Demand Response to Advertising in the Australian Meat Industry

Nicholas E. Piggott, James A. Chalfant, Julian M. Alston, and Garry R. Griffith

The implications of model specification choices for the measurement of demand response to advertising are examined using Australian data. Single-equation models versus complete systems and alternative corrections for autocorrelation are evaluated. Competing advertising efforts by two producer bodies are included. Across all specifications, the evidence on advertising effects is fairly consistent. In the preferred model, the only statistically significant effects of advertising are for Australian Meat and Livestock Corporation advertising (of beef and lamb) on the demand for beef (positive) and on the demand for chicken (negative). Australian Pork Corporation advertising does not have any statistically significant effects.

Key words: advertising, Almost Ideal demand system, Australian meat demand, autocorrelation corrections.

Sir Toby:	O knight, thou lack'st a cup of canary. When did I see thee so put down?
Sir Andrew:	Never in your life, I think, unless you see canary put me down. Me thinks sometimes I have no more wit than a Christian or an ordinary man has, but I am a great eater of beef, and I believe that does harm to my wit.
Sir Toby:	No question.

—William Shakespeare, *Twelfth Night*

It appears to be widely believed that consumer preferences shifted significantly away from red meat during the 1970s and 1980s, and the industry has responded by investing in generic advertising. Similar concerns about changes in consumption patterns have been raised in Australia, Canada, and the United States, and the industry has responded similarly in the three countries. Whether this advertising has been profitable from the industry viewpoint is an important question. In order to answer that question, we must first measure whether advertising has led to a statistically significant shift in demand.

In this study we consider the implications of specification choices for findings concerning the statistical and economic significance of demand response to advertising by two producer groups in Australia: pork producers (Australian Pork Corporation, APC) and beef and lamb producers (Australian Meat and Livestock Corporation, AMLC). In particular, we investigate whether the results from tests for advertising effects are sensitive to the choices of functional form for demand equations and whether single-equation models or a complete systems approach is used. Previous studies of demand response to generic meat advertising have generally found that advertising has been a profitable undertaking from the industry viewpoint; that is, not only did advertising increase demand, it increased demand enough to more than cover the costs of the advertising expenditure (e.g., Ball and Dewbre; Ward and Lambert). Those results were obtained using single-equation models with relatively simple functional forms for demand equations (typically linear in the levels or logarithms of prices and quantities),

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which are not fully consistent with demand theory. Such models are ad hoc, and need not correspond to the underlying data-generating processes, so the findings might suffer from what Leamer termed "fragility": the models may be misspecified and the measured demand response to advertising cast into doubt. More flexible functional forms can be used to attempt to reduce the bias from misspecification. An alternative to ad hoc single-equation models is to incorporate advertising variables into flexible functional forms and to estimate the demand equations as a system.

In the context of Australian meat demand, it is desirable to use a systems approach for two other reasons, apart from any econometric advantages. As shown by Piggott, Piggott, and Wright, the returns to a particular group of producers (say pork producers) from their own advertising activities will be modified by price feedback effects from markets for related commodities (e.g., other meats) when prices are endogenous in a supply-demand system. This will be so even when pork advertising directly affects only the demand for pork, if it indirectly affects the other demands through induced changes in pork prices. In addition, advertising one meat can have direct impacts on the demand for other meats that may be of interest. Using a systems approach makes it possible to measure both the direct and indirect cross-commodity effects of advertising in a consistent fashion. Further, demand systems based on flexible functional forms should approximate a wider range of underlying sets of preferences, with a smaller risk of specification bias arising from either incorrect functional forms or using a model that is not fully consistent with the theory of consumer demand.

The four contributions of this paper are as follows: (*a*) to incorporate advertising variables into the Almost Ideal demand system in a way that preserves the desirable theoretical properties of the model; (*b*) to compare the results from single-equation models, the Almost Ideal demand system, and the linear approximate Almost Ideal model; (*c*) to illustrate the implications of using Berndt and Savin's more general correction for autocorrelation in the context of a system of equations; and (*d*) to obtain evidence on the direct and cross-commodity effects of advertising among the different types of meat, and to evaluate the economic and statistical significance of the demand response to advertising.

Models Incorporating Advertising Effects

There have been many attempts to estimate demand relationships for meat in Australia (e.g., Alston and Chalfant 1991; Ball and Dewbre; Beggs; Cashin; Goddard and Griffith; Martin and Porter; Murray; Piggott). As in most of these studies, we treat meat as weakly separable, so that consumption of each meat depends only on group expenditure and the meat prices (and, perhaps, demand shift variables such as meat advertising). The group consists of beef, lamb, pork, and chicken. Alston and Chalfant's (1991) quarterly data on nominal average retail prices in cents per kilogram and quarterly per capita consumption (disappearance) in kilograms for the period 1978:3 to 1988:4 (forty-two observations) were used. Homogeneity was imposed in all of the models; adding-up and symmetry restrictions were imposed in the systems.[1,2]

Previous studies of demand response to advertising have found that advertising effects tend to persist, so that current consumption responds to advertising in previous periods. It was presumed that advertising effects may affect consumption for several quarters. Hence, in each demand equation, four quarterly observations (the current value and three lags) were included for real advertising expenditure ($A_{i,t}$) by each of the two producer groups, AMLC and APC. The advertising expenditure data are from Ball and Dewbre. They were computed as the sum of real advertising expenditures in each of three media (television, radio, and print) calculated as nominal advertising expenditure deflated by a price index for each medium. We do not have any data on advertising of chicken.

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[1] The data are consistent with the Generalized Axiom of Revealed Preference (GARP), which suggests that they are consistent with having been generated by the maximization of a stable, well-behaved utility function by a representative consumer. This result can be used to justify the assumption that the meat group is weakly separable, and supports the imposition of the theoretical restrictions on the demand equations. However, the stable preferences that "rationalize" the data may not be plausible (e.g., it might be required that beef is an inferior good), and the power of the test is unknown in this application (see Alston and Chalfant 1992). Further, the theoretical restrictions and the assumption of stable preferences might be incompatible with a particular functional form for demand. In this study we include trends and seasonal dummies and advertising variables as demand shifters, in spite of the GARP result that would suggest that these variables may not belong in the demand equations. One could argue that the GARP results imply that these variables will have small effects, but this requires faith in the power of GARP and a belief that the estimated parametric functional form is correct.

[2] All single-equation models were estimated using SAS as well as the OLS or the AUTO procedure in SHAZAM (White et al.); the systems were estimated using SAS.

Advertising Effects in the Almost Ideal Model

Recent studies have made progress with incorporating advertising variables in flexible demand models, including the translog, Rotterdam, and Almost Ideal demand systems (e.g., Baye, Jansen, and Lee; Cox; Duffy; Goddard and Amuah; Goddard and Griffith; Green, Carman, and McManus; and studies in Kinnucan, Thompson, and Chang). The Almost Ideal demand system (Deaton and Muellbauer) is used in the models below. It has been used previously to study the response of demand for food products to advertising, but not in the ways that are developed below, and typically only in its linear approximate form.

The equation for the budget share of the ith good is

$$S_{i,t} = \alpha_i + \sum_{j=1}^{n} \gamma_{ij} \ln P_{j,t} + \beta_i \ln\left(\frac{M_t}{P_t}\right)$$

where, in time t, $q_{i,t}$ = per capita consumption of meat i, $P_{i,t}$ = its price, $M_t = \Sigma_i P_{i,t} q_{i,t}$, reflecting the weak separability assumption, $S_{i,t} = P_{i,t} q_{i,t} / M_t$, and

$$\ln P_t = \alpha_0 + \sum_{k=1}^{n} \alpha_k \ln P_{k,t} + \frac{1}{2} \sum_{k=1}^{n} \sum_{j=1}^{n} \gamma_{kj} \ln P_{k,t} \ln P_{j,t}.$$

Almost always, P_t is approximated using Stone's price index (P_t^*), yielding the Linear Approximate (LA) Almost Ideal model.[3] Although more popular, the LA model is not integrable, and recent work (e.g., Buse) has called the linear approximation into question, so it may be better to estimate the original version instead. The two alternatives are tried below. Demand shifters are introduced as modifications of the "intercepts" (α_i's) as follows:[4]

$$\alpha_{i,t} = \alpha_{i0} + \sum_{j=1}^{2} \phi_{ij} \sum_{k=0}^{3} \omega_{jk} A_{j,t-k} + \tau_i T_t + \sum_{m=1}^{3} \theta_{im} QD_{m,t}$$

where $A_{j,t-k}$ is a measure of the real quarterly advertising expenditure by the jth producer group (j = 1 for AMLC; 2 for APC) lagged k = 0, 1, 2, or 3 quarters, T_t is a time trend set equal to 1 in 1978:3, and the QD_m's (m = 1, 2, or 3) are quarterly intercept dummies. Advertising is included as a free-form lag of four quarterly investments.[5] For each type of advertising, the lag weights sum to one ($\Sigma_k \omega_{jk} = 1$), and the shape of the lag profile is restricted to be the same across the equations (since the same values for each ω_{jk} appear in each share equation). Different advertising impacts are reflected in the different values for the ϕ_{ij} parameters across equations. Hence, the direct impact of advertising expenditure type j lagged k quarters on the share of meat type i is given by the product $\phi_{ij}\omega_{jk}$, and the total effect of an additional unit of advertising in the current quarter on consumption over a year (i.e., the current and subsequent three quarters) is measured by ϕ_{ij}.

In the conventional Almost Ideal model, the intercepts (α_{i0}) are constants, restricted to sum to one across the equations. We preserve those adding-up restrictions on the modified parameters.[6] To do so requires that the following restrictions hold:

$$\sum_{i=1}^{n} \alpha_{i0} = 1; \sum_{i=1}^{n} \phi_{ij} = 0 \forall j;$$

$$\sum_{i=1}^{n} \tau_i = 0; \text{ and } \sum_{i=1}^{n} \theta_{im} = 0 \forall m.$$

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[3] As shown by Moschini, the results may be sensitive to the units of prices used to construct Stone's price index. We followed conventional practice and did not make the correction recommended by Moschini. However, we did evaluate whether making such a correction would affect the results; it did not affect any qualitative conclusions.

[4] Note that the "intercepts" appear also in the equation for the price index, P_t, in the nonlinear version of the model.

[5] The choice of lag length did not seem to determine the results. Using the Akaike Information Criterion (e.g., Maddala), the four-quarter lag length was preferred over longer or shorter lags in three of the eight single-equation models. In three cases, using one less lag resulted in a very slight improvement in this criterion; in two cases one more lag represented an improvement. We settled on the specification with current value and three lags of advertising as a compromise, rather than report the best-fitting lag structure for each equation, for four related reasons. First, we avoid reporting different equations depending on which criterion is chosen (arbitrarily) for determining lag length. Second, a single lag structure for all equations is necessary for the systems, and is desirable in single-equation models for making comparisons with equations from systems. Third, our underlying theory is that advertising, if it has any effect, does so on perceptions about the goods making up the meats group; to model these demands jointly implies that the same measure of advertising should affect all demands. Finally, the results of interest, namely, the total effect of advertising in each demand equation, given by the sum of effects across the current and lagged values of advertising, are quite stable with respect to this choice.

[6] We could introduce demand shift variables as modifiers of any of the parameters—i.e., the intercept terms (the α_i's), the price coefficients (γ_{ij}'s), or the income coefficients (β_i's)—or any combination. Intercepts terms in share equations, unlike price coefficients, can easily be kept consistent with the underlying restrictions. As a general rule, the coefficients of any variables added to the share equations must sum to zero across the shares or the shares will not sum to one.

The other parametric restrictions implied by theory are unaffected by the modification of the intercepts. Symmetry implies that each $\gamma_{ij} = \gamma_{ji}$, and homogeneity requires $\Sigma_j \gamma_{ij} = 0$.

Advertising Effects in the Logarithmic Model

A logarithmic demand equation for the *i*th good, including advertising, time trends, and seasonal dummy variables as intercept shifters, assuming separability of the meat group, is given by

$$\ln q_{i,t} = \alpha_{i0} + \sum_{j=1}^{2} \phi_{ij} \sum_{k=0}^{3} \omega_{jk} A_{j,t-k} + \tau_i T_t + \sum_{m=1}^{3} \theta_{im} QD_{m,t} + \sum_{j=1}^{n} \gamma_{ij} \ln P_{j,t} + \beta_i \ln \left[\frac{M_t}{P_t^*} \right].$$

A difference from a more conventional logarithmic specification is the use of Stone's price index as the deflator for total expenditures. Dividing income by Stone's price index is equivalent to using the Slutsky equation to partition the uncompensated demand elasticities into compensated elasticities and income effects, and then collecting the terms involving income effects. Thus, the price coefficients are compensated elasticities, and imposing homogeneity requires that the price coefficients sum to zero within each equation (i.e., $\Sigma_j \gamma_{ij} = 0$).[7]

Noting that the right-hand side of the double-log model is identical to that of a share equation in the LA model, we also considered single-equation models with S_i as the dependent variable, instead of $\ln q_i$, and tested each against the other. Since these models are ad hoc, and may be misspecified in ways that mean that the results are fragile, specification tests for functional form were applied. Of course, the more important question, for present purposes, will be how the two models compare in terms of estimates of effectiveness of advertising.

Specification Tests of the Double-Log and Share Models

The fact that the share and double-log specifications have a common right-hand side makes it convenient to test each specification against a more general compound model that includes each model as a special case.[8] Consider the model

$$(1 - \lambda) \ln q_{i,t} + \lambda S_{i,t} = \alpha_{i0} + \sum_{j=1}^{2} \phi_{ij} \sum_{k=0}^{3} \omega_{jk} A_{j,t-k} + \tau_i T_t + \sum_{m=1}^{3} \theta_{im} QD_{m,t} + \sum_{j=1}^{n} \gamma_{ij} \ln P_{j,t} + \beta_i \ln \left[\frac{M_t}{P_t^*} \right].$$

The hypotheses $\lambda = 0$ (i.e., the double-log model is correct) or $\lambda = 1$ (i.e., the share model is correct) can be tested using variants of Davidson and MacKinnon's C- or P-tests for nonnested hypotheses. Details of these tests are described by Alston, Chalfant, and Piggott, who have shown that the tests are well behaved in data sets such as the one being used in this study.

Coefficient Estimates and Hypothesis Tests

In this section we discuss the results from estimating the models described above. We first consider the single-equation results, and then turn to the systems.

Results from the Single-Equation Logarithmic Models

Results for the double-log models are summarized in table 1. The models were first estimated without autocorrelation corrections, and we tried three variants of Ramsey's specification error test (RESET), in which predictions from the models ($\hat{y}$) were added to those models as regressors. Each model was reestimated with $\hat{y}^2$ added, with both $\hat{y}^2$ and $\hat{y}^3$, and finally, with $\hat{y}^2$, $\hat{y}^3$, and $\hat{y}^4$, and, in each case, the statistical significance of the added regressors was tested. Passing the RESET test then corresponds to an insignificant test statistic for all three specification tests.[9] Failing the RESET test suggests the model should be rejected, but does not imply a particular alternative. In addition, the models were estimated allowing for first-order autocorrelation.[10]

The double-log models generally performed

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[7] The elasticities themselves are largely unaffected by this reparameterization; i.e., the Slutsky equation can be used without bias.

[8] This is analogous to the nesting of the Rotterdam and LA models suggested by Alston and Chalfant (1993).

[9] Thursby has shown that RESET is robust to the presence of autocorrelation, in the context of performing the test using OLS when autocorrelation in the residuals also is present.

[10] Following Cashin, initially fourth-order autocorrelation corrections were tried as well, but these were found to be unnecessary.

Table 1. Compensated Double-Log Models of Demand for Meat in Australia (1978:3–1988:4)

	Beef ($i = 1$)	Lamb ($i = 2$)	Pork ($i = 3$)	Chicken ($i = 4$)
α_{i0}	–3.3481*	0.0795	1.1812*	0.6988
τ_i	–0.0054*	–0.00035	0.0044*	0.0077*
γ_{i1}	–0.4212*	0.6790*	0.4330*	0.3362
γ_{i2}	0.4313*	–1.2587*	0.1695	–0.1037
γ_{i3}	0.1330	0.1073	–0.8689*	0.2309
γ_{i4}	–0.1430	0.4724*	0.2664*	–0.4634*
β_i	1.8221*	0.4264*	0.1471*	0.1766
ϕ_{i1} (AMLC)	0.000191*	0.000048	–0.000074	–0.00034*
ϕ_{i2} (APC)	0.000262*	–0.000262	0.000101	–0.00033
ρ	0.4242*	0.1716	–0.1482	0.5997*
T_c	–4.5919*	0.5630	–2.4486*	2.9693*
T_p	–4.5246*	0.5278	–2.4346*	2.7986*
RESET	PASS	PASS	FAIL	PASS

Notes: The coefficient estimates are for the models corrected for first-order autocorrelation. The RESET results refer to uncorrected models. RESET = FAIL unless none of the three test statistics were statistically significant at the 95% confidence level. T_c and T_p are t-statistics for the test that the double-log model is correct against a share model alternative, using a critical value of 1.96. ρ is the first-order autocorrelation coefficient. Asterisks (*) are used to denote statistically significant coefficients ($\alpha = 0.05$). For brevity, seasonality coefficients and quarterly lag weights are not reported.

as might be expected. On the whole, the models fit the data well, although there is significant serial correlation in the equations for beef and chicken. The demand elasticities were generally in line with expectations and plausible. The compensated own-price elasticities were all negative (–0.4 for beef, –1.3 for lamb, –0.9 for pork, and –0.5 for chicken) and statistically significant. The cross-price elasticities generally support the view that the meats are all substitutes, with the strongest substitution effects being between beef and lamb, and between chicken and pork. The expenditure elasticities suggest that increases in meat expenditure will lead to increases in consumption of each meat type, with an increase in beef's share and a decrease in the shares of each of the other three meat types.

In relation to advertising effects, the results were mixed but plausible. AMLC advertising (of beef and lamb) had a statistically significant, positive effect on the demand for beef, and a statistically significant, negative effect on the demand for chicken. AMLC advertising did not have any statistically significant effects on demand for lamb or pork, although the estimated coefficients were of the expected positive and negative signs, respectively.

Although APC advertising (of pork) was not statistically significant in the pork or lamb equations, the estimated coefficients were of the expected positive and negative signs, respectively. Curiously, APC advertising did have a statistically significant, positive cross-effect on demand for beef (an unexpected result that seems anomalous), and a more plausible negative effect on demand for chicken (although the coefficient became statistically insignificant when autocorrelation corrections were made).[11]

Three of the four double-log models passed the RESET test. The double-log model was rejected for pork. This suggests that the double-log model may not be an appropriate specification for pork, but no particular alternative is implied. In addition, the double-log models were tested against a specific alternative, the share-dependent, single-equation version of the LA model, using the C- and P-tests. Significant t-statistics (denoted T_c and T_p, respectively) indicate rejection of the double-log model. The test against the share model indicated rejection

[11] In preliminary work with these models we tried different specifications of the advertising variables, including logarithms versus levels of advertising expenditures and fixed weights versus free-form weights. The use of levels of advertising seemed preferable, and is convenient for estimating a free-form lag structure. The models with free-form lags consistently outperformed those with fixed (linearly declining or all equal) weights. However, the measures of the total advertising effect were largely unaffected by the lag structure. The individual lag weights are not, for the most part, intrinsically interesting. Some were negative numbers, which may seem implausible (although interpreting individual coefficients in a free-form distributed lag is difficult). However, the statistically significant individual lag weights, when associated with significant effects of advertising on demand, were always plausible values of around 0.3 or 0.4.

of the double-log model for beef, pork, and chicken, but not for lamb, regardless of whether advertising was included or autocorrelation corrections were made.

In summary, only the lamb equation passed both the RESET and C- and P-tests. The lamb equation was also distinguished in that it was free from autocorrelation and it was the only meat for which the time trend in demand was not significant. The double-log models for the other meats showed some evidence of significant autocorrelation (beef and chicken), or of misspecification from the RESET test (pork), or the C- and P-tests (beef and chicken).

Results from the Single-Equation Version of the LA Model

A second set of single-equation results (table 2) was obtained by replacing the ln q_i's in the double-log models with budget shares to obtain single-equation versions of the LA model:

$$S_{i,t} = \alpha_{i0} + \sum_{j=1}^{2} \phi_{ij} \sum_{k=0}^{3} \omega_{jk} A_{j,t-k} + \tau_i T_t + \sum_{m=1}^{3} \theta_{im} QD_{m,t} + \sum_{j=1}^{n} \gamma_{ij} \ln P_{j,t} + \beta_i \ln \left[\frac{M_t}{P_t^*}\right].$$

Like the double-log models, these models were subjected to tests for autocorrelation and RESET tests, as well as against the specific double-log alternative.

The results from the share equations are comparable to the corresponding results from the double-log models. There is a significant negative trend and autocorrelation in the beef share equation. Both AMLC advertising and APC advertising are statistically significant, both having positive effects on demand for beef. The model passed the RESET test. The model with autocorrelation corrections is (marginally) rejected when tested against a double-log alternative.

The share models for both lamb and pork are plausible: autocorrelation corrections did not appear necessary; neither APC nor AMLC advertising is statistically significant, the models both pass the RESET test, and the share model is not rejected by the double-log alternative for either pork or lamb. There is a statistically significant positive trend in pork's share and an insignificant negative trend in lamb's share.

Finally, the share model for chicken fails all of the tests. The pattern of serial correlation is much the same as it was for the double-log model. Both AMLC advertising and APC advertising have statistically significant negative effects on chicken's share, although when the model is corrected for autocorrelation APC advertising is no longer statistically significant (as happened in the double-log model). The chicken model fails the RESET test and both the C- and P-tests reject the share model against a double-log alternative. There is a significant negative trend in chicken's share.

Table 2. Single-Equation Share Models of Demand for Meat in Australia (1978:3–1988:4)

	Beef ($i = 1$)	Lamb ($i = 2$)	Pork ($i = 3$)	Chicken ($i = 4$)
α_{i0}	−0.7741*	0.4979*	0.7981*	0.5090*
τ_i	−0.00236*	0.000083	0.00105*	0.00113*
γ_{i1}	0.0265	0.0174	−0.0266	−0.0200
γ_{i2}	0.1173*	−0.0666*	−0.0041	−0.0450*
γ_{i3}	−0.0229	−0.0056	0.0031	0.0020
γ_{i4}	−0.1209*	0.0548*	0.0275	0.0629*
β_i	0.3949*	−0.1023*	−0.1840*	−0.1135*
ϕ_{i1} (AMLC)	0.000084*	0.000002	−0.000024	−0.000047*
ϕ_{i2} (APC)	0.000108*	−0.000046	0.000008	−0.000053
ρ	0.3332*	0.1394	0.0134	0.6405*
T_c	2.1300*	−1.9184	0.7732	−4.3973*
T_p	2.0793*	−1.7530	0.8240	−4.2489*
RESET	PASS	PASS	PASS	FAIL

Notes: See notes to table 1; T_c and T_p are now the t-statistics for the test that the share model is correct against a double-log model alternative, using a critical value of 1.96.

Overall, the share-dependent models seem to perform slightly better than their logarithmic counterparts. With the exception of chicken, the share-dependent models pass the RESET test. The beef and chicken share equations are rejected by a double-log alternative using the C- and P-tests. Autocorrelation is comparable between the share and double-log models. The main results with respect to advertising seem similar between the two specifications; this is borne out later in the computation of elasticities and marginal revenues for demand response to advertising.

Autocorrelation Corrections in Demand Systems

We turn now to estimating systems of demand equations. Both the nonlinear Almost Ideal model and the LA model were estimated with and without advertising variables, and also with corrections for first-order autocorrelation. To consider general forms of autocorrelation, we assume the vector of errors in the system of equations is determined by $\mathbf{e}_t = \mathbf{R}\mathbf{e}_{t-1} + \mathbf{v}_t$ for $t = 2, \ldots, T$, where $\mathbf{v}_t$'s are independent $N(0, \Sigma)$ random vectors, and $\mathbf{R}$ is an n by n matrix of unknown parameters. As is well known, Berndt and Savin showed that, with $\mathbf{e}_{t-1}$ and $\mathbf{v}_t$ statistically independent, the adding-up property of shares ($\iota'\mathbf{S}_t = 1$) implies a restriction $\iota'\mathbf{R} = k$, where k is an unknown constant. Typically, this restriction has been imposed by also forcing $\mathbf{R}$ to be diagonal, hence, the common restriction that the autocorrelation coefficient, ρ, is the same for every share equation (e.g., Cashin). This is unlikely to be valid, if our single-equation results are any guide, and an advantage of using the full $\mathbf{R}$ matrix is that its diagonal elements need not be the same across share equations.

Berndt and Savin showed that maximum likelihood estimation of a system of $n - 1$ such equations satisfies invariance, provided that the $\mathbf{R}$ matrix is appropriately restricted. The restriction $\iota'\mathbf{R} = k$ can be transformed into a more tractable restriction of the form $\iota'\overline{\mathbf{R}} = 0$, where $\overline{\mathbf{R}}$ is an n by $n - 1$ matrix with elements $\overline{R}_{ij} = R_{ij} - R_{in}$ for $i = 1, \ldots, n$ and $j = 1, \ldots, n - 1$. Now define $\overline{\mathbf{R}}^*$ as the matrix formed by the first $n - 1$ rows of $\overline{\mathbf{R}}$. It is the elements of $\overline{\mathbf{R}}^*$, not $\overline{\mathbf{R}}$ or $\mathbf{R}$, that are obtained in estimation.[12]

Three alternative specifications of $\mathbf{R}$ are investigated here: (R1) no autocorrelation, $\mathbf{R} = 0$; (R2) a diagonal $\mathbf{R}$ matrix, with the diagonal elements restricted to be the same, as in the typical study; and (R3) the most general specification, relaxing the restrictions that the off-diagonal elements are zero and the diagonal elements are all the same. For each specification, we estimated the elements of $\overline{\mathbf{R}}^*$ and the other model parameters jointly using nonlinear iterated seemingly unrelated regressions. Under the assumption that the $\mathbf{v}_t$'s are normally distributed, the results are equivalent to maximum likelihood estimates (Berndt and Savin).

Results from the Almost Ideal Demand System

Table 3 contains results for six models: two specifications of the group price index and three different autocorrelation structures (R1, R2, and R3). The coefficients for chicken demand can be computed using the adding-up condition. Consider first the effect of the choice of the price index. The Almost Ideal model is theoretically preferred but more difficult to estimate. For any given autocorrelation structure, the Almost Ideal model has uniformly higher log-likelihood values than does the corresponding LA version, but there is no important difference in what the two models would imply about any particular economic hypothesis. Estimated expenditure and advertising coefficients, in particular, appear to be quite similar.

The different autocorrelation structures do result in different values for particular coefficients and might result in more important differences in conclusions about economic questions. The model with the most general autocorrelation structure (R3) is preferred; the hypotheses that the off-diagonal elements can be restricted to zero (R2) or the entire matrix can be restricted to zero (R1) were rejected. Interestingly, in the Almost Ideal model, the test of R1 against R2 fails to reject the hypothesis of no autocorrelation. In other words, taking the conventional approach of testing a model with zero autocorrelation (R1) against a model with the same first-order autocorrelation coefficient in every equation (R2) would lead to the conclusion that autocorrelation is not a problem in this model. However, by using the more general autocorrelation model, which most have not used, we rejected that conventional conclusion.

The different autocorrelation corrections change the estimated coefficients somewhat, even when they do not change any qualitative

[12] The estimated elements of $\overline{\mathbf{R}}^*$ can be used to recover the elements of $\mathbf{R}$ by using prior information in the form of zero restrictions or other information, as described by Berndt and Savin. However, actually solving for the individual R_{ij}'s may not be as important as simply knowing whether they are collectively statistically significant.

Table 3. Almost Ideal Demand System Results Under Alternative Specifications

	LA Model			Almost Ideal Model		
	R1	R2	R3	R1	R2	R3
Trends						
τ_1	−2.37E-03*	−2.32E-03*	−2.25E-03*	−2.30E-03*	−2.29E-03*	−2.28E-03*
τ_2	5.87E-05	1.78E-05	4.55E-05	3.91E-05	2.87E-05	5.06E-05
τ_3	1.08E-03*	1.09E-03*	9.95E-04*	1.07E-03*	1.08E-03*	1.00E-03*
Price effects						
γ_{11}	0.0618*	0.0579	0.1067*	2.3236*	2.2776*	0.0559
γ_{12}	0.0174	0.0173	0.0154	−0.3456*	−0.3523	0.0306
γ_{13}	−0.0321*	−0.0315	−0.0499*	−1.1205*	−1.1062*	−0.0277
γ_{22}	−0.0340*	−0.0347*	−0.0340*	0.0314	0.0331	−0.0393
γ_{23}	0.0074	0.0132	0.0052	0.1767*	0.1848*	−0.0002
γ_{33}	−0.0175	−0.0221	0.0285	0.5099*	0.5041*	0.0252
Expenditure effects						
β_1	0.3787*	0.3532*	0.3203*	0.3668*	0.3596*	0.3095*
β_2	−0.0693*	−0.0640*	−0.0686*	−0.0595*	−0.0605*	−0.0697*
β_3	−0.1856*	−0.1788*	−0.1496*	−0.1764*	−0.1742*	−0.1382*
Advertising effects						
ϕ_{11}	1.14E-04*	1.06E-04*	6.06E-05*	1.42E-04*	1.37E-04*	6.09E-05*
ϕ_{12}	4.45E-05	4.14E-05	3.14E-05	3.89E-05	3.97E-05	3.33E-05
ϕ_{21}	−1.64E-05	−1.23E-05	−8.93E-06	−2.31E-05	−2.12E-05	−9.89E-06
ϕ_{22}	−1.07E-05	−6.33E-06	−1.20E-05	−8.36E-06	−7.64E-06	−1.48E-05
ϕ_{31}	−2.26E-05	−2.45E-05	−4.29E-06	−3.31E-05*	−3.31E-05*	−3.76E-06
ϕ_{32}	2.53E-06	−4.21E-06	2.16E-05	−8.02E-06	−8.43E-06	2.45E-05
ϕ_{41}	−7.51E-05*	−6.96E-05*	−4.74E-05*	−8.63E-05*	−8.23E-05*	−4.72E-05*
ϕ_{42}	−3.13E-05	−3.09E-05	−4.11E-05	−2.25E-05	−2.36E-05	−4.30E-05
$\ln\mathcal{L}$	473.744	475.137	492.581	481.489	481.6983	496.7578
ρ		0.1900*			0.0807	

Notes: In order to save space, intercepts, seasonality parameters, advertising lag weights, and the elements of the autocorrelation matrix for R3, are not reported. 1 = AMLC advertising, 2 = APC advertising. The goods are 1 = beef, 2 = lamb, 3 = pork, 4 = chicken; $\ln\mathcal{L}$ is the value of the log-likelihood; ρ is the first-order autocorrelation coefficient in the system under R2.

conclusions. In all of the models, AMLC advertising has a statistically significant positive effect on beef demand and a statistically significant negative effect on chicken demand, while APC advertising effects are never statistically significant. In two models that were rejected, the Almost Ideal model with autocorrelation structures R1 and R2, AMLC advertising also had a statistically significant negative effect on pork demand (ϕ_{31}).

The estimated coefficients for demand response to advertising are mostly plausible and in accord with the results from the single-equation models, albeit mostly insignificant: AMLC advertising of beef and lamb would be expected to have positive coefficients in the beef equation and negative coefficients in the pork and chicken equations, but it would not be expected to have a negative coefficient in the lamb equation; APC advertising of pork would be expected to have a positive coefficient in the pork equation and negative coefficients in the equations for lamb and chicken, but it would not be expected to have a positive coefficient in the beef equation.

An Evaluation of Demand Response to Advertising

Elasticities of demand response to advertising measure the percentage change in consumption of the *i*th good in response to a 1% increase in the *j*th type of advertising expenditure (i.e., $\mu_{ij} = \partial \ln q_i/\partial \ln A_j$). An alternative measure of advertising effectiveness is the marginal revenue from advertising (i.e., $MR_{ij} = P_i \partial q_i / \partial A_j$), the increase in sales revenue that would result from a dollar increase in advertising expenditure, holding constant the price of the product. The marginal revenue from advertising must be greater than one for advertising to pay. The increase in revenue must also cover the opportunity cost of additional domestic sales (either foregone export sales or the marginal cost of production), so it would have to be greater than one. For a traded good, with advertising applied only to the domestic component but with no separation of the markets, the primary effect of advertising domestically is to draw product off the relatively elastic export market and onto the domestic market. In this case, it will be possible for the advertising to be profitable for the industry only if it leads to an increase in both domestic and export prices (when the markets are not separated) or an increase in either the domestic market price or the export market price (when the markets are separated). If the export demand facing Australia were perfectly elastic, it would be necessary to separate the markets for domestic advertising to be profitable.

In other words, it is not sufficient to have a statistically significant impact on domestic demand for the advertising to be profitable. Nor is it sufficient to have a marginal revenue from advertising greater than one, computed holding the product price constant. The advertising must also lead to a rise in the price of the product sufficient to cover any additional costs of production (Alston, Carman, and Chalfant). Whether it does will depend on the elasticity of supply, the price elasticity of total demand (which depends on the elasticities of domestic and export demand and the fraction exported), and the elasticity of total demand response to advertising (which is equal to the elasticity of domestic demand with respect to advertising multiplied by the fraction of output consumed domestically).[13] Even these parameters may not be sufficient to evaluate fully the economic effects of advertising when we consider that the interactions among the related meat markets, in both consumption and production, may call for an explicit multimarket analysis, as proposed by Piggott, Piggott, and Wright.

Such evaluations are well beyond the scope of the present paper, although they are a logical next step. For the present, we do report the advertising elasticities and the marginal revenue products to check whether the minimum necessary conditions for profitable advertising have been met. To do this requires taking account of the dynamic specification of the demand response to advertising. We have computed total or long-run elasticities reflecting, effectively, a permanent increase in advertising expenditure by using the four-quarter sum of the effects of an increase in advertising in a particular quarter (as measured by the coefficient ϕ_{ij} in each equation multiplied by the sum of the lag weights, one).

Advertising elasticities are reported in table 4, and the corresponding marginal revenues from advertising are reported in table 5. The asterisks denote those cases where the corresponding effect of advertising on demand was found to be statistically significant. Comparing the estimates across columns shows the effects of specification choices.

The elasticities and marginal revenues were virtually identical between single-equation models with expenditure shares versus logarithms of quantities as the dependent variable. Slightly greater differences emerged when cross-equation restrictions were imposed to go from single-equation share models to the LA model (the advertising effects measured using the linear approximation were almost identical to those from the Almost Ideal model reported in tables 4 and 5). In particular, the systems estimates restricted the lag structure for advertising effects to be the same across the different meats, as well as imposing the usual cross-equation restrictions, and each type of restriction could account for some of the differences between the single-equation and systems results. Nonetheless, the results were remarkably similar between the single-equation models and the systems.

Confining attention to the statistically significant effects, the results were more sensitive to choices about autocorrelation corrections for either system than between systems and single-equation models. In particular, the most general correction for autocorrelation resulted in somewhat smaller elasticities of demand response to AMLC advertising. Even so, the pattern of re-

[13] The marginal gross producer profit from advertising an exportable good on the domestic market, without separation of domestic and export markets, is approximately equal to

$$\Pi \approx Q \frac{dP}{dA} = \frac{k_D\left(\frac{\mu}{\iota}\right)}{\varepsilon - k_D\eta_D - (1 - k_D)\eta_E}$$

where k_D is the fraction consumed domestically, η_D is the domestic demand elasticity, η_E is the elasticity of export demand, ε is the elasticity of supply, μ is the elasticity of domestic demand response to advertising, and ι is the domestic advertising intensity (advertising expenditure as a fraction of the value of sales). Advertising is profitable, at the margin, if marginal gross profit (Π) is greater than one (the marginal cost of advertising).

sults is remarkably consistent. AMLC advertising has statistically significant positive effects on demand for beef (elasticities between 0.15 and 0.40) and negative effects on demand for chicken (elasticities between –0.05 and –0.10). APC advertising of pork was found to have a positive effect on beef demand that was statistically significant in the single equation models, but not in the systems.

The results suggest that AMLC advertising may have been profitable for the beef industry: the elasticities and marginal revenues from AMLC advertising are around 0.015 and 24:1, respectively, in the preferred Almost Ideal demand system model with the most general autocorrelation correction (the last column in tables 4 and 5). However, a necessary condition for profitability is that the export price is signifi-

Table 4. Elasticities of Demand Response to Advertising (μ_{ij}) for a Range of Specifications

	Single-Equation Models		Almost Ideal Demand System		
	Double-log	Share	R1	R2	R3
AMLC					
Beef	0.0310*	0.0266*	0.0372*	0.0358*	0.0157*
Lamb	0.0078	0.0020	–0.0206	–0.0187	–0.0080
Pork	–0.0121	–0.0191	–0.0171*	–0.0176*	0.0010
Chicken	–0.0552*	–0.0587*	–0.0959*	–0.0921*	–0.0540*
APC					
Beef	0.0220*	0.0178*	0.0053	0.0055	0.0038
Lamb	–0.0221	–0.0255	–0.0040	–0.0036	–0.0069
Pork	0.0085	0.0032	–0.0020	–0.0022	0.0122
Chicken	–0.0277	–0.0328	–0.0126	–0.0132	–0.0247

Notes: Single-equation models were corrected for first-order autocorrelation. Elasticities of demand response to advertising were calculated at every data point (1978:3 to 1988:4). The figures in the table are the sample means of the elasticity estimates. An asterisk denotes an elasticity for which the underlying response parameter in the demand equation was significantly different from zero.

Table 5. Marginal Revenues from Advertising (MR_{ij}) Under a Range of Specifications

	Single-Equation Models		Almost Ideal Demand System		
	Double-log	Share	R1	R2	R3
AMLC					
Beef	47.15*	41.81*	57.07*	55.06*	24.19*
Lamb	3.44	0.92	–9.26	–8.38	–3.55
Pork	–8.27	–11.75	–9.79*	–10.17*	0.87
Chicken	–23.16*	–23.58*	–38.02*	–36.50*	–21.51*
APC					
Beef	62.99*	52.26*	15.21	15.54	10.75
Lamb	–18.25	–22.25	–3.46	–3.08	–5.97
Pork	10.96	3.70	–2.13	–2.30	14.32
Chicken	–21.83	–25.52	–9.61	–10.16	–19.10

Notes: Single-equation models were corrected for first-order autocorrelation. Marginal revenues from advertising (holding prices constant) were calculated at every data point (1978:3 to 1988:4) using the equation in the text. The figures in the table are the marginal revenues evaluated at the last data point. An asterisk denotes a marginal revenue coefficient for which the underlying response parameter in the demand equation was significantly different from zero.

cantly responsive to Australian beef exports. From the point of view of the lamb industry, there is no evidence that the AMLC advertising has been profitable. Similarly, there is no persuasive evidence that the APC advertising has had any effect on demand for any of the meats.

Conclusion

The widespread and rising popularity of advertising as a component of the marketing mix chosen by producer groups, often funded through some mandatory check-off arrangements with the sanction of government, means that the nature of food demand response to advertising is becoming a matter of public interest, as well as a matter for private interest.[14] It is especially interesting when producer organizations fund advertising campaigns in direct competition with one another in a "beggar-thy-neighbor" fashion, as is the case in the Australian meat industry with the advertising investments undertaken by the AMLC and the APC. The question is often raised as to whether the two groups of producers would do better to agree to quit, even if it were in the interest of each to continue if they could not cooperate.

We have tried alternative functional forms and alternative dynamic specifications and various tests in an attempt to validate a preferred model specification, while paying attention to the desirability of preserving consistency with economic theory. A striking feature of the results is that the estimated advertising effects were not very sensitive to functional forms. The consistent result was that AMLC advertising was statistically significant in the equations for beef (a positive effect) and chicken (a negative effect) in every model.

Cross-commodity effects of advertising may be important. AMLC advertising had no effect on demand for lamb while increasing the demand for beef. Such an effect calls into question the desirability of a cooperative approach to advertising between producers of products that are close substitutes. A negative effect of AMLC advertising on chicken demand might be consistent with the objectives of the beef and lamb industry, and the effect could be economically important. Finally, APC pork advertising was usually found not to have any statistically significant effects. The measured responses were, however, in directions that would be consistent with a "beggar-thy-neighbor" approach to meat marketing strategy.

A complete economic evaluation of the benefits and costs of AMLC and APC promotion campaigns is beyond the scope of the present study, although we have indicated some issues that such an evaluation should address and our results provide the basis for such an evaluation. The missing additional information is on the values of the elasticities of supply and export demand for the various products, and their market shares. We suspect that, unless the Australian meat industry has a surprising degree of market power in international markets, the promotional campaigns may not have been profitable, but a more confident answer to that question remains the subject of further work.

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[14] Several of the mandatory check-offs for promotion in the United States have been challenged legally by producers who feel that they are of no benefit. Recently, the U.S. beef promotion program has been challenged by a lawsuit filed in the U.S. District Court in Kansas on 2 August 1994 (*Goetz* v. *United States of America*, Civ. Action No. 94-1299-FGT).

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