] Between 1987 and 1999/2000 the proportion of total food expenditure allocated to FAFH increased from 14 per cent to 23 per cent.[2] In particular, the quick-service sector is recognised as the fastest growing component. The sector is somewhat diverse in that its components include the branded fast food chains, ethnic takeaways and traditional chip shop takeaways (Mintel, 2001). The sector has outperformed the wider eating out market in recent times in terms of growing market share at the expense of full service options such as hotel and restaurant meals.

The main objective of this paper is to analyse the factors determining expenditure on quick-service meals by Irish households. One feature of this study is that while previous studies on FAFH in Europe have examined the entire market (Manrique and Jensen, 1998; Mihalopoulos and Demoussis, 2001), little work has been done on disaggregating the market into its diverse sub-sectors.[3] A second objective is to compare some methodological alternatives which can be used in estimating models of household expenditure at the micro level. A

[1] In keeping with most other studies in this area this paper classifies foods 'at home' and 'away from home' based on where the food was prepared or obtained, not where it was consumed (Lin *et al.*, 2001).
[2] Authors own calculations.
[3] Lazaridis (2002) disaggregated the Greek market into expenditure on restaurant meals, expenditure in coffee houses and expenditure on takeaway meals and canteens.

discussion of both parametric and semiparametric alternatives to estimating limited dependent variable models of the kind proposed in this paper is also provided.

The paper is structured into the following sections. Section 2 describes the data and variables used in this study while Section 3 discusses the methodology. Section 4 compares the performance of the various estimators considered and identifies the most appropriate one for use in this analysis. The results of the model are described in Section 5. The paper concludes with Section 6.

2. Data and Variables

The data used in this study are taken from the Irish Household Budget Surveys (HBS) of 1994/5 and 1999/2000, collected by the Central Statistics Office of Ireland (CSO) (CSO, 1997; 2002).[4] In the HBS each household maintains a detailed diary of household expenditure over a two week period. Data on the socio-economic characteristics of household members are also collected. The survey covered a random sample of 7,877 and 7,644 households in both urban and rural areas throughout the state in 1994 and 1999 respectively. Only households whose size is readily identifiable from the HBS are included in this analysis. The records range from single adult households to households with up to four adults with and without children, resulting in a sample size of 7,305 households in 1994 and 7,171 households in 1999.

The dependent variable is household expenditure on quick service meals adjusted for household size using EU adult equivalence scales.[5] The HBS does not report price or quantity data and as a result price is assumed constant across

[4] The 1994/5 HBS and the 1999/2000 HBS are hereafter referred to as 1994 and 1999.
[5] EU adult equivalence scales give the first adult a weight of 1, each other adult 0.7, and each child under 14 years a weight of 0.5.

each cross-section. As quick-service meals are a relatively homogeneous product it is unlikely that there would be much variation in price. Quality differences are also uncontrolled for due to data limitations.

The theory of household production (Becker, 1965) underpins much of the literature on FAFH consumption. The household is seen as a consuming and a producing unit. The value of the household's time is therefore important as households with a higher opportunity cost of time can be expected to outsource household production such as meal preparation by purchasing time-saving products such as FAFH. In this study, the number of workers employed in the labour force is used as a proxy for household time constraints.[6] Furthermore, the ever expanding commuter belt around urban areas in Ireland and the increase in the length of the average commute to work places greater time pressures on individuals commuting long distances compared with those that do not. A variable representing whether or not the household is a 'commuter' household is therefore also incorporated into this analysis as a second proxy for the opportunity cost of time of the household.

Recent evidence of Irish households' food expenditure patterns has indicated that there are two competing factors influencing expenditure decisions: health knowledge and convenience (Newman *et al.,* 2001; 2003). In particular, recent media exposure to the health consequences of obesity and diabetes may affect the demand for quick-service foods. FAFH has been found to have lower nutritional quality than food prepared at home across international studies (Binkley *et al.,* 2000; Burns *et al.,* 2001; Guthrie *et al.,* 2002). In this analysis, the behaviour of

[6] It can be hypothesized that the greater the proportion of household members in the labour force the less likely they will have time for food preparation in the home leading to a greater reliance on processed foods and other time-saving choices such as FAFH and quick-service meals.

households in relation to the purchase of goods such as alcohol and tobacco, products associated with known health risks, are used to proxy the health awareness of households. Furthermore, it is hypothesised that education levels may have a bearing on the decision to participate in the FAFH market: better educated household managers could be expected to have greater awareness of the health consequences of eating certain types of foods.

Previous empirical research has identified how specific economic and demographic characteristics of the household influence demand for FAFH. Household size, the presence of children, being resident in urban areas, age and income levels all have been found to be significant factors that should be considered in a study on FAFH expenditure.[7] In analysing food expenditure patterns of households on aggregate we are attempting to explain the expenditure decisions of the household manager, that is the individual primarily responsible for household activities including meal preparation. The HBS, however, does not indicate which household member this is. Following Stewart *et al.* (2004), the household manager for single households is defined as the survey respondent while for married couples the household manager is defined as the person who works the fewest hours outside of the home. This approach, while straightforward for households of one adult or a married couple, becomes ambiguous for households of multiple unmarried adults. In these cases the household manager is always the survey respondent.[8] Where individual characteristics are used to explain expenditure on FAFH they refer to characteristics of the household manager defined in this way rather than the head of household as has been the

[7] See for example, Byrne *et al.* (1998), McCracken and Brandt (1987) and Stewart *et al.* (2004) amongst others.

[8] In the HBS the head of household is the oldest person in the household and given that the completion of the expenditure diary is in itself a task indicative of household management the choice of the head of household as the household manager can easily be justified.

case in previous studies. All variables used in this analysis are described in Table 1 with descriptive statistics presented in Table 2.

INSERT TABLES 1 AND 2 ABOUT HERE

3. Methodology

Cross-sectional data on household expenditure patterns are complicated by the existence of zero observations on expenditure, implying that the relationship between the independent variables and the expenditure variables may be more complex than assumed in conventional regression models. The standard tobit model (Tobin, 1958) was originally developed to accommodate censoring in the dependent variable and was designed to overcome the bias associated with assuming a simple linear relationship between the dependent and independent variables in the presence of such censoring. The tobit model assumes that all zeros are attributable to standard corner solutions. Negative values of the dependent variable are assumed to exist but are considered to be unobservable and bunched at zero. By assuming that an unobservable latent framework generates the data the model incorporates this assumption into the modelling process.[9] The standard tobit model can be written as:

$$
\begin{array}{lll}
y_i* = x_i'\beta + u_i & u_i \sim N\left(0, \sigma^2\right) & i = 1,..........,n \\
y_i = y_i* & \text{if } y_i* > 0 & \\
y_i = 0 & \text{otherwise} &
\end{array}
\qquad (1)
$$

[9] In the case of expenditure on food items, such as FAFH, we assume that the household's decision to purchase a food item depends on the utility they derive from the consumption of that item. It may be the case that the household dislikes the food item to such an extent that they attach a negative level of utility to it. In such a case, all we would observe in the expenditure data are zero values. As such, the tobit model, which takes account of censoring of this kind in the dependent variable, is appropriate in this case.

where x_i is vector of explanatory variables corresponding to the ith household, y_i is the observed level of expenditure by the ith household and y_i^* is an unobserved continuous latent variable assumed to determine the value of y_i. The latent variable is only observed if it is greater than or equal to zero, however, it is allowed to take on negative values even though these values cannot be observed.

Standard estimators for these types of models are based on maximum likelihood estimation (MLE). MLE produces consistent estimates of the parameters of the tobit model under appropriate assumptions such as homoscedasticity and the normality of the error terms. The likelihood function of the tobit model can be written as (Tobin, 1958):

$$L(\beta,\sigma^2) = \prod_{y_i=0}\left[1-\Phi\left(\frac{x_i'\beta}{\sigma}\right)\right]\prod_{y_i=1}\left[\sigma^{-1}\phi\left(\frac{y_i - x_i'\beta}{\sigma}\right)\right] \qquad (2)$$

where Φ and ϕ refer to the standard normal cumulative distribution and density functions respectively. The consistency of MLE requires a complete and correct specification of a parametric family of the error distribution. If the model is misspecified, model assumptions must be relaxed, and estimators are needed which remain consistent under more general assumptions. Semiparametric estimators have been developed for this purpose.

Semiparametric estimators are hybrids of parametric and nonparametric approaches. They allow for a more general specification of the nuisance parameters, are more consistent than corresponding parametric models and are typically more precise than their nonparametric counterparts (Powell, 1994). However, if the parametric model is correctly specified, they are in general less efficient than the corresponding maximum likelihood estimator (Powell, 1994). Semiparametric estimators useful for cross-sectional type analyses include the censored least absolute deviation (CLAD) and symmetrically censored least squares (SCLS) estimators (Powell, 1984; 1986). Both are considered in this paper.

The CLAD estimator was first proposed by Powell (1984). The CLAD model is a median estimator. To make the estimator robust to misspecification problems the sample is reduced eliminating data points and observations that fall outside the uncensored region from the sample (the recensoring step). Least absolute deviations is then applied to the remaining observations (the regression step). Bootstrapping is used to compute the residuals. The CLAD estimator is robust to heteroscedasticity and non-normality and provides consistent estimates in the presence of censored data. The estimator, however, may be less efficient than its parametric alternative depending on the extent to which outliers are a problem in the dataset. A number of studies have compared the CLAD approach to MLE with varying results.[10] Blisard *et al.* (2004) applied CLAD to analyse household expenditure on fruits and vegetables finding it to be more robust to outliers than least squares estimation. Chay and Powell (2001), analysing the earnings gap between black and white households, found that semiparametric estimation produced much more precise estimates than MLE. In contrast, Berg and Kaempfer (2001) failed to reject a heteroscedasticity adjusted model estimated using MLE in favour of the CLAD approach in their application to cigarette demand in South Africa.

The SCLS estimator was also proposed by Powell (1986). It is built on the CLAD estimator but uses symmetric trimming. By assuming that the true

[10] A number of studies have also considered the Least Absolute Deviations approach, without censoring, as an alternative to the standard parametric models for dealing with censoring finding the former to be superior in most cases. See for example, Fullerton (1998), Yoo (2001) and Yoo *et al*. (2000).

dependent variable is symmetrically distributed around the regression function and that the observed dependent variable will have an asymmetric distribution, symmetry can be restored by symmetrically censoring the dependent variable. The coefficients can be estimated by least squares with observations falling outside the uncensored region purged, using the symmetrically trimmed data only to arrive at the SCLS estimator.[11] The motivation for using the symmetric trimming approach is that consistency will not be dependent on either homoscedasticity or the known distribution of the error terms (Powell, 1986). Yoo (2003) found that SCLS significantly outperformed standard parametric alternatives in dealing with zero expenditures on mobile communications in South Korea. Chay and Powell (2001) also used the SCLS estimation to analyse the black and white earnings gap however, they conclude that the CLAD estimator was superior due to its consistency under more general assumptions.

4. Model Comparison

INSERT TABLE 3 ABOUT HERE

The model of Irish households' quick-service expenditure is estimated using MLE, CLAD and SCLS estimation in econometrics software package STATA S/E Version 8.[12] Estimation of the tobit model using the 1994 data reveals that the semiparametric estimation performs poorly against MLE (Table 3). In particular, the results of the model are extremely different when estimated using SCLS compared with the other techniques. While the coefficients have mostly the same sign they diverge considerably in magnitude. For example the coefficient for income estimated using SCLS is 13.076, twice as large as the

[11] For consistency, bootstrapping is also used to compute the residuals of the SCLS model.
[12] The respective ado files were obtained at http://elsa.berkeley.edu/~kenchay. For consistency, the residuals of the tobit model estimated using MLE are computed using bootstrapping.

coefficient when estimated using MLE (5.1147). In addition, the estimated standard errors are much larger in the former resulting in what appears to be a very poorly specified model. This is due to the fact that SCLS estimation reduces the data to 1,744 observations through symmetric trimming, purging 5,561 observations in total. This suggests that the true distribution of the dependent variable is highly asymmetric and unlikely to be normal. CLAD estimation of the 1994 model performs slightly better with a sample of 2,381 after trimming, although standard errors are still larger than those produced using MLE.

Similarly, in 1999 the sample for the CLAD (3,764) and SCLS (3,379) are greatly reduced by trimming when compared with the full sample of 7,171, clearly indicating the presence of a significant asymmetry in the distribution of the dependent variable. The SCLS performs poorly again with standard errors larger in most cases than in both the CLAD and MLE approaches. However, the CLAD estimates have shown some improvements. Standard errors are lower for CLAD estimation than MLE for many of the estimates.

Newey (1987) proposed a version of the Hausman for testing between semiparametric and parametric alternatives of the censored tobit model. The null hypothesis for the test is that the parametric tobit model is consistent while the alternative is that it is inconsistent indicating the presence of misspecification problems such as heteroscedasticity or non-normality. The CLAD and SCLS estimators will be consistent under both the null and alternative hypotheses as they are robust to such misspecification errors. The test statistic is given by (Newey, 1987):

$$H = \left(\hat{\beta}_{sp} - \hat{\beta}_{ml}\right)' \left[V\left(\hat{\beta}_{sp}\right) - V\left(\hat{\beta}_{ml}\right)\right]^{-1} \left(\hat{\beta}_{sp} - \hat{\beta}_{ml}\right) \tag{3}$$

The statistic is distributed χ^2 with degrees of freedom equal to the number of parameters. The results of this test comparing the parametric model to the semiparametric models in each year are presented in Table 4.

INSERT TABLE 4 ABOUT HERE

In all cases the tobit model estimated using MLE is strongly rejected in favour of the semiparametric alternatives. However, as our discussion above revealed, the semiparametric estimates are far inferior to the MLE estimates in efficiency terms, particularly in 1994. We are therefore faced with two options as to how we should proceed: on the one hand we could proceed with the semiparametric estimates accepting the loss in efficiency; or on the other hand we could attempt to adjust the tobit model so that MLE yields consistent estimates. In this application, the extent of the efficiency loss in using the semiparametric approaches is so great that the analysis proceeds with the latter option.

The maximum likelihood estimates presented in Table 3, while more efficient than the semiparametric estimates are inconsistent as they are not robust to non-normality or heteroscedasticity. Further testing for these misspecification problems confirms our findings as both homoscedasticity and normality are rejected.[13] To correct for these problems multiplicative heteroscedasticity of the form given in Equation 4 is assumed (where z_i are the continuous variables of the model) and an inverse hyperbolic sine (IHS) transformation of the dependent variable (see Equation 5) is used to allow for non-normality (Burbidge *et al.*, 1998).[14] The log likelihood for the adjusted model is given in Equation 6.

$$\sigma_i = \sigma \exp\left(z_i{}'\alpha\right) \tag{4}$$

[13] The results of a Lagrange Multiplier test for heteroscedasticity, and Pagan and Vella's (1989) moments based test for non-normality are presented in the Appendix.
[14] See Newman *et al.* (2003) for an example of an application of the adjusted model.

$$T(\theta y_i) = \log\left[\theta y_i + \left(\theta^2 y_i^2 + 1\right)^{\frac{1}{2}}\right] \Big/ \theta = \sinh^{-1}(\theta y_i) \quad (5)$$

$$L(\beta,\sigma^2) = \prod_{y_i=0}\left[1-\Phi\left(\frac{x_i'\beta}{\sigma_i}\right)\right]\times \prod_{y_i=1}\left[\sigma_i^{-1}\phi\left(\frac{y_i - x_i'\beta}{\sigma_i}\right)\left(1+\theta^2 y_i^2\right)^{-\frac{1}{2}}\right] \quad (6)$$

Likelihood ratio tests imply that the adjusted tobit model is superior to the standard tobit model in modelling quick-service expenditure using Irish data (see Appendix). Adjusted for heteroscedasticity and non-normality the standard errors are dramatically lower than the unadjusted tobit and each of the semiparametric estimators for both of the datasets (Table 5).

5. Results

INSERT TABLE 5 ABOUT HERE

The results of the IHS heteroscedastic tobit model are presented in Table 5. Income has a positive effect on quick-service expenditure but at a decreasing rate. This is in line with previous results: as households earn more income they purchase more leisure activities, including dining amenities (Byrne *et al.*, 1998; McCracken and Brandt, 1987). At higher income levels, households move away from quick service consumption potentially toward other full-service options.

Household size has a significantly positive effect on quick-service expenditure though at a decreasing rate. Lazaridis (2002) found that family size had a positive effect on expenditure on takeaway meals and attributed this to the fact that the more members in the household the greater the probability that at least one member will spend on FAFH. Mintel (2004) found that households with four or more adults had the highest penetration for quick-service foods in Ireland, a result supportive of this finding.[15]

The age of the household manager has a negative and statistically significant effect on quick-service expenditure. The negative age effect is common to many studies on FAFH demand that have shown that older households display reduced levels of expenditures compared to younger households. Stewart *et al.* (2004) found that older U.S. household managers are less likely to frequent quick-service outlets and have a preference for full-service options as they age.

Being a female has a significant and positive effect on quick-service expenditure in 1999. Interestingly this result is in line with that of Mintel (2004) who found that while overall the percentage of adults eating fast foods and takeaways have increased, women have shown a more positive shift than men. It is likely that increasing female labour force participation rates are influencing this result.

Being a married couple has a negative effect on quick-service expenditure. This result supports the hypothesis that married couples would be more likely to eat food at home as against single couples due to economies of scale and possibly an added value associated with the family meal eating occasion. The results for single households are surprising as they indicate that being in such households has a negative effect on quick-service expenditure. This result is contrary to expectations given that the benefits of preparing one's meals diminish in smaller households. It is possible that an age effect is being observed here as single pensioners are one component of this category of households. In both sets of

[15] In contrast, it should be noted that Stewart *et al.* (2004) found that larger households spend less on all segments of FAFH, supporting their hypothesis that such households benefit from economies of scale in food preparation at home.

results the coefficient on the presence of children is insignificant throughout. One would have expected children to have a positive effect on quick-service expenditure given that advertising in the sector is heavily targeted at children. The results, however, do not bear this out.

Households resident in urban areas have a significant and positive effect on quick-service expenditure. This can be explained by the fact that towns provide greater access to a variety of quick-service outlets compared with rural areas. Jekanowski *et al.* (2001) showed that increased outlet density of fast food restaurants enhanced the convenience associated with fast foods and thus increased the growth in fast food expenditure. They suggested that most of the growth in fast food consumption is due to an increasing supply of convenience.

Home ownership also has a negative effect on quick-service expenditure. The act of owning a home can be considered indicative of a household investing time in household production, including meal preparation. Renters on the other hand, not having the same level of attachment to the home, and in all likelihood having more individualistic lifestyles, would be expected to favour quick-service as a convenience option. Manrique and Jensen (1998) found a similar result for Spanish households.

As expected the number of workers is positively related to quick-service expenditure, consistent with the hypothesis that households with high time values will eat out rather than at home to save time. This result is consistent with household production theory as households with higher time constraints would be expected to sacrifice time from household production to engage in other activities (Lazaridis, 2002). Similarly the coefficient on the commuter variable is positive and significant in both years. The positive relationship between commuters and the purchase of quick-service can be interpreted as a further demand for convenience products by commuters. With the ever expanding commuter belt in Ireland at present this will undoubtedly have a favourable impact on the quick-service industry into the future.

The result for the health awareness variable highlights how there are two competing forces influencing demand for quick-service: the demand for convenience competes with the demand for a healthier lifestyle. This variable is negative and significant in both survey years indicating that households with high levels of health awareness are less likely to consume quick-service foods. In addition, in 1999, having higher levels of education has a significantly negative effect on quick-service expenditure. This could reflect, as hypothesised, that higher levels of education are consistent with greater levels of health awareness. On the other hand, as found by Byrne *et al*. (1998), this result may reflect the fact that households with higher education levels may be more likely to consume FAFH at up-market facilities rather than at quick-service facilities. A similar result is found for the social class of the household manager with higher social classes spending less on quick-service.

The household appliances were included as they have been found to have significant effects on Irish expenditure on prepared meals (Newman *et al.* 2003). The results are as expected with freezer ownership, a proxy for close substitutes to FAFH expenditure such as prepared meals, having a negative effect on quick-service expenditure and microwave ownership, a proxy for convenience preferences in food preparation, having a positive effect. There is also some evidence of seasonality in the results supporting the inclusion of seasonal dummies.

6. Conclusion

The main objectives of this paper were to analyse the factors determining quick-service expenditure in Ireland and in so doing explore semiparametric alternatives to maximum likelihood estimation of tobit models. This study found that there was an efficiency/consistency trade-off when the semiparametric estimators were compared to the unadjusted tobit model. Despite the technical difficulties associated with estimating tobit models in the presence of heteroscedasticity and non-normality, the loss in efficiency suffered in the semiparametric approach was considered too great a sacrifice to make and so we proceeded with an adjusted model. While the findings of this study correspond with most others in terms of the outright rejection of the MLE approach (see for example Chay and Powell (2001) and Yoo (2003)), in this case the semiparametric approaches were not considered efficient enough to be used as an alternative. This may be due to the larger dataset, more censoring and a greater number of outliers characteristic of the sample used in this study compared with others. Many of the results of the variables examined have the expected sign and most are significant. Household income, household size and urban residents have significant and positive influences on quick-service expenditure. Older families, single households and married couples and households with higher levels of education and health awareness experience reduced expenditure levels.

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Table 1: Description of variables

Dependent Variable	*Description*
Quick Service	Per capita average weekly expenditure on quick-service ( )
Independent Variables	
Income	Proxied by per capita average total weekly household expenditure ( )
$Income^2$	Income squared
Age	Age of household manager (1-8)
Hhold	Number of persons in the household
$Hhold^2$	Household size squared
Workers	Number of persons in gainful employment outside the home
Social Class	Social1 = 1 for household manager categorised as higher professional, lower professional, employer or manager, 0 otherwise Social2 = 1 for household manager categorised as salaried employee and non-manual workers, 0 otherwise Base category = household manager categorised as manual worker, farmer, other agricultural worker or fishermen
Female	1 = Female household manager 0 = Male household manager
Single, married	Single=1 for single adult household with or without children, 0 otherwise Married=1 for married couple with or without children, 0 otherwise Base category = households with 2 or more adults with or without children
Education	1 = if household manager has Leaving Certificate or a higher level 0 = if household manager has less than Leaving Certificate education
Homeowner	1 = Household owns their own home 0 = Household does not own their own home
Urban	1 = Urban household 0 = Rural household
Children	1 = Children present 0 = No children present
Commuter	1 = A Household member is employed outside the home and incurs travelling expenses 0 = Household members are not in employment or do not incur travelling expenses
Health	1 = Household spends nothing on alcohol or tobacco during the survey period 0 = Household spends a positive amount on tobacco or alcohol during the survey period
Household Appliances	Freezer = 1 if household is in possession of a freezer, 0 otherwise Microwave = 1 if household is in possession of a microwave, 0 otherwise
Seasonal dummies	Spring = 1 if consumption occurred in Spring, 0 otherwise Summer = 1 if consumption occurred in Summer, 0 otherwise Autumn = 1 if consumption occurred in Autumn, 0 otherwise Base category = consumption occurred in Winter

Table 2: Summary statistics

	Mean ( )		*Std. Deviation*		*Maximum( )*		*% Zeros*	
Dependent	*1994*	*1999*	*1994*	*1999*	*1994*	*1999*	*1994*	*1999*
Quick Service	1.029	1.878	2.133	3.423	35.56	78.81	56%	50%
Independent - Continuous								
Income (ln)	4.907	5.274	0.614	0.679	7.039	8.401		
Income2 (ln)	24.455	28.271	6.054	7.119	49.559	70.569		
Age	5.050	5.274	1.725	1.657	8	8		
Hhold	2.932	2.904	1.674	1.535	11	12		
Hhold2	11.398	10.789	12.222	10.676	121	144		
Workers	1.016	1.160	0.839	0.911	4	4		
Independent - Discrete								
Education	0.360	0.458						
Social1	0.217	0.242						
Social2	0.216	0.273						
Urban	0.553	0.649						
Children	0.390	0.376						
Female	0.464	0.449						
Single	0.241	0.281						
Married	0.406	0.370						
Homeowner	0.789	0.832						
Commuter	0.268	0.349						
Health	0.201	0.188						
Freezer	0.216	0.294						
Microwave	0.457	0.704						
Spring	0.238	0.195						
Summer	0.260	0.368						
Autumn	0.431	0.304						

Table 3: Semiparametric and parametric estimation results: 1994, 1999

		1994			*1999*	
	CLAD	*SCLS*	*Tobit*	*CLAD*	*SCLS*	*Tobit*
Constant	-23.083***	-39.9982	-16.1292***	-34.150***	-45.194***	-23.2261***
	(5.8053)	(28.2465)	(3.9114)	(6.3429)	(0.0069)	(4.7973)
Income	7.7623***	13.076	5.1147***	11.3858***	15.4230***	7.4421***
	(2.0859)	(9.8102)	(1.4987)	(2.2385)	(3.6597)	(1.8403)
Income^2	-0.6142***	-1.0014	-0.3360**	-0.9010***	-1.2214***	-0.4932***
	(0.1911)	(0.8240)	(0.1467)	(0.1977)	(0.3170)	(0.1711)
Age	-0.4927***	-0.9349	-0.6611***	-0.7544***	-1.0146***	-1.1180***
	(0.1048)	(0.7211)	(0.0495)	(0.0922)	(0.1670)	(0.0771)
Hhold	0.6952*	0.1887	0.5492***	1.5758***	1.1589***	1.2973***
	(0.3857)	(1.5902)	(0.1445)	(0.3424)	(0.4483)	(0.2680)
Hhold^2	-0.0713*	-0.0209	-0.0478***	-0.1381***	-0.1041**	-0.1113***
	(0.0447)	(0.2323)	(0.0151)	(0.0082)	(0.0498)	(0.0247)
Workers	0.3209***	0.4222	0.4071***	0.2115*	0.2502*	0.4162***
	(0.0906)	(0.4265)	(0.0777)	(0.1098)	(0.1342)	(0.1047)
Female	0.0559	-0.1528	-0.1453	0.1130	-0.0096	0.0328
	(0.1547)	(0.3639)	(0.1282)	(0.1495)	(0.2181)	(0.1326)
Education	-0.2921*	-0.3094	-0.1869	-0.2382	-0.5027**	-0.4647***
	(0.1751)	(0.6389)	(0.1224)	(0.1628)	(0.2267)	(0.1950)
Single	-0.1876	-0.8195	-0.7047***	-0.5609	-1.0438**	-1.4785***
	(0.4581)	(1.4597)	(0.2087)	(0.3668)	(0.4798)	(0.4579)
Married	-0.6485***	-0.4292	-1.3399***	-0.8819***	-1.0048***	-1.8597***
	(0.1605)	(0.6069)	(0.1132)	(0.1555)	(0.2661)	(0.2503)
Social1	-0.1520	-0.7513	-0.5353***	-0.6743***	-0.6064**	-0.1683
	(0.1795)	(0.7843)	(0.1741)	(0.2107)	(0.3086)	(0.2079)
Social2	0.1340	0.0780	-0.0621	-0.2805*	-0.4061*	0.0513
	(0.1449)	(0.7941)	(0.1467)	(0.1596)	(0.2273)	(0.1576)
Urban	0.8822***	3.0017**	1.2199***	1.3033***	1.5911***	1.6910***
	(0.3242)	(1.3566)	(0.1228)	(0.1540)	(0.2083)	(0.1591)
Children	0.0117	0.1281	-0.0077	-0.6472**	-0.2761	-0.1070
	(0.2109)	(0.6191)	(0.1475)	(0.2567)	(0.2523)	(0.2685)
Homeowner	-0.3905*	-0.4539	-0.4624***	-0.4138*	-0.4086	-0.5083**
	(0.1905)	(0.5038)	(0.1662)	(0.2163)	(0.3092)	(0.2327)
Commuter	0.4807***	0.5872	0.5719***	0.6774***	1.1208***	0.9664***
	(0.1459)	(0.3978)	(0.0966)	(0.1338)	(0.1900)	(0.1816)
Health	-0.6414**	-1.3168	-0.8563***	-0.4330	-0.7172	-0.5856***
	(0.3307)	(0.8837)	(0.2014)	(0.2853)	(0.8723)	(0.1773)
Freezer	-0.1486	-0.2777	-0.1961	-0.2248	-0.3027*	-0.3062*
	(0.1701)	(0.4433)	(0.1211)	(0.1572)	(0.1635)	(0.1681)
Microwave	-0.1060	0.1032	0.1846*	0.0330	0.3602	0.2315
	(0.1366)	(0.2711)	(0.1076)	(0.1946)	(0.3566)	(0.1830)
Spring	0.1051	0.4882	0.0260	0.2679	0.4389*	-0.0211
	(0.1755)	(0.9908)	(0.1486)	(0.2234)	(0.2596)	(0.1936)
Summer	0.2573	0.4959	0.0814	0.3755*	0.4906**	0.1324
	(0.1356)	(0.5659)	(0.1268)	(0.2139)	(0.2255)	(0.2026)
Autumn	0.2497	0.3148	0.2446**	0.5504***	0.5634**	0.2704
	(0.1657)	(0.4334)	(0.1194)	(0.1955)	(0.2695)	(0.1668)

Notes: ***significant at the 1% level, **significant at the 5% level, *significant at the 10 % level. Standard errors are given in parentheses.

Table 4: Hausman test of consistency of parametric tobit model

	1994		1999	
	Test Statistic	*P-value*	*Test Statistic*	*P-value*
MLE vs. SCLS	330.88	0.0000	178.93	0.0000
MLE vs. CLAD	115.40	0.0000	170.31	0.0000

Table 5: IHS heteroscedastic tobit results: 1994, 1999

Variable	*1994 IHS Het Tobit*	*Heteroscedasticity Terms*	*1999 IHS Het Tobit*	*Heteroscedasticity Terms*
Constant	-5.8034***	1.0623***	-8.7794***	1.2913***
	(0.8715)	(0.0754)	(1.0907)	(0.0925)
Income	1.900***	-	2.7647***	
	(0.3386)		(0.4034)	
Income^2	-0.1371***	-	-0.1997***	
	(0.0332)		(0.0371)	
Age	-0.2072***		-0.2813***	-0.0174*
	(0.0145)		(0.0194)	(0.0099)
Hhold	0.3667***	-0.3736***	0.6932***	-0.3287***
	(0.0623)	(0.0405)	(0.0833)	(0.0379)
Hhold^2	-0.0336***	0.0311***	-0.0647***	0.0276***
	(0.0068)	(0.0053)	(0.0090)	(0.0049)
Workers	0.1189***		0.0812***	0.0208
	(0.0183)		(0.0225)	(0.0153)
Female	-0.0047		0.0719**	
	(0.0272)		(0.0364)	
Education	-0.0489		-0.1348**	
	(0.0306)		(0.0384)	
Single	-0.3983***		-0.4742***	
	(0.0752)		(0.1000)	
Married	-0.3419***		-0.4334***	
	(0.0343)		(0.0439)	
Social1	-0.1260**		-0.0842*	
	(0.0383)		(0.0475)	
Social2	-0.0247		0.0046	
	(0.0320)		(0.0394)	
Urban	0.2819***		0.4159***	
	(0.0308)		(0.0360)	
Children	-0.0153		-0.0777	
	(0.0441)		(0.0546)	
Homeowner	-0.1217**		-0.1555**	
	(0.0384)		(0.0499)	
Commuter	0.0877**		0.1350***	
	(0.0268)		(0.0352)	
Health	-0.2680***		-0.2055***	
	(0.0468)		(0.0581)	
Freezer	-0.0398		-0.0760**	
	(0.0296)		(0.0353)	
Microwave	0.0452		0.0884**	
	(0.0277)		(0.0429)	
Spring	-0.0196		0.0294	
	(0.0348)		(0.0570)	
Summer	0.0330		0.0703	
	(0.0298)		(0.0526)	
Autumn	0.0633**		0.1237**	
	(0.0290)		(0.0537)	
IHS Parameter	0.3812***		0.2830***	
	(0.0244)		(0.0925)	
Log Likelihood	-3589.40		-4362.95	

Notes: *** significant at 1% level, **significant at the 5% level, * significant at the 10% level. Standard errors are given in parentheses.

Appendix

Specification testing

	1994		*1999*	
	Test Statistic	*P-value*	*Test Statistic*	*P-value*
Lagrange Multiplier Test for Heteroscedasticity Ho: Homoscedasticity	4628.95	0.0000	2510.14	0.0000
Conditional Moments Test for Non-normality Ho: Normality	330.29	0.0000	234.55	0.0000
Likelihood Ratio Test for IHS Heteroscedastic Tobit Model Ho: Unadjusted Model	12,847.50	0.0000	15,539.48	0.0000