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A MODEL OF THE DISAGGREGATED DEMAND FOR MEAT IN AUSTRALIA

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The focus of this study is the estimation of the Australian demand for meat between 1967 and 1990, employing a demand systems approach which uses the linear approximate, almost ideal demand system (LA/AIDS) model. Two demand systems are estimated by maximum likelihood methods, one for aggregate types of meat and one for disaggregated meat products. After correcting for serial correlation in the two demand systems, restrictions from utility theory are imposed and tested for their appropriateness. By using a new data set on the Australian retail price and consumption of fresh pork, ham and bacon, the results from the disaggregated model provide the first estimates of the own-price, cross-price and expenditure elasticities for these commodities.

Introduction

One of the oldest and most important uses of econometrics is its application to households in the estimation of demand relationships. Moreover, empirical research into the consumption behaviour of households is characterised by a close relationship between economic theory and appropriate estimation techniques.

A demand systems approach is used in this paper to analyse the Australian demand for meat between 1967 and 1990, using a linear approximate (LA) version of the Almost Ideal Demand System (AIDS) of Deaton and Muellbauer (1980, 1980a).[1,2] The demand system is

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[1] Parametric studies using a demand systems approach and Australian data have been carried out by Murray (1984), Chalfant and Alston (1986), Beggs (1988) and Alston and Chalfant (1991). A nonparametric study was also conducted by Chalfant and Alston (1988). Moschini and Mielke (1989), among others, have recently used the LA/AIDS model to estimate quarterly meat consumption in the United States.

[2] Some of the major surveys of the voluminous literature on demand systems estimation can be found in Brown and Deaton (1972), Powell (1974), Barten (1977), Johnson, Hassan and Green (1984), Deaton and Muellbauer (1980a) and Blundell (1988).

estimated using maximum likelihood methods, and a number of asymptotic tests are carried out on the validity of the imposition of restrictions derived from utility theory. In particular, tests are conducted for both homogeneity and symmetry of the demand functions, and for homotheticity of the utility function. Two versions of the system are estimated and tested — an aggregate model using the budget shares of major types of meat as dependent variables, and a disaggregated model in which the demand for particular cuts of meat is emphasised. Such a disaggregation enables a more precise analysis of the demand interrelationships between the various types of meat.

The structure of the LA/AIDS models to be estimated is specified in the next section 'Specification of the LA/AIDS Model'. Then, in the section 'Maximum Likelihood Estimation', the estimation method is outlined. A full description of the sources and types of data used is given below in 'Description and Sources of the Data'. In the section headed 'Statistical Inference', the procedures for statistical inference are presented. Then the estimation results and important elasticity estimates are given, followed by concluding comments.

Specification of the LA/AIDS Model

Detailed derivations of the AIDS model are available in Deaton and Muellbauer (1980, 1980a). These authors demonstrate that the general form of the AIDS expressions for budget shares is:

$$(1) \qquad W_i = \alpha_i + \sum_j \gamma_{ij} \log p_j + \beta_i \log \left(\frac{x}{P}\right)$$

where W_i is the budget share of the ith good, x is total expenditure on the group of goods being analysed, p_j is the price of the jth good in the group, and P is a price index defined by:

$$(2) \qquad \log P = \alpha_0 + \sum_i \alpha_i \log p_i + \frac{1}{2} \sum_i \sum_j \gamma_{ij} \log p_i \log p_j .$$

Rather than use this inherently nonlinear price index, in this analysis and in most demand studies the linear approximate (LA/AIDS) form has been used, with P being approximated by Stone's (1953) geometric price index ($\log P = \Sigma_i W_i \log p_i$). In general, the results obtained with the AIDS model have been found to differ only slightly from those obtained when the LA/AIDS version is estimated (Deaton and Muellbauer 1980, Anderson and Blundell 1983).[3] In equation (1) above the ith budget share (W_i) is expressed in terms of prices and real expenditure on meat. The α_i intercept represents the budget share when all logarithmic prices and expenditure are zero; $\gamma_{ij} = \partial W_i / \partial \log p_j$; and $\beta_i = \partial W_i / \partial \log (x/P)$.

[3] However, by using Stone's index to obtain the advantages of linearity, other advantages are foregone, such as integrability.

The following set of restrictions are derived from utility theory and often imposed upon the LA/AIDS parameters:

$$(3) \qquad \sum_i \alpha_i = 1,\ \sum_i \beta_i = \sum_i \gamma_{ij} = 0 \qquad \text{adding-up}$$

$$(4) \qquad \sum_j \gamma_{ij} = 0 \qquad \text{homogeneity}$$

$$(5) \qquad \gamma_{ij} = \gamma_{ji} \qquad \text{Slutsky symmetry.}$$

The expenditure and price elasticities for the model are given by:

$$(6) \qquad \eta_i = 1 + \frac{\beta_i}{W_i}$$

$$(7) \qquad \varepsilon_{ii} = -1 + \frac{\gamma_{ii}}{W_i} - \beta_i$$

$$(8) \qquad \delta_{ii} = -1 + \frac{\gamma_{ii}}{W_i} + W_i$$

$$(9) \qquad \varepsilon_{ij} = \frac{\gamma_{ij}}{W_i} - \beta_i \left(\frac{W_j}{W_i}\right)$$

$$(10) \qquad \delta_{ij} = \frac{\gamma_{ij}}{W_i} + W_j$$

where η denotes the expenditure elasticities, ε denotes the Marshallian (uncompensated) price elasticities and δ denotes the Hicksian (compensated) price elasticities.[4] Finally, the Allen elasticities of substitution (σ) are given by

$$(11) \qquad \sigma_{ii} = \frac{\delta_{ii}}{W_i} = 1 + \frac{\gamma_{ii}}{W_i^{\,2}} - \frac{1}{W_i}$$

$$(12) \qquad \sigma_{ij} = \frac{\delta_{ij}}{W_j} = 1 + \frac{\gamma_{ij}}{W_i W_j} \qquad i \neq j\,.$$

Separability and aggregation

In this analysis both an aggregated (model A) and a disaggregated (model B) demand system for meat in Australia are estimated. The four types of meat included in model A are beef and veal, lamb, pork and chicken (to be referred to as the meat group for model A). The six types of meat included in model B are beef and veal, lamb, fresh pork, ham, bacon and chicken (to be referred to as the meat group for model B).

[4] Green and Alston (1990) present corrected formulae for calculating the LA/AIDS demand elasticities. The AIDS formulae used here, in which expenditure shares are treated as constant, have been found to approximate closely the LA/AIDS formulae.

In model B the data on consumption of pork (as used in model A) is disaggregated into fresh pork, ham and bacon to allow more detailed estimation of the expenditure and price elasticities for these cuts of pork.

The standard assumption is made here that the meat group is weakly separable from other food groups, as well as from all other commodity groups. Unless this assumption is made it would not be possible to limit the number of prices appearing in each equation to the prices of items within the meat group. It should also be noted that Alston and Chalfant (1987) found some support for this assumption of weak separability in their analysis of the demand for meat in Australia.

Given separability, the system of share equations arising from the LA/AIDS model includes as explanatory variables the total expenditure on the meat group and the prices of individual types of meats within the meat group. The real expenditure variable is the total expenditure on all meat, divided by Stone's (1953) geometric price index. Finally, the dependent variables of the system, the budget share for each meat, were obtained by dividing expenditure on each meat by the total expenditure on all meat.

Given the nature of market-level data, most studies using the LA/AIDS model have used aggregate commodities in the analysis. For example, in the case of meat they have modelled the demand for pork rather than the demand for fresh pork, bacon and ham. Until this study, such disaggregated data has not been available in Australia, and it is of interest to model the demand for these disaggregated meat products to analyse how the demand for aggregate pork reflects the more varied demand for its constituent products.

Maximum Likelihood Estimation

The general nonlinear form for the representation of the LA/AIDS model given in equation (1) is:

$$W_{it} = f_i(X_{it}, \beta) + u_{it}, \qquad t = 1,\ldots, T \qquad i=1,\ldots, n \tag{13}$$

where for each of the n regression equations: W_t is a (nx1) vector of budget shares; $f(X_t, \beta)$ is a (nx1) nonlinear vector function of the elements in X_t and β, where X_t is a (kx1) vector of explanatory variables and β is a (mx1) vector of unknown coefficients; and u_t is a (nx1) vector of disturbance terms. It is assumed that: $E(u_{it})=0$, $E(u_{is},u_{jt})=\sigma_{ij}$ if $s=t$, $E(u_{is},u_{jt})=0$ if $s\neq t$. Let $u_i=(u_{i1},u_{i2},\ldots,u_{iT})'$ be the error vectors for the n share equations, and let $u'=(u'_1,\ldots,u'_n)$ be the joint error vector for the n equations of the estimated system. Then, the variance-covariance matrix for the system is $\Omega = \Omega(\Sigma) = E[uu'] = \Sigma \otimes I_T$, where Σ is the nxn symmetric matrix, assumed to be positive definite, whose ijth element is σ_{ij}. Hence, it is assumed that u are homoscedastic error terms which are contemporaneously correlated, but with vanishing own- and cross-lag covariances.

As the sum of budget shares in the meat demand system equals one, the contemporaneous covariance matrix will be singular. However, a system with a non-singular covariance matrix can be obtained by estimating only n–1 of the n equations of the system. Barten (1969) has shown that given the absence of serial correlation in the disturbances, maximum likelihood estimates of parameters can be obtained by arbitrarily deleting an equation, and that these estimates are invariant to which equation is deleted. The parameters for the omitted share equation can be calculated from those of the included n–1 equations, using the adding-up restrictions of equation (3). In the present study, the remaining n–1 equations are subject to maximum likelihood estimation using the iterative, nonlinear seemingly unrelated regressions (SUR) framework, assuming normality of the error terms. It should be noted that if the iterative SUR estimator converges for a given sample size, then it converges to the maximum likelihood estimator (as initially conjectured by Kmenta and Gilbert 1968 and demonstrated by Dhrymes 1971). The maximum likelihood estimates of the meat demand systems were obtained using the nonlinear regression procedure of SHAZAM version 6.1 (White *et al.* 1988).

Description and Sources of the Data

The data for the aggregate analysis (model A) are 94 quarterly observations from 1967(1) to 1990(2), and for the disaggregated analysis (model B) the data are 34 quarterly observations from 1982(1) to 1990(2). Model A comprises three equations, the dependent variables being the budget shares for beef and veal, lamb, and pork, respectively. Model B comprises five equations, with the budget shares for beef and veal, lamb, fresh pork, ham, and bacon as dependent variables. In both models, the chicken share equation has been omitted from the system to prevent singularity of the covariance matrix. In common with other Australian studies of meat demand, fish is excluded owing to the unavailability of data. Mutton is excluded owing to the concerns expressed by Chalfant and Alston (1988) regarding the quality of this data.

Quarterly data on per capita consumption of meats are not directly available in Australia, so a series for apparent domestic consumption was calculated for each type of meat. Apparent per capita consumption was calculated as the difference between production and the carcase weight equivalent of net exports, plus the reduction in the quantity of frozen stocks. Consumption data on beef, lamb, pork (that is, all types of pigmeats) and chicken for the period 1967(1) to 1982(2) were derived from Griffith, Freshwater and Smith (1983). Consumption data for the period 1982(3) to 1990(2) were obtained from the Australian Bureau of Agricultural and Resource Economics (ABARE), and were derived using data supplied by the Australian Meat and Livestock Corporation and the Australian Bureau of Statistics (ABS 1990b, 1990c).[5]

For the disaggregated analysis the apparent per capita consumption of pigmeat was broken down into consumption of fresh pork, bacon and ham. Data on the consumption of fresh pork, bacon and ham were taken from the Australian Pork Corporation's (APC) quarterly household consumption surveys. The estimated household consumption of the three types of pigmeats was then multiplied by the estimated number of households in Australia (taken from ABS 1990) to derive the estimated quarterly national consumption data. As the APC data only accounts for household consumption of the three types of pigmeats, (while the data on beef, lamb and chicken also includes non-household consumption), it was decided to allocate apparent Australian consumption of all pork among the three types of pigmeats using the relative proportions of consumption of fresh pork, bacon and ham calculated from the APC data. These relative proportions were multiplied by the ABARE series for the apparent per capita consumption of pork (all pigmeats) to derive the data on the apparent per capita consumption of fresh pork, bacon and ham. Such a procedure implicitly assumes that the relative shares of the three types of pigmeats in non-household consumption are the same as for household consumption, but is necessitated by the paucity of data on non-household consumption of pork in Australia.

The explanatory variables used in both the aggregate and disaggregate analyses included the retail prices of beef, lamb, pork (all pigmeats), chicken, fresh pork, bacon and ham. Data on the retail prices of beef, lamb, pork and chicken were taken from ABARE (1990) and ABS (1990a), and are based on the average retail price of selected cuts (weighted by expenditure) in capital cities of each of the six Australian states. Data on the retail prices of fresh pork, bacon and ham were taken from the quarterly household consumption surveys of the APC, and are calculated on a similar basis to the ABARE data.

Statistical Inference

Since there is little point in testing restrictions on an invalid model, the first testing exercise was to find the appropriate specification of the demand system. Accordingly, a series of nested tests was carried out for the presence of autocorrelation in the disturbance terms of the estimated demand systems, followed by further tests on the applicability of the imposition of restrictions derived from utility theory.

In these tests, the LA/AIDS model of equation (1) with autocorrelation (as set out in equation (15) below) was taken as the 'correct'

[5] One caveat is that apparent consumption data may not necessarily reflect patterns of demand at the household level due to the inclusion in it of demand by the meat manufacturing sector and the catering industry. However, exact data on the manufacturing and catering shares of total apparent consumption are not available for Australia, and in any case are small relative to the household share of total consumption.

specification, with: $i=1,...,3$ (beef, lamb and pork) and $t=1,...,94$ (the 'aggregate' model A); and $i=1,...,5$ (beef, lamb, fresh pork, ham and bacon) and $t=1,...,34$ (the 'disaggregated' model B). In both models, quarterly additive intercept dummies were included to capture seasonality.

Correlation between disturbances for different equations at a given point in time would be expected, given that these disturbances are likely to reflect common unmeasurable factors. Further, as noted by Judge *et. al.* (1985), given that the observations on any equation in a system will frequently be observations over time, it is reasonable to introduce some of the time series assumptions for the disturbance vector in single equations. Hence, disturbances can be both contemporaneously[6] and serially correlated.

Serial correlation

As noted under 'Maximum Likelihood Estimation' above, in the absence of serial correlation the deletion of an arbitrary equation to ensure non-singularity of the covariance matrix results in the parameter estimates of the included equations being invariant to the equation omitted. However, when u_t is autocorrelated, (and it is likely to be for time series data) then Berndt and Savin's (1975) solution for the linear model needs to be implemented to enable maximum likelihood estimation of the nonlinear system.

The revised n–1 equation system is:

$$(14) \qquad W_t' = f'(X_t, \beta) + u_t' \qquad t = \rho + 1, ..., T$$

and the error vector can be written (given $\Sigma_i u_{it} = 0$) as:

$$(15) \qquad u_t' = R_1 u_{t-1}' + R_2 u_{t-2}' + \ldots + R_\rho u_{t-\rho}' + v_t' \qquad t = \rho+1, ..., T$$

where: $v'_t \sim N_{n-1}(0, \Omega)$; the $'$ superscript denotes a vector with the nth element deleted; β is the vector of coefficients for the first n–1 equations; and R_i, $i=1,...,\rho$ are square matrices of autoregressive parameters of order n–1. Given the above configuration, the maximum likelihood parameter estimates will be invariant to whichever is the nth equation deleted, and parameter estimates for the deleted equation can be constructed in the same manner as for the case in which $R_i=0$, $i=1,...,\rho$. Assuming no autocorrelation across equations (that is, each R_i, $i=1,...,\rho$, is diagonal), Berndt and Savin (1975) have shown that it is necessary to restrict the R_i so that each diagonal element is the same

[6] Breusch and Pagan's (1980) Lagrange Multiplier test for contemporaneous covariances in the meat demand systems was carried out for both models A and B. The null hypothesis that all covariances were equal to zero was rejected for both models, and so given the presence of cross-equation restrictions, a systems approach will give more efficient parameter estimates than OLS estimation of each equation.

for the adding-up condition to be preserved. This restriction is imposed here.

It is well known that the Durbin-Watson statistic, the autocorrelation function and partial autocorrelation function statistics are of limited value in testing for autocorrelation outside the single-equation context. However, they are often erroneously used as diagnostic devices by researchers in testing for the presence of autocorrelation in demand systems.

Bewley (1986) has argued that analysis of a demand system, with or without cross-equation restrictions, requires that a system approach be followed in testing for autocorrelation. Accordingly, asymptotic test statistics are used here to test for the presence of autocorrelation in the demand systems, keeping in mind the caveat that the specification errors from applying asymptotically valid tests of inference in small-sample settings may not be trivial. Moreover, the reliability of such tests is aiso dependent on the accuracy of the approximation provided by the flexible functional form used in the estimation.

Maximum likelihood estimation allows the maximised likelihood values to be used to make formal comparisons between versions of the model which embody different restrictions. An appropriate method of testing for vector autocorrelation is to perform a sequence of Likelihood Ratio (LR) tests by estimating the system under the null hypothesis of independently distributed disturbances and under the alternative, a diagonal vector autoregressive (AR) process of order ρ.[7] That is:

$$H_o\colon\ u'_t = R_1 u'_{t-1} + R_2 u'_{t-2} + \ldots + R_\rho u'_{t-\rho} + v'_t$$

against the alternative

$$H_1\colon\ u'_t = v'_t$$

where $t=\rho+1,\ldots,T$ for all $n-1$ share equations, and again the $'$ superscript denotes a vector with the nth element deleted. The above LR statistic is asymptotically distributed as $\chi^2_{.05,\rho}$ with ρ being the order of the diagonal R_ρ matrix of autoregressive parameters (Judge *et al.*, 1985).

The LA/AIDS model was estimated with homogeneity and Slutsky symmetry imposed. The LR statistics were then calculated for H_0 against each of the following four alternatives:

$$R_1 \neq 0;\ R_2 \neq 0;\ R_3 \neq 0;\ R_4 \neq 0,$$

which would appear sufficient given the quarterly nature of the data used in the analysis. The results are given in Table 1.

[7] This statistic is formed using the values for the log of the likelihood functions for the restricted ($L(\theta^*)$) and unrestricted ($L(\theta)$) models, as: $LR = -2[\log L(\theta^*) - \log L(\theta)]$. The LR test statistic is asymptotically distributed as a χ^2 variable with degrees of freedom equal to the number of restrictions.

For both models A and B, the calculated LR statistics for all four of the above representations of the error structure greatly exceeded their respective critical values, with the largest value of the LR statistic being for $\rho=4$ for both models, (that is R_4). Hence the hypothesis of independently distributed disturbances is rejected for both models.[8]

Tests of demand restrictions

Evidence from work by Laitinen (1978), Meisner (1979), Bera, Byron and Jarque (1981) and Bewley (1983) indicates that asymptotic test statistics tend to over-reject restrictions derived from utility theory when imposed on demand systems in finite samples. Using Monte Carlo experiments on LR tests of functional form restrictions, Wales (1984) also found that the LR test rejected the null hypothesis too often.

All of the above researchers noted that as sample size increases, unless an adjustment is made to the significance level, the test becomes biased in type I error. Bera, Byron and Jarque (1981) and Bewley (1986) have argued that a size correction of $(T-k)/T$, (where T is the number of observations and k the number of explanatory variables in each equation), to the uncorrected asymptotic test statistic had excellent small-sample properties when the number of equations was less than or equal to 5 (as for both models A and B). Accordingly, the adjusted LR statistic (LR^*) is given by: $LR^* = ((T-k)/T)LR$.

Given the presence of autoregressive disturbances, asymptotic tests of the validity of the demand restrictions were used, in particular the LR^* test statistic. The maintained model was diagonal vector autoregressive, without the imposition of homogeneity and symmetry of the demand functions, and the results of the tests of demand restrictions for models A and B are given in Table 1.[9] For model A the hypothesis of homogeneity is not rejected at the 1 per cent level (it is rejected at the 5 per cent level), that of symmetry conditional on homogeneity is not rejected at either level, while the hypothesis of

[8] As in the single-equation case, the presence of autocorrelation implies a misspecification of some type. Both Deaton and Muellbauer (1980) and Blanciforti and Green (1983) have argued that the imposition of homogeneity may account for the presence of autocorrelation, because of the neglect of the dynamic aspects of consumption which occurs with the use of static demand equations. While *ad-hoc* formulations of dynamic LA/AIDS models can be found in the literature (see Blanciforti and Green 1983, Ray 1984 and Alston and Chalfant 1986) it is not clear that the homogeneity restrictions continue to hold in such models. An exception is Kneebone *et.al.* (1990), which does provide a theoretically-consistent specification of a dynamic AIDS model. Note also that specification error due to incorrect functional form can result in the imposition of autocorrelation, which in previous studies has lead people to assert the presence of habit persistence in consumption or introduce lagged dependent variables in demand equations (see Alston and Chalfant 1991a).

[9] Note that in an m-equation demand system there will in general be m homogeneity restrictions and $(m(m-1))/2$ symmetry restrictions.

TABLE 1
Results of Tests for Demand Restrictions

	Restriction Tested	Value Log Likelihood	LR* Value[c]	Critical Value (5 per cent)	Critical Value (1 per cent)
Model A	Static Model[a]	895.9418			
	Diagonal AR(1)	948.8720	105.86	$\chi^2_{.05,1} = 3.84$	
	Diagonal AR(2)	961.3855	130.88	$\chi^2_{.05,2} = 5.99$	
	Diagonal AR(3)	978.3008	164.72	$\chi^2_{.05,3} = 7.81$	
	Diagonal AR(4)	984.0316	176.18	$\chi^2_{.05,4} = 9.49$	
	Maintained Model[b]	992.6371			
	Homogeneity	986.5864	10.67	$\chi^2_{.05,3} = 7.81$	$\chi^2_{.01,3} = 11.3$
	Symmetry	984.0316	4.51	$\chi^2_{.05,6} = 12.60$	$\chi^2_{.01,6} = 16.8$
	Homotheticity	944.3295	70.01	$\chi^2_{.05,3} = 7.81$	$\chi^2_{.01,3} = 11.3$
Model B	Static Model[a]	657.8756			
	Diagonal AR(1)	671.4662	27.18	$\chi^2_{.05,1} = 3.84$	
	Diagonal AR(2)	679.6495	43.54	$\chi^2_{.05,2} = 5.99$	
	Diagonal AR(3)	682.5530	49.35	$\chi^2_{.05,3} = 7.81$	
	Diagonal AR(4)	688.7273	61.70	$\chi^2_{.05,4} = 9.49$	

Maintained Model[b]	716.3066			
Homogeneity	694.9761	28.44	$\chi^2_{.05,5} = 11.1$	$\chi^2_{.01,5} = 15.1$
Symmetry	688.7273	8.33	$\chi^2_{.05,10} = 18.3$	$\chi^2_{.01,10} = 23.2$
Homotheticity	647.8713	54.47	$\chi^2_{.05,5} = 11.1$	$\chi^2_{.01,5} = 15.1$

[a] Vector autocorrelation is tested by estimating both models A and B under the null hypothesis of independent disturbances, that is the static (or unrestricted) model, and in turn under each of the following alternative hypotheses: diagonal vector autoregressive AR(1), AR(2), AR(3) and AR(4) (restricted models).

[b] Note that the maintained model for both models A and B is diagonal vector AR(4), without the imposition of restrictions derived from utility theory.

[c] The LR test statistic is used in testing for vector autocorrelation.

homotheticity of the utility function is rejected at both levels, as expected.[10] Similar results are found for model B, except that the hypothesis of homogeneity is rejected at both levels of significance. On the basis of the above testing procedure, it is argued that the preferred specification for models A and B is diagonal AR(4), with homogeneity and symmetry of the demand functions imposed.

Estimation Results and Elasticity Estimates

Parameter estimates for models A and B are presented in Tables 2 and 3, respectively. These have been derived using equation (1), the error structure as given by equation (15) and the results obtained in 'Statistical Inference', above.[11]

TABLE 2
Estimated Parameters, LA/AIDS for Model A, 1967(1) to 1990(2)

Parameter[a]	Estimate	Standard Error
αB	-0.6480^b	0.1071
αL	0.3420^b	0.0666
αP	0.7370^b	0.0666
αC	0.5689	—
γBB	0.0564^b	0.0293
γBL	0.0370^b	0.0161
γBP	-0.0344^b	0.0169
γBC	-0.0591	—
γLL	-0.0570^b	0.0155
γLP	0.0222	0.0122
γLC	-0.0022	—
γPP	0.0030	0.0182
γPC	0.0092	—
γCC	0.0521	—
βB	0.3431^b	0.0300
βL	-0.0686^b	0.0179
βP	-0.1556^b	0.0192
βC	-0.1188	—
Bs1	0.0531^b	0.0039
Bs2	0.0410^b	0.0046

Cont'd

[10] Note that the imposition of Slutsky symmetry automatically imposes homogeneity in the LA/AIDS model. Accordingly, the unrestricted model for testing the hypothesis of symmetry alone had homogeneity imposed.

[11] Hausman (1978) tests were carried out to examine the validity of the assumed exogeneity of the price regressors in each of the equations of both demand systems. The hypothesis of price exogeneity was not rejected for any of the equations in either models A or B. This study also follows many other studies in the demand literature in assuming exogenous meat expenditures. La France (1991) has recently pointed out the potential problems associated with this assumption.

TABLE 2 (continued)

Parameter[a]	Estimate	Standard Error
Bs3	0.0298[b]	0.0040
Ls1	–0.0095[b]	0.0023
Ls2	–0.0091[b]	0.0027
Ls3	0.0010	0.0024
Ps1	–0.0376[b]	0.0024
Ps2	–0.0296[b]	0.0027
Ps3	–0.0287[b]	0.0025
ρ1	0.4277[b]	0.0685
ρ2	0.0813	0.0660
ρ3	0.1823[b]	0.0649
ρ4	0.2539[b]	0.0615

[a] Asymptotic standard errors are given in column 3, and (–) indicates the parameters of the chicken equation (which was left out for estimation) were derived using the adding-up restrictions of equation (3). Notation is: B=beef, L=lamb, P=pork, C=chicken. Note that s1, s2 and s3 are the quarterly seasonal dummy variables. The autocorrelation coefficients are denoted by ρ.

[b] Significant at the 5 per cent level.

TABLE 3
Estimated Parameters, LA/AIDS for Model B, 1982(1) to 1990(2)

Parameter[a]	Estimate	Standard Error
αB	–0.0111	0.1220
αL	0.2899[b]	0.0664
αF	0.1411[b]	0.0243
αH	0.3923[b]	0.0295
αN	0.0967[b]	0.0345
αC	0.0911	—
γBB	0.2033	0.1277
γBL	–0.0251	0.0625
γBF	0.0031	0.0214
γBH	–0.0251	0.0246
γBN	0.0017	0.0339
γBC	–0.1579	—
γLL	–0.0052	0.0411
γLF	0.0155	0.0124
γLH	–0.0205	0.0141
γLN	–0.0032	0.0163
γLC	0.0385	—
γFF	–0.0165	0.0106
γFH	0.0383[b]	0.0083
γFN	–0.0118	0.0125
γFC	–0.0286	—
γHH	–0.0420[b]	0.0130

Cont'd

TABLE 3 (continued)

Parameter[a]	Estimate	Standard Error
γHN	0.0221	0.0139
γHC	0.0272	—
γNN	0.0009	0.0218
γNC	-0.0097	—
γCC	0.1302	—
βB	0.2016[b]	0.0359
βL	-0.0406[b]	0.0210
βF	-0.0461[b]	0.0080
βH	-0.0987[b]	0.0083
βN	-0.0349[b]	0.0113
βC	0.0187	—
Bs1	-0.0612[b]	0.0140
Bs2	-0.0998[b]	0.0164
Bs3	-0.1029[b]	0.1936
Ls1	-0.0023	0.0080
Ls2	0.0035	0.0094
Ls3	0.0135	0.0111
Fs1	0.0143[b]	0.0031
Fs2	0.0262[b]	0.0036
Fs3	0.0236[b]	0.0043
Hs1	-0.0226[b]	0.0032
Hs2	-0.0247[b]	0.0038
Hs3	-0.0202[b]	0.0046
Ns1	0.0311[b]	0.0045
Ns2	0.0408[b]	0.5209
Ns3	0.0415[b]	0.0061
ρ1	0.3936[b]	0.1115
ρ2	0.3870[b]	0.1197
ρ3	0.1690	0.1127
ρ4	-0.0538	0.1097

[a] Asymptotic standard errors are given in column 3, and (—) indicates the parameters of the chicken equation (which was left out for estimation) were derived using the adding-up restrictions of equation (3). Notation is: B=beef, L=lamb, C=chicken, H=ham, N=bacon, F=fresh pork. Note that s1, s2 and s3 are the quarterly seasonal dummy variables. The autocorrelation coefficients are denoted by ρ.

[b] Significant at the 5 per cent level.

Using equations (6), (7) and (9), the own-price, cross-price and expenditure elasticities were calculated for both models. These elasticities are reported in Table 4 (for model A) and Table 5 (for model B), evaluated at the mid-points of the periods under analysis, 1978(4) and 1985(4), respectively. Elasticities were also calculated for other quarters, but showed little variation over the respective periods of interest.

TABLE 4
Model A: Uncompensated Expenditure and Price Elasticities[a] for Meat, Australia, 1978(4)

Meat i	ε_{iB}	ε_{iL}	ε_{iP}	ε_{iC}	η_i
Beef (B)	−1.235	−0.023	−0.196	−0.194	1.650
Lamb (L)	0.507	−1.326	0.249	0.044	0.525
Pork (P)	0.236	0.221	−0.829	0.143	0.228
Chicken (C)	0.027	0.118	0.262	−0.469	0.061

[a] Elasticities are calculated from equations (6), (7) and (9) for expenditure (η_i), and uncompensated (Marshallian) own- (ε_{ii}) and cross-price (ε_{ij}) elasticities, respectively.

Note that η_i is the elasticity of demand for meat i with respect to expenditure on all meat, not total expenditure on all (meat and non-meat) goods. Note also that the uncompensated price elasticities are calculated for a given level of expenditure on all meat, and that ε_{ij} reflects the elasticity of demand for meat i with respect to changes in the price of meat j. All elasticities are calculated at the mid-point of the period analysed, 1978(4).

TABLE 5
Model B: Uncompensated Expenditure and Price Elasticities[a] for Meat, Australia, 1985(4)

Meat i	ε_{iB}	ε_{iL}	ε_{iF}	ε_{iH}	ε_{iN}	ε_{iC}	η_i
Beef (B)	−0.822	−0.113	−0.019	−0.102	−0.016	−0.356	1.376
Lamb (L)	−0.018	−0.989	0.103	−0.082	0.001	0.260	0.769
Fresh Pork (F)	0.419	0.356	−1.202	0.680	−0.141	−0.316	0.305
Ham (H)	−0.250	−0.164	0.249	−1.185	0.185	0.294	0.333
Bacon (N)	0.384	0.053	−0.179	0.512	−0.948	−0.073	0.344
Chicken (C)	1.074	0.993	−0.111	1.166	0.176	−0.227	1.114

[a] Elasticities are calculated from equations (6), (7) and (9) for expenditure (η_i), and uncompensated (Marshallian) own- (ε_{ii}) and cross-price (ε_{ij}) elasticities, respectively.

Note that η_i is the elasticity of demand for meat i with respect to expenditure on all meat, not total expenditure on all (meat and non-meat) goods. Note also that the uncompensated price elasticities are calculated for a given level of expenditure on all meat, and that ε_{ij} reflects the elasticity of demand for meat i with respect to changes in the price of meat j. All elasticities are calculated at the mid-point of the period analysed, 1985(4).

The uncompensated (Marshallian) own-price elasticities reported in Table 4 are all of the appropriate sign, and indicate that the demand for beef and lamb is price elastic, while that for pork and chicken is less elastic. The results reported in Table 5 for the disaggregated analysis of model B are also all of the appropriate sign, with fresh pork and ham being price elastic, and beef, lamb, bacon and chicken being less elastic. As noted by Murray (1984) and Fisher (1979), the relatively low values for ε_{CC} in both models is consistent with a high proportion of chicken consumption in Australia being for fast food and/or reserved for special occasions. The magnitude of the own-price elasticity estimates for beef, lamb, pork and chicken are close to values obtained from previous studies (Chalfant and Alston 1986, Alston and Chalfant 1991). Those for fresh pork, ham and bacon have not been reported or published before in the literature.

The own-price elasticity result for aggregate pork reflects the underlying elasticities of its constituent products. The results in Tables 4 and 5 reveal that the own-price elasticity for pork is smaller in absolute value than the own-price elasticities of the three cuts of pork (fresh pork, ham and bacon). It seems that the own-price response for pork in the aggregate is reduced by the substitution between fresh pork, bacon and ham. Eales and Unnevehr (1988) have a similar finding for the demand for beef and chicken in the United States.

Two general results for the uncompensated cross-price elasticities reported in Tables 4 and 5 are the large differences in ε_{ij} and ε_{ji} in many cases, and the finding (contrary to *a priori* expectations) of a large number of negative uncompensated cross-price elasticities, particularly in model B. The latter result is most pronounced in the beef equations of both models A and B. This result could reflect either strong income effects or inadequacies in the data set for the consumption of beef, perhaps due to a large divergence between apparent consumption as measured by the data and actual consumption.

While pork and chicken are found to be gross substitutes in model A (Table 4), for model B (Table 5) chicken and fresh pork are found to be gross complements. Such complementary relationships can arise because of correlation in the quantities of chicken and fresh pork supplied, as producers of both types of meats rely heavily on cereal grains as an input to production. Reductions in supply of both meats (and hence increases in price) would then tend to occur together, and be measured as negative cross-price elasticities (Fisher 1979).

Of greater interest than the uncompensated cross-price elasticities is the substitution in consumption between meats in terms of their partial elasticities of substitution. The Allen elasticities of substitution (σ) are presented in Tables 6 and 7, and are estimated using the formulae for σ_{ii} and σ_{ij} given in equations (11) and (12), respectively.

The results for model A (see Table 6) are in accord with theory and imply that all pairs of goods are Hicks-Allen substitutes, and are close to the estimates found by Chalfant and Alston (1986). In Table 7 the results for model B are given, and imply again that all pairs of goods are Hicks-Allen substitutes, except for beef and chicken, fresh pork and chicken and fresh pork and bacon. The unexpected results for chicken have also been found by Martin and Porter (1985), and may reflect inadequacies in the data set for the consumption of chicken similar to those mentioned for beef above.

TABLE 6
Model A: Allen-Uzawa Partial Elasticities of Substitution[a] for Meat, Australia, 1978(4)

Meat i	σ_{iB}	σ_{iL}	σ_{iP}	σ_{iC}
Beef (B)	–0.694	1.486	0.676	0.113
Lamb (L)		–8.650	1.762	0.876
Pork (P)			–3.882	1.361
Chicken (C)				–3.643

[a] Elasticities calculated from equations (11 and (12) for own-price (σ_{ii}) and cross-price (σ_{ij}) Allen elasticities of substitution, respectively.

All elasticities are calculated at the mid-point of the period analysed, 1978(4).

TABLE 7
Model B: Allen-Uzawa Partial Elasticities of Substitution[a] for Meat, Australia, 1985(4)

Meat i	σ_{iB}	σ_{iL}	σ_{iF}	σ_{iH}	σ_{iN}	σ_{iC}
Beef (B)	–0.157	0.734	1.087	0.683	1.061	–0.790
Lamb (L)		–4.841	2.325	0.214	0.649	2.333
Fresh Pork (F)			–17.792	4.901	–2.353	–1.619
Ham (H)				–7.676	3.808	2.117
Bacon (N)					–17.433	0.280
Chicken (C)						–0.263

[a] Elasticities calculated from equations (11 and (12) for own-price (σ_{ii}) and cross-price (σ_{ij}) Allen elasticities of substitution, respectively.

All elasticities are calculated at the mid-point of the period analysed, 1985(4).

The expenditure elasticities reported in Tables 4 and 5 conformed to *a priori* expectations.[12] All meats in model A are normal goods, with beef having an expenditure elasticity of greater than one. In model B all meats are again normal goods, with beef being joined by chicken in having an expenditure elasticity in excess of one. It is also worth noting that the expenditure elasticity for aggregate pork is less than those for fresh pork, bacon and ham.

Conclusions

In this study the LA/AIDS model has been used to analyse the demand for meat in Australia. Two demand systems were estimated by maximum likelihood methods, one for aggregate types of meat and one for disaggregated meat products. Asymptotic tests for the presence of serial correlation in the error structure of the equations of the two demand systems found the static model with independent residuals to be inadequate for both of the models estimated. As a result, alternative vector autoregressive error structures were explored.

A number of nested asymptotic tests were also carried out to determine the desirability of imposing demand restrictions derived from utility theory. The hypothesis of symmetry conditional on homogeneity was not rejected for both models. As always, caution should be exhibited in drawing inferences from the results of parametric tests, as they are actually tests of the *joint* hypothesis that the assumed functional form is correct *and* of the validity of the hypothesis being tested.[13]

Using a new data set on the Australian retail price and consumption of fresh pork, ham and bacon, the first own-price, cross-price and expenditure elasticity estimates for these commodities were obtained from the disaggregated model. In both models the expenditure, uncompensated own-price and uncompensated cross-price elasticities were generally all of the appropriate sign, and will provide information on the strength of demand interrelationships between the types of meat included in these models.

In future research with this and similar disaggregated data sets, further work will be conducted on introducing dynamics and more flexible parametric and nonparametric functional forms. This work should prove fruitful as estimation of the demand for meat by the LA/AIDS model (which assumes constant marginal budget shares) could have induced autocorrelated residuals if there are changing

[12] These are expenditure elasticities for individual types of meat with respect to *expenditure on all meat*, rather than with respect to *total expenditures (or income)*. To convert the expenditure elasticities to the latter base from the former, Blanciforti and Green (1983) suggest the expenditure elasticities calculated here be multiplied by an estimate of the all meats expenditure elasticity with respect to total expenditure (or income), which for Australia has been estimated to be about 0.4 (see Richardson 1976).

[13] For further details see the papers by Alston and Chalfant (1991 and 1991a).

marginal budget shares for particular meats owing, for example, to changes in tastes. Nevertheless, the results of this analysis do indicate the demand interrelationships which exist between a group of previously unstudied disaggregated cuts of meat and the traditionally-studied group of aggregate meats. The results also highlight how the demand for aggregate pork is influenced by the underlying demand for its constituent products.

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