The Demand for Organic, Integrated-Agriculture, and Conventional Fresh Vegetables: A Censored Inverse Almost Ideal Demand System

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Abstract

The Inverse Almost Ideal Demand System is employed for the empirical analysis of the demand for organic, integrated-agriculture, and conventional fresh vegetables, using a cross section data surveyed in Rethymno, Greece during the 2005-06 period. The demand system is estimated by employing the Amemiya-Tobin model by Wales and Woodland for the estimation of censored equation systems, which ensures that the adding-up restriction is satisfied for both the latent and the observed expenditure shares. The problem regarding the logarithm of quantities when zero purchases are reported, is resolved in a theoretically consistent way that allows full-sample estimation and yields unbiased parameter estimates. The empirical results suggest that integrated-agriculture fresh vegetables are luxury goods, whereas the cross-quantity uncompensated flexibilities indicate that consumers are not regular buyers of any of the three types of fresh vegetables. Both groups of consumers who currently buy integrated-agriculture vegetables and those who buy conventional vegetables can be easily induced to buy organic vegetables.

Introduction

Organic agriculture can be defined as a production system based on locally or farm-produced renewable inputs in preference to external ones, and aims to promote and enhance ecosystem health (FAO). The importance of the organic farming sector in the European Union is reflected in the recent reforms of the European Union's (EU) Common Agricultural Policy (CAP), and in respective Regulations. Since the 1992 CAP Reform, organic farming has been assigned an important role in the enhancement of environmental protection throughout the EU, while EU Regulations 2078/92 and 2092/91 provided specific incentives for conversion to and maintenance of organic farming and established organic products as distinctively different from their conventional counterparts (provision of standards and certification). However, higher prices received, is the most important incentive for farmers to convert to organic agriculture (Burton, Rigby, and Young, 1999, 2003; O'Riordan and Cobb). Farmers receive higher prices for organic products when consumers believe that there is a quality premium available in organic product's attributes (Loureiro, McCluskey, and Mittelhammer; Boland and Schroeder). On the consumer's side, the demand for organic products in Europe exceeds supply, despite that, on average, the price of organic products is twice that of conventionally grown food (Sylvander and Le Floc'h Wadel).

Although consumers' acceptance of organic products is vital for the growth of the organic farming sector, most studies on consumer demand examine consumer attitudes, identify their motivation for purchasing organic products, and elicit willingness to pay for organic products relative to their conventional counterparts. On the other hand, there is only one empirical study for organic products that employs traditional demand analysis in order to provide estimates of consumers' responsiveness to price changes. Specifically, Thompson and Kidwell collected data on prices and cosmetic effects for five organic and conventional produce items, as well on consumers' socio-economic and demographic characteristics, in order to estimate the choice between organic and conventional produce, and the choice between two stores for shopping. Using a two-equation probit model, Thompson and Kidwell provided estimates of the elasticities (or conditional effects) of the goods' defects, difference in prices, type of store, and socio-economic and demographic characteristics. However, due to the lack of data on expenditures on organic and conventional produce, their resulting econometric model did not allow the derivation of measures of the interrelationships between organic and conventional produce (*i.e.*, cross-price elasticities).

The limited number of empirical studies on consumer demand for organic products is due to the fact that organic products are relatively new compared to their conventional counterparts, and therefore to the paucity of sufficient historical data on retail prices and consumption. In this context, the aim of the present paper is to provide empirical evidence of the consumption of both organic and integrated-agriculture fresh vegetables using a cross-section of data[1]. Specifically, the Inverse Almost Ideal Demand System (IAIDS) of Eales and Unnevehr, and Moschini and Vissa is employed for the empirical analysis of household demand for organic, integrated-agriculture and conventional fresh vegetables in Crete, Greece. The choice of an inverse instead of a direct demand system, apart of the lack of sufficient time-series data on organic consumption, is based on the nature of the goods in question. Inverse demand systems are often employed in the case of quickly perishable foods, agricultural, and fishery products for which quantities cannot adjust in the short-run. The underlying assumption is that since supply of such commodities may be fixed during short-intervals, price must adjust so that the available quantity is consumed. Fresh vegetables are produced subject to biological lags, they are quickly perishable commodities and cannot be stored. As a result the supply of fresh vegetables is highly inelastic during short-intervals. It is, therefore, more reasonable to employ an inverse demand system; that is, a demand system for which quantities are taken as predetermined (*i.e.*, exogenous) while it is prices that adjust so that the available quantity is consumed.

The use of cross-section data in our analysis, however, is not without complications. It is common in micro-level analyses of consumer demand for many households to report zero purchases of certain commodities. The presence of these zero observations gives rise to two problems. Firstly, in the case of an inverse demand system, such as the IAIDS, where expenditure shares are functions of the logarithm of the quantities purchased, the logarithm of zero purchases, when reported, cannot be defined. One way to deal with this problem would be to estimate the IAIDS system only for the households that report positive purchases, but this may result in sample selection bias. Another way is full-sample estimation by assigning some small positive number or unity to the zero purchases. However, this approach has also serious drawbacks: it is not independent of the units of measurement of the respective explanatory variable(s), and if there is a large number of households in the sample that report zero purchases then the resulting parameter estimates may be biased. In order to overcome this problem, we employ an approach which allows full-sample estimation and results in efficient and unbiased estimates.

The second problem related to the presence of zero purchases is that standard systems estimation methods, e.g., seemingly unrelated regression or maximum likelihood, lead to biased parameter estimates. Two main approaches have emerged in the literature for the estimation of micro-level demand systems. The first approach is the Kuhn-Tucker model of Wales and Woodland, and its dual model proposed by Lee and Pitt. The Kuhn-Tucker model of Wales and Woodland assumes that preferences are random over the population. It starts with the maximisation of a random direct utility function, subject to budget and non-negativity constraints, and then, the standard Kuhn-Tucker conditions are used for the derivation of the demand equations. Lee and Pitt extended the Kuhn-Tucker approach to a dual form, taking the maximisation of the indirect utility function as a starting point and using Roy's identity for the derivation of the demand equations. Their approach is based on the use of virtual (reservation) price relationships – which are shown to be dual to the Kuhn-Tucker conditions – in order to identify corner solutions and define demand regime switching.

The second approach for the estimation of censored demand systems was proposed also by Wales and Woodland and is a non-trivial modification of Amemiya's extension of the tobit model (Tobin) for a system of equations. This model assumes that preferences are non-random and the non-negativity restriction for the observed shares is incorporated by assuming that the observed expenditure shares are the sum of the utility maximizing shares (the latent shares) and a random disturbance term which follows a truncated normal distribution. In the Amemiya-Tobin model, the adding-up constraints hold for the latent expenditure shares but not for the observed (*i.e.*, censored) expenditure shares. In order to impose adding-up for the observed shares, Wales and Woodland proposed a mapping of the latent to the observed shares which specifies each positive observed expenditure share as the ratio of the respective latent share to the sum of the positive latent shares. The model then generates a density for expenditure shares which has the form of a partially-integrated mixed discrete-continuous multivariate distribution, *i.e.*, it is a continuous *pdf* with respect to the positive observed shares and a discrete probability mass with respect to the zero observed shares.

The Amemiya-Tobin model by Wales and Woodland is the approach adopted in the present paper to account for the presence of zero purchases in our sample. This approach has also been employed by Dong, Gould, and Keiser, for the estimation of Mexican household demand for 12 food categories.[2] The advantage of this model over the Kuhn-Tucker model and its dual lies in that it can be applied in any demand system specification. On the contrary, the applicability of the latter models is quite limited as it is difficult to solve the Kuhn-Tucker conditions or the virtual price relationships for direct or indirect utility functions underlying many demand systems, such as the IAIDS. A difficulty in the application of the Amemiya-Tobin model of Wales and Woodland (but also of the Kuhn-Tucker model and its dual) lies in the requirement for evaluation of multiple probability integrals in the likelihood function, a task that is difficult when there are many goods in the demand system. The use of two-step estimators instead, which offer simplified procedures for the estimation of censored demand systems (e.g., Heien and Wessells; Shonkwiler and Yen; Perali and Chavas), results in parameter estimates that, although consistent, lack in efficiency relative to the maximum likelihood estimators of Wales and Woodland, and Lee and Pitt. Moreover, the problem of adding-up of the observed shares is not adequately addressed.[3]

The rest of the paper as organized as follows. In the following section, the IAIDS model is presented and discussed. Then, we illustrate and discuss the approach adopted for tackling the problem that arises when the explanatory variables in the IAIDS model, which are expressed in logarithms, can also take on zero values. In the same section, the Amemiya-Tobin model by Wales and Woodland, and the methodology of derivation of expected expenditure shares needed for the computation of flexibilities are also presented. Description of the data and descriptive statistics are presented in the following section. The empirical results are then presented and discussed, and the last section summarises and concludes the paper.

Theoretical Model

As mentioned at the outset, the present empirical study employs the IAIDS model of Eales and Unnevehr, and Moschini and Vissa for the estimation of demand conditions for organic, integrated-agriculture and conventional fresh vegetables[4]. The IAIDS model is derived in a manner similar to that of the Almost Ideal Demand System (AIDS) of Deaton and Muellbuer by employing a logarithmic distance function to represent preferences (instead of the log-cost function of the AIDS model). The distance function is an alternative representation of preferences and allows the derivation of theoretically consistent inverse demands. It is a scalar measure of the amount by which an arbitrary quantity vector *q* must be scaled up or down, along a ray radiating from the origin, to that quantity vector that just attains a target utility level *u*, *i.e.*,

$$D\left(u,\mathbf{q}\right) \equiv \max_{\delta>0}\left\{\delta \,\middle|\, u\left(\mathbf{q}/\delta\right) \geq u\right\} \tag{1}$$

where *u* is a utility level, **q** is an arbitrary vector of quantities consumed, and *δ* is a scalar. The distance function is dual to the cost function and possesses mathematical properties with respect to quantities that are the same as (reciprocal to) the properties of the cost function with respect to prices (utility); *i.e.*, it is linear homogeneous in quantities, concave and non-decreasing in quantities, and decreasing in utility (Cornes).

Deaton and Muellbauer used a cost function representing PIGLOG preferences as a starting point for the derivation of the Almost Ideal Demand System (direct AIDS). Eales and Unnevehr applied the PIGLOG parameterization of the cost function to the distance function, and specified the following logarithmic distance function as a starting point for the derivation of the IAIDS model.

$$ln\,D(u,\mathbf{q}) = (1-u)\,ln\,\alpha(\mathbf{q}) + u\,ln\,b(\mathbf{q}) \tag{2}$$

where $ln\,\alpha(\mathbf{q})$ and $ln\,b(\mathbf{q})$ are specified in a manner analogous to that employed in the AIDS, *i.e.*,

$$ln\,\alpha(\mathbf{q}) = \alpha_0 + \sum_j \alpha_j\, ln\,q_j + 0.5\sum_i\sum_j \gamma_{ij}^*\, ln\,q_i\, ln\,q_j \tag{3.a}$$

$$ln\,b(\mathbf{q}) = \beta_0 \prod_j q_j^{-\beta_j} + ln\,\alpha(\mathbf{q}) \tag{3.b}$$

where $i, j = 1,\ldots,N$ indicates commodities consumed. Hence, the IAIDS logarithmic distance function is written as:[5]

$$ln\,D(u,\mathbf{q}) = \alpha_0 + \sum_j \alpha_j\, ln\,q_j + 0.5\sum_i\sum_j \gamma_{ij}^*\, ln\,q_i\, ln\,q_j + u\beta_0 \prod_j q_j^{-\beta_j} \tag{4}$$

Application of the Shephard-Hanoch Lemma to relation (4) yields the compensated inverse demands for good *j*. Inversion of the distance function at the optimum yields the direct utility function which may be used to uncompensate these inverse demand equations. Then the following system of inverse demands is derived:

$$w_i = \alpha_i + \sum_j \gamma_{ij}\, ln\,q_j + \beta_i\, ln\,Q \tag{5}$$

where $\gamma_{ij} = (1/2)\left(\gamma_{ij}^* + \gamma_{ji}^*\right)$, and Q is a quantity index defined by

$$lnQ = \alpha_0 + \sum_j \alpha_j \, ln q_j + 0.5 \sum_i \sum_j \gamma_{ij} \, ln q_i \, ln q_j \qquad (6)$$

The above system is non-linear in its parameters. A linear approximation to the IAIDS model is derived by replacing the IAIDS quantity index (6), with a quantity index analogous to a log-linear analogue to the *Laspeyres* price index[6]:

$$lnQ = \sum_j w_{j0} \, ln q_j \qquad (7)$$

where w_{j0} is the mean expenditure share for good *j*.

In order to account for household heterogeneity, variables involving household characteristics must also be included in the model. A common way to include socio-demographic variables which preserves the linearity in the model given by equations (5) and (7) is by augmenting the α_i terms so that

$$\alpha_i = \alpha_{i0} + \sum_k \zeta_{ik} Z_k \qquad k = 1, \ldots, K \qquad (8)$$

Restrictions similar to that for the direct AIDS share equations apply also to the IAIDS; the system represented by relations (5) and (7) must add-up to total expenditure, be homogeneous of degree zero in quantities and satisfy symmetry. These imply adding-up restrictions $\sum_i \alpha_i = 1$, $\sum_i \gamma_{ij} = 0$, $\sum_i \beta_i = 0$, and $\sum_i \zeta_{i\kappa} = 0$; homogeneity restrictions $\sum_j \gamma_{ij} = 0$; and symmetry restrictions $\gamma_{ij} = \gamma_{ji}$ on the parameters of the IAIDS model. However, the adding-up restrictions ensure that the budget constraint is satisfied only in the absence of censoring. The problem of adding-up in a censored demand system is dealt with in the next section.

Empirical Model

As mentioned in the introductory section, a common problem encountered by the use of the IAIDS system in cross section data of consumer demand is that the logarithm of zero consumption cannot be defined. To overcome this problem we redefine the IAIDS by including a 0-1 dummy variable indicating non-zero consumption of good *j*.[7] Consider an IAIDS model for *N* goods and *T* households. Suppose that the first *s* households buy all the N goods, while T-s households buy only the first N-1 goods and report zero purchases for the Nth good. If the quantity index given by equation (7) is used instead of the true IAIDS index, then the IAIDS expenditure share equation is given by:

$$s \text{ households: } w_i = \alpha_i + \sum_{j=1}^{N-1} \gamma_{ij}\, ln\, q_j + \gamma_{iN}\, ln\, q_N + \beta_i \left(\sum_{j=1}^{N-1} w_{j0} lnq_j + w_{N0} lnq_N \right) \quad (9)$$

$$T\text{-}s \text{ households: } w_i = \tilde{\alpha}_i + \sum_{j=1}^{N-1} \gamma_{ij}\, ln\, q_j + \beta_i \sum_{j=1}^{N-1} w_{j0}\, ln\, q_j \quad (10)$$

In the above equations, the quantity coefficients are the same both for the s and the T-s households, but the constant parameters are not necessarily the same. However, as it is, the model given by equations (9) and (10) represents two demand systems. To deal with this, a 0-1 dummy variable is included in the model, and the explanatory variable for the N^{th} good, q_N, is redefined so that

$$\text{if } q_N > 0 : D_N = 0 \quad \text{and} \quad lnq_N^* = ln\left(max\left\{q_N, D_N\right\}\right) = lnq_N$$

$$\text{if } q_N = 0 : D_N = 1 \quad \text{and} \quad lnq_N^* = ln\left(max\left\{q_N, D_N\right\}\right) = ln(1) = 0$$

Hence, if household behaviour is described by equations (9) and (10), the model to be estimated is:

$$\text{All } T \text{ households} \quad \begin{aligned} w_i &= \alpha_i + \left(\tilde{\alpha}_i - \alpha_i\right) D_N + \sum_{j=1}^{N-1} \gamma_{ij}\, ln\, q_j + \gamma_{iN}\, ln\, q_N^* \\ &+ \beta_i \left(\sum_{j=1}^{N-1} w_j lnq_j + w_{N0} lnq_N^* \right) \end{aligned} \quad (11)$$

Thus, if $q_N > 0$, then $D_N = 0$ and $lnq_N^* = lnq_N$, and the expenditure share equation given by (11) is equivalent to the one given by equation (9). On the other hand, if $q_N = 0$, then $D_N = 1$ and $lnq_N^* = 0$, and the expenditure share equation (11) is equivalent to the one given by equation (10). In conclusion, in the case of an inverse demand system where the explanatory variables involving purchased quantities are expressed in logarithms, this technique is very convenient. It allows the same system of expenditure share equations to be estimated for all households, either reporting zero purchases or reporting positive purchases, without resulting in biased estimates. Moreover, the additional constant parameter introduced by the dummy variable whenever zero purchases are reported acts only as a demand shifter, leaving the slope of the demand functions and the own- and cross-quantity flexibilities unaffected.

A second problem that needs to be addressed in the presence of zero purchases, irrespectively to which demand system (direct or inverse) is adopted, is that of censoring of the expenditure shares. In the presence of censoring, if standard system estimation methods are employed, the parameter estimates will be biased. To account for the presence of censoring, we employ the Amemiya-Tobin approach by Wales and Woodland for the estimation of censored demand systems. In order to present this approach, we will use matrix notation and define the IAID system of latent expenditure share equations as:

$$\mathbf{w}^* = \mathbf{xb} + \boldsymbol{\varepsilon} \qquad \boldsymbol{\varepsilon} \mid \mathbf{x} \sim N(\mathbf{0}, \boldsymbol{\Sigma}) \tag{12}$$

where $\mathbf{w}^*$ is a vector of the IAIDS latent expenditure shares, $\mathbf{x}$ is a vector of the IAIDS explanatory variables, $\mathbf{b}$ is a vector all demand parameters, and $\boldsymbol{\varepsilon}$ is a vector of error terms assumed to be distributed as multivariate normal with mean $\mathbf{0}$ and variance-covariance matrix $\boldsymbol{\Sigma}$. The adding-up restrictions for the IAIDS parameters imply that the latent expenditure shares satisfy the budget constraint, that is, $\sum_i w_i^* = 1$ and $\sum_i \varepsilon_i = 0$. As a consequence, the joint normal density function of the latent expenditure shares is degenerate and one of them may be dropped, during estimation, as redundant.

However, one must ensure that the observed (i.e., censored) expenditure shares, $\mathbf{w}$, also satisfy the budget constraint. To ensure this, Wales and Woodland used the following mapping of the latent to the observed expenditure shares:

$$w_i = \frac{w_i^*}{\sum_{j \in \mathbf{J}} w_j^*} \qquad \text{if } w_i^* > 0 \tag{13.a}$$

$$w_i = 0 \qquad \text{if } w_i^* \le 0 \tag{13.b}$$

where w_i^* and w_i are the latent and observed expenditure shares of good *i*, respectively, and **J** is the set of positive latent expenditure shares. Equation (13.a) defines each observed expenditure share as the ratio of the respective latent share to the sum the positive latent expenditure shares, thus forcing the observed expenditure shares to add-up to unity. Assuming that the first *m* goods are purchased, while consumption for the remaining *N*-*m* goods is zero, and dropping one of the latent shares as redundant, say, the last of the positive shares, w_N^*, Wales and Woodland provide the following expression for the probability density/mass of $\mathbf{w} = \left(w_1,\ldots,w_m,0,\ldots,0\right)$:

$$f\left(w_1,\ldots,w_m,w_{m+1},\ldots,w_N;\mathbf{xb},\mathbf{\Sigma}\right)$$
$$= \begin{cases} \int_{w_1}^{\infty}\int_{\alpha_{m+1}}^{0}\cdots\int_{\alpha_{N-1}}^{0} g^*\left(w_1^*,\frac{w_1^* w_2}{w_1},\ldots,\frac{w_1^* w_m}{w_1},w_{m+1}^*,\right. & \\ \left.\ldots,w_{N-1}^*;\mathbf{xb},\mathbf{\Sigma}\right) J\left(\mathbf{w}\right) dw_{N-1}^*\ldots dw_{m+1}^* dw_1^* & 1 \le m < N \\ g^*\left(w_1^*,\ldots\ldots,w_{N-1}^*\right) & m = N \end{cases} \tag{14}$$

where $g^*\left(\cdot\right)$ is the conditional joint normal p.d.f. of the first *N*-1 latent shares w_i^*, and

$$J\left(\mathbf{w}\right) = \left[1 + \left(w_2/w_1\right)^2 + \ldots + \left(w_m/w_1\right)^2\right]^{1/2} \tag{15}$$

$$a_{m+1} = 1 - \sum_{i=1}^{m} w_1^* \frac{w_i}{w_1} \tag{16}$$

$$a_l = a_{m+1} - \sum_{i=m+1}^{l-1} w_i^* \qquad l = m+2,\ldots,N-1 \tag{17}$$

The term $J(\mathbf{w})$ is the Jacobian between the latent and the observed shares and can be ignored in the likelihood function as it does not depend on the parameters of the model (Wales and Woodland). The log-likelihood for this model is the sum of the logarithm of expressions in (14) over all households, where each household is associated with one of the two expressions.

Once the model given by equations (12), (13.a), and (13.b) is estimated, the economic interpretation of the parameter estimates requires computation of the consumption scale, uncompensated, and compensated (*Antonelli*) flexibilities, based on the unconditional expectation of the observed shares. In order to derive the expected observed shares, we follow the simulation procedure suggested by Dong, Gould, and Keiser: we start by simulating the error terms in (12), then substitute the simulated error terms in (12) to derive the simulated latent shares, and, finally, we use the simulated latent shares and the mapping rule of Wales and Woodland in order to compute the expected observed shares. Specifically, the procedure of Dong, Gould and Keiser starts with the estimation of *R* replicates of the (Nx1) error term, $\boldsymbol{\varepsilon}$ in (12). The simulated latent shares are then calculated by

$$\mathbf{w}_r^* = \overline{\mathbf{x}}\tilde{\mathbf{b}} + \boldsymbol{\varepsilon}_r \qquad (18)$$

where $\overline{\mathbf{x}}$ is the vector of the exogenous variables evaluated at the sample means, $\tilde{\mathbf{b}}$ is the vector of the maximum likelihood parameter estimates, and $\boldsymbol{\varepsilon}_r$ is the *r*th replicate of the error term vector, $\boldsymbol{\varepsilon}$. The *r*th replicate of the *i*th observed share is derived according to the mapping rule of Wales and Woodland, i.e.,

$$w_{ir} = \frac{w_{ir}^*}{\sum_{j \in J} w_{jr}^*} \qquad \text{if } w_{ir}^* > 0 \qquad (19.a)$$

$$w_{ir} = 0 \qquad \text{if } w_{ir}^* \leq 0 \qquad (19.b)$$

Hence, the expected value of the *i*th observed share can be computed as the average of its *R* replicates, that is,

$$E(w_i) = \frac{1}{R}\sum_{r=1}^{R} w_{ir} \qquad (20)$$

Dong, Gould, and Keiser substitute the expected observed shares in an arch elasticity formula in order to calculate the simulated uncompensated elasticities for the direct AIDS. In the present paper, we used the IAIDS system as a starting point and derived the following formula for the calculation of the simulated uncompensated flexibilities:

$$f_{ij} = -\delta_{ij} + \frac{1}{E(w_i)}\frac{1}{R}$$

$$\sum_{r=1}^{h}\left\{\frac{\left[\left(\gamma_{ij} + \beta_i\left(w_{jr}^* - \beta_j \, lnQ\right)\right)\left(1 - \sum_l w_{lr}^*\right) - w_{ir}^* \sum_m \left(\gamma_{mj} + \beta_m\left(w_{jr}^* - \beta_j \, lnQ\right)\right)\right]}{\left(\sum_m w_{mr}^*\right)^2}\right\} \quad (21.a)$$

where $i, j = 1,\ldots,N$ indicate goods in the demand system; $r = 1,\ldots,h$ indicates replications for which the simulated latent share of good *i*, w_{ir}^*, is positive; *m* indicates goods the simulated latent share of which is positive; *l* indicates goods the simulated latent share of which is non-positive; δ_{ij} is the *Kronecker delta* ($\delta_{ij} = 1$ for $i = j$; $\delta_{ij} = 0$ for $i \neq j$); $E(w_i) = (1/R)\sum_{r=1}^{h}\left(w_{ir}^* / \sum_m w_{mr}^*\right)$; *Q* is the quantity index defined by relation (7) and; β_j, and γ_{ij} are the maximum likelihood estimates of the IAIDS parameters[8]. The simulated consumption scale flexibilities, f_i, and the simulated compensated flexibilities, f_{ij}^*, can be computed using the following relations, respectively[9]

$$f_i = \sum_j f_{ij} \quad (21.b)$$

$$f_{ij}^* = f_{ij} - E(w_j) f_i \quad (21.c)$$

Data

The present empirical analysis employs a IAIDS system for three fresh vegetable groups: (*a*) organic fresh vegetables, (*b*) integrated-agriculture fresh vegetables, and (*c*) conventional fresh vegetables. To this aim, cross-sectional data were collected on quantities purchased, retail prices and several socio-demographic characteristics, via questionnaires answered by consumers of fresh vegetables, in Rethymno, Greece. The collection of the data took place at several super-markets and grocery stores selling both conventional and organic and/or integrated-agriculture fresh vegetables. In total 171 questionnaires were randomly collected and used in the present analysis.

Each of the three fresh vegetable groups in the demand system includes fresh vegetables that can be easily found in Rethymno both as conventional ones and as organic and/or integrated-agriculture ones (tomatoes, cucumbers, and sweet peppers), and other fresh vegetables. Aggregation of the quantities included in these three groups was done using *Divisia* indices with expenditure shares serving as weights. Quantities are measured in kilograms and refer to two-weeks figures. Retail prices were collected on the spot and are in Euros per kilogram. Finally, two variables involving household characteristics were chosen to be included in our analysis: one demographic variable namely, the household size, and a 0-1 dummy for the existence of information about organic vegetables.

The descriptive statistics for the household data are summarized in Tables 1 and 2. Table 1 focuses on the mean and standard deviation of the expenditure shares and quantities consumed, for the three fresh vegetable groups, and of the household characteristics included in our analysis. Conventional fresh vegetables are associated with the highest average share, as 121 households (71%) in our sample reported consumption of this commodity group, whereas 85 (50%) and 53 (31.0%) households reported consumption of organic and integrated-agriculture fresh vegetables, respectively. Finally, 106 households in our sample reported that they were informed about organic vegetables, while 65 households reported otherwise. Table 2 characterises the purchase patterns in our data: the majority of the households in our sample (105 households (61.4%)) purchased only one type of fresh vegetables, while 44 households (25.7%) purchased only two types of fresh vegetables and 22 households (12.9%) purchased all three types of fresh vegetables.

Empirical Results

The IAIDS model for 3 fresh vegetable categories and 2 socio-demographic variables was modified so that the problem of the logarithm of explanatory variables taking the value of zero is accounted for. As zero purchases have been reported for all the three commodities, three dummy variables taking the values of zero and one were included in our model. Thus, the final IAID system of latent expenditure share equations is given by:

$$w_i^* = \alpha_{i0} + \sum_j \delta_{ij} D_j + \sum_j \gamma_{ij} \, ln \, q_j^* + \beta_i \, ln \, Q + \sum_k \zeta_{ik} Z_k + \varepsilon_i , \tag{22}$$

where

$$ln \, Q = \sum_j w_{j0} \, ln \, q_j^* , \tag{23}$$

D_j indicates commodity-specific dummies taking the value of zero if $q_j > 0$ and the value of one if $q_j = 0$, $lnq_j^* = ln\left(\max\left\{q_j, D_j\right\}\right)$, and Z_k indicates variables involving household characteristics, namely, household size, and existence of information about organic vegetables. The introduction of commodity-specific dummies and the budget constraint now require that the adding-up restriction $\sum_i \delta_{ij} = 0$ be imposed on the δ_{ij} parameters.

The parameters of the IAIDS model were estimated using the GAUSS software system, with the adding-up, homogeneity and symmetry restrictions imposed. The maximum likelihood parameter estimates along with standard errors are displayed in Table 3[10]. In total, 32 out of the 36 parameters were statistically significant. Specifically, all the own- and cross-quantity coefficients, γ_{ij}, were statistically significant at the 1% level of significance. In addition, two out of the three total consumption coefficients, β_i, were found to be statistically different from zero at the 1% and 5% level of significance, respectively.

Interpretation of the γ_{ij} and β_i parameter estimates is provided by the (simulated) consumption scale, uncompensated and compensated flexibilities presented in Tables 4-6. The scale flexibilities, shown in Table 4, measure the change in the expenditure-normalised price of good *i* (i.e., the consumer's marginal valuation of good *i*) in response to a scale expansion in the consumption bundle. For example, the scale flexibility for organic vegetables is -1.0111 which indicates that if the consumption of all vegetables is increased by 1% then the expenditure-normalised price of organic vegetables would be decreased by about 1.01%. Scale flexibilities can be used to classify goods as luxuries or necessities. In our case, integrated-agriculture vegetables are classified as luxuries $(f_i > -1)$, while organic and conventional vegetables are necessities $(f_i < -1)$.

The compensated (*Antonelli*) flexibilities are reported in Table 6. As Barten and Bettendorf point out, compensated flexibilities are imperfect measures of the interrelationships between goods because the negative-semidefiniteness of the *Antonelli* matrix and the homogeneity restriction lead to dominance of complementarity in the *Antonelli* matrix (*i.e.* dominance of positive cross-quantity effects). Here, all cross-quantity compensated flexibilities are positive, implying that organic, integrated-agriculture and conventional vegetables are *net q-complements*, and not *net q-substitutes* as was *a priori* expected[11]. For this reason, we will not discuss the compensated flexibilities, and turn to the analysis of uncompensated flexibilities instead. As shown in Table 5, all own-quantity uncompensated flexibilities are lower than minus one. Small responses of normalised prices to own-quantity changes suggest that, in terms of a direct demand system, the goods in the consumption bundle are price elastic. A good reason for this is that the goods in question are very good substitutes for one another, and, consequently, the demand for a certain good is commonly quite elastic.

All cross-quantity uncompensated flexibilities are negative indicating that the goods in the system are *gross q-substitutes*, as expected. As far as the substitution effects between organic and integrated-agriculture vegetables are concerned, it would take a decrease of 0.12% in the expenditure-normalised price of organic vegetables to induce consumers to absorb a 1% increase in the quantity of integrated-agriculture vegetables. On the other hand, a marginal increase in the consumption of organic vegetables requires that the expenditure-normalised price of integrated-agriculture vegetables be decreased by 0.19%. This suggests that it is easier to induce consumers currently buying organic vegetables to "revert" to integrated-agriculture vegetables, but it is relatively difficult to induce consumers currently buying integrated-agriculture vegetables to "revert" to organic vegetables. Organic vegetables are more expensive than integrated-agriculture vegetables. Given the consumers' budget, it will take a relatively higher decrease in the expenditure-normalised price of integrated-agriculture vegetables for the consumers of these products to be able to save money and buy the (relatively more expensive) organic vegetables. On the other hand, consumers currently buying organic vegetables can be easier induced to buy integrated-agriculture vegetables. It could be that the higher price of organic vegetables (compared to the integrated-agriculture vegetables) acts, from the consumers' point of view, as an indicator that organic vegetables are "better" than the integrated-agriculture vegetables. Thus, if prices of organic vegetables decline to reach those of integrated-agriculture vegetables, integrated-agriculture vegetables might be though of to be just as good as the organic ones.

Regarding the substitution effects between conventional and integrated-agriculture vegetables, the response of the expenditure-normalised price of conventional vegetables to a marginal increase in the consumption of integrated-agriculture vegetables is -0.08%. This is lower, in absolute value, than the response of the expenditure-normalised price of integrated-agriculture vegetables to a marginal increase in the consumption of conventional vegetables (-0.18%). Hence, the magnitude of these two cross-quantity flexibilities suggests that, despite the higher price of integrated agriculture vegetables, it is it is relatively easier to induce consumers currently buying conventional vegetables to buy integrated-agriculture vegetables than it is to induce consumers currently buying integrated-agriculture vegetables to buy conventional vegetables. A reason for this could be that the higher price of integrated-agriculture vegetables (compared to the conventional vegetables) acts, from the consumers' point of view, as an indicator that integrated-agriculture vegetables are "better" than the conventional vegetables. Thus, if prices of conventional vegetables decline, the difference between the prices of the two products will increase, and consumers currently buying conventional vegetables might think of it as a degradation in their quality, and "revert" to integrated-agriculture vegetables. Finally, as far as the interrelationships between organic and conventional vegetables are concerned, despite the higher price of organic vegetables, it is relatively more easy to induce consumer who currently buy conventional vegetable to buy organic ones, than it is to induce consumers currently buying organic vegetables to buy conventional ones

In total, the small magnitude of the cross-quantity uncompensated flexibilities indicates that there are high substitution possibilities between different types of fresh vegetables, that is, the consumers in our sample are not regular buyers of any of the three types of fresh vegetables examined here. In particular, as far as the local organic-farming sector is concerned, expansion of this sector may increase competition among producers and lead to price reductions that will induce consumers to absorb the available quantity.

Summary and Conclusions

The present study aims to contribute to the literature of empirical studies of consumer demand for organic products. Specifically, cross-sectional data were employed for the estimation of the demand for organic, integrated-agriculture and conventional fresh vegetables. Since quantities of goods, such as fresh vegetables, can be fixed in the short-run due to biological production lags and product perishability, it is more reasonable to employ an inverse demand system. In particular, the Inverse Almost Ideal Demand System of Eales and Unnevehr, and Moschini and Vissa was employed. In addition, the Amemiya-Tobin model by Wales and Woodland was used for accounting for the occurrence of zero purchases in our sample, and the censored IAIDS system was estimated using maximum likelihood.

Another contribution of the present paper lies in accounting, in a theoretically consistent way, for the problem introduced in inverse demand systems in which consumed quantities are expressed in logarithms, when zero consumption is reported. Our censored inverse demand system was modified by using an approach which has been applied in the empirical production analysis literature, and which allows full-sample estimation and derivation of efficient and unbiased parameter estimates.

Expected observed shares were computed via simulations and used for the computation of flexibilities. Scale flexibilities suggest that integrated-agriculture fresh vegetables are luxury goods, while organic and conventional fresh vegetables are necessities. The small magnitude of the own-quantity uncompensated flexibilities implies that the goods in our demand system are own-quantity inflexible, and, hence, own-price elastic, a result explained by their being very good substitutes for one another. Finally, cross-quantity uncompensated flexibilities indicate that increases in the supplied quantity of organic vegetables, can be absorbed by both consumers who currently buy integrated-agriculture vegetables and those who buy conventional vegetables, suggesting that expansion of the local organic-farming market can be successful.

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Table 1. Summary Statistics of the Data.

Variable		Households Consuming	Mean	Standard Deviation
Expenditure Shares (in %)				
Organic vegetables		0.50	0.33	0.39
Households consuming			0.67	0.28
Integrated vegetables		0.31	0.19	0.35
Households consuming			0.61	0.38
Conventional vegetables		0.71	0.48	0.43
Households consuming			0.68	0.35
Quantities Purchased (in kgs.)				
Organic vegetables			1.21	1.79
Households consuming			2.44	1.86
Integrated vegetables			0.72	1.45
Households consuming			2.32	1.75
Conventional vegetables			2.22	2.36
Households consuming			3.14	2.23
Household Characteristics				
Household size (no. of persons)			2.96	1.40
Information about Organics				
Yes (no. of households)	106			
No (no. of households)	65			

Table 2. Distribution of Purchases

Vegetables Purchased	No of Households	% of Households
Only organic vegetables	25	14.6
Only integrated vegetables	21	12.3
Only conventional vegetables	59	34.5
Both organic and integrated vegetables	4	2.3
Both organic and conventional vegetables	34	19.9
Both integrated and conventional vegetables	6	3.5
All vegetables purchased	22	12.9
Total	171	100.0

Table 3. Parameter Estimates of the IAIDS for Fresh Vegetables in Crete, Greece.

	Organic	Integrated	Conventional
Intercept (α_{i0})	0.5065 (0.0214)*	0.0481 (0.0242)**	0.4454 (0.0228)*
Household characteristics (ζ_{ik})			
Household size	0.0463 (0.0097)*	-0.0622 (0.0099)*	0.0160 (0.0136)
Information	-0.1076 (0.0116)*	0.1512 (0.0103)*	-0.0435 (0.0156)*
Quantities (γ_{ij})			
Organic	0.2250 (0.0060)*		
Integrated	-0.1149 (0.0070)*	0.2249 (0.0134)*	
Conventional	-0.1101 (0.0069)*	-0.1100 (0.0093)*	0.2202 (0.0111)*
Total consumption (β_i)	-0.0164 (0.0163)	0.0625 (0.0205)*	-0.0461 (0.0211)**
Commodity specific dummies (δ_{ij})			
Organic	-0.6788 (0.0155)*	0.1278 (0.0180)*	0.3344 (0.0093)*
Integrated	0.1674 (0.0176)*	-0.4528 (0.0197)*	0.2289 (0.0160)*
Conventional	0.5114 (0.0287)*	0.3250 (0.0245)*	-0.5633 (0.0193)*
Error correlation (ρ_{ij})			
Integrated	-0.5836 (0.0788)*		
Conventional	-0.9989 (0.0005)*	-0.8597 (0.1271)*	

Note: Asymptotic standard errors in parentheses. *(**) Indicate significance level at the 1(5) percent.

Table 4. Simulated Consumption Scale Flexibilities.

Vegetables	Scale Flexibilities
Organic	-1.0111
Integrated	-0.8814
Conventional	-1.0308

Table 5. Simulated Uncompensated Flexibilities.

Vegetables	Organic	Integrated	Conventional
Organic	-0.7136	-0.1243	-0.1732
Integrated	-0.1855	-0.5203	-0.1756
Conventional	-0.1170	-0.0759	-0.8379

Table 6. Simulated Compensated (Antonelli) Flexibilities.

Vegetables	Organic	Integrated	Conventional
Organic	-0.3921	0.0418	0.3503
Integrated	0.0947	-0.3754	0.2807
Conventional	0.2107	0.0935	-0.3042

Endnotes

[1] Integrated-agriculture vegetables are produced with the use of restricted amount of chemicals and pesticides. As such, integrated-agriculture vegetables are different from organic and conventional vegetables, since they are not free from pesticides and chemicals, nor is the use of the latter unrestricted.

[2] The demand system adopted by Dong, Gould, and Keiser is a direct AIDS, and the problem of estimating a demand system with a large number of goods was dealt with by the use of simulated maximum likelihood. The small number of goods in the present paper does not require the use of a simulation procedure.

[3] Under these two-step estimators, the problem of adding-up of observed expenditures shares is dealt with by dropping one share equation during estimation. However, the parameter estimates are not invariant to the equation dropped (see, for example, Yen, Lin, and Smallwood).

[4] The list of empirical studies adopting inverse demand systems for the estimation of the demand for food, agricultural and fishery products is small but growing. Inverse demand systems have been applied to the demand for fish (Barten and Bettendorf; Eales, Durham, and Wessels; Fousekis and Karagiannis; Beach and Holt; Holt and Bishop; Moro and Sckokai), to the demand for meat (Eales and Unnevehr; Eales; Holt and Goodwin; Holt), to the demand for fruits (Brown, Lee and Seale), to the demand for vegetables (Rickersten), and to the demand for composite food and nonfood commodities (Huang). For the rationale of the use of inverse demand systems in food demand see, for example, Barten and Bettendorf, and Eales and Unnevehr.

[5] As Eales and Unnevehr note, there is no closed-form solution for the dual of the AIDS logarithmic cost function. Hence, the following logarithmic distance function of the IAIDS is not the dual of the AIDS logarithmic cost function.

[6] The respective log-linear analogue of the Laspeyres price index is given by $ln\, P = \sum_j w_{j0}\, ln\, p_j$, where p_j is the price of good *j*. Buse and Moschini have shown that this price index provides a better approximation to the 'true' direct AIDS index than the widely used Stone price index does.

[7] Battese, Malik, and Gill drawing from the findings reported by Battese utilized a similar approach in the context of the stochastic production frontier model when zero input use is reported by farmers in the sample.

[8] Dong, Gould, and Keiser use the following formula for the calculation of the price elasticities for the direct AIDS: $\eta_j = -\delta_j + \frac{\Delta E(W)}{\Delta P_j} \frac{P + \Delta P_j/2}{E(W) + \Delta E(W)/2}$, where η_j is the elasticity vector with respect to a small change on price j, ΔP_j, δ_j is a vector of 0's with the jth element 1, $E(W)$ is the vector of the simulated observed shares, and $\Delta E(W)$ is the change of the simulated $E(W)$ given the change of price, ΔP_j.

[9] Relations (21.b) and (21.c) were derived by Anderson. Relation (21.b) is the restriction of homogeneity zero of the inverse demands, expressed in terms of flexibilities, while relation (21.c) is the analogue of the Slutsky equation for inverse demand systems.

[10] In order to accommodate heteroscedasticity of the random errors, we estimated the IAIDS model under the assumption that the diagonal elements of the variance-covariance matrix of the error terms depend, exponentially or additively, on one or both the demographic variables. We then carried out a Langrange ratio test, which indicated that a homoscedastic rather than a heteroscedastic error structure should be employed. Therefore, the parameter estimates and the simulated flexibilities derived from the homoscedastic specification are the ones presented in Tables 3-6.

[11] Dominance of complementarity in the *Antonelli* matrix has been reported in several studies of consumer demand for goods that are *a priori* expected to be substitutes, such as different categories of meats (e.g., Eales and Unnevehr; Holt) and fish (e.g., Barten and Bettendorf; Fousekis and Karagiannis).