The Silence of the Lambdas: A Test of the Almost Ideal and Rotterdam Models

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During the past decade, the Linear Approximate (LA) Almost Ideal Demand System and the Rotterdam model have been adopted by agricultural economists as the demand systems of choice in most applications. The apparent explanation is that the two models are both (second-order) locally flexible and compatible with demand theory, they have identical data requirements and are equally parsimonious with respect to parameters, and both are linear in the parameters. While the two models are thus equally attractive in most respects, and indeed appear very similar in structure, they lead to different results in some applications. This article develops a test of each against the other. In an illustrative application to U.S. meat demand, the Almost Ideal model is rejected while the Rotterdam model is not.

Key words: Almost Ideal Demand System, demand systems, flexible functional forms, Rotterdam model.

Two demand systems have come to prominence in agricultural economics: the Almost Ideal Demand System and the Rotterdam model. In the comparatively short time since the Almost Ideal model was introduced by Deaton and Muellbauer (1980a, 1980b), it has been widely adopted by agricultural economists, to the point that it now appears to be the most popular of all demand systems.[1]

Its popularity can be ascribed to two properties: first, it is as flexible as other locally flexible functional forms (such as the Translog) but has the added advantage of being compatible with aggregation over consumers; second, and more importantly, it is relatively easy to estimate and interpret (largely due to the use of an approximation to the original Almost Ideal model). Almost always, empirical applications use the "Linear Approximate" (LA) version of the Almost Ideal model, in which the Almost Ideal price index is replaced with Stone's geometric price index, as suggested by Deaton and Muellbauer (1980a). The approximation is not integrable but is used in the hope that it provides a reasonable approximation to the true demand system (whether that system is of the Almost Ideal or some other form).

The Rotterdam model was first proposed by Barten (1964) and Theil (1965), prior to the development of so-called flexible functional forms and the advent of duality theory. It had been thought to be unduly restrictive, and this may explain why it has been used less often than the LA model, at least in the recent agricultural economics literature. However, the Rotterdam model is now known to be as flexible as any other locally flexible functional form (Mountain).[2] Thus, its popularity is rising, and we predict that it will be the main alternative to the LA model in the next few years.

The two models are similar in many respects. They are (second-order) locally flexible func-

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[1] Recent applications of the Almost Ideal model include, for example, Chalfant; Chalfant, Gray and White; Eales and Unnevehr; and Moschini and Meilke. Studies that have used the Rotterdam model include, for example, Capps and Schmitz; Penm; and Theil and Clements. A few studies have used both models, including Alston and Chalfant (1991a, 1991b); Barten (1992); Brester and Wohlgenant; and Piggott. Further examples are cited in the working paper version of this article.

[2] See also Barnett (1979, 1984), Byron, and the discussion in Deaton and Muellbauer (1980b, p. 73).

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tional forms, have identical data requirements, are equally parsimonious with respect to numbers of parameters, and are linear in parameters. Because they each have all of these characteristics, and most alternatives do not, these two models are likely to continue to be chosen more often than any others. Economic theory does not provide a basis for choosing *ex ante* between the two models, and provides only a limited basis for *ex post* discrimination (such as when one model violates the law of demand or another strong prior belief). They are difficult to compare using simple goodness-of-fit measures, because the dependent variables are different in the two systems. In a typical demand study, only one functional form is tried, so the choice between the Almost Ideal and Rotterdam models is likely to be made arbitrarily in advance.

Making the right choice may be important. The two models lead to different results in some applications. For instance, Alston and Chalfant (1991a) found that the Rotterdam model did not reject stable preferences, while the LA model implied statistically significant structural changes in demand. Piggott tried both models, augmented with advertising variables, with the same consumption and price data, and found greater advertising effects in the LA model than in the Rotterdam model. Alston and Chalfant (1991b) found contradictory results concerning trends when comparing these two models with Canadian meat consumption data. In contrast, Brester and Wohlgenant found that elasticity estimates did not vary too much between a first-differenced LA model and a Rotterdam model of U.S. meat demand.

None of these studies has attempted a statistical test for the "correct" model. Below we develop a test of each of these two models against the other. The fact that the two models can yield different results implies that testing each specification is appropriate. As with any pair of nonnested models, there are four possible outcomes from such a test. One may find that either model is rejected in favor of the other, that either is an acceptable representation of the data, or that both are rejected as incomplete. In an illustrative application to U.S. meat demand, the Almost Ideal model is rejected but the Rotterdam model is not.

A Compound-Model Approach

Suppose we have two alternative models in which the right hand sides are identical but the dependent variable differs:

$$\text{Model 1: } y = f(x)$$

$$\text{Model 2: } z = f(x)$$

For instance, both linear and logarithmic dependent variables could be of interest, where $z = \ln(y)$. In this case, the Box-Cox transformation can be used to nest both alternatives, and it is possible to test each against the more general alternative. More generally, however, it will not be possible to nest two competing models conveniently in such a simple alternative model. Sometimes nonnested hypothesis testing procedures may suffice, but such procedures seem to be better adapted to different right-hand sides. However, by estimating the compound model

$$\lambda y + (1 - \lambda)z = f(x)$$

a test of the hypothesis that $\lambda = 0$ is a test of the null hypothesis that model 2 is correct. By switching the roles of model 1 and model 2, we can test the hypothesis that model 1 is correct. A similar motivation underlies various nonnested tests (e.g. Davidson and MacKinnon), but when the two models have different right-hand sides, the compound model may not be feasible to estimate.

The Models

In this section, it is shown that the right-hand side of a first-differenced version of the LA model is virtually identical to that of the Rotterdam model, even though the dependent variables differ substantially. Moreover, the two dependent variables are not easily compared by conventional means, because they involve different transformations of the underlying consumption data. Thus, the choice between the LA and Rotterdam models fits the compound model framework described above.

The linear approximate (LA) version of the Almost Ideal demand system is given by

$$(1) \qquad s_i = \alpha_i + \sum_{j=1}^{n} \gamma_{ij} \ln p_j + \beta_i \ln (x/P)$$

where s_i denotes the budget share of good i ($i = 1, \ldots, n$), p_j the price of good j, and x the total expenditure on the n goods. P is given by Stone's geometric price index:

$$(2) \qquad P = \prod_{j=1}^{n} p_j^{s_j}.$$

Budget shares thus depend on relative prices and on a measure of real income.

In several studies, the LA model has been estimated in first-differenced form (e.g. Deaton and Muellbauer 1980a, Eales and Unnevehr, Moschini and Meilke, Alston and Chalfant 1991b, and Brester and Wohlgenant). In first-differenced form, the LA model becomes

$$(3)\quad \Delta s_i = \sum_{j=1}^{n} \gamma_{ij}\Delta \ln p_j + \beta_i \Delta \ln (x/P)$$
$$= \sum_{j=1}^{n} \gamma_{ij}\Delta \ln p_j + \beta_i [\Delta \ln x - \Delta \ln P]$$

where Δ denotes the first-difference operator. The first difference of Stone's index in (3) may be decomposed into three components:

$$(4)\quad \Delta \ln P = \sum_{j=1}^{n} s_j \cdot \Delta \ln p_j + \sum_{j=1}^{n} \Delta s_j \cdot \ln p_j - \sum_{j=1}^{n} \Delta s_j \cdot \Delta \ln p_j$$

The third term is generally negligible. The second term is also likely to be quite small; in the context of time-series data, shares usually do not change much from one observation to the next.[3] Substituting the first term of $\Delta \ln P$ from (4) into the first-differenced LA model in (3) yields

$$(5)\quad \Delta s_i \approx \sum_{j=1}^{n} \gamma_{ij}\Delta \ln p_j + \beta_i \left[\Delta \ln x - \sum_{j=1}^{n} s_j \Delta \ln p_j\right].$$

These two versions of the first-differenced LA model—equations (3) and (5)—differ only in the approximations to the price index in the "income term" and thus seem likely to produce approximately the same estimates.

Equation (5), in particular, is quite similar to the Rotterdam model, a point first made by Deaton and Muellbauer (1980a). The Rotterdam model has a different dependent variable but essentially the same right-hand side; any differences are in the specification of the income term. The dependent variable is given by $\bar{s}_i \Delta \ln q_i$, where q_i denotes the quantity consumed of good i and $\bar{s}_i$ denotes the average of $s_{i,t}$ and $s_{i,t-1}$. For instance, the absolute price version of the Rotterdam Model (e.g. Theil and Clements, p. 25) is given by

$$(6)\quad \bar{s}_i \Delta \ln q_i = \sum_{j=1}^{n} \gamma_{ij}\Delta \ln p_j + \beta_i DQ$$

where

$$(7)\quad DQ = \sum_{i=1}^{n} \bar{s}_i \Delta \ln q_i.$$

DQ thus plays the role of the real income term; it is referred to by Theil and Clements as a "finite change version of the Divisia volume index." It is approximately equal to

$$(8)\quad DQ^* = \Delta \ln x - \Delta \ln P^*$$

(e.g. Theil 1971, p. 332, or Theil and Clements, p. 22), where

$$(9)\quad \Delta \ln P^* = \sum_{j=1}^{n} \bar{s}_j \cdot \Delta \ln p_j.$$

The similarity of $\Delta \ln P^*$ to the first-difference of Stone's price index in (4) is evident. It is the same as the first and largest term of $\Delta \ln P$, except that a moving average of budget shares has been substituted for the current values of budget shares.[4]

On the right-hand side, it is only in the real income terms that the first-differenced LA model can be distinguished from the Rotterdam model. If DQ^* from (8) is used in the Rotterdam model, then algebraically the only remaining differences are those between $\Delta \ln P$ and $\Delta \ln P^*$. The differences involve the use of $\bar{s}_j$ instead of s_j in $\Delta \ln P^*$ and the deleted Δs_j terms from $\Delta \ln P$ in the LA model.

Alternative Specifications of the Real Income Term

Before deriving a specification test for these two models, it is worth considering how much dif-

[3] In the empirical work reported below, the mean absolute value of the first term was more than six times that of the second term, while the mean absolute value of the second term was nearly six times that of the third term. While such comparisons are necessarily specific to a particular data set, they are at least indicative of what one can expect in the typical meat demand application. We return to this point when discussing Barten's test in the penultimate section of the paper.

[4] To avoid simultaneity of the price index and the dependent variables (shares), Eales and Unnevehr suggested using lagged budget shares in Stone's index in an LA model. The Rotterdam version may thus be viewed as a compromise between Eales and Unnevehr's approach and the use of current shares in Stone's index. As a practical matter, it does not seem to make much difference empirically whether s_i or $\bar{s}_i$ (or, for that matter, the lagged value of s_i) is used in the Rotterdam specification.

ference is made by varying the form of the real income terms. We first compare the two specifications of the Rotterdam model, then several alternative versions of the first-differenced LA model. The estimates reported below are obtained by fitting alternative models to Moschini and Meilke's data, which consist of quarterly per capita consumption and retail prices of beef, chicken, pork, and fish in the United States, for the years 1967–88.

Rotterdam Model

Table 1 shows estimates of two alternative models of the Rotterdam form. Model I is estimated with the Divisia volume index (DQ) playing the role of real income, while model II is estimated in terms of expenditures and prices (using DQ^*). Quarterly dummy variables (D_j) and intercepts were included to capture possible trends or seasonality in the dependent variables not accounted for by the model. The θ_{ij}'s, for seasonal effects, are restricted to sum to zero within each equation. Beef, chicken, pork, and fish are denoted by 1, 2, 3, and 4, respectively. Weak separability, homogeneity, and adding-up are treated as maintained hypotheses and imposed on all models in this study. The two models estimated are

$$\text{Model I: } \bar{s}_i \Delta \ln q_i = \tau_i + \sum_{j=1}^{4} \theta_{ij} D_j + \sum_{j=1}^{4} \gamma_{ij} \Delta \ln p_j + \beta_i DQ$$

and

$$\text{Model II: } \bar{s}_i \Delta \ln q_i = \tau_i + \sum_{j=1}^{4} \theta_{ij} D_j + \sum_{j=1}^{4} \gamma_{ij} \Delta \ln p_j + \beta_i DQ^*$$

$$= \tau_i + \sum_{j=1}^{4} \theta_{ij} D_j + \sum_{j=1}^{4} \gamma_{ij} \Delta \ln p_j + \beta_i \left[\Delta \ln x - \sum_{j=1}^{4} \bar{s}_j \cdot \Delta \ln p_j \right].$$

Table 1. Parameter Estimates for Rotterdam Models I and II

	Model I		Model II	
	Estimate	Std. Error	Estimate	Std. Error
τ_1	−0.00048	0.00067	−0.00048	0.00065
θ_{11}	0.00868*	0.00132	0.00867*	0.00128
θ_{12}	0.00467*	0.00126	0.00467*	0.00122
θ_{13}	0.00365*	0.00145	0.00365*	0.00137
β_1	0.63137*	0.03078	0.63110*	0.03095
γ_{11}	−0.17403*	0.01679	−0.17414*	0.01647
γ_{12}	0.16223*	0.01411	0.16215*	0.01386
γ_{13}	0.00659	0.00658	0.00675	0.00631
τ_2	−0.00015	0.00064	−0.00015	0.00061
θ_{21}	−0.01081*	0.00126	−0.01082*	0.00122
θ_{22}	−0.01299*	0.00120	−0.01299*	0.00114
θ_{23}	−0.00106	0.00138	−0.00105	0.00132
β_2	0.29677*	0.02950	0.29642*	0.02892
γ_{22}	−0.18030*	0.01479	−0.18040*	0.01496
γ_{23}	0.00707	0.00642	0.00713	0.00645
τ_3	0.00075*	0.00027	0.00075*	0.00025
θ_{31}	0.00192*	0.00055	0.00193*	0.00052
θ_{32}	0.00734*	0.00050	0.00734*	0.00050
θ_{33}	−0.00180*	0.00058	−0.00181*	0.00055
β_3	0.01791	0.01338	0.01813	0.01302
γ_{33}	−0.01618*	0.00547	−0.01613*	0.00559
$\ln L$	1143.79		1143.89	

* denotes significance at the 0.05 level, based on asymptotic t-ratios.

As evident from table 1, it makes virtually no difference, in terms of the parameter estimates, which alternative is used. Such a result is not surprising: DQ and $\Delta\ln(x/P^*)$ have a correlation coefficient of one, rounded to five decimal digits. Thus, $\Delta\ln(x/P^*)$, defined as DQ^* in (9), can be substituted for the Divisia volume index, DQ, in the Rotterdam model, and approximately the same fit is obtained. Estimates of parameters and their standard errors differ little between the two alternatives.[5] This finding is advantageous, since using DQ^* rather than DQ makes the model closer in functional form to the LA model.

Almost Ideal Models

Consider now four variations of the first-differenced LA model:

$$\text{Model III: } \Delta s_i = \tau_i + \sum_{j=1}^{4}\theta_{ij}D_j + \sum_{j=1}^{4}\gamma_{ij}\Delta\ln p_j + \beta_i DQ$$

$$\text{Model IV: } \Delta s_i = \tau_i + \sum_{j=1}^{4}\theta_{ij}D_j + \sum_{j=1}^{4}\gamma_{ij}\Delta\ln p_j + \beta_i DQ^*$$

$$= \tau_i + \sum_{j=1}^{4}\theta_{ij}D_j + \sum_{j=1}^{4}\gamma_{ij}\Delta\ln p_j + \beta_i\left[\Delta\ln x - \sum_{j=1}^{4}\bar{s}_j\Delta\ln p_j\right]$$

$$\text{Model V: } \Delta s_i = \tau_i + \sum_{j=1}^{4}\theta_{ij}D_j + \sum_{j=1}^{4}\gamma_{ij}\Delta\ln p_j + \beta_i\left[\Delta\ln x - \sum_{j=1}^{4}s_j\Delta\ln p_j\right]$$

$$\text{Model VI: } \Delta s_i = \tau_i + \sum_{j=1}^{4}\theta_{ij}D_j + \sum_{j=1}^{4}\gamma_{ij}\Delta\ln p_j + \beta_i\Delta\ln(x/P)$$

$$= \tau_i + \sum_{j=1}^{4}\theta_{ij}D_j + \sum_{j=1}^{4}\gamma_{ij}\Delta\ln p_j + \beta_i[\Delta\ln x - \Delta\ln P]$$

Model III corresponds to the absolute-price version of the Rotterdam model, as defined in model I, with the dependent variable changed to the first difference of s_i. Model IV features DQ^* instead of DQ and thus corresponds to model II. To obtain model V, one simply uses the actual budget share (s_i) in the income term, instead of the average budget share ($\bar{s}_i$) used in model IV. Finally, model VI is the exact first-differenced LA model from (3). These four models would be expected to yield quite similar parameter estimates, and did so, as can be seen in table 2.

The first three sets of parameter estimates are very similar. This reinforces the conclusion about real income terms in the Rotterdam models (I and II) and also implies that the choice between s_i and $\bar{s}_i$ in the real income term is unimportant (as shown by the similarity of results from models IV and V). Estimates of model VI are similar, but differences between it and models III, IV, or V are generally greater than are differences among models III through V. Using the exact expression for $\Delta\ln P$ seems to have its greatest effect on the estimated β_i terms; it has little effect on estimated trends or seasonal effects. Thus, models III, IV, and V are not perfect substitutes for the first-differenced LA model in VI. However, given the similarity of results, they certainly appear indicative of what model VI will yield.

We did not anticipate that conclusions about an economic hypothesis, such as symmetry, would be affected much by choice of any particular one of these four models. However, symmetry was rejected in model VI but not in models III through V. It is interesting that model VI, using the standard specification of the income term of the LA model, also had the smallest log-likelihood value of all four models. Model VI again had the smallest log-likelihood value when symmetry was not imposed, but the differences between VI and III–V were smaller than when symmetry was imposed. Since symmetry need not hold with market-level data, the rejection of symmetry cannot be used to reject the

[5] We repeated the comparison between models with the symmetry restriction relaxed, and found the same degree of similarity between the two models. This may be because symmetry proved to be a mild restriction in the Rotterdam model: a likelihood-ratio test did not imply rejection of the symmetry hypothesis.

Table 2. Parameter Estimates for Almost Ideal Models III–VI

	Model III	Model IV	Model V	Model VI
τ_1	−0.00047	−0.00047	−0.00043	−0.00047
θ_{11}	0.00869*	0.00869*	0.00869*	0.00834*
θ_{12}	0.00477*	0.00477*	0.00479*	0.00417*
θ_{13}	0.00359*	0.00359*	0.00356*	0.00402*
β_1	0.13064*	0.13067*	0.13081*	0.11803*
γ_{11}	0.07876*	0.07874*	0.07832*	0.08664*
γ_{12}	0.01992	0.01992	0.02009	0.01925
γ_{13}	−0.04887*	−0.04884*	−0.04882*	−0.05536*
τ_2	−0.00025	−0.00025	−0.00025	−0.00027
θ_{21}	−0.01061*	−0.01061*	−0.01059*	−0.01035*
θ_{22}	−0.01278*	−0.01278*	−0.01278*	−0.01279*
θ_{23}	−0.00137	−0.00137	−0.00138	−0.00150
β_2	0.01897	0.01890	0.01941	0.03213
γ_{22}	0.02414	0.02413	0.02449	0.02456
γ_{23}	−0.02453*	−0.02451*	−0.02488*	−0.02514*
τ_3	0.00076*	0.00076*	0.00073*	0.00076*
θ_{31}	0.00174*	0.00174*	0.00172*	0.00195*
θ_{32}	0.00727*	0.00727*	0.00726*	0.00768*
θ_{33}	−0.00172*	−0.00172*	−0.00169*	−0.00201*
β_3	−0.08909*	−0.08906*	−0.08961*	−0.08112*
γ_{33}	0.08386*	0.08382*	0.08407*	0.08980*
ln L	1139.71	1139.74	1139.63	1133.77

* denotes significance at the 0.05 level, based on asymptotic t-ratios.

standard Almost Ideal specification in favor of the Rotterdam model. It could be argued, however—at least if we choose to maintain the symmetry restriction—that rejection of symmetry in the LA model lends support to our result below that the LA model should be rejected.

Specification Tests

Consider now the following compound model,

$$(10) \quad (1 - \lambda_1)\bar{s}_i \cdot \Delta\ln q_i + \lambda_1 \Delta s_i = \tau_i + \sum_{j=1}^{4} \theta_{ij} D_j + \sum_{j=1}^{4} \gamma_{ij} \Delta\ln p_j + \beta_i \Delta\ln(x/P^*)$$

a linear combination of models II and IV. If $\lambda_1 = 0$, equation (10) reduces to model II, the Rotterdam; if $\lambda_1 = 1$, equation (10) reduces to model IV, one approximation to the first-differenced LA model. A test of the hypothesis that $\lambda_1 = 0$ can be interpreted as a test of the hypothesis that the Rotterdam is the correct specification. Subject to the qualification that model IV is not exactly the first-differenced LA model, finding λ_1 to be near one would be evidence against the validity of the Rotterdam model in favor of the LA model.

The LA can be tested directly as well. In the alternative compound model

$$(11) \quad (1 - \lambda_2)\Delta s_i + \lambda_2 \bar{s}_i \cdot \Delta\ln q_i = \tau_i + \sum_{j=1}^{4} \theta_{ij} D_j + \sum_{j=1}^{4} \gamma_{ij} \Delta\ln p_j + \beta_i \Delta\ln(x/P).$$

$\lambda_2 = 0$ implies that the LA model is correct, while λ_2 near 1 is evidence against the LA in the direction of the Rotterdam. The right-hand side of (11) is based on model VI, the first-differenced LA. Testing $\lambda_2 = 0$ in (11) is a better approach to testing the LA model than is a test of $\lambda_1 = 1$ in (10), since Stone's price index is used in (11) and the LA model thus appears as the null hypothesis. As we noted above, imposing $\lambda_1 = 1$ in (10) yields model IV, not model VI, so there is a possibility of rejecting the LA simply because an approximation to Stone's price index has been used.

A Test of the Rotterdam Specification

The first compound model contains model II as a special case, that is, when λ_1 is restricted to zero. Parameter estimates and their standard errors are shown in table 3. The first two columns of results pertain to the estimated model and, for comparison, parameter estimates obtained earlier under the restriction that $\lambda_1 = 0$ are shown. The λ_1 test does not reject the Rotterdam model. In other words, imposing the Rotterdam model

Table 3. Testing the Rotterdam Specification

	Estimate	Std. error	Restr. est.
λ_1	0.057899	0.33108	0.0
τ_1	−0.00048322	0.00065433	−0.00048410
θ_{11}	0.0086721*	0.0013041	0.0086712
θ_{12}	0.0046783*	0.0011884	0.0046724
θ_{13}	0.0036476*	0.0013359	0.0036512
β_1	0.60212*	0.16885	0.63110
γ_{11}	−0.15950	0.085016	−0.17414
γ_{12}	0.15392*	0.049046	0.16215
γ_{13}	0.0035282	0.019293	0.0067507
τ_2	−0.00015426	0.00061939	−0.00014830
θ_{21}	−0.010809*	0.0012544	−0.010822
θ_{22}	−0.012981*	0.0011280	−0.012993
θ_{23}	−0.0010690	0.0013102	−0.0010504
β_2	0.28036*	0.096109	0.29642
γ_{22}	−0.16856*	0.069318	−0.18040
γ_{23}	0.0052989	0.012145	0.0071258
τ_3	0.00075212*	0.00026108	0.00075177
θ_{31}	0.0019169*	0.00052153	0.0019277
θ_{32}	0.0073322*	0.00048809	0.0073356
θ_{33}	−0.0018069*	0.00055986	−0.0018120
β_3	0.011928	0.037612	0.018133
γ_{33}	−0.010348	0.033516	−0.016134
ln L	1143.91		1143.89

* denotes significance at the 0.05 level, based on asymptotic t-ratios.

as a restriction on the compound model is supported by these data. The estimated value of λ_1 is 0.06, with a standard error of 0.33, so we cannot reject the null hypothesis that λ_1 is zero and the Rotterdam model is correct. Even if it were statistically significant, $\lambda_1 = 0.06$ would seem "close" to the Rotterdam case.

A Test of the Almost Ideal Specification

To test the validity of the LA model, equation (11) is estimated. Since the alternative model now is not exactly of the Rotterdam form—just as earlier the alternative was not exactly the LA—there seems to be no reason to expect $\lambda_2 = 1 - \lambda_1$. Now, a test of the null hypothesis that $\lambda_2 = 0$ is a test that the LA model is correct; finding evidence that λ_2 is not zero is evidence against the null hypothesis. A rejection in the direction of $\lambda_2 = 1$ can be interpreted as evidence that the Rotterdam model may be the more appropriate hypothesis.

Estimates of the second compound model are shown in table 4, along with the estimates obtained earlier for $\lambda_2 = 0$ (model VI). The λ_2 test rejects the LA model. In other words, imposing the LA model as a restriction on the compound model is not supported by these data. The estimated value of λ_2 is 0.36 with a much smaller standard error (0.10), so we can reject the null hypothesis that λ_2 is zero and that the LA model is correct.

Differences in Estimated Elasticities

A purely statistical test of one model against another is of interest for model selection. We showed that model choice matters when economic hypotheses, such as symmetry, are to be tested. It is also worth considering the extent to which conclusions about elasticities are affected by model choice. Table 5 shows the means of estimated elasticities from the Rotterdam model (model II) and the first-differenced LA model (model VI). The elasticities were calculated by assuming that the relevant income variable is expenditure on the group, as is common in demand systems in which a subset of all commodities has been treated as weakly separable. Hence, the elasticities should be interpreted with some care. There is virtually no difference between the two models in mean elasticities—a somewhat surprising result, because the models do differ in terms of consistency with the data. Apparently, observed differences between the two models are at least partly in coefficients that do not affect elasticities, such as the seasonal coefficients. We noted earlier that Brester and

Table 4. Testing the Almost Ideal Specification

	Estimate	Std. Error	Restr. Est.
λ_2	0.35997*	0.097569	0.0
τ_1	−0.00043202	0.00074397	−0.00046836
θ_{11}	0.0086097*	0.0014389	0.0083420
θ_{12}	0.0034464*	0.0013323	0.0041700
θ_{13}	0.0039673*	0.0014636	0.0040188
β_1	0.31936*	0.066130	0.11803
γ_{11}	0.010378	0.028790	0.086640
γ_{12}	0.073111*	0.020614	0.019246
γ_{13}	−0.00042506*	0.0084730	−0.055359
τ_2	−0.00023392	0.00060806	−0.00027154
θ_{21}	−0.010348*	0.0012380	−0.010351
θ_{22}	−0.013210*	0.0011394	−0.012788
θ_{23}	−0.0013726	0.0012436	−0.0015005
β_2	0.13973*	0.043035	0.032130
γ_{22}	−0.049192	0.024704	0.024563
γ_{23}	−0.015681*	0.0074318	−0.025135
τ_3	0.00059525	0.00030941	0.00075974
θ_{31}	0.0017794*	0.00062447	0.0019493
θ_{32}	0.0075707*	0.00055656	0.0076787
θ_{33}	−0.0013913*	0.00064193	−0.0020104
β_3	−0.057852*	0.017233	−0.081115
γ_{33}	0.044362*	0.013993	0.089804
ln L	1140.44		1133.77

* denotes significance at the 0.05 level, based on asymptotic t-ratios.

Table 5. Estimated Elasticities: Rotterdam and Almost Ideal Specifications

Elasticity	Rotterdam (Model II)	Almost Ideal (Model VI)
	Mean values of expenditure elasticities	
η_1	1.25	1.23
η_2	1.06	1.11
η_3	0.17	0.23
η_4	0.52	0.34
	Mean values of uncompensated price elasticities	
η_{11}	−0.98	−0.95
η_{12}	−0.03	−0.03
η_{13}	−0.12	−0.13
η_{14}	−0.12	−0.13
η_{21}	0.04	0.01
η_{22}	−0.94	−0.94
η_{23}	−0.09	−0.10
η_{24}	−0.07	−0.08
η_{31}	−0.02	−0.14
η_{32}	0.02	−0.02
η_{33}	−0.17	−0.07
η_{34}	0.00	−0.01
η_{41}	−0.21	−0.15
η_{42}	−0.04	0.01
η_{43}	−0.03	−0.02
η_{44}	−0.23	−0.18

Wohlgenant observed a similar result (i.e., only minor differences in elasticities between the LA and Rotterdam models). They used the first-differenced LA model, as opposed to one estimated in levels of shares. When Piggott, and Alston and Chalfant reported more important differences between elasticities from the two models, the results were based on LA models that were not first-differenced. Thus, it appears that the first-differenced LA model is closer to the Rotterdam model than is the LA model estimated in levels of shares.

Barten's Test

A referee has pointed out that Barten (1992) has proposed another way of testing the Rotterdam model against an Almost Ideal alternative. Barten's model is given by

$$\bar{s}_i \Delta \ln q_i = \beta_i DQ^* + \sum_{j=1}^{n} \gamma_{ij} \Delta \ln p_j + \delta_1 \bar{s}_i DQ^* - \delta_2 \bar{s}_i \Delta \ln (p_i/P)$$

where all variables are as defined above. A test of the joint hypothesis that $\delta_1 = 0$ and $\delta_2 = 0$ in this model is a test of the Rotterdam model against the general alternative. That alternative includes our model IV as a special case (but not the first-differenced LA model) when $\delta_1 = \delta_2 = 1$. As in our test of the Rotterdam model using λ_1 and (10), then, the "added model" is only approximately the LA model. It is only an approximation (*i*) because it uses the moving average of shares, rather than current shares; (*ii*) because some higher order terms are dropped in the use of a discrete approximation to the relative change in shares; and (*iii*) because the Divisia index P^*, rather than Stone's price index P, is used. Of these factors, our results suggest the most important is likely to be the use of P^* rather than P—for instance, symmetry was not rejected in model IV (using P^*) but it was rejected in model VI (using P).

Barten's form thus seems less well suited to testing the LA model than is (11), but the two tests seem likely to give similar results for most data sets. Barten's has an advantage over our approach in that the separate cases $\delta_1 = 1$, $\delta_2 = 0$ or $\delta_1 = 0$, $\delta_2 = 1$, which Barten discusses, can be considered. Our test considers the situation in which either the Almost Ideal or Rotterdam model is of interest, not an intermediate case between them.

For comparison, we estimated Barten's model and obtained results similar to those in tables 3 and 4, leading to the same conclusions. The joint hypothesis $\delta_1 = 0$ and $\delta_2 = 0$ (i.e. the Rotterdam model) was not rejected, while the hypothesis that both parameters are equal to one was rejected. Thus, even though the first-differenced LA model and the approximation using P^* (that appears as a special case in Barten's model and that we estimated as model IV) yield different conclusions about symmetry, the LA is rejected for this data set whether we use our approach or Barten's. Further experience may indicate situations in which Barten's test and ours are likely to bring different results and in which one test is preferred to the other. It seems more likely that, due to different specifications of the real income term, the alternative tests would differ more about the LA model than the Rotterdam model, since the latter appears as a special case in both approaches.

Conclusion

Although the linear approximation to the Almost Ideal Demand System has become dominant in food demand applications, recently there has been a renewal of interest in the Rotterdam model, especially in the demand for alcoholic beverages. Both models are attractive for a variety of reasons, including (local) flexibility, compatibility with demand theory, ease of use, familiarity, and plausibility. Because the two models are the most heavily used by agricultural economists, and because in some cases the choice between the two matters, it is important to be able to distinguish econometrically between them.

We have developed a test for such a purpose. In an application to the demand for meat, the results were conclusive: the test rejected the LA model but not the Rotterdam model. This is not to be interpreted as evidence that the Rotterdam model is superior in any general way. Other data sets could yield opposite conclusions, or could lead to rejecting both models or neither. Inasmuch as the test is easy to apply, we recommend it where there is interest in the LA model, the Rotterdam, or both. The same testing philosophy might be extended to other situations in demand analysis, and to other types of applications, where alternative models involve the same (or similar) right-hand sides but different dependent variables.

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[Footnotes]

[2] **Theoretical Foundations for the Rotterdam Model**
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