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SPENDING ON ALCOHOL: EVIDENCE FROM THE FAMILY EXPENDITURE SURVEY 1970–1983*

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Household expenditure on alcoholic drink is substantial, constituting around $7\frac{1}{2}\%$ of total household spending in the United Kingdom. This expenditure is of policy interest for two major reasons. First, taxes on beer, wines and spirits provide a major source of government revenue, accounting for some 10% of indirect tax receipts. How is the share of consumer spending devoted to alcohol likely to be affected by changes in income or in occupational or demographic structure? Second, there is concern on social and health grounds with the level of consumption and its distribution. Our data[1] for the examination, in this paper, of the factors governing the pattern of consumption are drawn from a pooling of cross-sections from the *Family Expenditure Survey* (FES) for the years 1970 to 1983. In Fig. 1 we show the distribution of shares of alcohol in total household spending for households in the sub-sample of the FES from 1970–83 used here. The distribution has a mode close to zero and a long tail to the right, which extends well beyond the range shown in the diagram: the top percentile begins at a share of 27%. It seems unlikely that this could be explained by price and income variation alone and suggests that careful attention to stochastic distributional assumptions is required. Moreover, unlike many goods, there is a sizeable proportion of households – around one fifth – who have zero recorded expenditure. (These observations are not shown in the diagram.)

In this paper we attempt to model how observed characteristics affect the share of spending and the incidence of zero expenditure. The characteristics include occupation and economic activity of the head of household, the household composition, age, and region. The model that we seek to estimate has a single equation[2] and is a generalisation of the Tobit. It is described in Section I. The estimates are discussed in Section II. We concentrate here on total spending on alcohol; in a companion paper (Atkinson *et al.* 1989*b*) we present separate results for beer, wine, and spirits. The implications of the results in terms of the determinants of alcohol spending are the subject of

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[1] We provide in Atkinson *et al.* (1989*a*) a more extended discussion of the crucial issues of data and the interpretation of results, including comparisons with those of other researchers. Unfortunately pressures of space have meant that the discussion of these issues in this version has had to be curtailed.

[2] Our focus on one equation is because we wish to experiment extensively with the specification of characteristics and the distributional assumptions. Notice that a one-equation model can provide a useful framework for tax-revenue analysis using a composite tax rate for other goods and the cross-elasticity of demand for the composite commodity (see Stern, 1987).

Section III. The conclusions are summarised in Section IV. The main analytical contributions of the paper lie in the detail of the specifications of characteristics and the use of the gamma distribution to model stochastic terms (giving the gamma-Tobit). Together they can provide substantial insights into the way in which alcohol consumption and its response to changes in variables vary within the population, although as we shall show, and this constitutes a crucial part of the argument of the paper, the results from this type of model need careful interpretation.

I. MODELLING ALCOHOL EXPENDITURE

The variable to be explained is taken throughout to be the share of total household expenditure on alcohol, denoted by w. The basic formulation that we have used is the following:[3]

$$w = \alpha + \beta \log(m/\pi) + \delta[\log(m/\pi)]^2 + \gamma \log(p/\pi), \tag{1}$$

where

w is the share of expenditure on alcohol
m is total household expenditure
p is the price of alcohol
π is the retail price index

and α, β, γ, δ are coefficients which may depend on household characteristics but not on m/π or p/π. We have a simple (restricted) quadratic approximation in the space of w, $\log m$, $\log p$ and household characteristic variables.

Using x to denote the quantity purchased, so that $w = p\,x/m$, we have the expenditure elasticity

$$(m/x)\,\partial x/\partial m = 1 + [\beta + 2\delta \log(m/\pi)]/w. \tag{2}$$

The effect of price variation is represented by the own-price term, with other prices entering only via the price index π. If, as we shall do in interpreting the results, we ignore the contribution of the price of alcohol to the overall retail price index, then the own-price elasticity is:

$$-(p/x)\,\partial x/\partial p = 1 - \gamma/w. \tag{3}$$

We experimented with the inclusion of other prices, particularly that of tobacco, but the simple form with p and π proved superior in terms of likelihood values. However, whether or not there is a smoker in the household does seem to be of importance and is included amongst the characteristics.

Equations (2) and (3) remind us that the estimated variation of the elasticity within the population is strongly influenced by functional form (and the form plays an important role in tax analysis too, see Atkinson and Stern, 1980). An alternative functional form based on the linear expenditure system was also investigated but did not show advantages over that discussed here (see Atkinson *et al.* 1989*a*).

[3] See, for example, Working (1943), Leser (1963) and Deaton and Muellbauer (1980).

Characteristics

The vector $\mathbf{z}$ represents characteristics, which are listed in the Data Appendix and include the number of household members of different types, the region, the age of the household head, etc. One possible baseline when considering the impact of $\mathbf{z}$ on the shares of spending is to assume that it leads to a simple shift in the constant term, α, in equation (1). Families with children allocate a larger fraction of their expenditure to, say, food, and a lower fraction to alcohol. It is however clear that the demographic variables may have a differential impact on the share at different expenditure levels. This would mean that $\mathbf{z}$ would affect the value of β. Similarly, $\mathbf{z}$ may enter the determination of γ. The latter would have the important consequence that the price elasticity could vary across households directly, rather than the price elasticity being simply a transform of the share (see equation (3)). We therefore consider in our empirical work the interaction between the characteristic variables, $\mathbf{z}$, and the income and price terms (see Pollak and Wales, 1981, for a discussion of these issues).

Treatment of Stochastic Term and Zero Expenditures

The observed spending may depart from that predicted for a particular household using a demand equation such as (1) for several different reasons including random transitory deviations of the actual spending from its 'normal' level (e.g. infrequency of purchase), a systematic unobserved difference for that particular household (which we call a *fixed effect*), errors in measurement and so on. This is typically treated by adding to the equation a stochastic term ϵ, assumed to be independent and identically distributed across individuals, with mean zero and variance σ^2, to form the variable w^*:

$$w^* = w + \epsilon. \tag{4}$$

There is however no reason in general to expect there to be a zero probability that the right-hand side of (4) is negative.[4]

The best-known method of treating the possibility that expenditures may be zero is to assume that w^* is censored at zero. The observed w^{**} is related to w^* by:

$$w^{**} = \max(0, w^*). \tag{5}$$

In effect the density is 'piled up' at zero. With the additional assumption that ϵ is normally distributed, this gives the Tobit model. We have then a probability $\Phi(w/\sigma)$ of observing a positive expenditure share, where Φ denotes the cumulative distribution function of the unit normal, and a probability $1-\Phi(w/\sigma)$ of observing zero expenditure.

The Tobit model is restrictive. In particular, the probability of zero

[4] If the expenditure on alcohol deviates from that predicted this may imply that total expenditure may deviate in the same direction and the potential positive correlation may cause a problem of consistency (see Keen, 1986). However, it is possible that higher expenditure on one good is accompanied by lower expenditure on another so that the overall effect on expenditure from errors for many goods may be small. This is the assumption adopted in this paper. The problem is considered explicitly in the system treatment of Blundell *et al.* (1988).

consumption is bound tightly to the determination of the amount consumed by those who have a positive share. For example, it is conceivable that, among those households that consume alcohol, the share is an increasing function of the price (where demand is relatively inelastic), so that p enters positively in w^*, but that the probability of drinking at all is reduced by a price rise. For this reason we investigated two alternatives to the Tobit model.

The first is the 'double hurdle' model which performed satisfactorily in our investigation of tobacco consumption (Atkinson *et al.* 1984*a* and 1990). This was investigated by Deaton and Irish (1984), in a simple p-Tobit form, who found it to be problematic. Further elaborations in our research did not prove successful.

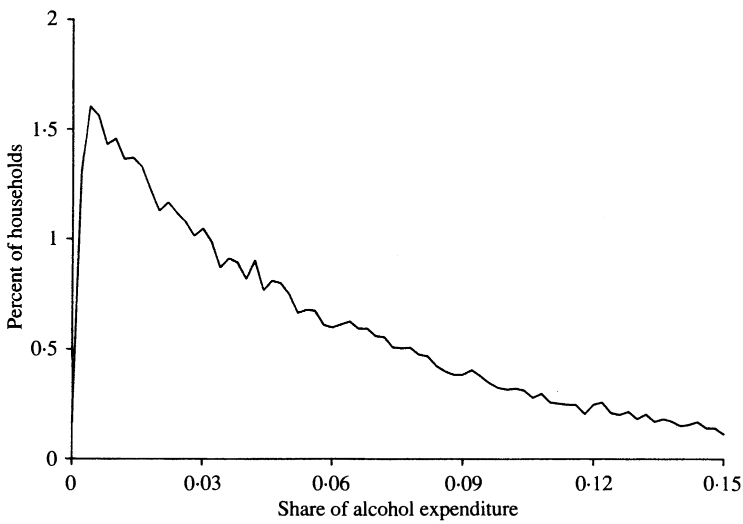

Fig. 1. Frequency distribution for share of alcohol expenditure.

The second variation on the Tobit model retains the one-stage model, but allows more flexibility in the distribution of ϵ. The distribution of observed shares in Fig. 1 certainly suggests that the distributional assumption may need to be re-considered. It is conceivable that the skewed distribution of shares might result from the distribution of explanatory variables, but it may also be that the assumption of normality for ϵ is inappropriate. We have used here the gamma distribution, where the density is:

$$
\begin{aligned}
f(\epsilon) &= A\,[\exp(-\epsilon/\kappa\sigma)]\,(\kappa\epsilon/\sigma+1)^{(1/\kappa^2)-1} && \text{if} \quad \kappa\epsilon/\sigma+1>0 \\
&= 0 && \text{otherwise} \qquad (6)
\end{aligned}
$$

where A is a constant. The mean is zero, the variance is σ^2 and the skewness[5] is given by 2κ.

[5] This is obtained from the standard form of the gamma distribution by a linear transformation designed to obtain a variate with mean 0 and variance σ^2. The 'shape index' α of the standard form (see Johnson and Kotz, 1970) is equal to $1/\kappa^2$. For further details of the gamma-Tobit model and its estimation, see Gomulka (1986).

The gamma distribution allows a variety of shapes for the density function as κ varies for fixed σ with the normal distribution represented by the special case $\kappa = 0$. Thus the Tobit is nested in the gamma-Tobit and we have a straightforward test of the Tobit model. In Fig. 2 we sketch the density function for $0 < \kappa < 1$, which provides a distribution which is positively skewed with a strong weight in the upper tail and thus may be appropriate for the case under consideration (the estimated value of κ falls in this range).

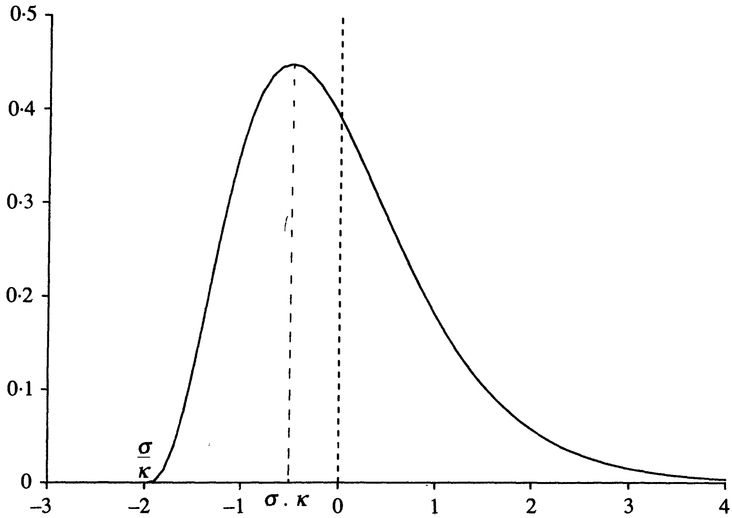

Fig. 2. Gamma frequency distribution for kappa = 0·5.

II. ESTIMATES OF THE MODEL

The data are from the *Family Expenditure Surveys* and are described, and variables defined, in the Appendix. Our sample is drawn from the surveys covering the period April 1970 to December 1983. After excluding those households where the head is retired and unoccupied (and a small number of peculiar cases – see the Appendix) we have 68,854 households.

We begin with the Tobit model for two reasons. First, it provides a well-known benchmark against which to assess the performance of the gamma-Tobit model developed for this paper. Second, it provides a relatively simple laboratory within which to explore the treatment of household characteristics. Maximum likelihood estimation of the gamma-Tobit with nearly 70,000 observations imposes a substantial computational burden – even with a CRAY computer. Accordingly, detailed explorations of the model were conducted using the Tobit and (occasionally) dropping the regional variables, which generally did not affect the behaviour of other variables.

In considering model selection it should be noted that t- and χ^2-values may be expected to increase with sample size. The Schwartz (1978) criterion for the F-test suggests a critical value of $\log N$ where N is the number of observations, correspondingly $q \log N$ for χ^2 where there are q degrees of freedom, or $\sqrt{\log N}$

for the t-test where there is just one degree of freedom: in this case log N is 11·14 and $\sqrt{\log N}$ is 3·34. We accordingly take in general 3·34 for our critical t-value and 11·14 g for χ^2 values.

These criteria led us to the Tobit equation presented in the first column of Table 1. This has two interaction terms with price (including that with time on account of its intrinsic interest) and six interaction terms with income. The implications of the estimated coefficients are discussed in Section IV, but we may note that the squared expenditure term remains highly significant even with the interaction terms. The coefficients on log (m/π) and its square indicate that the share of spending on alcohol first rises with total spending and then falls.

Table 1

Estimated Coefficients for the Tobit and the Gamma-Tobit Models

		Gamma-Tobit	
	Tobit	'Comparable'	'Preferred'
constant	−0·103	−0·0986	−0·0995
	(0·011)	(0·0065)	(0·0063)
log (m/π)	0·104	0·0787	0·0793
	(0·006)	(0·0034)	(0·0033)
$[\log (m/\pi)]^2$	−0·0187	−0·0107	−0·0107
	(0·0008)	(0·0005)	(0·0005)
log (p/π)	−0·0601	−0·00909	−0·00757
	(0·0099)	(0·00465)	(0·00222)
TOBACCO	0·0194	0·00858	0·00857
	(0·0006)	(0·00035)	(0·00035)
TIME	0·000980	0·000310	0·000267
	(0·000155)	(0·000069)	(0·000032)
DQRT2	0·00346	0·00152	0·00153
	(0·00072)	(0·00037)	(0·00037)
DQRT3	0·00582	0·00295	0·00291
	(0·00072)	(0·00035)	(0·00035)
DQRT4	0·00978	0·00353	0·00353
	(0·00072)	(0·00034)	(0·00034)
AGE	−0·0136	−0·00892	−0·00891
	(0·0014)	(0·00078)	(0·00075)
OCC1	−0·0263	−0·00653	−0·00641
	(0·0042)	(0·00232)	(0·00233)
OCC2	−0·00529	−0·00145	−0·00145
	(0·00089)	(0·00043)	(0·00043)
FORCES	−0·00648	−0·00318	—
	(0·00286)	(0·00144)	—
DUNP	−0·0368	−0·00331	—
	(0·0069)	(0·00400)	—
DOEARNS	0·00478	0·00389	0·00390
	(0·00060)	(0·00032)	(0·00032)
OWN	−0·0134	−0·00243	−0·00243
	(0·0006)	(0·00026)	(0·00025)
NMEN	0·0551	0·0203	0·0203
	(0·0034)	(0·0011)	(0·0011)

Table 1 (*cont.*)

	Tobit	Gamma-Tobit 'Comparable'	Gamma-Tobit 'Preferred'
NOTMEN	−0·0919	−0·0213	−0·0214
	(0·0032)	(0·0015)	(0·0015)
NCHLDRN	−0·0336	−0·00809	−0·00809
	(0·0019)	(0·00090)	(0·00087)
TIME × log (p/π)	0·00352	0·000397	—
	(0·00142)	(0·000610)	—
NCHLDRN × log (p/π)	0·0153	−0·000968	—
	(0·0036)	(0·00196)	—
AGE × log (m/π)	0·00257	0·00167	0·00167
	(0·00041)	(0·00021)	(0·00021)
OCC1 × log (m/π)	0·00545	0·00143	0·00140
	(0·00116)	(0·00062)	(0·00062)
DUNP × log (m/π)	0·00970	0·000770	—
	(0·00217)	(0·00118)	—
NMEN × log (m/π)	−0·00736	−0·00386	−0·00385
	(0·00093)	(0·00030)	(0·00030)
NOTMEN × log (m/π)	0·0220	0·00539	0·00542
	(0·0009)	(0·00038)	(0·00037)
NCHLDRN × log (m/π)	0·00773	0·00142	0·00145
	(0·00051)	(0·00023)	(0·00023)
Northern	0·0198	0·0105	0·0105
	(0·0012)	(0·0005)	(0·0005)
Yorks/Humberside	0·0150	0·0080	0·0080
	(0·0011)	(0·0005)	(0·0005)
East Midlands	0·0098	0·0068	0·0068
	(0·0012)	(0·0006)	(0·0006)
East Anglia	−0·0043	0·0010	0·0011
	(0·0015)	(0·0006)	(0·0006)
South East	−0·0016	0·0029	0·0029
	(0·0009)	(0·0005)	(0·0005)
South West	0·0009	0·0032	0·0032
	(0·0012)	(0·0005)	(0·0005)
Wales	0·0099	0·0059	0·0059
	(0·0014)	(0·0006)	(0·0006)
West Midlands	0·0101	0·0079	0·0078
	(0·0011)	(0·0005)	(0·0005)
North Western	0·0150	0·0081	0·0081
	(0·0010)	(0·0004)	(0·0004)
Scotland	0·0103	0·0055	0·0055
	(0·0011)	(0·0005)	(0·0005)
Northern Ireland	−0·0127	−0·0109	−0·0110
	(0·0021)	(0·0015)	(0·0015)
σ	0·0635	0·0565	0·0564
	(0·0002)	(0·0002)	(0·0002)
κ	—	0·788	0·787
		(0·006)	(0·006)
log likelihood	114,330·8	121,816·7	121,813·3

Note: The coefficients of dummy variables are given to 4 decimal points (a movement in the last place corresponding to a rise from, say, 5 % to 5·01 %); the coefficients of other variables, including the interaction terms, are given to 3 significant figures.

The Gamma-Tobit

Our investigation of the double-hurdle model pointed to the limiting case where only the second hurdle is relevant and hence to the Tobit model (Deaton and Irish, 1984, encountered similar difficulties in their estimation of the p-Tobit). The 'gamma-Tobit', on the other hand, provides a strikingly better fit to the data than the Tobit. The results for total spending on alcohol are shown in the second column of Table 1. The model in columns 1 and 2 is the same, apart from the extra skewness parameter κ which enters the gamma distribution. Hence the models are nested and the comparison of (twice the log) likelihoods is distributed as chi-square with one degree of freedom. We see that the extension from the normal to the gamma distribution provides a chi-square of 14,972, which is a huge increase and the Tobit would be resoundingly rejected against the gamma-Tobit. The value of σ falls from 0·0635 for Tobit to 0·0565 for the gamma-Tobit, which may be compared with a fall from 0·0694 to 0·0635 on the introduction of all the explanatory variables into the Tobit.

The skewness parameter κ is highly significant and takes a value of 0·788, which is 131 standard errors from the value of 0 which gives the normal distribution. From Fig. 2, we see that the corresponding random variable has a zero probability of values below $-\sigma/\kappa$, which has an estimated value of $-0{\cdot}072$. Thus if the measured variables predict a share of $\hat{w}$ then the right-hand side of the regression equation has zero probability of falling below $\hat{w}-0{\cdot}072$. The mode is at $-\kappa\sigma$, which is equal to $-0{\cdot}045$. Notice that a number of coefficients which are significant in the Tobit are no longer significant in the gamma-Tobit. We present a restricted gamma-Tobit in the third column of Table 1 and take it as a basis for further discussion.

To investigate stability we split the sample into two periods of approximately equal length (1970–6 and 1977–83) and estimated the gamma-Tobit separately for the two periods. A comparison of the sum of the likelihoods for the split samples and the likelihood for the pooled sample allows a test of the hypothesis that all coefficients are unchanged across the two periods. This is not rejected – twice the difference in the log-likelihoods is equal to 186·6 and the relevant critical value ($q \log N$, see above discussion based on the Schwartz criterion) is 389·9, whereas the conventional size of the chi-square statistic the hypothesis of no change in the model would be rejected. Further, if we look at individual coefficients we see that around two-thirds of those for 1970–6 lie within two standard errors of those for 1977–83. The coefficients involving price appear more unstable but this might be due to the fact that the variation of price is closely associated with time and is importantly reduced when we take shorter periods.[6]

[6] As a check on specification we calculated 'robust' standard errors using the covariance matrix defined by White (1982, p. 5). They were quite close to the classical standard errors reported here, with the differences being even smaller for gamma-Tobit than Tobit. This proximity provides some reassurance on model specification and we take it as an indication that heteroscedasticity problems are not serious.

III. INTERPRETATION OF THE RESULTS

A central reason for estimating the model is to help understand the effects of a change in a variable of interest, say expenditure or price, on the alcohol share. To examine this more closely we write (1) as

$$w = \mathbf{k}.\mathbf{X} \tag{1a}$$

where $\mathbf{X}$ is the vector of explanatory variables in (1) and $\mathbf{k}$ the vector of coefficients. Then (4) may be written

$$w^* = \mathbf{k}.\mathbf{X} + \epsilon \tag{4a}$$

Where we have variables entering non-linearly into $\mathbf{X}$ and also censoring, estimated responses are dependent on the levels of the variables and on ϵ, and there will be variation in response across households associated with both elements. This has the important implication that we have to consider the interpretation to be given to the stochastic term. This term may arise for several reasons, and they have different implications. Here we consider only two, but they serve to illustrate the possibilities. The first is that there is an individual household *fixed effect*, equal to the difference between the observed and predicted share. In this case, the effect of a change in X_i is calculated conditional on the vector $\mathbf{X}$ and the specific value of ϵ_h observed for household h. The second case is where there is a *random transitory component* and the relevant variable for the household is the mean spending conditional on $\mathbf{X}$. In other words, we form the average over possible outcomes for ϵ.

These considerations assume particular significance in the present context, since the censoring at zero involves a nonlinearity in an essential manner. The fact that the observed w^{**} is given by (5) means that the effect of variations in X_i depends on the interpretation of the stochastic term – even without any of the X_i entering non-linearly. With a *fixed effect* interpretation, then for a household with a strictly positive share (referred to later as a 'drinking household') the impact on the share of marginal variations in X_i is given simply by k_i, and the income and price elasticities are calculated using the actual values of the spending share. For households with a zero share, the value of ϵ cannot be calculated (only an upper bound), but marginal changes in X_i have no effect, and the price and income elasticities are zero.

With a *random transitory component* interpretation, we calculate the mean conditional on $\mathbf{X}$. If there are no values of ϵ such that w^* is negative, then the mean is $\mathbf{k}.\mathbf{X}$ (since the mean of ϵ is zero). On the other hand, suppose that there are values of ϵ such that w^* is negative. The condition for this in the case of the gamma-Tobit is that $\mathbf{k}.\mathbf{X} < \sigma/\kappa$. Then the mean conditional on $\mathbf{X}$ is given by

$$\tilde{w}^{**} = \mathrm{E}(w^{**}\,|\,\mathbf{X}) = \int_{-\mathbf{k}.\mathbf{X}}^{\infty} (\mathbf{k}.\mathbf{X} + \epsilon) f(\epsilon)\, d\epsilon \tag{7}$$

where f is the probability density function of the random term ϵ and hence

$$\frac{\partial \tilde{w}^{**}}{\partial X_i} = k_i\left[1 - F\left(-\frac{\mathbf{k}.\mathbf{X}}{\sigma}\right)\right] \tag{8}$$

where F is the distribution function for ϵ normalised to have unit variance. For the Tobit, F is the standard normal and for the gamma-Tobit we have the distribution function for the gamma distribution.

The term in the square bracket in (8) is the probability that the share is positive. In other words, the derivative of the mean share (conditional on $\mathbf{X}$) with respect to an explanatory variable is the coefficient multiplied by the probability that the observed share is positive. Notice however that the k_i are scaled by the same factor in (8) so that k_i/k_j still correctly measures the relative impact on the share of marginal changes in X_i and X_j. In order to calculate the price elasticity (the elasticity of the mean quantity conditional on $\mathbf{X}$), we make use of (3) to obtain

$$1-\gamma[1-F(-\mathbf{k}.\mathbf{X}/\sigma)]/\tilde{w}^{**} \tag{8a}$$

To calculate the expenditure elasticity, we make use of (2) to obtain

$$1+[1-F(-\mathbf{k}.\mathbf{X}/\sigma)]\,[\beta+2\delta\log(m/\pi)]/\tilde{w}^{**} \tag{8b}$$

If in the case of gamma-Tobit $\mathbf{k}.\mathbf{X} > \sigma/\kappa$ then $-\mathbf{k}.\mathbf{X}/\sigma$ lies to the left of the point where the density becomes positive (see (6) and Fig. 2), therefore $F(-\mathbf{k}.\mathbf{X}/\sigma) = 0$ and the first term in square brackets, the probability that the share is positive, is equal to 1. There is then no censoring and the expressions (7)–(8b) are the same as for the fixed effect interpretation.

The interpretation of the estimated equations is not therefore as straightforward as may appear at first sight. Even in the absence of non-linearities in the model explaining w^*, the coefficients in the tables of estimates provide direct answers as to the effect of the different variables only in the case of the fixed effect interpretation of the stochastic term or for a household not subject to censoring. It is not therefore sufficient simply to present the estimated coefficients, leaving the remainder of the task of interpretation to be carried out by the reader.

We can now discuss in more detail the results obtained from the estimated model, concentrating on our preferred gamma-Tobit equation, as represented by the third column of Table 1, distinguishing the two different approaches to the stochastic term just indicated.

Fixed Effect Interpretation

From our preferred equation in the third column of Table 2, we can see that the coefficients on both the interaction terms and the square of log expenditure are highly significant, so that the response of the share of spending to an increase in income varies across observations, depending on the vector $\mathbf{X}$. The range of variation of β in our sample is in fact from 0·061 to 0·124. The same variation across the sample applies to the expenditure elasticity, given by equation (2), which further depends (via w) on the value of ϵ, as does the price elasticity (see equation (3)). At this point, we adopt the fixed effect interpretation, where the share w is taken as that observed. The elasticity is calculated for households with positive shares, referred to as 'drinking households'.

The distribution of the expenditure elasticities for drinking households is given in Fig. 3 (see also Table A. 2 (*a*)). There is a concentration in the range above 1: around 55% of drinkers have an expenditure elasticity between 1·0 and 1·4 and a further 20% between 1·4 and 2·0. On the other hand, for about 10% the calculated elasticity is greater than 3. The interaction with characteristics contributes to this variation within the sample although, as we warned in Section I, a major role is played by the functional form.

The price elasticity for the uncensored share is given by $1-\gamma/w$. With the estimated coefficient on log price of $-0{\cdot}00757$ the elasticity is a declining function of the share, lighter drinkers having a higher price elasticity. With a share of 2%, the elasticity is 1·38, whereas with a share of 5% the elasticity is 1·15. The distribution of the price elasticity for drinking households is shown in Fig. 4 (see also Table A. 3 (*a*)). Over 70% of drinkers in the sample have an elasticity between 1·0 and 1·3, although there are some 10% with a calculated elasticity greater than 2·0. Given that we do not find significant price interaction effects in the gamma-Tobit model, the observed distribution of price elasticities is generated solely by the relationship with the share, and hence with expenditure. It must be recognised that this is essentially a result of our choice of functional form.

The compensated elasticity is equal to the uncompensated elasticity (both elasticities defined to be positive where quantity falls with price) minus the marginal propensity to consume. The first of these is greater than unity (γ is negative and see (3)), and we have only to check that the marginal propensity is less than unity to ensure negativity of the compensated own-price effect. This is indeed satisfied everywhere in the sample.

Random Transitory Component Interpretation

We now consider the interpretation of the stochastic term as arising from random transitory components, with the variable of interest being the mean share conditional on $\mathbf{X}$: the effects of changes in explanatory variables on the expected share conditional on $\mathbf{X}$ are calculated using (7)–(8*b*). The second frequency distributions in Figs. 3 and 4 show the expenditure and price elasticities (respectively) under the 'random transitory' interpretation (see also Tables A. 2 (*b*) and A. 3 (*b*)). Around 80% of the expenditure elasticities fall between 1·1 and 1·5 so that the distribution is much more tightly bunched than for the fixed effect interpretation. The price elasticities in Fig. 4 similarly have a much smaller range of variation in the 'random transitory' case than in the fixed effect case, with more than 99% falling between 1·11 to 1·15. These are even more tightly bunched than the expenditure elasticities since (i) there appears to be little variation in γ across the sample, in contrast to β, (ii) for the expenditure elasticity we have a term involving $\delta \log(m/\pi)$ and (iii) β is generally larger than γ so has a bigger multiplicative effect on the reciprocal of the share.

The expressions for the expenditure and price elasticities (8*a*) and (8*b*) indicate why the distribution is more tightly bunched with the random transitory component interpretation. In the case of the price elasticity, for example, the probability of a positive value $[1-F(-\mathbf{k}.\mathbf{X}/\sigma)]$ rises with $\mathbf{k}.\mathbf{X}$

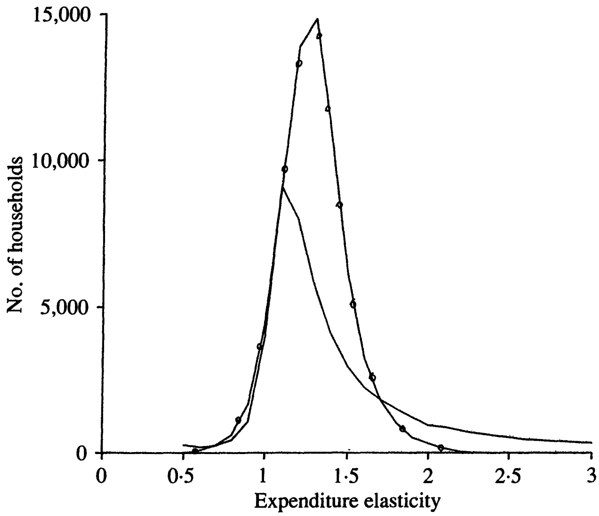

Fig. 3. Frequencies of expenditure elasticities. ———, Fixed effects; —⊖—, random effects.

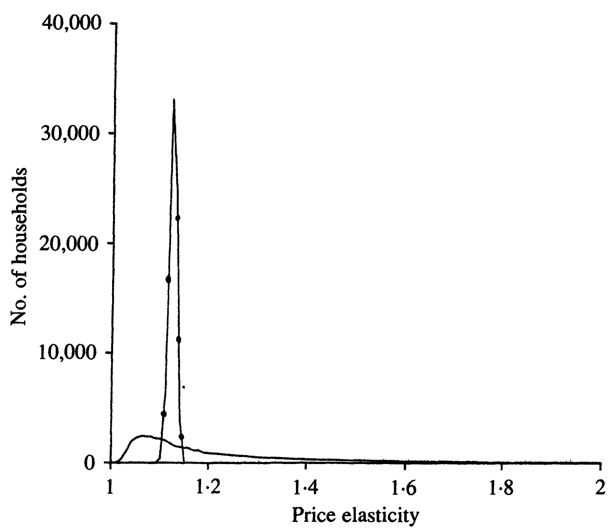

Fig. 4. Frequencies of price elasticities. ———, Fixed effects; —⊖—, random effects.

and so tends to moderate the fall in the elasticity arising from the rise in $\hat{w}^{**}$ (remember that γ is negative).

Estimated Coefficients in Preferred Version of Gamma-Tobit

The coefficients of income and price have been discussed above. The coefficients for *TIME* and the quarter dummies suggest that their influence is quite small. The regions all show a higher share than the base case, London, with the exception of Northern Ireland. The highest positive effect is for the North, where the estimated coefficient indicates that the share is a whole percentage point higher, whereas Northern Ireland is more than a percentage point lower (under the fixed effect interpretation). Of the household composition variables, a positive effect is associated with the number of men, *NMEN*. The addition of an extra woman, on the other hand, reduces the share, as does the presence of children although to a lesser degree. The effect of age is also relevant. A household with the mean expenditure headed by a person aged 20 has a share which is around one percentage point higher than a similar household headed by someone aged 50. It should be noted that the coefficients of the interaction terms have in all cases the reverse sign from that of the characteristic variable, indicating that the differences decline with the level of household spending.

Estimates for the Population as a Whole

The first reason why our results cannot be directly extrapolated to the population as a whole is that we have, by design, excluded households headed by a person who is retired or unoccupied. The households covered are only some 70% of the total responding to the FES. Such restriction of the sample is common: for example, Blundell *et al.* (1988) exclude households headed by a person aged 60 or over (or under 18) or by someone who is self-employed. The reason for restricting the sample is typically to ensure a more homogeneous group, but this means that the findings cannot be assumed to apply to those excluded from the analysis. For the 30% of households not covered, the price responses may be quite different.

The second reason is that there is significant non-response to the FES, and the survey does not cover the non-household population (such as those living in army barracks or in hotels). As we see in the Data Appendix, there is some reason to suppose that those omitted from the sample are, on average, heavier drinkers than those who are included. If this meant simply that they have a large constant term in (1) then the analysis of the implications for the aggregate share would be unchanged, but if the larger share arises multiplicatively, then simple grossing-up of the sample would underestimate the effects of changing explanatory variables. This is not something we can pursue in the absence of further information about the consumption responses of those not responding to the FES; we have, for example, no way of knowing whether they differ in, say, their price or expenditure elasticities of demand.

IV. CONCLUSIONS

We have investigated a model of the determination of the share of household expenditure on alcohol using FES data for the United Kingdom 1970–83, paying particular attention to the introduction of variations in characteristics across the population, and the interpretation of the stochastic term. Our concern has been as much with the approach to modelling and interpretation as with the substantive results, but we have reached the following conclusions:

(*a*) The Gamma-Tobit extension provides a more satisfactory fit to the data than the standard Tobit model. The estimate of κ indicates positive skewness of the distribution of the stochastic term and is highly significant, and reflects the long upper tail in the data. This appears consistent with what we know about the upper tail in the drinking distribution. Comparison with our experience with modelling tobacco spending suggests that zero expenditures may need different treatment for different items.

(*b*) There are significant household composition effects on the share of spending on alcohol, with an additional man increasing the share, and additional women and children reducing the share. Together with a decline in the share with age, these generate a 'life-cycle' of alcohol consumption.

(*c*) The household characteristics, and occupation, appear to interact with expenditure; the differences between the predicted shares for different household types tending to be smaller at higher expenditure levels.

(*d*) There are significant regional differences, with the coefficient for the North of England being the highest and for Northern Ireland the lowest.

(*e*) There is a small but significant upward trend in the alcohol share over time.

(*f*) The calculation of expenditure and price elasticities depends on the interpretation given to the stochastic term. (It was also emphasised that their variation with the share of spending was heavily dependent on our choice of functional form.) If it is treated as a fixed household effect, then the distributions of expenditure and price elasticities are quite spread. If the stochastic term represents random transitory variation, and attention is focused on mean demand conditional on the explanatory variables, then the distribution of expenditure elasticities is concentrated on the range 1–1·5 and the distribution of price elasticities is tightly concentrated around 1·12.

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Data Appendix

The FES is a survey of a random sample of private households in the United Kingdom, which has been in continuous operation since 1957. The effective annual sample is around 11,000, of which about 70% agree to cooperate fully. As part of the survey, each spender in the household maintains a detailed record of expenditure during 14 consecutive days. These 'diary' records provide the basis for the expenditure data, and the alcoholic drink item includes beer, shandy, cider, wines, and spirits.

Taking the FES households as a whole, the share of alcohol in total expenditure is around $4\frac{1}{2}$–5%. This falls quite a long way short of the proportion indicated by the national accounts ($7\frac{1}{2}$%), and this discrepancy has been a matter for concern. After careful study the view of the OPCS (Kemsley *et al.* 1980, p. 53) is that the major reason for the discrepancy lies in the omission of a minority of heavy drinkers from the survey. If this is correct, then to the extent that the reasons for exclusion are not such as to cause sample selection bias, our analysis of individual household expenditure by those responding to the FES is not affected. For further discussion of some of these issues, see Atkinson *et al.* (1989*a*). Note, however, that the behaviour of heavy drinkers may be rather different from the rest of the population. Accordingly we focus our attention in the interpretation of results in Section III on distributions of responses within the sample rather than on aggregate expenditure.

In this paper we are concerned with the expenditure of households where the head is neither retired nor unoccupied and where there is strictly positive total household expenditure (i.e. ignoring the small number of households which report zero spending for all goods). This selection is intended to ensure a more homogenous group for analysis. We also excluded at the outset the top and bottom 0·05% in the distribution of total household expenditure on all items, in order to avoid giving undue weight to peculiar observations (69 cases in all). The data are drawn from the surveys covering the period April 1970–December 1983. This yielded a total sample of 68,854 households, or 71·1% of all households covered by the FES in that period. For 19·3% of the sample, the recorded expenditure on alcohol is zero. The share for those with strictly positive expenditure has a mean of 6·1%, and a median of 4·4%, indicating positive skewness.

Table A. 1

Means and standard deviations of the variables

	Mean	S.D.	Minimum	Maximum
Share of alcohol				
All households	0·0496	0·0590	—	0·742
Drinking households	0·0615	0·0598	—	—
Income and price				
$\log (m/\pi)$	3·408	0·537	1·344	5·436
$[\log (m/\pi)]^2$	11·902	3·656	1·807	29·550
$\log (p/\pi)$	−0·129	0·057	−0·228	0·000
TOBACCO	0·667	0·471	—	—
TIME	6·810	3·947	0·021	13·729
DQRT 2	0·247	—	—	—
DQRT 3	0·258	—	—	—
DQRT 4	0·257	—	—	—
Characteristics				
AGE	4·303	1·298	1·6	9·5
OCC 1	0·289	—	—	—
OCC 2	0·100	—	—	—
FORCES	0·008	—	—	—
DUNP	0·050	—	—	—
DOEARNS	0·582	—	—	—
OWN	0·556	—	—	—
Household composition				
NMEN	1·040	0·503	0	6
NOTMEN	1·149	0·575	0	6
NCHLDRN	0·967	1·188	0	10
Regions				
Northern	0·061	—	—	—
Yorks/Humberside	0·090	—	—	—
East Midlands	0·068	—	—	—
East Anglia	0·036	—	—	—
South East	0·184	—	—	—
South West	0·070	—	—	—
Wales	0·048	—	—	—
West Midlands	0·097	—	—	—
North Western	0·114	—	—	—
Scotland	0·092	—	—	—
Northern Ireland	0·017	—	—	—
Interaction terms				
$TIME \times \log (p/\pi)$	−0·979	0·611	−2·024	0
$NCHLDRN \times \log (p/\pi)$	−0·124	0·173	−1·911	0
$AGE \times \log (m/\pi)$	14·582	4·693	2·363	43·973
$OCC\,1 \times \log (m/\pi)$	1·045	1·664	0	5·436
$DUNP \times \log (m/\pi)$	0·153	0·676	0	5·251
$NMEN \times \log (m/\pi)$	3·650	2·063	0	28·448
$NOTMEN \times \log (m/\pi)$	4·015	2·344	0	27·078
$NCHLDRN \times \log (m/\pi)$	3·400	4·249	0	43·957

The variables included from the FES are:

Expenditure and Price. The effect of expenditure is represented by total expenditure[7] deflated by the retail price index (all items), denoted by (m/π). The effect of price is represented by the alcohol component of the retail price index, expressed relative to that for all items, denoted by (p/π). The monthly figures were obtained from the *Department of Employment Gazette* and interpolated to the week in question.

Region. There is evidence from the national drinking surveys that average consumption differs by region: in particular that of men in Northern Ireland was about 28% less than in the rest of the United Kingdom (Wilson, 1981, Table 2). We include dummy variables for each of the twelve standard regions, the excluded region being Greater London.

Age. The national drinking surveys show a declining consumption with age (Wilson, 1981, Table 3). The variable used is the age of the head of household in years (divided by 10).

Occupation. There are a number of reasons to expect alcohol consumption to be related to the nature of employment, and we represent this by the occupation of the head of household:

OCC 1 Professional, technical, administrative, managerial and teachers

OCC 2 Clerical workers: e.g. clerks, commercial travellers, agents and shop assistants

FORCES Members of HM Forces.

These groups account for around 40% of the sample. The remainder are manual workers.

Employment Situation. It is sometimes suggested that the unemployed may have a different level of consumption, either, in one direction, because they have more leisure, or, in the other direction, because they lack the network of friends at work. Also, the employment status of other household members may be relevant. We introduce dummy variables *DUNP* where the head of household is unemployed and *DOEARNS* where there are 2 or more earners (employed or self-employed) in the household.

Household Structure. The level of consumption is likely to be affected by household size and there may also be a specific effect on alcohol of the presence of certain types of person. We therefore include separate variables for the number of people in the household of particular types:

NMEN, defined as the number of males 18 or over or under 18 and married,

NOTMEN, the number of other household members aged 16 and over, and

NCHLDRN, the number of children.

Owner-Occupation. In our earlier analysis of a commodity demand system (Atkinson and Stern, 1980), we found that alcohol expenditure was negatively related to owner-occupation. The interpretation of this variable ($OWN = 1, 0$) is open to debate, but may be seen as an expression of tastes.

Tobacco. Alcohol and tobacco consumption may be linked. We included a variable (*TOBACCO*) which is 1 if household expenditure on tobacco is positive and zero otherwise. The price of tobacco proved insignificant.

The means and the standard deviations of the explanatory variables used are given in Table A.1. There is in addition a *TIME* variable, measured in years from the second week of April 1970, and a dummy variable for each of the three quarters of the year apart from the first.

[7] Total household expenditure is taken from the FES, including 'non-diary' items, but excludes the imputed rent on owner-occupier or rent-free dwellings, replacing this by the actual cash outgoings. This exclusion is motivated by the limitations to the measurement of imputed rent but is not entirely satisfactory. Any owner-occupied household with no mortgage, for example, would have an understated expenditure and an overstated share on alcohol. This would suggest a positive coefficient on the owner-occupation zero-one variable which is included. We take total expenditure rather than disposable income for a number of reasons, including the fact that expenditure relates to the fourteen days commencing with the interview, whereas the income data relate to the previous pay period or earlier (investment income, for example, covers the previous 12 months). Total expenditure is defined as the FES code P378 less the imputed rent and plus the actual mortgage payments (for all owner-occupiers).

Table A.2

Distribution of Expenditure Elasticities

(a) Fixed Effect Interpretation: Drinking Households

Expenditure elasticity	Frequency	%	Cumulative frequency	Cumulative %
< 0	429	0·8	429	0·8
0	112	0·2	541	1·0
0·5	276	0·5	817	1·5
0·6	191	0·3	1,008	1·8
0·7	262	0·5	1,270	2·3
0·8	439	0·8	1,709	3·1
0·9	1,076	1·9	2,785	5·0
1	4,043	7·3	6,828	12·3
1·1	9,101	16·4	15,929	28·7
1·2	7,985	14·4	23,914	43·0
1·3	5,650	10·2	29,564	53·2
1·4	4,034	7·3	33,598	60·5
1·5	2,949	5·3	36,547	65·8
1·6	2,241	4·0	38,788	69·8
1·7	1,825	3·3	40,613	73·1
1·8	1,497	2·7	42,110	75·8
1·9	1,221	2·2	43,331	78·0
2	939	1·7	44,270	79·7
2·1	883	1·6	45,153	81·3
2·2	775	1·4	45,928	82·7
2·3	670	1·2	46,598	83·9
2·4	600	1·1	47,198	85·0
2·5	1,376	2·5	48,574	87·4
3	1,727	3·1	50,301	90·5
3·5	1,061	1·9	51,362	92·5
4	1,036	1·9	52,398	94·3
5	925	1·7	53,323	96·0
6	586	1·1	53,909	97·0
7	358	0·6	54,267	97·7
8	281	0·5	54,548	98·2
9	219	0·4	54,767	98·6
10	83	0·1	54,850	98·7
> 10	705	1·3	55,555	100.0

Notes:

(i) The expenditure elasticity is calculated using equation (2), estimates from the third column in Table 1, and the observed values of variables including w^h. Note that β depends on *AGE*, *OCC*1, *NMEN*, *NOTMEN*, *NCHLDRN*.

(ii) The figures in percentage columns give numbers in the preceding column as a percentage of the total number of households with a positive spending share (which are 80·7% of the total).

Table A.2

(b) Random Transitory Component Interpretation: All Households

Expenditure elasticity	Frequency	%	Cumulative frequency	Cumulative %
0·2	1	0·0	1	0·0
0·3	1	0·0	2	0·0
0·4	1	0·0	3	0·0
0·5	15	0·0	18	0·0
0·6	89	0·1	107	0·2
0·7	261	0·4	368	0·5
0·8	617	0·9	985	1·4
0·9	1,673	2·4	2,658	3·9
1·0	4,438	6·4	7,096	10·3
1·1	8,986	13·1	16,082	23·4
1·2	13,890	20·2	29,972	43·5
1·3	14,866	21·6	44,838	65·1
1·4	10,712	15·6	55,550	80·7
1·5	6,095	8·9	61,645	89·5
1·6	3,260	4·7	64,905	94·3
1·7	1,819	2·6	66,724	96·9
1·8	1,029	1·5	67,753	98·4
1·9	539	0·8	68,292	99·2
2·0	329	0·5	68,621	99·7
2·1	162	0·2	68,783	99·9
2·2	56	0·1	68,839	100·0
2·3	14	0·0	68,853	100·0
2·4	1	0·0	68,854	100·0

Table A.3
Distribution of Price Elasticities
(a) Fixed Effect Interpretation: Drinking Households

Price elasticity	Frequency	%	Cumulative frequency	Cumulative %
1	4,558	8·2	4,558	8·2
1·1	20,215	36·4	24,773	44·6
1·2	9,642	17·4	34,415	61·9
1·3	5,224	9·4	39,639	71·4
1·4	3,159	5·7	42,798	77·0
1·5	2,317	4·2	45,115	81·2
1·6	1,562	2·8	46,677	84·0
1·7	1,201	2·2	47,878	86·2
1·8	1,004	1·8	48,882	88·0
1·9	741	1·3	49,623	89·3
2	610	1·1	50,233	90·4
2·1	522	0·9	50,755	91·4
2·2	465	0·8	51,220	92·2
2·3	397	0·7	51,617	92·9
2·4	339	0·6	51,956	93·5
2·5	741	1·3	52,697	94·9
3	897	1·6	53,594	96·5
3·5	502	0·9	54,096	97·4
4	376	0·7	54,472	98·1
4·5	225	0·4	54,697	98·5
5	177	0·3	54,874	98·8
5·5	138	0·2	55,012	99·0
6	114	0·2	55,126	99·2
6·5	68	0·1	55,194	99·4
7	68	0·1	55,262	99·5
7·5	49	0·1	55,311	99·6
8	35	0·1	55,346	99·6
8·5	31	0·1	55,377	99·7
9	24	0·0	55,401	99·7
9·5	21	0·0	55,422	99·8
10	8	0·0	55,430	99·8
> 10	125	0·2	55,555	100·0

Notes:
(i) The uncensored price elasticity is $1-\gamma/w^h$ (see equation (3)) where we use the observed value for w^h. We use $\gamma = -0{\cdot}00757$ from the third column in Table 1.
(ii) See note (ii) for Table A.2 (*a*).

(b) Random Transitory Effect Interpretation: All Households

Price elasticity	Frequency	%	Cumulative frequency	Cumulative %
1·09	25	0·0	25	0·0
1·10	397	0·6	422	0·6
1·11	6,274	9·1	6,696	9·7
1·12	33,101	48·1	39,797	57·8
1·13	24,798	36·0	64,595	93·8
1·14	4,079	5·9	68,674	99·7
1·15	180	0·3	68,854	100·0