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Consumer willingness to pay for traditional food products

Péter Balogh [a], Dániel Békési [b], Matthew Gorton [c,*], József Popp [a], Péter Lengyel [a]

[a] Faculty of Economics and Business, University of Debrecen, Böszörményi u. 138, H-4032 Debrecen, Hungary
[b] Department of Economics, Vienna University of Economics and Business, Welthandelsplatz 1, A-1020 Vienna, Austria
[c] Marketing, Operations and Systems Subject Group, Newcastle University Business School, 5 Barrack Road, Newcastle upon Tyne, UK

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ABSTRACT

Reflecting the growing interest from both consumers and policymakers, and building on recent developments in Willingness to Pay (WTP) methodologies, we evaluate consumer preferences for an archetypal traditional food product. Specifically we draw on stated preference data from a discrete choice experiment, considering the traditional Hungarian mangalitza salami. A WTP space specification of the generalized multinomial logit model is employed, which accounts for not only heterogeneity in preferences but also differences in the scale of the idiosyncratic error term. Results indicate that traditional food products can command a substantial premium, albeit contingent on effective quality certification, authentic product composition and effective choice of retail outlet. Promising consumer segments and policy implications are identified.

Introduction

EU farmers confront worsening terms of trade and declining real incomes, and generally remain dependent on direct payments and other subsidies for survival (European Commission, 2014). In 2012, subsidies accounted for over 50 per cent of EU farmers' net income, and farm incomes are lower in absolute and relative terms in the New Member States (NMS) from Central and Eastern Europe (European Commission, 2014). Rural areas in the NMS are more dependent on agriculture as a source of income and employment, with opportunities for gainful employment in the non-farm rural economy relatively scarce (Davidova et al., 2013).

To boost competitiveness and profitability, the EU seeks to stimulate enhanced value-added production, drawing on its reputation for quality goods (European Parliament and the Council of the European Union, 2012). One potential type of quality goods are Traditional Food Products (TFPs). A traditional food may be classified as: 'a product [...] made accurately in a specific way according to the gastronomic heritage, [...] and known because of its sensory proprieties and associated with a certain local area, region or country' (Guerrero et al., 2009, p. 348). These goods generally possess positive images due to superior taste, nostalgia and/or ethnocentrism (Almli et al., 2011; Vanhonacker et al., 2010). However, the ability of TFPs to contribute to improved farm incomes, without recourse to subsidies, depends on whether consumers are willing to pay a premium for them compared to cheaper alternatives. In other words, with TFPs not receiving any direct, supplementary subsidies, additional value added has to come on the demand side but the willingness of consumers to pay for such goods, and specific attributes that may be attached to them, remains unclear.

The paper addresses this central question, building on recent advances in Willingness to Pay (WTP) methodologies, which are applied to an exemplary case of a Traditional Food Product (TFP) – that of Hungarian mangalitza salami. Mangalitza salami is an ideal product for exploring WTP for a TFP as the main motivation for its purchase in Hungary, as discussed below, is its indigenous origin and heritage. Data collection occurred in the Northern Great Plain of Hungary, a lagging region, which is characterized by a relatively high dependence on agriculture and real farm incomes below the EU average (MARD, 2011). The study seeks to understand consumer perceptions of value and identify promising segments for targeting. Specifically, using a Discrete Choice Experiment (DCE), explicitly accounting for unobserved heterogeneity in correlated WTP coefficients and observable demographic/socio-economic characteristics, our model is estimated in WTP space. Train and Weeks (2005) advocate this approach and reparametrize the random parameter (mixed) logit model (RPL) by defining the distribution of WTP directly. Nevertheless, despite the clear advantages of the WTP space framework, it has been used, notwithstanding some notable exceptions (e.g. Balcombe et al., 2010), infrequently in the food policy literature. As a result most previous food-related WTP studies assume that the price coefficient is fixed across consumers, so that

* Corresponding author.
E-mail address: matthew.gorton@newcastle.ac.uk (M. Gorton).

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the moments of WTP are equal to the moments of the non-monetary attribute coefficient scaled by the price coefficient. However, this is an unnecessarily strong assumption of homogeneous price sensitivities. Moreover, as Train and Weeks (2005) note, a fixed cost coefficient implies that the scale parameter, and consequently the variance of unobserved utility or the degree of certainty in decisions, is the same for all respondents. Hence, in such models, potential scale heterogeneity across decision-makers may be falsely attributed to variations in WTP.

Greene and Hensher (2010) demonstrate that the WTP space model is nested within the recently developed Generalized Multinomial Logit Model (GMNL) of Fiebig et al. (2010). This framework considers not only random preferences, but is also able to decouple preference heterogeneity from scale heterogeneity, which is related to differences in the variance of the error term (i.e. differences in the degree of randomness in the decision-making process). This is particularly relevant for studies that use stated preference data, as consumers may interpret and process choice tasks differently, so that the level of certainty regarding their choices may vary. Hence, applying the GMNL model, we demonstrate how best to account for preference and scale heterogeneity and also take the demographic characteristics of respondents into consideration. There are some notable works that follow a similar path, for instance Balcombe et al. (2009) estimate consumers' WTP for pesticide reductions. Dealing with baskets of goods rather than a specific product, Balcombe et al. (2010) examine WTP for reductions in various nutrients (fat, sugar salt). Zanoli et al. (2013) investigate Italian consumers' WTP for beef attributes while Campbell and Doherty (2013) consider opportunities for adding value to chicken meat. However, demographic variables are not incorporated into the analysis of Balcombe et al. (2010) and Campbell and Doherty (2013), so that the authors do not account for observable preference heterogeneity. Zanoli et al. (2013) attempt to trace such effects by comparing conditional parameter estimates of WTP between different demographic groups. Instead of that, following Train (2009), we apply a GMNL specification that includes interaction terms for WTP measures and demographic variables. Consequently, significant differences in monetary valuations of product traits can also be explained on the basis of observable respondent-specific attributes, making recommendations regarding market segmentation, targeting and positioning easier to define.

As a further important feature, unlike most other WTP space studies (with the notable exceptions of Balcombe et al., 2009, 2010), we allow for correlations between the random WTP coefficients of different attributes, which yield a more realistic picture of consumer preferences for particular food products (in this case mangalitza salami). However, unlike Balcombe et al. (2010), we focus on a specific product instead of an aggregated basket of goods. Particular attention is given to the effect of quality certification, retail channel, and the share of mangalitza in the salami, taking into account the relationships between consumer valuations of these product traits. This leads to set of practical recommendations for marketers and policy makers regarding promising consumer segments and strategies for improving the added value of TFPs.

The paper is structured as follows. It begins with an overview of the current literature on consumer attitude toward traditional foods and a description of the mangalitza case. Section 'Choice experimental design and data' documents the data and the design of the DCE. After that, the econometric models are specified. Section 'Results' presents the results prior to the discussion of conclusions.

Traditional Food Products (TFPs) and the mangalitza case

While there are few official definitions, the European Parliament and the Council of the European Union (2012) identifies TFPs as those with "proven usage on the domestic market for a period that allows transmission between generations; this period is to be at least 30 years". Consumers perceive that TFPs are: anchored in the past (Guerrero et al., 2009), tied to specific localities, regions or countries and typically evoke strong memories of childhood (Cerjak et al., 2014; Rudawska, 2014). Moreover, they regard knowledge as to how to produce and consume TFPs as being passed from one generation to the next, usually in a domestic setting or by artisans (Guerrero et al., 2009). TFPs possess also distinctive sensory merits (Molnár et al., 2011) which are generally evaluated positively (Almli et al., 2011). Importantly, consumers judge the merit of a particular TFP in terms of its authenticity (Tregear et al., 1998), with those perceived as genuine forming part of an area's gastronomic heritage (Guerrero et al., 2009).

Pieniak et al. (2009) model the relationships between food choice motivations, as defined by Steptoe et al. (1995), and attitude toward and consumption of TFPs. They found the importance placed on familiarity and the natural content of food to be positively associated with attitude to, and consumption of, TFPs. Conversely, the importance consumers placed on convenience and weight control were negatively related to attitude and consumption of TFPs. No significant relationships, however, were established between the degree to which consumers' valued sensory qualities, price sensitivity and attitude toward and consumption of TFPs. This may reflect the heterogeneity of TFPs as a category, which limits the degree to which general attitudes map on to purchases of specific goods.

TFPs typically, but not universally, have strong associations with a particular origin and locality (Verbeke et al., 2016). Some TFPs in the EU are thus suitable for and already are protected under one of three main designation schemes: Protected Designation of Origin (PDO), Protected Geographical Indication (PGI) and Traditional Speciality Guaranteed (TSG). These quality schemes seek to protect producers and consumers from inferior, copycat goods so that only members of an approved consortium can use a registered name. By 2015, there were approximately 1200 products, across the EU, which had confirmed PDO, PGI or TSG status and several high profile TFPs benefit from the legislation (De Roest and Menghi, 2000). However, the impact of the schemes in many countries is limited by poor consumer and producer awareness of the labels and a lack of understanding of the differences between the schemes (Gorton and Tregear, 2008; Tregear et al., 1998). In this environment, privately owned brand names are often more important quality signals to consumers than designation labels (Kizos and Vakoufaris, 2011; Tregear et al., 2007). In other cases, national schemes of certification or alternative quality labels may be of greater salience (De Pelsmacker et al., 2005). Moreover, not all TFPs are suitable for PDO, PGI or TSG designation – for instance where production is very diffuse or the good is now regarded as generic.

Studies modeling consumer preferences for geographical indications and other quality labels highlight their heterogeneity (Resano et al., 2012; van Ittersum et al., 2007). An important challenge is thus to understand the socio-economic and demographic determinants of such heterogeneity. This is best attempted in relation to a specific TFP; as given the heterogeneity of the traditional foods category, general assertions may be of limited value in understanding consumer choices related to specific products. As a result Molnár et al. (2011) calls for research that goes beyond 'general consumer perceptions and preferences relating to traditional foods as a food product category' (p. 237).

Mangalitza represents an ideal product for investigating consumer behavior relating to TFPs, as its appeal rests on its long, distinctive history and status as part of Hungary's gastronomic heritage. While it was the most common swine breed in Hungary until the latter half of the Nineteenth Century, by the late 1970s it had almost completely disappeared due to its inferior feed conversion,

meat/fat ratio and reproductive performance compared to commercial white breeds (Egerszegi et al., 2003). By 1990, there were only approximately 200 mangalitza pigs left, with the breed on the verge of extinction (Pocsai, 2012a). However, interest in the breed revived and by 2011 there were approximately 6200 sows and 115 controlled mangalitza breeders in Hungary (Pocsai, 2012a). It does not possess any protected status at European level (e.g. PDO, TSG) but there is co-ordination at the domestic level via the National Association of Mangalitza Breeders (NAMB). The NAMB certifies mangalitza pigs, officially guaranteeing the origin of genuine mangalitza products. However, in recent years the fortunes of mangalitza producers have been unstable and their commercial viability hinges on a sufficient number of consumers paying a price premium that offsets the breed's poorer feed conversion ratio and reproductive rate. Understanding consumers' WTP for mangalitza products and the best strategies for realizing this, in terms of product composition, labeling and retail channels is thus of paramount concern. This is addressed in the experiment presented below.

Choice experiment design and data

Design

Prior to the experiment, focus groups with experts from the Hungarian pork sector and with potential and actual buyers were conducted in order to identify salient product attributes and their levels. Table 1 presents the resulting set of attributes and levels, which relate to retail prices, quality certification, share of mangalitza meat in the salami and retail channels. All prices were expressed in the local currency – Hungarian Forints (HUF).

Credence quality attributes are impossible to assess at the point of purchase and after consumption (Darby and Karni, 1973). This is relevant in this case as, unfortunately; deception regarding mangalitza products is not uncommon. As the typical retail price for these goods is substantially above the price of regular pork, the incentive for counterfeiting and fraud is relatively strong (Pocsai, 2012b). In order to prevent market failure stemming from information asymmetry and to gain consumer trust, third-party certification, that ensures compliance with respective standards, is thus critical for product differentiation. In the case of mangalitza third-party quality certification is undertaken by NAMB.

The share of mangalitza meat in the salami was considered an important aspect of product differentiation. Including non-mangalitza salami would have been feasible; however, in order to keep the experiment as traceable as possible, this alternative was embedded into the 'no purchase' option (opt-out). Rather the share of mangalitza meat in the salami was included (levels of 50%; 75%; 100%). Finally, as the nature of the retail format (farmers' market; butcher/small independent retailers; hyper/supermarket) may have a significant effect on consumers' purchase decisions (De Pelsmacker et al., 2005; Vecchio, 2010), this was also taken into consideration. As in most of the literature, no actual monetary transactions occurred in our experiment. This might give rise to a hypothetical bias; however, evidence presented by Lusk and Schroeder (2004) suggests that in the case of estimating marginal WTP values, this bias is negligible, as results do not differ significantly from those of non-hypothetical settings.

In order to avoid the base levels being confounded with the intercept (opt-out), we use effects coding instead of dummy variable coding. As suggested by Bech and Gyrd-Hansen (2005), the base levels (no certification, 50% mangalitza share and farmers' market, for each of the three attributes) were set equal to the negative sum of the estimated coefficients of the other levels within an attribute. Consequently, effects for all levels can be estimated.

After defining the attributes and their levels, the full factorial design (all possible combinations of the attribute levels) consisted of $2^1 \times 3^2 \times 4^1 = 72$ hypothetical purchase scenarios (choice cards). It was infeasible for subjects to rate all possible combinations of these scenarios. Therefore a fractional factorial orthogonal main effects design (Kuhfeld et al., 1994) was created, where this number was reduced to 16. In the experiment, participants were presented with 8 choice sets (e.g. Table 2). Each set was augmented by the 'no purchase' option in order to resemble realistic choice occasions and to prevent forced choices that would otherwise overstate the likelihood of choosing any of the alternatives (Hensher et al., 2005).

The orthogonal main effects design assumes no interaction effects between attributes (i.e. their levels are independent from each other). As Louviere et al. (2000) and Lusk et al. (2003) suggest, the main effects typically explain a substantial portion of variance in the model and setups using this specification tend to have a good predictive ability.

Data

The questionnaire was piloted with 50 randomly selected consumers and subsequently modified. The final survey was carried out between August and October 2012 in the North Great Plain region of Hungary. This is the most significant region for breeding mangalitza pigs – nearly 40% of the total number of sows is bred here and it is also the location of NAMB's headquarters. In total, 309 individuals fully completed the questionnaire. Data collection occurred face to face.

We collected data in six towns and 18 villages in the region. In towns we intercepted consumers at five hyper- and supermarkets (Auchan, Tesco, Interspar/Spar, Match), seven farmers' markets and 11 butchers/small stores. In villages intercepts occurred at 23 butchers/small stores and 18 farmers' markets. From the middle of the 2000s onwards, hyper- and supermarkets in Hungary have sold mangalitza products, marketing them as derived from an ancient Hungarian but world-famous breed of pig.

Table 1
Mangalitza salami attributes, attribute levels and coding in the DCE.

Attribute	Level	Effects coding	
Retail price of mangalitza salami	1500 HUF/kg 2000 HUF/kg 2500 HUF/kg 3000 HUF/kg	Continuous variable	
Certification by NAMB	No	−1	–
	Yes	1	–
Share of mangalitza in the salami	50%	−1	−1
	75%	1	0
	100%	0	1
Retail channel	Farmers' market	−1	−1
	Butcher/small store	1	0
	Hyper-/supermarket	0	1

Table 2
Example of a set of choice cards.

	Product 'A'	Product 'B'	None of these products
Price (HUF/kg)	1500	2500	
Certification by NAMB	Yes	Yes	
Share of mangalitza	75%	50%	
Retail outlet	Hyper/supermarket	Butcher/small store	
Your choice	□	□	□

The questionnaire consisted of three parts. Firstly, respondents reported their actual pork buying habits (products purchased and retail outlets patronized) and whether they had consumed mangalitza products previously (variable *Experience*). Secondly, respondents completed the DCE. The third part focused on demographic/socio-economic questions. Respondents were selected using quota sampling, so as to be representative (regarding age, gender and residence) of the Northern Great Plain region of Hungary. In order to ensure discretion and maximize response rate, data on income was not revealed directly, but respondents sorted themselves into groups after being informed of the national mean level of monthly gross income per capita. Table 3 summarizes the characteristics of the sample. These characteristics are coded with the following base categories: male, lowest level of education, lowest income, and no experience.

Econometric methods

The random parameter (mixed) logit model (RPL)

The econometric analysis of the choice data is based on random utility theory. Individuals choose the alternative that gives them the highest possible level of utility from the available choice set. Given i.i.d. type I extreme value distributed error terms (ε), with the scale parameter being normalized to 1 (and thus their variance to $\Pi^2/6$), this specification is referred to as the standard Multinomial Logit Model (MNL) of McFadden (1974). Notwithstanding the obvious advantages of this formulation, like the straightforward interpretation and easy derivation of a simple closed-form expression for the choice probabilities, it is accompanied by some rather strong and restrictive assumptions. However, these can be relaxed. Mixed logit models (McFadden and Train, 2000) can capture unobserved preference heterogeneity, whereas the MNL assumes homogeneous preferences for product attributes. Second, the RPL relaxes the Independence of Irrelevant Alternatives (IIA) assumption by allowing for flexible variance-covariance structures for the unobservable portion of utility. Appendix A provides a detailed description of the method.

Table 3
Survey descriptive statistics.

Demographic/socio-economic measures	Sample (n = 309)	Northern Great Plain region[a]
Gender (%)		
Female	56.3	52.1
Male	43.7	47.9
Age of head (mean)	40.1	40.2
Residence		
Urban	62.8	68.1
Rural	37.2	31.9
Highest level of education achieved (%)		
(1): Elementary	8.8	–
(2): Secondary	51.1	–
(3): University	40.1	–
Monthly gross income (%)		
(1): Substantially below average	39.6	–
(2): Below average	18.4	–
(3): Average	25.2	–
(4): Above average	16.8	–
Experience – previously consumed mangalitza (%)		
Yes	51.5	–
No	48.5	–

http://www.ksh.hu/nepszamlalas/tables_regional_00?lang=en.
[a] Data of the Hungarian Central Statistical Office (KSH) 2011.

The generalized multinomial logit model (GMNL)

Earlier research recognizes that consumer behavior may depend not only on heterogeneity in preferences but also on differences in the scale of the idiosyncratic error term (Louviere et al., 2002). Scale heterogeneity might be interpreted as the variation of randomness in the decision-making process over respondents, i.e. the variance of the error term (and hence the degree of certainty) may differ across individual decision-makers. This is especially relevant for stated preference data, where respondents could interpret choice situations differently and pay varying levels of attention to the task presented (Train and Weeks, 2005). To tackle this, Fiebig et al. (2010) proposed the generalized multinomial logit model (GMNL). Unlike previous model types, where the scale of the error term is normalized to 1, the GMNL model attempts to decouple preference and scale heterogeneity by nesting the RPL specification within a more general framework. Appendix B details the method in greater depth.

Results

Before proceeding to the model estimations, we provide a brief overview of results regarding pork, and specifically mangaliza, buying habits as well as motives for purchase. The survey revealed that fresh pork is bought most often at butchers/small stores (51%), then in hyper-/supermarkets (29%) and farmers' markets (15%), with 5% of respondents not buying fresh pork. Multiple retailers account for a larger share of the market for processed pork products like salami, being the main outlet for purchase for 49% of respondents, compared to 34% and 8% for butchers/small stores and farmers' markets respectively (9% of the respondents never buy processed pork products). Regarding frequency of consumption, 43% of respondents consume fresh meat pork/pork products once a week or more often, 45% once a month, 5% once in every six months or less often and 7% not at all.

The survey elicited information regarding respondents' mangalitza consumption, observing that 51% of respondents (n = 159) regularly purchase or consume mangalitza products, with the remainder never having sampled mangalitza products. Most consumption is in the traditional, processed form (salami) rather than fresh meat. Overall, 31% of regular mangalitza consumers eat this TFP every month, 36% consume it approximately twice yearly and 23% eat it less than once every six months. A small proportion of respondents (9%) consume mangalitza once a week or more frequently. Questions regarding motives, amongst those who regularly purchase mangalitza products, reveal that most (n = 76) prefer this type of meat due to its indigenous origin and heritage, followed by perceived health benefits (n = 69) and perceived superior tastiness (n = 64). The most important self-reported reason for purchase, therefore relates to mangaliza's geographical origin and its status as a TFP.

Table 4 presents the results of three model specifications. The first column details the simple MNL model with all parameters being fixed. In the second, the RPL model is estimated with a price coefficient being lognormally distributed. Finally, in the GMNL-WTP-space model, the price coefficient is normalized to -1 and the attribute coefficients are to be interpreted as WTP values. The RPL and the GMNL-WTP-space models were estimated using the user-written Stata module of Gu et al. (2013). All models, where simulation was necessary, were estimated by using 1000 Halton draws. The upper part of Table 4 contains the means and standard deviations of the random (respondent-specific) coefficients. The scale is normalized to 1 in the first two models and only estimated for the GMNL specification. The bottom part shows the socio-economic interaction effects with the means of the

Table 4
Model estimates.

	MNL	RPL(lognormal price)		GMNL-WTP-space	
	Mean	Mean	SD	Mean	SD
Price/1000	−1.079***	−0.538***	0.927***	−1 (norm.)	–
	(0.081)	(0.109)	(0.050)	–	–
OptOut	−3.043***	−7.351***	5.379***	−6.251***	2.419***
	(0.411)	(1.051)	(0.525)	(0.892)	(0.300)
Certification	0.493**	1.048**	1.744***	0.504**	0.682***
	(0.228)	(0.490)	(0.199)	(0.203)	(0.062)
Share75	0.254	0.694**	0.228	0.236	0.298***
	(0.186)	(0.289)	(0.150)	(0.152)	(0.066)
Share100	0.481***	0.819**	0.887***	0.421**	0.305***
	(0.172)	(0.397)	(0.133)	(0.166)	(0.056)
Hypermarket	−0.771***	−1.499***	0.959***	−0.641***	0.458***
	(0.215)	(0.389)	(0.154)	(0.187)	(0.072)
Butcher	0.376**	0.921***	0.538***	0.291**	0.236***
	(0.191)	(0.341)	(0.139)	(0.130)	(0.059)
Scale (mean)	–	–	–	2.199***	–
	–	–	–	(0.145)	–
Scale heter. (τ)	–	–	–	0.557***	–
	–	–	–	(0.062)	–
OptOut × z					
Income4	0.224	−1.198*	–	−0.454	–
	(0.306)	(0.717)	–	(0.655)	–
Experience	−0.368*	−0.495	–	−0.666	–
	(0.191)	(0.469)	–	(0.437)	–
Certification × z					
Income4	−0.316*	−0.734*	–	−0.247	–
	(0.176)	(0.403)	-	(0.159)	-
Education3	0.449**	0.761	–	0.365*	–
	(0.217)	(0.469)	–	(0.197)	–
Share75 × z					
Experience	0.210**	0.324*	–	0.190**	–
	(0.100)	(0.174)	–	(0.079)	–
Hypermarket × z					
Income2	0.409**	0.558**	–	0.335***	–
	(0.165)	(0.279)	–	(0.130)	–
Income3	0.297*	0.366	–	0.224*	–
	(0.156)	(0.265)	–	(0.120)	–
Income4	0.324	0.547*	–	0.210	–
	(0.201)	(0.318)	–	(0.150)	–
Butcher × z					
Education3	−0.319*	−0.595*	–	−0.146	–
	(0.181)	(0.318)	–	(0.125)	–
No. of obs.: 7416	Pseudo R^2	LL	AIC	BIC	
MNL	0.161	−2278.597	4655.194	4993.852	
RPL (lognormal price coeff.)	0.315	−1861.091	3876.181	4408.358	
GMNL-WTP-space	0.293	−1918.748	3979.497	4470.206	
Base levels of effects codes (means)					
No certification: −0.493 (MNL); −1.048 (RPL); −0.504 (GMNL)					
Share50: −0.735 (MNL); −1.513 (RPL); −0.657 (GMNL)					
Farmers' market: 0.395 (MNL); 0.578 (RPL); 0.351 (GMNL)					

(Clustered, robust Std. Err.).
* Denote statistical significance at the 10% level.
** Denote statistical significance at the 5% level.
*** Denote statistical significance at the 1% level.

coefficients. Similar to other DCE studies using effects coding (e.g. Collins et al., 2012), we report the base levels of utility coefficients. Due to effects coding, these are the negative sum of the other levels of the given attribute. Consequently, unlike in the case of dummy coding, these are not confounded with the ASC (alternative specific constant) or with each other.

There is clear evidence against opt-out; on average, consumers have a strong preference for mangalitza salami. Nearly all attribute (and WTP) coefficients – for almost all specifications – are statistically significant for the base demographic segment justifying the appropriate choice of product characteristics. On the other hand, some attribute means are not significantly different from zero. In the simple MNL model, this could lead to the conclusion that respondents do not think that such a product trait is important regarding their choice decisions. However, this would be inappropriate, as all the standard deviations of the coefficients in the GMNL model are significantly different from zero. Consumers simply weight the attributes differently, and averaging these weights across respondents leads to statistical insignificance. The RPL and GMNL specifications account for unobserved preference heterogeneity which is obscured in the MNL results (Hu, 2006).

The random parameters are defined as being correlated with each other. Table 6 presents their full variance-covariance matrix for the GMNL-WTP-space model. All random models were first estimated with uncorrelated coefficients. Compared to these, the correlated versions, presented in Table 4, provide a better fit to the data. Besides preference heterogeneity, we also find statistically significant scale heterogeneity; therefore the assumption of identical scales across individuals should be rejected. Concerning model fit, Wald tests reveal the joint significance of the explanatory variables in each model.

Comparing the specifications, it can be seen that the RPL model fits the data better than the other two approaches. This is similar to Train and Weeks' (2005) results but from the marketer's standpoint, where existing moments (especially means) of WTP distributions are required, in contrast to approaches with simulated ratios of coefficients, our GMNL-WTP-space model yields much more plausible results. These are revealed in the values reported in Table 5 since there are extremely high differences between the mean and median estimated by the RPL model. This reflects that in spite of the model's good fit, the estimated distributions are heavily skewed. This approach also yields standard deviations in excess of those of the GMNL-WTP-space model. After examining the outliers it can be stated that the percentiles, the minimum and the maximum values are not in an economically plausible range. These indexes are observed for the GMNL-WTP-space model in a significantly lower range.

Following Hole and Kolstad (2012), we use simulation with 10^7 random draws from the attribute coefficient distributions of the RPL model (which are then divided by draws from the distribution of the price coefficient) and WTP distributions of the GMNL model. Table 5 details the WTP distributions for the base demographic segment regarding the respective model variants. Regarding the three models, results are identical in signs but Table 5 reveals that the three specifications lead to quantitatively different WTP values. This is relevant for marketers and policy makers seeking to justify a price premium for these attributes. As a result of effects coding, it is possible to calculate the monetary valuation of all attribute levels. In the case of a binary attribute like certification, the absence of the attribute (i.e. "no certification") is associated with the negative of the coefficient. Hence the utility difference between a product with and without certification is the coefficient of certification multiplied by 2. Generally, the WTP to switch from one level of an attribute to another is the difference in the corresponding coefficients (Collins et al., 2012).

On average, individuals evaluate the presence of quality certification positively and are willing to pay a range of HUF914 (€3.05)–HUF1884 (€6.28) more for a product with a certificate as compared to one without (i.e. $WTP_{certification}$–$WTP_{no\ certification}$). This is in line

Table 5
WTP estimates (base segment).

Attribute/level	Model	Mean	Median	SD.	CV	Percentile (5%)	Percentile (95%)	Min.	Max.
Certification	RPL	0.942	0.451	2.648	2.811	−1.463	4.868	−124.619	264.666
	GMNL	0.504	0.504	0.682	1.352	−0.617	1.626	−3.115	4.185
	MNL	0.457	0.457	0	0	0.457	0.457	0.457	0.457
No certification	RPL	−0.942	−0.451	2.648	−2.811	−4.868	1.463	−264.666	124.619
	GMNL	−0.504	−0.504	0.682	1.352	−1.626	0.617	−4.185	3.115
	MNL	−0.457	−0.457	0	0	−0.457	−0.457	−0.457	−0.457
Share50	RPL	−1.359	−0.769	2.029	−1.493	−4.662	−0.002	−173.128	32.197
	GMNL	−0.657	−0.657	0.426	−0.649	−1.358	0.044	−2.857	1.543
	MNL	−0.680	−0.680	0	0	−0.680	−0.680	−0.680	−0.680
Share75	RPL	0.623	0.384	0.793	1.273	0.070	1.947	−1.908	67.470
	GMNL	0.236[a]	0.236	0.298	1.261	−0.253	0.726	−1.452	1.758
	MNL	0.235[a]	0.235[a]	0	0	0.235[a]	0.235[a]	0.235[a]	0.235[a]
Share100	RPL	0.736	0.376	1.495	2.032	−0.432	3.060	−46.540	109.541
	GMNL	0.421	0.421	0.305	0.725	−0.081	0.923	−1.092	1.993
	MNL	0.445	0.445	0	0	0.445	0.445	0.445	0.445
Hypermarket	RPL	−1.347	−0.755	2.059	−1.529	−4.680	0.033	−193.496	38.370
	GMNL	−0.641	−0.641	0.458	−0.714	−1.395	0.112	−3.271	1.686
	MNL	−0.715	−0.715	0	0	−0.715	−0.715	−0.715	−0.715
Butcher	RPL	0.827	0.471	1.220	1.474	0.013	2.812	−17.079	112.246
	GMNL	0.291	0.291	0.236	0.811	−0.097	0.679	−0.935	1.525
	MNL	0.349	0.349	0	0	0.349	0.349	0.349	0.349
Farmers' market	RPL	0.519	0.244	1.638	3.154	−1.024	2.906	−70.682	164.398
	GMNL	0.351	0.351	0.515	1.470	−0.497	1.198	−2.553	3.286
	MNL	0.366	0.366	0	0	0.366	0.366	0.366	0.366

[a] The mean of the attribute coefficient was not significantly different than zero.

with the literature on the value of third party certification for ensuring the veracity of credence quality attributes (van Ittersum et al., 2007; Zanoli et al., 2013). Underpinning the finding regarding the general preference for this product, salami made from purebred mangalitza is more attractive to consumers relative to crossbreed alternatives.

Interestingly, concerning the place of purchase, hyper- and supermarkets are less preferred compared to farmers' markets and smaller butchers. Decision-makers value the attribute of farmers' markets at HUF366/€1.22 (MNL), HUF519/€1.73 (RPL) and HUF351/€1.17 (GMNL), hence they are willing to pay a premium of HUF1081 (€3.60), HUF1866 (€6.22) and HUF992 (€3.31) compared to hyper- or supermarkets in the respective models (i.e. $WTP_{farmers'market}$–$WTP_{hyper\text{-}supermarket}$). This is in keeping with how consumers tend to trust and prefer smaller, local producers for speciality food goods (Kneafsey, 2012; Tregear and Ness, 2005; Vecchio, 2010).

In the case of quantifying the implicit value of product attributes, differences across respondents are significant. Moreover, for each attribute, there are consumers who evaluate them contrary to the population means (i.e. 23%, 21%, 8%, 8% and 11% of respondents for the GMNL attribute coefficients in Table 4, respectively). For market segmentation, these buyers should be accounted for; instead of just dealing with the mean WTP values, this heterogeneity should be taken into consideration. Enhanced understanding of consumer monetary valuation is a major advantage of random coefficient models compared to the MNL which assumes preference homogeneity. Furthermore, as differences in WTP are rather substantial between the RPL and GMNL specifications, the role of scale heterogeneity appears important.

Next, we discuss the relationships between attribute valuations. Focusing on the model variant that yields the most plausible results, namely the GMNL-WTP-space specification, Table 6 details the variance-covariance matrix of WTP coefficients. There is a statistically significant positive association between the valuations of the presence of quality certification and 75%-share mangalitza, i.e. those consumers who exhibit a higher WTP for certified salami are more likely to value the 75%-share alternative as well. At the same time, these individuals are prepared to pay less for hyper- and supermarket products, which may be explained by the quality-consciousness of these respondents and a belief in small, local outlets. At the same time, those consumers, who value butchers and independent, small stores more, have a lower WTP for hyper- and supermarkets, which is not surprising, given the competing and contrasting features of these retail channels.

Considering the link to opting out, a less positive utility from choosing mangalitza salami (i.e. a larger OptOut coefficient) is associated with higher WTP for the presence of certification and for salami made from purebred mangalitza. Individuals who are less inclined to choose the product, and are also less experienced (Table 4) place greater emphasis on quality certification. This is an important policy finding concerning the marketing of TFPs. Certification, by decreasing information asymmetry, has a stronger effect on consumers who are *ceteris paribus* less decisive in their choices and who have relatively weaker preferences for this particular good. For expanding the customer base of TFPs, it is thus vital to highlight the presence of quality certification on packaging.

As for the socio-economic interaction terms, experience, income, and the level of education seem to significantly influence consumer preferences and WTP. In the first model, use-experience, as intuitively expected, has a statistically significant negative effect on opting out. Respondents who have already tried mangalitza products are less likely to forgo the consumption of mangalitza salami. This suggests that stimulating product trial has an important spill-over effect on WTP. Consumers with more experience also prefer the 75%-share alternative more compared to their inexperienced counterparts. This might also underpin the positive effect of prior consumption on current preferences. Higher-income respondents seem to dislike hypermarkets less than lower-income individuals. Wealthier decision-makers are likelier to own a car, which makes out-of-town hypermarkets more accessible.

Finally, in keeping with Nayga et al. (1998) and Drichoutis et al. (2005), our results reveal that individuals with university

Table 6
Variance-covariance matrix of WTP coefficients – GMNL-WTP-space model.

	β_{OptOut}	$\beta^*_{Certification}$	$\beta^*_{Share75}$	$\beta^*_{Share100}$	$\beta^*_{Hypermarket}$	$\beta^*_{Butcher}$
β_{OptOut}	5.85*** (1.452); [1]	1.00*** (0.235); [0.60]	0.00 (0.184); [0.00]	0.36** (0.177); [0.49]	−0.32 (0.225); [−0.29]	0.17 (0.142); [0.29]
$\beta^*_{Certification}$		0.46*** (0.084); [1]	0.12*** (0.043); [0.59]	0.05 (0.046); [0.24]	−0.09* (0.047); [−0.29]	0.02 (0.032); [0.14]
$\beta^*_{Share75}$			0.09** (0.039); [1]	0.02 (0.023); [0.18]	−0.09*** (0.035); [−0.66]	0.02 (0.019); [0.31]
$\beta^*_{Share100}$				0.09*** (0.034); [1]	−0.02 (0.034); [−0.11]	−0.02 (0.025); [−0.23]
$\beta^*_{Hypermarket}$					0.21*** (0.066); [1]	−0.09*** (0.034); [−0.86]
$\beta^*_{Butcher}$						0.06** (0.028); [1]

(Clustered, robust Std. Err.); [Correlation coefficients].
* Denote statistical significance at the 10% level.
** Denote statistical significance at the 5% level.
*** Denote statistical significance at the 1% level.

education place greater importance on quality certification. Such consumers are more likely to comprehend and better interpret this additional information. Hence a viable option to widen the reach of such quality certification would be to formulate its core message as clearly and plainly as possible. All in all, these socio-economic/demographic measures are of great relevance, and – unlike in many similar works that fail to account for such effects in WTP measures – their analysis contributes to a better understanding of consumer behavior and identifying strategies for improving the marketing of TFPs.

Conclusion

We investigate consumer monetary valuations (WTP) of attributes for a specific Traditional Food Product (TFP) drawing on the case of Hungarian mangalitza salami. The retail price, quality certification, retail channel and meat source (pure versus cross-breed) are identified as important attributes in consumers' decisions. Results indicate that the proposed model specification proves to be especially useful from a policymakers and marketer's perspective, as it yields – unlike simulated ratios of coefficients – easily interpretable and plausible WTP distributions. The Discrete Choice Experiment (DCE) reveals that generally consumers prefer quality certification, products from farmers' markets and salami made from entirely mangalitza meat. However, all the WTP values of the tested product attributes exhibit considerable variation across respondents.

Three key policy implications can be drawn. Firstly, TFPs can command a substantial price premium compared against mainstream alternative products. This is possible even in lagging EU regions and evidenced both by the DCE results and actual retail mark-ups. There is thus an opportunity for TFPs to increase the added value of farm production, through 'demand-side' innovation rather than relying on subsidies. Secondly, the choice of retail outlet matters, with, overall, butchers and farmers' markets commanding a substantial premium compared to hyper- and supermarkets. Several initiatives seek to promote short supply chains and farmers' markets in Hungary and elsewhere (Benedek and Balázs, 2015; Kneafsey, 2012). However, while these can help improve returns to some farmers, such outlets remain marginal with multiple retailers possessing a dominant and increasing market share throughout the EU, including in its New Member States (Dries et al., 2004; Euromonitor, 2014). Producers thus face a trade off in selecting marketing channels. While independent butchers and farmers' markets offer the prospect of higher mark ups, this may not, in the case butchers, always transmit into higher producer prices and throughput remains small. Multiple retailers offer opportunities for far greater sales but lower price premiums. Thirdly, effective certification and regulatory systems are vital to realize higher mark ups and protect the integrity of TFPs. This is evident in the mangalitiza case, i.e. even where certification is at the local level and not backed by a PDO/PGI official designation. Certification is particularly important for increasing the customer base – as inexperienced consumers and those who have relatively weaker preferences for the good place greater emphasis on quality certification. Unfortunately, many quality labels possess inadequate regulatory systems (European Court of Auditors, 2011), resulting from inexperience and limited resources. There is a consequent need thus to share experiences between successful TFPs, which command substantial premiums and possess robust regulatory systems, with those less well developed.

The analysis identifies the importance of consumers' prior experience – individuals who have not consumed mangalitza previously are significantly less likely to choose the product. This highlights both a market strength and weakness of TFPs. Consumers associate TFPs with habits and heritage passed from one generation to another, which can generate a substantial price premium. However, traditions cannot easily be exported, so the TFP outside of its 'area of influence', without the emotional attachment of past experience, is likely to be perceived as another conventional, competing product within a crowded marketplace. While new traditions may be created and imported, for instance Verbeke et al. (2016) notes the example of couscous in France, the conversion of a non-traditional to a TFP by its very nature takes time. Rapid growth in sales of a TFP outside of its customary area of influence will thus have to depend far less on an appeal to tradition.

From a methodological perspective, the results demonstrate how such analysis can be extended to other markets to generate insights into consumer preferences for specific product attributes. Producers and retailers can also benefit from the model specification of correlated coefficients, as it reveals the interconnected nature of these monetary valuations. The presented framework may be applied readily to other food product categories, as coefficient and scale heterogeneity are very likely to be important features of consumer behavior in other cases as well. As far as further research is concerned, the profiling of consumer segments could be enriched by considering psychological traits as well as the demographic/socio-economic characteristics detailed here. Comparisons between hypothetical and non-hypothetical settings

in choice experiments, as well as with actual purchase behavior would also enhance our understanding.

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Appendix A. Description of the RPL method

The utility respondent n receives from alternative j is given by Eq. (1) ($n = 1, \ldots, N$; $j = 1, \ldots, J$) with a part, which is observable to the researcher and a random one known only by the individual.

$$U_{nj} = V_{nj} + \varepsilon_{nj} \tag{1}$$

The systematic component V_{nj} is a linear function of the mangalitza salami attributes in vector x_{nj} and the vector of utility weights (β) for each attribute.

$$U_{nj} = \beta' x_{nj} + \varepsilon_{nj} \tag{2}$$

In Eq. (3), β_n is partitioned into a mean part ($\bar{\beta}$) and a person n-specific deviation (η_n).

$$U_{nj} = \beta_n' x_{nj} + \varepsilon_{nj} = (\bar{\beta}' + \eta_n') x_{nj} + \varepsilon_{nj} \tag{3}$$

Following Train (2009), the probability of individual n's observed choice sequence $y_n = \{y_{n1}, \ldots, y_{nT}\}$ from J alternatives described by the vector of product attributes x is:

$$P(y_n|x_n, b, W) = \int_\beta \prod_{t=1}^{T} \left\{ [exp(\beta' x_{ny_{nt}t})] \Big/ \left[\sum_{j \in J} exp(\beta' x_{njt}) \right] \right\} \cdot \phi(\beta|b, W) d\beta \tag{4}$$

This mixed logit formula is a weighted average of the MNL probability calculated at different values of β. The weight is the probability density (ϕ) of β over respondents with mean b and variance-covariance matrix W.

Appendix B. Description of the GMNL-WTP-space method

Starting from the simple logit model (for the sake of traceability, suppressing the choice occasion index, t) with the scale parameter (σ) of the error term (ε) made explicit and being respondent-specific, we obtain:

$$\widetilde{U}_{nj} = \beta_n' x_{nj} + \varepsilon_{nj}/\sigma_n \tag{5}$$

where $\varepsilon_{nj}/\sigma_n$ has the variance of $\Pi^2/(6\sigma_n^2)$. Simply rewriting the above expression by multiplying both sides by σ_n, the following equivalent formulation is obtained, which is referred to as the GMNL-II specification by Fiebig et al. (2010):

$$U_{nj} = (\sigma_n \beta_n') x_{nj} + \varepsilon_{nj} \tag{6}$$

As β and σ cannot be separately identified, Fiebig et al. (2010) propose the following general specification of the scale parameter: $\sigma_n = exp(\bar{\sigma} + \theta' z_n + \tau \varepsilon_{0n})$, where $\varepsilon_{0n} \sim N(0, 1)$ and z_n is a vector of individual characteristics, with $\bar{\sigma} = -\tau^2/2$ so that $E(\sigma_n) = 1$, when $\theta = 0$. Note that the GMNL model reduces to the RPL specification if $\tau = \theta = 0$, as τ provides a measure of scale heterogeneity; and to MNL if $\tau = \theta = var(\eta_n) = 0$.

This framework also nests the model of Train and Weeks (2005). As detailed by Greene and Hensher (2010), the GMNL model can be reparametrized to estimate taste parameters in WTP space. First, separating the price variable (p) and its coefficient ($\beta_{p,n}$), we obtain:

$$U_{nj} = \sigma_n(-\beta_{p,n} p + \beta_n' x_{nj}) + \varepsilon_{nj} = \sigma_n \beta_{p,n}(-p + (\beta_n'/\beta_{p,n}) x_{nj}) + \varepsilon_{nj} \tag{7}$$

Normalizing the price coefficient ($\beta_{p,n}$) of $-p$ to 1 yields the WTP space specification in Eq. (8), where $\beta_n'^*$ directly gives the individual-specific WTP estimates.

$$U_{nj} = \sigma_n(-p + \beta_n'^* x_{nj}) + \varepsilon_{nj} \tag{8}$$

This formulation bypasses the necessity of specifying the distribution of the ratio of two random coefficients, as in traditional preference-space models, which might lead to skewed and dubious WTP distributions.

Fiebig et al. (2010) note that the model performs relatively poorly if the alternative-specific constant (in our case, the opt-out alternative) is scaled, because it is fundamentally different from observed attributes. Hence in the final specification, vector $\boldsymbol{x}$ includes only observed attributes of the product from Table 1, which are absent in case that the opt-out option is chosen (Eq. (9)). In Eq. (10), the opt out coefficient ($\beta_{optout,n}$) is split into three parts: the component $\beta_{0,j}$, which is constant across respondents – i.e. the mean coefficient for the base demographic segment; γ_j, in order to account for observed heterogeneity in the mean coefficient regarding preferences for mangalitza salami in general – captured by demographic variables (z); and the individual-specific deviation, $\eta_{0,n,j}$. The same specification applies to the WTP coefficients ($\beta_n'^*$), where μ_j is the vector of the demographic effects that influence the mean of WTP.

$$U_{nj} = \beta_{optout,n} + \sigma_n(-p + \beta_n'^* x_{nj}) + \varepsilon_{nj} \tag{9}$$

$$U_{nj} = (\beta_{0,j} + \gamma_j' z_n + \eta_{0,n,j}) + \sigma_n[-p + (\beta_j'^* + \mu_j' z_n + \eta_{n,j}'^*) x_{nj}] + \varepsilon_{nj} \tag{10}$$

We assume that the entire vector ($\beta_{optout,n}$, $\beta_n'^*$) has a multivariate normal distribution with correlated coefficients. This is appropriate, as there are no clear expectations on the signs of these coefficients and allowing for correlations can shed more light on the structure of preferences for different attributes. As there is no closed-form expression for its likelihood function, this final GMNL specification, similar to RPL, can be estimated by using simulated maximum likelihood methods.

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