

Functional foods as differentiated products: the Italian yogurt market

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Abstract

In spite of the growing consumers' interest for functional foods, the knowledge regarding the demand for these products and their profitability is limited. Adapting the LA/AIDS (Linear Approximated–Almost Ideal Demand System) model by means of Pinkse, Slade and Brett's distance metric method (2002), this article studies demand, substitution pattern, and profitability of conventional and functional alternatives inside the yogurt category in Italy. Results indicate that, in the yogurt market, functional alternatives' demand is often less elastic than that of their conventional counterparts, that brand loyalty plays a key role, and that the profitability of the functional alternatives is, on average, larger than that of conventional ones.

Keywords: functional foods, differentiated products, distance metric, yogurt

JEL classification: L11, L13, L66

1. Introduction

Thanks to the increased consumers' interest for health and wellness, the market for functional food products (claiming to provide health benefits beyond the traditional nutrients they contain)¹ has experienced strong growth. During the period 2004–2007, the sales of fortified and functional packaged goods in western Europe have experienced a rate of growth exceeding 10 per cent ([The Economist, 2009](#)), reaching a value of approximately USD 0.8 billion in 2006 ([Datamonitor, 2007](#)).

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1 The European Commission's Concerted Action on Functional Food Science in Europe (FuFoSE) states that '[...] a food product can only be considered functional if together with the basic nutritional impact it has beneficial effects on one or more functions of the human organism thus either improving the general and physical conditions or/and decreasing the risk of the evolution of diseases' ([Diplock et al., 1999](#)). See [Siró et al. \(2008\)](#) for a summary of other definitions.

Academics have shown a strong interest for this phenomenon (see Siró *et al.*, 2008 for a review of the literature on functional foods). Many studies have investigated aspects related to the demand for functional products, such as consumers' willingness to pay for food with health-enhancing features (e.g. West *et al.*, 2002; Markosyan, McCluskey and Wahl, 2009) or consumers' attitude towards them,² using survey data rather than data on actual consumer purchases. Surprisingly, only few studies have tried to assess the demand for these products and to characterise their markets using actual sales data: Yuan, Capps and Nayga (2009), for example, investigate sales cannibalisation due to the presence of a functional alternative in a given category (orange juice), while Bonanno (2011) analyses how health-related socio-demographic characteristics impact the odds of success of functional versus conventional yogurts in Italy.

Furthermore, no study so far has assessed functional foods' market performance and the extent of their differentiation. Functional foods are used by food manufacturers to attract new customers and to revitalise mature segments (Heasman and Mellentin, 2001), and their higher margins necessary to recover (i) the large R&D costs incurred in the development of the functional attributes,³ (ii) high marketing costs, and (iii) diseconomies of scope which may arise from excessive product-line length (Draganska and Jain, 2005) and from the failure to support the already existing core products (Herath *et al.*, 2008). If functional foods' manufacturers fail in their differentiation strategy, as prices increase consumers may be less likely to purchase the more pricey functional alternatives and more likely to switch to the conventional ones. Thus, understanding consumers' purchasing patterns for functional and conventional alternatives is key to understand the likelihood of success of these products and their profitability.

The lack of characterisation of the market of functional products (and of their performance) is surprising, particularly in the light of the rapid transformation it is experiencing. For example, in Europe, functional foods are under scrutiny due to the implementation of Regulation (EC) No. 1924/2006, regulating food products' health claims. The European Food Safety Authority (EFSA) has set up a panel of experts (i.e. the panel on Dietetic Products, Nutrition and Allergies – NDA) in charge of reviewing the validity of each product's health claim, whose truthfulness has to be supported by scientific evidence. In November 2009, the EFSA announced its first decisions on 523 claims – two-thirds of which were negative. Since then, food industry

2 According to this literature, consumers with a positive attitude towards functional foods have also a clear understanding and a positive perception of the health benefits they provide. For example, Verbeke (2005) shows that, among Belgian consumers, believing in the health benefits of functional foods is the main positive determinant of their acceptance. Using a cross-national sample, Labrecque *et al.* (2006) found that health, beliefs regarding health-related benefits, and credibility of information are the main positive determinants of the acceptance of these products.

3 Menrad (2003) reports that Unilever invested more than USD 50 million to develop Nestlé Lc1, a functional yogurt, and Becel®, a proactive margarine; this sum is considerably higher than the general estimated cost of developing a new food product (USD 2 million).

pundits have been concerned that the lack of transparency of EFSA's protocols could generate a climate of uncertainty jeopardising innovation and growth of the European food industry (Starling, 2009). Such uncertainty has not spared large companies: for example, Danone (which shared the support that the Yoghurt and Live Fermented Milks Association gave the legislation) withdrew in April 2009 two health claim (Article 13.5) submissions: a digestive health claim for Activia and an immunity claim for Actimel (drinkable), seeking further guidance from EFSA about scientific requirements. In August 2009, the company submitted an Article 14 disease reduction claim for Actimel and, in November of the same year, an Article 13.5 (health claim) for Activia, which were both denied (Starling, 2010).

This study contributes to the understanding and characterisation of functional foods' markets by pursuing the following objectives: (i) to characterise the demand for functional and conventional products inside one product category; (ii) to investigate the patterns and the determinants of consumers' switching between conventional and functional alternatives; and (iii) to provide an empirical assessment of the profitability of these products. To achieve these goals, Deaton and Muellbauer's (1980) Linear Approximated–Almost Ideal Demand System (LA/AIDS) model is modified following Rojas and Peterson's (2008) adaptation of Pinkse, Slade and Brett's (2002) Distance Metric (DM) method, and the model applied to a scanner database of yogurt purchases in 16 Italian regions, encompassing 18 conventional and 12 functional alternatives.⁴ The DM method builds on the concept that products more distant in the attribute space are less likely to be substitutes to one another. This method allows for a flexible substitution pattern while keeping the analysis tractable (e.g. only one equation needs to be estimated, even when a large number of alternatives are considered, and the number of the estimated parameters is heavily reduced). Although other frameworks exist that allow for flexible substitution patterns and ease of tractability (e.g. the random parameter logit by Berry (1994) and Berry, Levinsohn and Pakes (1995)), combining the DM method with the LA/AIDS model permits to quantify the role of product attributes on price-driven switching patterns. The Italian yogurt market is chosen as a case study since large yogurt manufacturers operating in this market (Danone, Parmalat and Nestlé) have heavily invested in adding new lines of functional products.⁵

The results show that the demand for functional yogurts in Italy is on average less price-elastic than that for conventional ones, and that brand loyalty plays a major role in this market. Also, although price does not appear as a strong determinant of switching between conventional and functional yogurts, intra-brand switching between functional and conventional yogurts is more likely than inter-brand, suggesting that the different functional

4 Each product is identified as a combination of brand (vendor), flavour, fat content, and the presence (or absence) of the functional attribute.

5 See Bonanno (2011) for a description of the different functional alternatives sold in the Italian yogurt market.

attributes existing across brands, as well as effective marketing campaigns, may lead to substantial switching costs. Lastly, the results suggest that, in most cases, functional yogurts generate higher short-run margins than their conventional counterparts.

2. The model

In the discussion that follows, the demand side of the model is treated first, highlighting the rationale and some of the technical details of incorporating the DM method into an LA/AIDS model. Then, illustrating the supply-side of the model, emphasis is given to the specific assumptions regarding the hypothesised market structures (Nash–Bertrand equilibrium) and on the characterisation of the derived measures of profitability.

2.1. The demand side

The demand for yogurts in Italy is modelled following the LA/AIDS model developed by [Deaton and Muellbauer \(1980\)](#). Although each Italian region is treated as a separate market, a regional subscript is omitted to simplify the notation; however, the illustration that follows holds for each one of the geographic markets considered. Let $j \in (1, \dots, J)$ and $t \in (1, \dots, T)$ be product and time indexes, respectively. Let q_{jt} be the retail-level quantity demanded for product j at time t and p_{jt} its price; the total expenditure for all yogurt types at time t is $x_t = \sum_j q_{jt}p_{jt}$, so that

$$w_{jt} = \alpha_j + \sum_{k=1}^J \gamma_{jk} \log p_{kt} + \beta_j \log \frac{x_t}{P_t^L} + e_{jt}, \quad (1)$$

where $w_{jt} = q_{jt}p_{jt}/x_t$ is product j 's expenditure share at time t , $\log P_t^L$ is [Moschini's \(1995\)](#) Laspeyres-type price index ($\log P_t^L \approx \sum_{j=1}^J w_j^0 \log p_{jt}$, where $w_j^0 = T^{-1} \sum_{t=1}^T w_{jt}$), the α , γ , and β are parameters to be estimated and e_{jt} is an error term. After imposing all the restrictions dictated by theory,⁶ the estimation of an LA/AIDS demand system will consist of $J - 1$ equations producing $J(J - 1)/2$ cross-price parameters, which, for large J , is an unmanageable task.

To circumvent this issue, each cross-price parameter γ_{jk} is assumed to be a function of the distance in the attribute space between products j and k . This approach, the DM method, was originally developed by [Pinkse, Slade and Brett \(2002\)](#) to analyse spatial price competition in the US wholesale gasoline market. It has been used first in demand analysis by [Pinkse and Slade \(2004\)](#), and [Slade \(2004\)](#), who applied it to 'representative consumer's' linear demand

6 For the AIDS to be consistent with the properties of a well-behaved demand system, the conditions of symmetry, homogeneity, and adding-up need to hold: $\gamma_{jk} = \gamma_{kj}, \forall j, k; \sum_{k=1}^J \alpha_j = 1; \sum_{k=1}^J \beta_j = 0$; and $\sum_{k=1}^J \gamma_{kj} = 0$.

systems based on a normalised-quadratic indirect-utility function, and then adapted to the LA/AIDS model by Rojas (2008), Rojas and Peterson (2008), and Pofahl and Richards (2009).⁷

In this application of the DM method (which will be referred to as DM-LA/AIDS), let Z_j^C and Z_j^D be sets representing product j 's attributes, measured in continuous space (calories, fat content etc.) and in discrete space (brand, flavours, presence of a functional attribute), respectively. Let δ_{jk}^C and δ_{jk}^D be measures of closeness between product j and k , function of continuous and discrete attributes, respectively. As in Rojas (2008), and in Pofahl and Richards (2009), δ_{jk}^C is specified as function of the Euclidean distance in characteristics space between product j and k ⁸, which in this context is defined as:

$$\delta_{jk}^C = \frac{1}{1 + 2\sqrt{\sum_l (z_{jl}^C - z_{kl}^C)^2}}, \quad (2)$$

where z_{jl}^C (z_{kl}^C) is the l -th continuous attribute of product j (k). Let z_{jl}^D be an indicator variable such that $z_{jl}^D = \{1 \text{ if product } j \text{ shows characteristic } l; 0 \text{ otherwise}\}$. The expression for δ_{jk}^D is:

$$\delta_{jk}^D = \begin{cases} 1 & \text{if } |z_{jl}^D - z_{kl}^D| = 0 \\ 0 & \text{if } |z_{jl}^D - z_{kl}^D| = 1 \end{cases}. \quad (3)$$

Given the closeness measures δ_{jk}^C and δ_{jk}^D , the cross-price parameter portion of the LA/AIDS is reformulated as follows:

$$\sum_{k=1}^J \gamma_{jk} \log p_{kt} = \gamma_{jj} \log p_j + \lambda_j \sum_{k \neq j}^J \delta_{jk}^C \log p_{kt} + \varphi_j \sum_{k \neq j}^J \delta_{jk}^D \log p_{kt}, \quad (4)$$

which gives $\gamma_{j1} = \lambda_j \delta_{j1}^C + \varphi_j \delta_{j1}^D, \dots, \gamma_{jn} = \lambda_j \delta_{jn}^C + \varphi_j \delta_{jn}^D$, where φ_j and λ_j are parameters to be estimated. Instrumentally to the purpose of this analysis and for ease of interpretation, the closeness measures δ_{jk}^C and δ_{jk}^D are used additively.

7 Also, both indirect and direct attempts to use the DM approach in the context of discrete choice models exist. For example, Richards, Hamilton and Patterson (2010) apply the DM method to the estimation of price conduct parameters, while estimating a nested-logit demand model. Differently, Richards *et al.*'s (2011) model incorporates a spatial weighting matrix in the 'mean utility' term of their multi-nomial logit model to allow for 'spatial lag' (in an attribute-space sense) in consumers' choice. Neither approach captures directly the role of distance in attributes on the substitution patterns, one of the goals of the current analysis.

8 Pinkse, Slade and Brett (2002) treat distance functions non-parametrically. However, Pinkse and Slade (2004) show that both parametric specification and semi-parametric specification of the model lead to similar results.

Following [Rojas \(2008\)](#) and [Rojas and Peterson \(2008\)](#), symmetry is imposed to the cross-price parameters by assuming $\lambda_1 = \lambda_2 = \dots = \lambda_J = \lambda$ and $\varphi_1 = \varphi_2 = \dots = \varphi_J = \varphi$. Since $\delta_{jk}^C = \delta_{kj}^C$ and $\delta_{jk}^D = \delta_{kj}^D$, one has $\lambda\delta_{jk}^C + \varphi\delta_{jk}^D = \gamma_{jk} = \gamma_{kj} = \lambda\delta_{kj}^C + \varphi\delta_{kj}^D$, which reduces the total number of cross-price parameters to be estimated from $J(J-1)/2$ to 2. Additionally, the number of equations to be estimated is reduced from $J-1$ to 1 by virtue of interacting product attributes with own-price, intercept, and expenditure coefficients.⁹ Imposing $\alpha_j = \alpha_0 + \sum_i \alpha_i z_{ji}^\alpha$, $\gamma_{jj} = \gamma_0 + \sum_l \gamma_l z_{jl}^\gamma$, and $\beta_j = \beta_0 + \sum_m \beta_m z_{jm}^\beta$, where z_j^α , z_j^γ , and z_j^β are subsets of product j 's attributes, the specification of the DM-LA/AIDS model is:

$$w_{jt} = \alpha_0 + \sum_i \alpha_i z_{ji}^\alpha + \log p_{jt} \left(\gamma_0 + \sum_l \gamma_l z_{jl}^\gamma \right) + \lambda \sum_{k \neq j}^J \delta_{jk}^C \log p_{kt} \\ + \sum_d \varphi^D \sum_{k \neq j}^J \delta_{jk}^D \log p_{kt} + \log \frac{x_t}{P_t^L} \left(\beta_0 + \sum_m \beta_m z_{jm}^\beta \right) + e_{jt}, \quad (5)$$

where the discrete closeness measures are indexed with the superscript D , $D = \{\text{brand (Br); flavour (Fl); functional (H); drinkable (Dr)}\}$, indicating, respectively, the discrete characteristics z_j^{Br} , z_j^{Fl} , z_j^{H} , and z_j^{Dr} .

The sign and magnitude of the estimated φ^D characterise the structure of consumers' switching motivated by a price increase. A positive φ^{Br} would, for example, suggest that consumers are more likely to respond to a price increase by switching to another alternative produced by the same manufacturer, i.e. that brand (vendor) loyalty plays a role in this market. Similarly, if the coefficient associated with closeness in the functional attribute φ^{H} is positive, consumers are more likely to switch within either conventional or functional yogurts than between them. If instead $\varphi^{\text{H}} < 0$, consumers are more likely to switch from a functional to a conventional yogurt (or vice versa). The other coefficients have similar interpretations.

Marshallian price elasticities are calculated as:

$$\eta_{jk} = \begin{cases} -1 + \frac{\gamma_0 + \sum_l \gamma_l z_{jl}^\gamma}{\bar{w}_j} - \left(\beta_0 + \sum_m \beta_m z_{jm}^\beta \right) & \text{if } j = k \\ \frac{\lambda\delta_{jk}^C + \sum_d \varphi^D \delta_{jk}^D}{\bar{w}_j} - \left(\beta_0 + \sum_m \beta_m z_{jm}^\beta \right) \frac{\bar{w}_k}{\bar{w}_j} & \text{if } j \neq k \end{cases}, \quad (6)$$

where $\bar{w}_j(\bar{w}_k)$ is product j 's (k 's) expenditure share measured at the sample

⁹ [Pinkse and Slade \(2004\)](#) originally proposed the interaction of product characteristics with own-price and intercept's coefficients, warning that it may increase the risk of multi-collinearity. For this reason, some researchers have estimated the full set of simultaneous equations ([Pofahl and Richards, 2009](#)).

means. Comparing own- and cross-price elasticities for functional and conventional yogurts will help characterise the role of price on consumers' acceptance of functional products in the presence of conventional alternatives.

2.2. The supply side

Let Y_n be the set of yogurts produced by manufacturer n . Assume manufacturer n maximises its profits by jointly setting prices for all the products it produces:

$$\max_{p_j} \pi_n = \sum_{j \in Y_n} q_j(p_j, p_{-j}; z_j, z_{-j})[p_j - c_j] - F_j, \quad (7)$$

where $q_j(p_j, p_{-j}; z_j, z_{-j})$ is the quantity demanded of yogurt j , function of own prices and product characteristics as well as other products; c_j is product j 's (constant) short-run marginal cost and F_j its fixed cost.

For the illustration that follows, it is assumed that yogurt manufacturers in the Italian market play a Bertrand game, and that prices are the outcome of a Nash equilibrium of such game. Assuming the existence of a Nash–Bertrand equilibrium is fairly common in the analyses of profit margins for firms operating in differentiated product markets. Since seminal works such as Hausman, Leonard and Zona (1994), countless applications using this assumption can be found in the literature.¹⁰ Among the many studies that use models based (at least in part) on the Nash–Bertrand assumption are Berry, Levinshon and Pakes (1995), investigating profit margins in the US car market; Nevo (2001), analysing the performance of ready-to-eat breakfast cereals manufacturers in the US; Slade (2004), testing the existence of unilateral versus coordinated effects in UK brewing; Di Giacomo (2008), evaluating the effect of new product introductions in the Italian yogurt market, and Rojas (2008), considering Bertrand–Nash as a possible structure of the competition among beer manufacturers in the US.

Under the Nash–Bertrand assumption, the optimisation problem in equation (7) leads to a vector of first-order conditions (FOCs) which can be expressed as:

$$p - c = -\Omega^{-1}q(p, z), \quad (8)$$

where $q(p, z)$ is a vector of quantity demanded, and each element of the matrix Ω is defined as

¹⁰ As Nevo (1998) illustrates, there are two main advantages of assuming that prices are the equilibrium outcome of a Bertrand game: first, it gives a finite number of possible outcomes for which PCMs can be easily calculated; second, it avoids problems of parameter identification arising when assessing explicitly the games played by the firms in the market.

$\Omega_{jk} = \Omega_{jk}^* \Delta_{jk}$, where

$$\Omega_{jk}^* = \begin{cases} 1 & \text{if } k, j \in Y_n, \\ 0 & \text{otherwise} \end{cases}, \quad \text{and} \quad \Delta_{jk} = \frac{\partial q_j(p, z)}{\partial p_k}. \quad (9)$$

In the context of a multi-product Nash–Bertrand equilibrium, Ω^* represents the ownership matrix, while the elements of Δ are partial derivatives of demand with respect to the vector of prices. Equation (8) defines implicitly the price–cost margin (PCM) of each product $j \in Y_n$. Following [Rojas \(2008\)](#), one can obtain different values of the PCMs by combining the estimated parameters of the DM-LA/AIDS with different structures of Ω^* . Two scenarios are considered here. The first assumes that the price of each yogurt is the outcome of a single-product Nash–Bertrand equilibrium ($\Omega_{jk}^* = 1 \forall j = k$, 0 otherwise) which will provide a lower bound to the estimated short-run profitability. The second assumes manufacturers setting independently the prices for the three product-lines, conventional spoonable, drinkable functional, and spoonable functional ($\Omega_{jk}^* = 1$ and $z_j^{Dr} - z_k^{Dr} = 0$ and $z_j^{Dr} - z_k^{Dr} = 0$).

The reader should notice that the estimated PCMs obtained from equation (8) will be short-run margins and therefore do not account for fixed costs. As such, the estimated margins will be upper bound estimates of the profitability of the yogurts sold in the Italian market. Given the larger R&D costs associated with functional foods' development ([Menrad, 2003](#)), it is conceivable that the PCMs obtained from equation (8) will overestimate the profitability of functional products more than that of conventional ones.

3. Data, model specification, and estimation

Equation (5) is estimated using primarily a scanner database provided by the Food Marketing Policy Center at the University of Connecticut, supplied originally by Information Resources, Inc. ([IRI, 2004–2005](#)). The data include 24 monthly observations of yogurt sales (quantities and values) for the period January 2004–December 2005 in hyper- and super-markets located in 16 Italian IRI regions,¹¹ for a total of 384 market combinations. Thirty products¹² are identified by vendor (Danone, Granarolo, Nestlé, Müller, and Parmalat, referred below as brands), flavour (plain,

11 IRI regions are defined consistently with the political boundaries of the Italian regions except for Piedmont and Val d'Aosta, Basilicata and Calabria, and Abruzzo and Molise. Trentino Alto Adige was excluded due to the strong presence of regional brands.

12 The products chosen belong to firms operating nationally, to exclude regional and local brands. A 0.5 per cent expenditure share threshold was used for the inclusion of a product in the analysis. Such value was calculated as the ratio of a product's national average value of monthly sales over the sum of the national average value of monthly sales for all the products considered. The sub-categories are identified by combination of fat content, flavour, and 'health' content (functional and conventional).

fruit, and other flavours), fat content (skim and whole),¹³ drinkable versus non-drinkable, and by the presence of a functional attribute, for a total of 11,520 observations. Volume and value of sales are used to calculate prices in EUR/kg.

Attributes information was collected from the manufacturers' websites or, when not available, from www.ciao.it, a website where Italian consumers share opinions on purchase experiences, often reporting the nutritional content of food products as they appear on the nutritional labels.¹⁴ The continuous product characteristics used in the analysis are protein, sugar, fat content and calories, referred to 100 g of product.¹⁵ Table 1 presents a summary statistics of the data for the 30 products included in the analysis, including product characteristics, price and expenditure shares. Product characteristics from the IRI data include average volume per unit (kg/unit) and a proxy for market coverage (average number of items per store). Additionally, monthly and regional dummies are included in the estimation to capture seasonal variation in yogurt consumption and unobservables across regions, respectively.

One challenge in estimating equation (5) was to maintain a flexible model specification while trying to mitigate the risk of multi-collinearity resulting from using the same product characteristics to shift the model's parameters. In the first place, average volume per unit and market coverage were chosen as intercept shifters (z_j^α) so that one could limit the use of physical product characteristics as own-price and expenditure shifters (z_j^γ and z_j^β , respectively).¹⁶ Fat content and calories were used in a mutually exclusive way as part of either z_j^γ or z_j^β , because of the large correlation of these two variables (0.87).

Thirty-two model specifications were estimated: 2 'full' specifications and 30 restricted ones. In the two 'full' specifications, the vectors of discrete product characteristics brand (i.e. vendor: Danone, Granarolo, Müller, and Parmalat) and flavour indicators (fruit, other flavours, and plain)¹⁷ are part of both z_j^γ and z_j^β : while in the first (second) specification fat (protein and sugar) content shifts the log-price (expenditure) parameter, protein and

13 "Skim" includes low-fat and fat-free yogurts; "whole" includes all the other yogurt alternatives.

14 The accuracy of the postings was evaluated by cross-checking available nutritional information at www.ciao.it and at manufacturers' websites, which, in the cases considered, were found to be accurate.

15 Continuous attributes measured in different units should be rescaled (or their distance normalised to 1) so that all the attributes have the same weight in determining closeness. Since, in this analysis, all continuous attributes are expressed in the same units, no rescaling is needed.

16 Attempts to use physical product characteristics, brand, and flavour indicators as intercept shifters resulted in very large variance inflation factor (VIF) whose average value for this specification was 606.7. Alternatively, the use of product fixed-effects as intercept shifters was also considered, as it could mitigate problems of unobserved heterogeneity (as in [Rojas and Peterson, 2008](#)). As most of the product characteristics are already included as part of either z_j^γ or z_j^β (or both), adding product dummies as intercept shifters resulted to be redundant, confirmed by a large average VIF (792.3).

17 As the information in the data does not allow to observe the flavours of drinkable yogurts, "Drinkable" becomes the excluded category when using flavour indicators.

Table 1. Product characteristics, average price, and expenditure shares by product

Brand	Flavour	Fat content	Calories (Cal/100 g)	Proteins (g/100 g)	Sugar (g/100 g)	Fat (g/100 g)	Price (EUR/kg)	Exp. share (per cent)
Conventional								
Danone	Plain	Skim	49	6.1	5.0	0.1	4.41	1.15
		Whole	99	3.3	12.5	3.7	4.37	1.35
	Fruit	Skim	52	4.1	7.9	0.1	4.40	12.11
	Others ^a	Skim	58	4.4	8.9	0.1	5.24	2.87
Granarolo	Plain	Skim	39	4.7	4.0	0.1	3.81	0.84
		Whole	68	3.5	3.5	4.0	3.50	0.94
	Fruit	Skim	75	3.9	13.7	0.1	4.02	1.49
		Whole	103	3.2	12.5	4.1	4.17	9.15
	Others	Whole	117	3.7	15.1	4.3	4.38	3.02
	Plain	Whole	109	5.1	11.3	4.5	2.91	4.08
Müller	Fruit	Skim	76	4.6	13.4	0.1	3.94	1.08
		Whole	111	2.9	16.1	3.6	3.37	10.10
	Others	Whole	118	4.4	15.8	4.4	3.45	2.62
	Fruit	Skim	40	4.2	5.6	0.1	4.05	1.63
Nestlé	Others	Skim	73	4.3	13.4	0.2	4.86	0.57
	Fruit	Skim	59	5.2	9.4	0.1	3.47	1.68
Parmalat	Fruit	Whole	109	3.4	15.5	3.7	3.19	6.24
	Others	Whole	119	3.3	15.4	4.7	3.45	0.85

Functional								
Danone	Plain	Skim	48	4.9	6.1	0.1	4.98	1.07
		Whole	72	4.2	5.1	3.5	4.96	1.40
	Fruit	Skim	52	4.4	7.5	0.1	5.32	1.08
		Whole	104	3.7	13.6	3.4	5.32	3.70
Parmalat	Others	Whole	103	3.8	13.5	3.3	5.31	7.67
	Fruit	Whole	103	3.1	14.0	3.8	4.91	0.77
	Others	Whole	106	3.1	14.0	4.2	5.01	1.02
Functional/drinkable								
Danone		Skim	29	2.7	3.7	0.1	5.55	3.79
		Whole	73	2.7	11.8	1.2	5.54	11.00
Granarolo		Whole	77	3	12	1.9	5.30	1.2
Nestlé		Skim	62	2.7	12.7	0.1	5.29	1.55
		Whole	77	2.6	14.5	0.9	5.21	3.98

Source: Calories, protein, sugar, and fat content come from nutritional labels collected from various sources. Price and 'Exp. Share' are obtained from IRI Infoscan data: January 2004–December 2005 averages.

^a'Others' indicate 'other flavours'.

sugar (fat) content shifts the category expenditure (log-price) parameter. The 30 restricted specifications were obtained excluding sequentially brand and flavour indicators from z_j^γ and z_j^β . Model selection was performed by monitoring the average VIF¹⁸ to reduce the presence of multi-collinearity, the significance and sign of the key estimated parameters, and magnitude, sign, and significance of estimated elasticities.

Estimates of the parameters of equation (5) can be biased if prices are correlated with demand shocks unaccounted for by the other variables in the model. Endogeneity was detected in the ‘full’ models using *C* statistics, obtained as the difference of two Sargan statistics (Hayashi, 2000: 232). To ensure unbiasedness of the estimates, an instrumental variable estimation method (Generalised Method of Moments – GMM) was adopted using variables related to yogurt manufacturing and retail costs as instruments for price. Such instruments are: farm-level milk price (national, monthly, EUR/l) price of cream at the origin (national, monthly, EUR/kg) interacted with fat content, farm-level national price of fruit (national, monthly, EUR/kg), taken all from the DATIMA database by ISMEA (Institute for the Study of Agricultural Markets); the producer price index for the dairy industry (national, monthly) by ISTAT (the National Institute of Statistics); the European import price (CIF) of sugar (monthly, USD/lb) by Index Mundi; retail workers’ per capita earnings (regional, annual, thousand EUR) by OIC (Italian Observatory of Commerce – Ministry of Economic Development); the industrial price of heating oil (national, monthly, EUR/hl) by DGERM (Office of the Director General of Energy and Mineral Resources – Ministry of Economic Development); and the commercial price of electricity at the source (regional, monthly, EUR/Mw) by the Manager of the Energy Market (GME).¹⁹

The instruments’ orthogonality was evaluated using Hansen’s (1982) *J*-statistic, while problems of weak instruments were ruled out using Staiger and Stock’s (1997) ‘rule of thumb’ (the value of the *F*-statistics for the joint significance of the instruments’ parameters in the first-stage equation exceeds 10 in all models). Following Blundell and Robin (2000) and Dhar, Chavas and Gould (2003), category expenditure is also treated as endogenous, and it is instrumented by regressing it on median household income from ISTAT, its squared term, a (monthly) time trend, and region dummies.²⁰ The different specifications of equation (5) were estimated in STATA v. 11.²¹

18 As product characteristics used to create closeness measures are the same as those used to shift own-price and deflated category expenditure, some level of correlation among the variables in the model should be expected even in the most restricted models.

19 Instrument’s data were collected from on-line sources and available from the author.

20 The use of a first-stage regression to account for expenditure endogeneity was motivated by the fact that time trend is part of the instruments proposed by Dhar, Chavas and Gould (2003) and monthly dummies are used to capture seasonal variations in the demand for yogurt in equation (5).

21 Durbin–Watson tests for the first-order autocorrelation of the error terms were performed for all model specifications. Most of the calculated *d* values were close to 2, which for a large dimension of the cross-sectional component of the panel indicates absence of serial correlation (Bhargava, Franzini and Narendranathan, 1982: pp. 536 and 537, Table II).

Table 2. Estimated parameters and related statistics

Variables	Model specifications and restrictions					
	Full model		Rest1		Rest2	
	No restrictions		z_j^γ : no flavours; z_j^β : no brands		z_j^γ : no flavours, no brands; z_j^β : no brands	
	Parameters	Standard errors	Parameters	Standard errors	Parameters	Standard errors
$\log p_j$	-0.0218	(0.0271)	-0.0326***	(0.0026)	-0.0465***	(0.0026)
$\log p_j \times \text{fat}$	-0.0023***	(0.0003)	-0.0020***	(0.0003)	-0.0032***	(0.0002)
$\log p_j \times \text{functional}$	0.0024	(0.0078)	0.0146***	(0.0025)	0.0204***	(0.0027)
$\log p_j \times \text{others}$	-0.0151	(0.0217)				
$\log p_j \times \text{fruit}$	0.0026	(0.0209)				
$\log p_j \times \text{plain}$	-0.0579***	(0.0206)				
$\log p_j \times \text{Danone}$	-0.0056	(0.0072)	-0.0039	(0.0043)		
$\log p_j \times \text{Granarolo}$	-0.0378***	(0.0066)	-0.0100***	(0.0010)		
$\log p_j \times \text{Müller}$	-0.0056	(0.0072)	-0.0005	(0.0010)		
$\log p_j \times \text{Parmalat}$	0.0269***	(0.0058)	-0.0014**	(0.0006)		
Closeness fat/sugar/protein	0.0017***	(0.0004)	0.0027***	(0.0004)	0.0002	(0.0005)
Closeness brand	0.0056***	(0.0007)	0.0031***	(0.0006)	0.0032***	(0.0001)
Closeness flavour	-0.0036***	(0.0008)	0.0042***	(0.0006)	0.0046***	(0.0006)
Closeness functional	-0.0007	(0.0012)	0.0002	(0.0004)	0.0008**	(0.0004)
Closeness drink	0.0024*	(0.0013)	-0.0010***	(0.0001)	-0.0008***	(0.0001)
$\log(x_i/P_i^L)$	-0.0023***	(0.0007)	-0.0031***	(0.0005)	-0.0038***	(0.0005)
$\log(x_i/P_i^L) \times \text{protein}$	6.15×10^{-4} ***	(3.8×10^{-5})	4.7×10^{-4} ***	(3.1×10^{-5})	3.5×10^{-4} ***	(3.5×10^{-5})
$\log(x_i/P_i^L) \times \text{sugar}$	2.93×10^{-5} ***	(7.8×10^{-6})	-2.84×10^{-5} ***	(6.7×10^{-6})	2.9×10^{-5} ***	(6.4×10^{-6})

(continued)

Table 2. (continued)

Variables	Model specifications and restrictions					
	Full model		Rest1		Rest2	
	No restrictions		z_j^γ : no flavours; z_j^β : no brands		z_j^γ : no flavours, no brands; z_j^β : no brands	
	Parameters	Standard errors	Parameters	Standard errors	Parameters	Standard errors
$\text{Log}(x_i/P_i^L) \times \text{functional}$	0.0004	(0.0006)	−0.0008***	(0.0003)	−0.0008***	(0.0003)
$\text{Log}(x_i/P_i^L) \times \text{others}$	−0.0034***	(0.0003)	0.0008***	(0.0001)	0.0010***	(0.0001)
$\text{Log}(x_i/P_i^L) \times \text{fruit}$	−0.0047***	(0.0004)	−0.0016***	(0.0003)	−0.0016***	(0.0003)
$\text{Log}(x_i/P_i^L) \times \text{plain}$	0.0017***	(0.0001)	0.0018***	(0.0001)	0.0018***	(0.0001)
$\text{Log}(x_i/P_i^L) \times \text{Danone}$	−0.0016***	(0.0006)				
$\text{Log}(x_i/P_i^L) \times \text{Granarolo}$	0.0026***	(0.0007)				
$\text{Log}(x_i/P_i^L) \times \text{Müller}$	0.0004	(0.0008)				
$\text{Log}(x_i/P_i^L) \times \text{Parmalat}$	−0.0030***	(0.0007)				
Average vol. unit	0.0546***	(0.0030)	0.0478***	(0.0029)	0.0467***	(0.0030)
Coverage	0.0083***	(0.0001)	0.0079***	(0.0001)	0.0077***	(0.0001)
Constant	−0.0243	(0.0242)	−0.0273*	(0.0142)	0.0007	(0.0143)
<i>R</i> -squared	0.7871		0.7580		0.7440	
Hansen <i>J</i> -test ($\chi^2_{(8)}$)	11.1531	($p = 0.1932$)	8.7963	($p = 0.3598$)	9.9644	($p = 0.2675$)
<i>F</i> -test significance instruments	$F_{(9,11458)} = 60.30$		$F_{(9,11465)} = 1585.9$		$F_{(9,11469)} = 1487.1$	
Mean VIF	106.24		26.19		17.20	

Note: Standard errors in parentheses. The coefficients of monthly dummies and regional fixed-effects are omitted for brevity.

*10 per cent significance level.

**5 10 per cent significance level.

***1 per cent significance level.

4. Empirical results

The results summarised below, and presented in Table 2, are for three specifications of equation (5); in those specifications, fat content shifts the own-price parameter, while protein and sugar content shift the expenditure parameter. Besides the results of a 'full' model specification (left columns), the results of two restricted specification are reported: one where brand indicators shift the own-price parameter and flavour indicators shift expenditure's (Rest1) and one where no discrete characteristics shift the own-price parameter and only flavour indicators shift expenditure parameter (Rest2). The restricted models produce VIF much smaller than the full model; the average VIFs are 106.24 for the full model and 26.19 and 17.11, respectively, for the two restricted models.²²

As discussed below, the estimated elasticities from the two restricted models are relatively close to one another; most restricted models whose average VIF is below 30 presented elasticities in the same range as those that will be discussed below, while models showing larger VIF presented highly unstable elasticities.²³ Similar specifications where protein content and sugar content shift own-price parameters while fat content shifts the expenditure parameter are qualitatively similar to those in Table 2, although showing larger values of VIF, and produce in most cases non-statistically significant elasticity estimates.²⁴

In terms of model performance, the full model's R -squared is 0.7871; the restricted models do not show considerable loss in explanatory power: the R -squared are, respectively, 0.7580 and 0.7440. The results in Table 2 are GMM estimates obtained using the instruments described above; OLS results are omitted for brevity and available upon request. The p -values associated with Hansen's (1982) J -statistic for the orthogonality of the over-identifying instruments vary from 0.19 for the full model to 0.36 for the specification Rest1, suggesting that the instruments used are orthogonal

22 Among the model specifications where fat content shifted the log price parameter while sugar and protein content shifted the expenditure parameter, the two specifications illustrated are not those showing the lowest VIF. Three other specifications presented average VIFs of 12.57, 12.48, and 4.34, respectively. These specifications present more restrictions than those reported in the text (no discrete characteristics were interacted with $\log p_{jt}$; $\log(x_{it}/P_t^L)$ was interacted, respectively, with brand indicators, flavour indicators, and no discrete characteristics) and their discussion is excluded. The estimated parameters of these specifications generate values of demand elasticities in line with those discussed in the main text.

23 As suggested by an anonymous reviewer, an attempt was made to estimate the full set of simultaneous equations as in Pofahl and Richards (2009) to avoid problems of multi-collinearity. SUR results showed values of own-price elasticities ranging from -0.29 (which, being smaller in magnitude than 1 conflicts with profit maximisation) to -3.59 (average -1.97). Attempts to correct for price endogeneity via 3SLS were made, using the instruments illustrated in the text, without obtaining sensible improvements of the results. The limited number of data points available to estimate the model in this fashion, coinciding with the number of market combinations (i.e. 384), and perhaps the limited variability of the instruments (some of which are at the national level) are probably the causes of not obtaining valid 3SLS estimates.

24 Full sets of result for the other model specifications are excluded for space limitation and are available upon request to the author.

to the error terms. Furthermore, the large values of the F -statistics for the joint significance of the instruments in the first-stage regression indicate that the results are free from weak instrument problems.

4.1. Estimated coefficients

The own-price coefficient, not statistically significant in the full model, is instead negative and significant at the 1 per cent level in the restricted models, showing a difference in magnitude across specifications Rest1 and Rest2 (-0.0326 and -0.0465 , respectively).²⁵ The coefficients associated with the interaction of log-price with fat content are negative and significant at the 1 per cent level in all models, varying from -0.0020 to -0.0032 , suggesting that, on average, Italian yogurt consumers are likely to show lower willingness to pay for products with higher fat content. The interaction of the functional indicator with log price generates positive coefficients in all specifications; although the coefficient is not statistically significant in the full model, it shows significance at the 1 per cent level in the other specifications (the estimates ranging from 0.0146 to 0.0204). This result indicates that, everything else constant, Italian consumers are less price-sensitive for functional yogurts than for conventional ones, corroborating previous findings that, on average, consumers tend to show a higher willingness to pay for products that carry a functional attribute (for example, West *et al.*, 2002; Markosyan, McCluskey and Wahl, 2009).

The role of flavours as shifters of the own-price parameters is weak, the only statistically significant coefficient in the full specification being that of plain yogurts; similar patterns are observed in other model specifications (all showing higher VIFs than those where brand indicators are interacted with log price), and are therefore not shown. The interactions of log price with vendor indicators suggest that Italian yogurt consumers are more price-sensitive for Granarolo's products (Granarolo's coefficients is negative and significant in the full model and Rest1); however, they tend not to show different price sensitiveness for Danone's and Müller's compared with Nestlé's (the excluded vendor) since the coefficients are not statistically significant in either specification. The direction of consumers' price sensitivity with respect to Parmalat is unclear.

The behaviour of the cross-price closeness measures (i.e. the weighted sum of the log prices of the other products in the market) is similar across specifications, although with some differences.

The estimated coefficient associated with δ_{jk}^C is positive and significant at the 1 per cent level in two out of three specifications, varying from 0.0018 in the full model to 0.0027 (Rest1), providing evidence that consumers respond to price

25 Although OLS results are not reported due to brevity, it should be noted that the estimated own-price parameters in the absence of instrumentation are smaller than those presented in Table 2; that of the Rest1 specification is -0.0303 , while for Rest2 specification, it is -0.0366 , which indicates that endogeneity bias may be a more severe issue in the more restricted model specifications.

increases by switching to products with similar nutritional profiles. Among discrete closeness measures, closeness in brands emerges as the strongest determinant of substitution, suggesting that, when motivated by a price change, Italian yogurt consumers tend to switch within products of the same manufacturer, or in other words, that this market is characterised by a substantial level of brand loyalty (the estimated coefficients are all significant at the 1 per cent level and vary from 0.0056 of the full model to 0.0031 of Rest1). The role of closeness in flavour is unclear: its coefficient is negative and significant in the full model and positive and significant in the restricted ones. Closeness in functional attribute does not appear to be a strong determinant of substitution, as its coefficients are not statistically significant, suggesting that price changes may not be a large motivator of switching between functional and conventional yogurts. Lastly, closeness in drinkable attribute does affect the substitution across yogurts; while the coefficient is positive but marginally significant in the full specification, its parameter is negative and statistically significant in the restricted ones, although small in magnitude (-0.0010 to -0.0008), suggesting that, as price of a non-drinkable yogurt increases, Italian consumers will likely switch to a drinkable one (and vice versa).

Concluding the illustration of the estimated parameters, the unshifted expenditure coefficient is negative in all models. Most of the product characteristics used as expenditure shifters generate statistically significant parameters whose signs are consistent across specifications. In particular, the positive and statistically significant coefficients for the interactions of protein and sugar (respectively) indicate that, as the total expenditure for the yogurt category increases, those products having larger (smaller) protein (sugar) content will tend to show larger expenditure shares. In other words, expenditure elasticities (not calculated for brevity) will be larger (smaller) for yogurts containing more protein (sugar). The parameters associated with discrete product characteristics interacted with the deflated log expenditure have similar interpretation. The coefficients of the demand intercept's shifters (average volume per unit and coverage) are positive, significant at the 1 per cent level and show similar magnitude across model specification, indicating that the yogurts sold in larger sizes and whose distribution is more intense tend to show larger expenditure shares.

4.2. Own-price elasticities

Estimates of own-price elasticities obtained using the estimated parameters of the restricted models and equation (6) are reported in Table 3; as the elasticities obtained from the parameters of those model specifications showing large VIFs show unreliably wide ranges, no elasticities for the full model are discussed. The range of elasticities presented in Table 3 varies from -1.22 to -6.86 (average value -3.14), and -1.27 to -9.38 (average value -3.76) for the Rest1 and Rest2 specifications, respectively. In all cases, Nestlé/conventional/other flavours/skim is the product showing the largest elasticity, while Danone/functional/drinkable/whole shows the smallest. The

Table 3. Estimated own-price elasticities

Vendor	Flavour	Fat content	Rest1	Rest2
Conventional				
Danone	Plain	Skim	−4.2008	−5.0873
		Whole	−4.2792	−5.3410
	Fruit	Skim	−1.3011	−1.3847
		Others	−2.2792	−2.6337
Granarolo	Plain	Skim	−6.0955	−6.5751
		Whole	−6.3719	−7.2661
	Fruit	Skim	−3.8698	−4.1403
		Whole	−1.5503	−1.6430
Müller	Others	Whole	−2.6781	−2.9656
		Plain	−2.0392	−2.4938
	Fruit	Skim	−4.0539	−5.2970
		Whole	−1.3955	−1.5682
Nestlé	Others	Whole	−2.5942	−3.2924
		Skim	−3.0225	−3.8901
	Fruit	Skim	−6.8638	−9.3831
Parmalat	Fruit	Skim	−3.0453	−3.8051
		Whole	−1.6716	−1.9439
	Others	Whole	−6.0752	−8.1619
Functional				
Danone	Plain	Skim	−3.0612	−3.4711
		Whole	−3.0680	−3.6555
	Fruit	Skim	−3.0483	−3.4563
		Whole	−1.7736	−1.9929
Parmalat	Others	Whole	−1.3711	−1.4755
		Fruit	−4.5340	−5.9832
	Others	Whole	−3.7478	−4.8847
Functional/drinkable				
Danone		Skim	−1.5791	−1.6943
		Whole	−1.2178	−1.2688
Granarolo		Whole	−3.6485	−3.6735
Nestlé		Skim	−2.1794	−2.7170
		Whole	−1.4960	−1.7272

Note: ‘Others’ indicates ‘other flavours’.
All estimated own-price elasticities are significant at the 1 per cent level.

magnitudes of the elasticities are consistent with other studies where consumers can substitute between numerous products (see, for example, [Villas-Boas, 2007](#); [Pofahl and Richards, 2009](#)); furthermore, the estimated values from specification Rest1 appear close to those of other brand-level demand analysis for yogurts. [Di Giacomo \(2008\)](#), who presents values of brand-level elasticities of demand for yogurt in Italy ranging from −0.88 to −2.66, reports also that an alternative specification of her model produced elasticities whose average was −3.17, close to the average values obtained here. [Bonanno’s \(2011\)](#) estimated own-price elasticity of demand for yogurt subcategories

in Italy (a higher level of aggregation than that used here) varies between -2 and -4.74 . In the US yogurt market, Draganska and Jain (2006) estimate own-price elasticities in a range between -2.45 and -6.25 , Villas-Boas (2007) presents an average value of -5.64 , while Richards *et al.* (2011) estimate a narrower range between -1.26 and -4.73 .

Overall, five patterns emerge:

1. *Functional versus conventional*. On average, functional alternatives show lower own-price elasticities than their conventional counterparts, although the opposite emerges for fruit-flavoured yogurts. This pattern does not necessarily hold if one considers values across brands: for example, Müller/plain/whole elasticities are smaller than Danone's plain/functional ones.
2. *Drinkable*. Functional drinkable yogurts show own-price elasticities of demand below the average values, with the exception of Granarolo's.
3. *Brand (vendor)*. The demand for Danone's yogurts tends to be less elastic than that for other brands, across flavours, fat content, and functional properties, with the exception of plain/conventional yogurts, where the 'whole' alternative by Müller shows the lowest magnitude of elasticity among plain yogurts.
4. *Flavours*. The demand for fruit-flavoured yogurts shows (on average) higher values of elasticity than that for other flavours and plain, for both conventional and functional yogurts alike.
5. *Fat content*. No unique trend emerges with respect to fat content and elasticities; for plain yogurts, the values are similar, while for fruit-flavoured and drinkable yogurts, whole alternatives' demand is less elastic than that for skim ones.

4.3. Cross-price elasticities

The discussion of the cross-price elasticities will consider the estimated parameters of the Rest1 specification only due to brevity, as those obtained from specification Rest2 are similar. The illustration is divided in two parts: cross-price elasticities between conventional and functional yogurts are discussed first (both intra- and inter-brand), whereas those among functional yogurts are discussed later.

Negative signs emerge for cross-price elasticities between functional and conventional yogurts (and some between functional ones) with different flavours and fat content, produced by different manufacturers (in many instances not statistically different from 0). This suggests that products located far from one another in the characteristic space are not likely to be seen as substitutes, as suggested by Kadiyali, Vilcassim and Chintagunta (1996). An alternative explanation, as suggested by Betancourt (2006), is that as households purchase multiple items during each shopping trip, different household members may prefer different alternatives belonging to the same category (i.e. different types of yogurt), resulting in some alternatives behaving as complements instead of substitutes.

Table 4. Selected own- and cross-price elasticities: Danone (specification Rest1)

		Conventional				Functional							
		Plain		Fruit Skim	Others Skim	Plain		Fruit		Others Whole	Drinkable		
		Skim	Whole			Skim	Whole	Skim	Whole		Skim	Whole	
Conventional													
Plain	Skim	−4.20	0.61	0.25	0.25	0.62	0.61	0.24	0.23	0.23	0.32	0.31	
	Whole	0.52	−4.28	0.18	0.21	0.51	0.52	0.20	0.23	0.21	0.27	0.25	
Fruit	Skim	0.02	0.02	−1.30	0.03	0.03	0.02	0.06	0.06	0.02	0.03	0.03	
Others	Skim	0.10	0.10	0.11	−2.28	0.11	0.10	0.12	0.10	0.24	0.13	0.13	
Functional													
Plain	Skim	0.67	0.65	0.27	0.28	−3.06	0.67	0.31	0.27	0.27	0.36	0.36	
	Whole	0.50	0.50	0.20	0.20	0.51	−3.07	0.21	0.21	0.21	0.28	0.28	
Fruit	Skim	0.26	0.26	0.73	0.32	0.31	0.28	−3.05	0.67	0.29	0.38	0.39	
	Whole	0.07	0.09	0.19	0.07	0.08	0.08	0.19	−1.78	0.10	0.10	0.11	
Others	Whole	0.03	0.04	0.03	0.09	0.04	0.04	0.04	0.05	−1.38	0.05	0.05	
Drink	Skim	0.10	0.10	0.11	0.10	0.10	0.10	0.10	0.10	0.11	−1.58	0.08	
	Whole	0.03	0.04	0.03	0.03	0.04	0.04	0.04	0.04	0.04	0.03	−1.22	

Note: ‘Others’ indicates ‘other flavours’.

All estimated elasticities presented in this table are significant at the 1 per cent level.

Intra-brand cross-price elasticities are illustrated using Danone as an example: the values are reported in Table 4. The notation η_{FC} (η_{CF}) indicates cross-price elasticity of demand for a functional (conventional) yogurt to a conventional (functional) one. The estimated cross-price elasticities for Danone's yogurts are positive and statistically significant at the 1 per cent level, with η_{FC} and η_{CF} showing similar magnitude for plain yogurts; for fruit-flavoured yogurts, the η_{CF} are smaller than η_{FC} , while the opposite is observed for other flavours, suggesting that Danone's plain functional yogurts may not be seen as different enough from the conventional ones to justify asymmetric cross-price elasticities. For the fruit-flavoured alternatives instead, differentiation is more marked, while for other flavours it appears weak, with consumers being more likely to purchase conventional yogurts if the price of a functional one increases than vice versa. Interestingly, the cross-price elasticities between functional drinkable yogurts and non-drinkable ones (both conventional and functional) are asymmetric, showing that, if motivated by a price increase, Italian consumers are more likely to switch from drinkable to non-drinkable yogurts.

A detailed discussion of inter-brand cross-price elasticities is omitted for brevity. Overall, three trends emerge: first, on average, cross-price elasticities among plain yogurts are larger than those among 'other flavours' and fruit-flavoured yogurts, whose average values are 0.42, 0.31, and 0.24, respectively. Second, values of cross-price elasticities are rather small for yogurts produced by different manufacturers and with different fat content (particularly η_{FC} and η_{CF}), some of them being negative. Third, cross-price elasticities for yogurts of the same flavour produced by the same manufacturers tend to be large (for example, the cross-price elasticity of Granarolo conventional/plain/skim to whole is 0.85, while the η_{CF} of Parmalat fruit/whole to skim is 0.90).

The values of cross-price elasticities among functional products, reported in Table 5, follow trends similar to those discussed above. In particular, functional yogurts of different flavours produced by different manufacturers show some negative (although not statistically different from zero) cross-price elasticities. Those produced by the same manufacturer or of the same flavour show large, positive, and statistically significant cross-price elasticities. Also, on average, cross-price elasticities for non-drinkable yogurts to drinkable ones have larger values than their counterparts.

4.4. Price–cost margins

The short-run PCMs calculated under the single-product Nash–Bertrand and the multi-product portfolio pricing for specification Rest1 are reported in Table 6. On average, PCMs increase from conventional to non-drinkable functional to drinkable ones; also, as expected, margins calculated under single-product Nash–Bertrand are lower than the portfolio pricing ones.

The average PCMs for conventional yogurts are 37.21 per cent under single-product Bertrand and 45.27 per cent under portfolio pricing, while

Table 5. Selected own- and cross-price elasticities: functional yogurts (specification Rest1)

			Non-drinkable							Drinkable				
			Danone					Parmalat		Danone		Nestlé		Granarolo
			Plain		Fruit		Others Whole	Fruit Whole	Others Whole					
			Skim	Whole	Skim	Whole								
Non-drinkable														
Danone	Plain	Skim	−3.06***	0.67***	0.31***	0.27***	0.27***	−0.03	−0.03	0.36***	0.36***	0.07*	0.07*	0.07*
		Whole	0.51***	−3.07***	0.21***	0.21***	0.21***	−0.01	−0.01	0.28***	0.28***	0.06*	0.05*	0.05**
	Fruit	Skim	0.31***	0.28***	−3.05***	0.67***	0.29***	0.37***	−0.02	0.38***	0.39***	0.08**	0.08**	0.08**
		Whole	0.08***	0.08***	0.19***	−1.78***	0.10***	0.12***	0.00	0.10***	0.11***	0.03**	0.02***	0.02***
Parmalat	Others	Whole	0.04***	0.04***	0.04***	0.05***	−1.38***	0.00	0.06***	0.05***	0.05***	0.01**	0.01**	0.01**
		Whole	−0.04	−0.03	0.52***	0.57***	0.00	−4.54***	0.48***	0.09**	0.11***	0.13**	0.11**	0.11***
	Others	Whole	−0.03	−0.03	−0.03	0.00	0.40***	0.36***	−3.75***	0.06**	0.05***	0.09**	0.08**	0.07**
Drinkable														
Danone		Skim	0.12***	0.10***	0.10***	0.10***	0.10***	0.11*	0.02*	−1.58***	0.08***	0.00	0.00	0.00
		Whole	0.04***	0.04***	0.04***	0.04***	0.04***	0.04**	0.01**	0.03***	−1.22***	0.00	0.00	0.00
Nestlé		Skim	0.11**	0.06**	0.07**	0.07**	0.08***	0.07**	0.08**	−0.02	0.01	−2.86***	0.00***	0.00
		Whole	0.09*	0.05*	0.05**	0.05**	0.05***	0.05**	0.05**	−0.02	0.00	0.00***	−2.18***	0.21
Granarolo		Whole	0.03**	0.02**	0.02**	0.02***	0.02***	0.02***	0.02**	−0.01	0.00	0.00	0.08	−1.50***

Note: ‘Others’ indicates ‘other flavours’.
 *10 per cent significance level.
 **5 per cent significance level.
 ***1 per cent significance level.

Table 6. Short-run price-cost margins (per cent)

Vendor	Flavour	Fat content	Single-product	Multi-product
Conventional				
Danone	Plain	Skim	23.81	36.30
		Whole	23.34	35.65
Granarolo	Fruit	Skim	76.75	79.03
		Others	43.83	50.61
	Plain	Skim	20.15	35.17
		Whole	18.64	32.43
Müller	Fruit	Skim	30.90	51.00
		Whole	68.89	74.29
	Others	Whole	42.14	49.79
		Plain	43.94	48.24
Nestlé	Fruit	Skim	20.28	33.28
		Whole	66.85	69.96
	Others	Whole	33.78	39.59
		Skim	33.08	34.11
Parmalat	Fruit	Skim	14.56	17.17
		Whole	32.82	41.95
	Others	Whole	59.61	63.48
			16.44	22.78
Functional				
Danone	Plain	Skim	32.66	66.86
		Whole	32.59	58.85
	Fruit	Skim	32.79	69.66
		Whole	56.22	73.45
Parmalat	Others	Whole	72.67	80.45
		Fruit	22.03	25.28
	Others	Whole	26.64	28.95
Functional/drinkable				
Danone		Skim	63.31	67.37
		Whole	81.81	83.40
Granarolo		Whole	34.93	34.93
Nestlé		Skim	45.78	52.79
		Whole	66.58	69.33

Note: 'Others' indicates 'other flavours'.

the average PCMs for functional, non-drinkable yogurts are 39.73 and 57.64 per cent, respectively, under the two different pricing scenarios. The average PCMs for drinkable functional products are 58.48 and 61.56 per cent for single-product Bertrand and portfolio pricing, respectively.

On average, Danone is the brand with the largest profitability. Among the conventional ones, the largest PCMs are obtained for Danone/fruit/skim (76.65 and 79.13 per cent for the two pricing scenarios); among non-drinkable functional ones, the largest profitability is obtained for Danone other flavours/whole (72.67 and 80.45 per cent); and for Danone/whole among drinkable ones (81.81 and 83.40 per cent).

Although the estimated PCMs for some functional alternatives reach values above 80 per cent, the reader should remember that they represent short-run profit margins, which are upper bounds to the actual products' profitability. Also, such values can be overestimated in the case of functional yogurts more than in that of conventional ones: as larger development and advertising costs characterise functional products, the results presented above are not able to capture in full whether functional yogurts sold in the Italian market are in fact profitable in the long run. However, the values presented in the bottom part of Table 6 may be seen as an *upper bound* to the measure of the profitability of functional yogurts *after* long-run investment costs are recovered.

5. Concluding remarks

As consumers' interest for nutraceutical food products grows, food manufacturers may see the development of functional products as an opportunity to revive mature markets. Despite many studies focused on understanding consumers' acceptance of functional foods, they have disregarded several dimensions (brand loyalty, switching behaviour between conventional and functional alternatives), which may play a major role in their increasingly complex markets.

This article has analysed the demand for functional and conventional products, and their profitability, using the Italian yogurt market as a case study via a flexible and parsimonious methodology, the DM method, applied to the LA/AIDS model. Results show that brand loyalty plays an important role in the Italian yogurt market and that the success of functional products is heavily influenced by it. Results also show that consumers of functional yogurts tend to be less price-sensitive than those of conventional ones, and that superior *short-run* performances are associated with the presence of a functional attribute.

Furthermore, both intra- and inter-brand substitution patterns across functional and conventional yogurts favour the former in most cases, supporting the existing evidence that consumers show remarkable interest for functional products. The results indicate that consumers buying non-drinkable yogurts may not be likely to switch to drinkable ones as price changes (and vice versa), suggesting that the success of drinkable yogurts may be due to an increase in their consumer base. Lastly, the results indicate that, as 'switching' between functional and conventional products produced by different manufacturers appears unlikely, yogurt manufacturers operating in the Italian market could expand their consumers' base via introducing new functional products, successfully avoiding sales cannibalisation.

This analysis does however show limitations. In the first place, although the analytical framework used (i.e. a 'representative consumer' DM-LA/AIDS model) served well the purposes of the present work, it does not allow to infer on other features which may characterise the Italian functional yogurt market (for example, a mixed-logit could allow to investigate consumers' heterogeneous responses to functional attributes). Second, the legitimacy of the results relies on the assumptions made regarding the product characteristics

used as determinants of product closeness and as parameters' shifters; such assumptions implicitly rule out that other (unobservable) features (such as texture) may impact substitution patterns and/or own-price responsiveness. As a result, our findings apply to those consumers whose preferences for yogurt do not depend on such unobserved characteristics.

Furthermore, these findings describe a scenario prior to the implementation of Regulation (EC) No. 1924/2006, and are unlikely to be representative of a highly regulated functional foods' market. As the implementation of the regulation has so far resulted in many health claims being denied, scholars should assess whether potential losses in profitability by functional food manufacturers could in fact result in what pundits have referred to as an 'innovation wasteland' (Starling, 2009). To this end, more research is needed to understand the long-run profitability of food manufacturers' strategic investments in functional products. Furthermore, an analytical framework similar to that implemented in this article could be used to evaluate the overall economic impact of Reg. (EC) No. 1924/2006, assessing potential welfare changes for both consumers and manufacturers via counterfactual analysis, simulating changes in the market equilibrium conditions and social welfare due to the disappearance of some of the functional alternatives.

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