Exploring scale effects of best/worst rank ordered choice data to estimate visitors' benefits from alpine transhumance

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Abstract

In many environmental valuation applications standard sample sizes for choice modelling surveys are impractical to achieve. One option to improve data quality is to use more in-depth surveys conducted on fewer respondents. This is certainly the case for studies on mountain visitation. We report on a study using high quality rank-ordered data in which the ranking of alternatives is elicited by means of the best-worst approach to alternative selection. The resulting exploded choice model involved the collection of 64 responses per person which were used to study the willingness to pay for external benefits produced to visitors by policies keeping in place the artifacts of alpine transhumance. The context of study is Val di Genova, a valley with summer pastures located in the North Eastern Alps where we study visitors WTP for pasture landscape, biodiversity, historical heritage and the up-keep of in-situ milk transformation. We find good evidence in support of this approach and find reasonable estimates of mean WTP, which appear theoretically valid.

Keywords: best/worst alternative selection, rank ordered choice models, heteroskedastic logit, non-market valuation

JEL classification: C25, H41, Q26, Q51

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I. Introduction

The purpose of this paper is to report the results of a multi-attribute stated preference study based on a survey of visitors to Val Genova, in the North Eastern Alps. The survey was designed to overcome the practical difficulties related to the particular good under valuation, while at the same time addressing several modelling issues at the forefront of stated choice research for non-market valuation. In particular, the stated choice approach we used needed to efficiently extract high quality information from respondents. Respondents in these locations are hard to intercept and approach. When they are located, they typically are in the midst of a mountain visit, and so have a high marginal cost of time making them reluctant to engage in surveys, leading to high numbers of refusals. Thus, the stated choice survey was designed as an exercise that would provide a full ranking of the alternatives in each choice set by posing most favourite/least favourite questions. Rank-ordered choices are well known (Hausmann and Ruud 1987) to provide analysts with richer preference information than simply asking a respondent to state their favourite alternative and/or provide partial rankings. In particular, the experimental design approach that we used was specifically designed to get as close as possible to estimating respondent-specific preferences (i.e., a specific model for each respondent). To do this we addressed a number of issues, such as a) estimation of scale differences across choices by the same respondents using a heteroskedastic logit model; b) incorporation of additional intra-personal taste heterogeneity with correlation; c) assessment of best/worst question approach for choosing alternatives as a way to derive high quality complete choice set rankings; and d) derivation and validation of respondent-specific WTP estimates.

The rest of the paper is organized as follows. Section II presents some background information on the external benefits produced by the transhumance system in the Alps. Section III describes the survey design and the data. Section IV describes the methods for data analysis and value derivation. Section V reports and discusses the results, while the last section (VI) presents the conclusions.

II. Transhumance in the Alps and its external benefits

Alpine transhumance is the practice of moving livestock, mostly cattle, from the bottom of alpine valleys (lowlands) to the greener mountain pastures (highlands) during summers. This practice is more than 2000 years old (Bätzing 2003) and it characterizes this region so well that the term used to define mountain grazing ("alpage" in French, "Alp" in Swiss German, "alpe" in some Italian regions) gave its name to the whole mountain chain: the Alps. It also created a set of common property rights on land use designed to ensure sustainable use of pastures all over the alpine area (Stevenson, 1991), and a characteristic landscape made up of meadows and pasture clearings in areas where woodland otherwise would have prevailed had nature taken its course.

In Trentino, in the eastern Italian Alps, these clearings usually host the "malghe". These are typical alpine cottages built with local stones and used to shelter both humans and livestock. Apart from the lodgings of the livestock keeper and family, the malghe host dairy facilities to transform milk into less perishable and more transportable forms: cheese and butter. Alpine pastures and buildings represent a central element of the archetypal alpine landscape and are responsible for typical impressions that tourists have of the Alps (Paulsch et al. 2003). Today these landscapes are relics of what was once an important part of both the Roman and the German agricultural systems in the Alps (Bätzing 2003). They produce a bundle of external benefits which is now valued by the collective more than the agricultural produce they were designed to produce (diary and meat). These local external benefits derive from landscape amenity creation, and various other functions: ecological (i.e. biodiversity conservation, soil erosion protection and snow cover stabilisation, carbon sequestration, water filtration) socio-economic, historical-cultural. All of which contribute to attracting tourists and hence to the economic vitality of the area.

Most of these functions contribute to the overall social benefits produced by farmers and livestock keepers in these areas, but their importance varies with context and site-specific conditions. Landscape amenities creation was defined as the most prominent non-agricultural use of pasture and grassland (Wytrzens and Mayer 1999), but other functions now are growing in importance (Lehmann and Hediger 2004).

The economic valuation of these external benefits has now become more important because the practice of mountain summer grazing is no longer economically viable without some extra financial support. Due to increasing costs and difficulties encountered by the dairy sector in mountain areas (Cozzi and Bizzotto 2004), much of this high altitude meadowland actually is no longer grazed.[1] Hence, grassland areas are being gradually taken over by the natural process of reforestation throughout the alpine region (Tasser and Tappeiner 2002). The risk of losing most of the ecological and socio-economic functions of grazed pastures is growing.

Some local authorities in the alpine region have undertaken to financially support the practice of transhumance on alpine pasture with specific agro-environmental payments (i.e. grazing premiums) and a compensation allowance under the EU Rural Development Policy 2007-2013. Even if the support to the so-called "second pillar" seems uncontroversial in EU countries, general economic conditions and budget pressures could lead to reduction of such a policy in 2013. In Austria several tourist-intensive communities have already required additional voluntary compensation payments from visitors to farmers for the provision of landscape services (Hackl et al. 2007). Such measures were advocated and studied early on in this area using contingent valuation to study visitors' willingness to pay (WTP) (Pruckner 1995).

Since a large part of the benefits produced by Alpine transhumance are enjoyed by visitors, the issue of adequate estimation of WTP for access to these alpine areas is of policy relevance. Especially considering that this could become an important source of local revenue in the near future. However, previous stated preference work on WTP ignored the multi-attribute nature of these benefits. These—instead—are important because they can be linked to specific policies that can be modular in nature. In fact, different recreationally important features can be maintained in place by specific policy actions. Hence the motivation for our multi-attribute stated preference study.

III. Survey design and data

Survey design followed the recommended five steps for a Choice Experiment: selection of attributes, definition of levels, choice of the experimental design, construction of choice sets, measurement of preferences. The relevant functions/attributes identified in the literature as impacted by abandonment were tested in two focus groups, the first organized with experts (managers and workers of the natural park Adamello-Brenta and agricultural engineers), and the second with a sample of the local population. From all attributes proposed, the two focus groups identified four attributes as being important to people's preferences for alpine pastures: 1) alpine pasture landscape (abbreviated as AL), 2) biodiversity conservation (abbreviated as BD), 3) cultural-historical function (abbreviated as ST) and 4) conservation of the typical modes of production of dairy products (butter and cheese, abbreviated as TR). These are known to have superior organoleptic and nutritional properties (Hauswirth et al. 2003).

Levels of attributes linked to different management options were determined by experts and scientists. To better illustrate the policy outcomes to respondents in terms of attribute and levels some show cards were prepared with 3 pictures, each accompanied by simple descriptions that were read aloud. All non-monetary attributes had two levels of policy action, while a third level was exclusively associated with the "abandonment" option (no action). Four levels of access fee were determined from preliminary results of a repeated dichotomous choice CV study in which 223 visitors to the valley were asked about their maximum and minimum WTP for both a maximum and a minimum combination of the attributes using a payment card. The payment form proposed to respondents was an access fee to the valley to be paid per person and per visit. This was clearly stated in the survey and the respondent's understanding of this detail was ascertained during the questionnaire completion.

Experimental design and sample

Given our 4×2^4 factorial structure, we constructed a design in 16 choice sets each with 4 alternatives, to which the "abandonment" alternative was added, for a total of five alternatives per choice set. This design was 100% efficient for the estimation of a main effects only indirect utility function and conditional logit model, under the null hypothesis of no information about the parameters, and other assumptions given in Street and Burgess (2007). The profiles identified by the experimental design were grouped into 16 ranking tasks. The design is reported in Table 1. We note thought that the efficiency of the design for a full rank-ordered exploded logit model depends on the coefficient values, including those of the scale parameter for heteroskedastic specifications. Consequently the a-priori efficiency of the design is of quite difficult evaluation.

The response task was framed as a sequential choice process, where respondents were asked to choose the most preferred alternative out of 5, than the less favoured out of 4, the second best out of the remaining 3, the second worst out of 2. As our objective was to investigate use values, the target population was defined as the total number of users of alpine pastures in the survey location, the Val Genova, as they will be affected by changes in pasture management. The sampling frame population were all the visitors to this valley. A list for this kind of population does not exist, so we used a non-list sampling procedure, carrying out an intercept survey on site, collecting data with face-to-face interviews.

Coverage error was reduced by the fact that there is only one entrance to the valley. Respondents were sampled either as they arrived or left the valley, using a systematic sampling probabilistic design by drawing randomly at a rate of one out of 5. A preliminary pilot study of 15 randomly selected visitors was carried out on site to test the questionnaire. Data for the final survey were collected from 1st August till 15th September 2008 on all week days by three trained interviewers. The number of respondents approached by interviewers was 146, 39 of whom either declined or did not complete the survey; so the final sample for estimation included 107 completed questionnaires, a response rate of 74%. This sample size was considered adequate for the purpose of the study and the number of visitors expected to reach the site of Malga Bedole in August (personal communication with Park management). As each respondent completed ranking tasks, average completion time for an interview was 32 minutes.

IV. Theory and methods

Background

A full ranking of alternatives in a choice set can be seen as a sequence of discrete choices each based on a gradually decreasing set of alternatives. Using this basic intuition, rank-ordered choice data have long been analyzed using random utility models and “exploded” multinomial logit, or similar specifications (e.g., Luce and Suppes 1965). But changes in choice set composition may impact the certainty associated with choices. Hausman and Ruud (1987) were the first to address the issue of the Gumbel error scale varying across choice tasks defined by each rank. The confounding effect of scale on identification of taste intensities in logit models was more broadly addressed by Swait and Louviere (1993). More recently scale variation has been linked to systematic changes in ex-post measures of choice complexity (Swait and Adamowicz 2001, DeShazo and Fermo 2002) and to respondents' abilities to perform cognitively (e.g. level of education by Scarpa et al. 2003). In mixed logit formulations specifying utility in WTP-space has been shown to allow researchers to isolate the effects of individual-specific scales from WTP estimates (Train and Weeks 2005, Scarpa et al. 2008). Caparros et al. (2008) showed that once scale differences are accounted for and the same experimental design is applied, preference estimates derived from preferred choice models are consistent with those obtained from first rankings. Robustness of value estimates from ranking experiments on lotteries also was studied by Bateman et al. (2007), who showed that preference reversal can be attenuated with adequate analysis. In our case we are interested in capturing the effects on scale of choice context structure and the preference elicitation method. In terms of the latter we focus on the best/worst approach and its effect on the scale parameter of the independent Gumbel error.

Gumbel scale parameters must be strictly greater than zero. Thus, a useful specification is an exponential with a linear index defined by a sum of terms, each of which is the product of a coefficient and a variable that can impact scale: $\lambda=\exp(\sum_k \theta_k)$.

Best/worst elicitation

Respondents faced with a request from an interviewer to provide a full preference ranking of a given set of alternatives either can be left to their own devices as to how to achieve do this, or given specific instructions as to how to think to provide a ranking. Recent advances in applied conjoint analysis (Auger et al. 2007; Flynn et al. 2007; Louviere and Islam 2008) suggest that obtaining a ranking from a reiterated set of best/worst choices has significant advantages in terms of cognitive effort. However, this approach has not been investigated in the context of non-market valuation. In view of results reported in the extant literature, we adopted this method and asked respondents to make best/worst choices in our study.

Exploded logit models from full rankings: some considerations

We used an "exploded logit" analysis of a full ranking of five alternatives, which implies a sequence of four discrete choices. The first preferred choice is a selection out of 5 alternatives and relates to our specification of the scale parameter $\lambda=\exp(\sum_k\theta_k)$, with k = 2, 3, 4, 5, via coefficient θ_5 and an dummy-coded indicator function for that choice made in the context of five alternatives. The second preferred choice is a selection from the remaining 4 and (coefficient θ_4 with an indicator function for four alternatives). The third preferred choice is from the remaining 3 and relates to the scale effect coefficient θ_3; finally the fourth preferred choice is from the remaining two alternatives, implied to be the least favorite of the original set, and represent the baseline for the scale effects.

Some caution is needed in interpreting the mapping between best/worst choice events and the structure of the resulting exploded logit. This mapping is summarized in Table 2. In a best/worst approach the full ranking is obtained in two steps. In the first step respondents are asked to indicate the best alternative (ranked 1st out of five) and then the worst of the remaining four. This is ranked 5th out of the original five, but it is selected out of four because the selection of first best alternative reduced the choice set by one. In the second step the exercise is reiterated with the remaining 3 unranked alternatives, the best of which defines the alternative with rank 2 out of the original five (but is selected out of 3), while the worst is that with rank 4 (but is selected out of 2), leaving out the one that by implication is ranked 3 of the original five.

In terms of the exploded logit formulation, the first round of best/worst contributes to the creation of the favorite choice in a choice set with the initial 5 alternatives, and the choice of the worst in the smallest of the 4 choice sets, the one with 2 alternatives. Because one expects there to be relatively less uncertainty in choosing the best and the worst alternatives in the first round (because these stand out from the rest), one should expect the scale associated with these choices to be systematically larger than the others, which implies $\theta_5>\theta_j$, where $j<5$ and that $\theta_2>\theta_3$ and $\theta_2>\theta_4$. This should be the case despite the fact that choice occurs in a context with five alternatives, which prior work suggests may be linked to larger variance and smaller scale (Caussade et al. 2005). Recall that our θ coefficients are indexed to the number of alternatives in the exploded logit resulting from the full ranking.

Alternatively, one might expect a smaller scale to be associated with choice of best and worst alternatives in the first round, because these choices are made from the largest choice sets. So, there may be two contrasting effects: 1) a potential increase in variance due to more alternatives in the choice set from which a selection occurs, and 2) a potential decrease in variance due to ease of identifying the best and worst in the set. Which of the two prevails is an empirical question that we address in our study.

One hypothesis consistent with the fact that the first round of best/worst results in choices with different certainty than the second round may be formulated in terms of the relative magnitude of the scale parameter associated with tasks involving 5 and 2 alternatives in the exploded logit as these are determined by the first best/worst and the second worst. Similarly, choice tasks with 3 and 4 alternatives are determined by the second best and the residual alternative. If the scale for the last choice set in the exploded logit is set equal to 1 for identification purposes, so that $\theta_2=0$ and that $\lambda_k = \exp(\Sigma_k\theta_k)$, we can formulate the following:

$$H_0: \theta_k=0 \qquad \text{and} \qquad H_k: \theta_k \neq 0 \qquad \text{for } k=3,4,5$$

In other words, if the coefficients in the exponential representing the scale are jointly different from the baseline of 0, there is difference in the certainty of choice between best and worst decisions in the two best/worst rounds. Note that only θ_3, θ_4 and θ_5 can be identified, referring to the choice context in which the residual alternative (θ_3), the best in the second (θ_4) and first round (θ_5) are identified, respectively. The two worst alternatives from the first and second round, respectively, remain as the reference scale and are set to 1.

In particular, if coefficients related to the second round of best, θ_4, and the residual, θ_3, are negative and significant, it would mean that the difference is in the expected direction, with the first round of best/worst associated with more precise (less uncertain) choices.

Heteroskedasticity along the panel

There is another obvious cause of heteroskedasticity in a sequence of 16 full rankings by the same respondent, which is the order in the sequence of rankings of the panel in which the choice occurs. Previous results (Caussade et al. 2005, Bateman et al. 2008) suggest that the order in the sequence should have a gradually higher effect on scale (often attributed to "learning effects"), reaching a peak and then declining when "fatigue effects" kick in, over-riding the learning effect. So, from Model 3 and higher we also account for order of the ranking in the scale parameter using the set of parameters, ρ_t, setting ρ_1=0. So that, $\lambda_{k=2,t=1}$=exp(0)=1 is the new baseline for these models, given that θ_2=0 for identification purposes.

Taste heterogeneity

Although the exponential specification of the scale parameter of the Gumbel error is best suited to focus on intra-personal scale variation, it ignores inter-personal taste heterogeneity. In other words, respondents might differ in the way they evaluate the same policy attributes related to alpine pastures. This can be addressed by extending the model in a typical random parameter panel format.

Let us denote individual respondents by n=1,…107; ranking task by t= 1,…,16; the number of alternatives in the choice context k = 3,4,5; and the generic alternative by j=1,…,5. The generic utility structure is given by:

$$U_{ntkj} = \lambda_{kt}\,(V_{ntkj} + \varepsilon_{ntkj}) = \exp(\Sigma_k\theta_k+\Sigma_t\rho_t) \times (\boldsymbol{\beta}_n'\mathbf{x}_t + \varepsilon_{ntkj}) \times \delta_k,$$

where $\delta_k = 0$ if the alternative is the favorite one from the previous ranking, and hence is unavailable in the choice context k, 1 otherwise; $\exp(\Sigma_k\theta_k+\Sigma_t\rho_t) = \lambda_{kt}$ and refers to the scale of the Gumbel error in the choice context and the order indicator; $\boldsymbol{\beta}_n$ is the individual specific vector of taste parameters $\{\beta_{al}, \beta_{bd}, \beta_{pr}, \beta_{sq}, \beta_{st}, \beta_{tr}, \beta_1, \beta_2, \beta_3\}$ associated with the attribute vector $\mathbf{x}_t$ ={*AL, BD, PR, SQ, ST, TR, Asc_Left*, *Asc_MdLeft*, *Asc_MdRight*,} that defines each alpine pasture policy scenario. Finally, ε_{ntkj} is the error assumed to be i.i.d. Gumbel with unitary scale.

Defining the log-sum of the exponentials of the indirect utilities at each choice context as $e_k=\Sigma_j \exp(\lambda_k V_{ntkj})$, the probability of observing a full ranking of five alternatives in each ranking task is a product of 4 logit probabilities:

$$\Pr(U_{ntk1} > U_{ntk2} > U_{ntk3} > U_{ntk4} > U_{ntk5}) = \exp(\lambda_k V_{nt1j})/e_1 \times \exp(\lambda_k V_{nt2j})/e_2 \times \exp(\lambda_k V_{nt3j})/e_3 \times \exp(\lambda_k V_{nt4j})/e_4$$

The probability of the 16 full rankings made by each respondent conditional on their taste vector $\boldsymbol{\beta}_n$ is the product of 16 of the above.

$$\Pr(y_n|\boldsymbol{\beta}_n) = \Pi_t \Pr(U_{ntk1} > U_{ntk2} > U_{ntk3} > U_{ntk4} > U_{ntk5})$$

If the model is a panel mixed logit the unconditional probability is approximated in estimation by numerically integrating out the taste variation by simulation using quasi-random draws, which makes this integral a weighted average of probabilities (Train 2003). Taking the log of such a weighted average gives the contribution of each respondent to the sample log-likelihood. Estimation proceeds by maximizing the sample log-likelihood over the set of parameter values θ_k, ρ_t and $\boldsymbol{\beta}$ for the fixed parameter case, or for panel mixed logit model for θ_k, ρ_t and the mean $\boldsymbol{\beta}_\mu$ and variance-covariance $\boldsymbol{\Omega}$ (hyperparameters) regulating the random behavior of $\boldsymbol{\beta}_n$. In estimating the off-diagonal elements of $\boldsymbol{\Omega}$, given a multivariate normal assumption for $\boldsymbol{\beta}_n$, the components of the Cholesky matrix were estimated $\{\eta_{bd,al}, \eta_{st,al}, \eta_{st,bd}, \eta_{tr,al}, \eta_{tr,bd}, \eta_{tr,st}\}$ from which the variance-covariance matrix, $\boldsymbol{\Omega}$, can be derived. In Model 4 we assume a $\boldsymbol{\Omega}$ with a diagonal structure implying independence of $\boldsymbol{\beta}_n$. In Model 5 we also allow for a full set of off-diagonal elements, thereby allowing for correlation.

Individual specific WTPs and their determinants

One of the main issues with hypothetical surveys is that of validation (Bishop et al. 1995). Since this study adopts an in-depth survey of visitors it makes sense to explore the relationship existing between posterior estimates of marginal WTP and respondents' socio-economic and attitudinal variables. In a panel mixed logit analysis tastes are assumed distributed across respondents. Consequently, the estimation provides population distributions of taste, which implies a population distribution of marginal WTPs. However, a better estimate of these distributions can be obtained for each respondent in the sample by conditioning on the observed choice. Any given pattern of observed discrete choices will give rise to the same distribution with attendant mean and variance. The conditional moments (mean and variance) of marginal WTPs can be computed using the estimator proposed by Greene et al. (2005):

$$\hat{E}\left[WTP_{att,n}\right]=\frac{\frac{1}{R}\sum_r \frac{-\hat{\beta}^r_{att,n}}{\hat{\beta}^r_{price,n}} L\left(\hat{\beta}^r_n|\mathbf{y}_n,\mathbf{x}_n\right)}{\frac{1}{R}\sum_r L\left(\hat{\beta}^r_n|\mathbf{y}_n,\mathbf{x}_n\right)}$$

where $L(.)$ is the posterior likelihood of the individual respondents and the $\boldsymbol{\beta}_n$ are drawn from the multivariate normal computed at the MSL estimates $\hat{\boldsymbol{\beta}},\hat{\boldsymbol{\Omega}}$.

Using the estimates from the best fitting model (Model 5), for each of the four attributes and for each respondent we obtain a set of 107 conditional means of marginal WTP distributions. We then regress these estimated means on socio-economic covariates in the form of a 4-period panel to account for the fact that marginal WTP for attributes are correlated within the same respondent. We first use fixed effects per respondents in OLS regression, and then random effects. With these regressions we explore the determinants of the means of the marginal WTPs. A valid study should produce a reasonable pattern of sensitivity of estimated individual means of marginal WTP to key socio-economic determinants, thereby enhancing the theoretical validity of the hypothetical survey.

V. Results

Heteroskedastic fixed parameter logit models

We present three models (1-3) with fixed parameters (Table 3). In all estimations the attribute coding was effect-coded with the highest level coded as 1, the lower level coded as -1 and abandonment coded as 0 and also identified by the status-quo constant. This also implies that the marginal WTP formula is going to be $2\times\beta_k/\beta_\$$. The first model is a simple MNL model with an intercept for the status-quo (β_{sq}) referring to the situation without payment, which implies a gradual move towards abandonment. We record a negative estimated value for the status quo coefficient indicating that most people prefer some policy action. This is not surprising as all policies offered some improvement on the status quo and the price range was not prohibitive. All policy attributes have coefficient estimates with positive sign and are significant. These estimates imply reasonable marginal WTPs that ranked first the transformation of milk in-situ, second the conservation of grazing areas, third the historical heritage, and last, but not far from third, biodiversity issues.

Model 2 differs from Model 1 as it includes a) dummy variables for order of appearance of alternative from left to right β_1, β_2, and β_3, capturing differences with respect to the list of generic alternatives, and b) order effects (ρ_t) in the scale parameter. Estimates of marginal WTPs are of similar order as Model 1. This model provides evidence of a "left-to-right" bias (Dobel et al. 2007) of some significance as the coefficients of the alternative specific constants for the 1st, 2nd and 3rd alternatives (from left to right) are individually significant. There also is evidence of an order effect as many of the ρ_t coefficients are individually significant.

Model 3 is similar to Model 2, except that it adds the effects of the number of alternatives in the exploded logit (θ_k) in the specification for the Gumbel error scale parameter, two of which are highly significant. In particular, the results indicate that choices associated with the first round of best/worst have lower uncertainty, as the values of θ_3 and θ_4 are negative and significant. This evidence supports the use of best/worst elicitation mechanisms in rank-ordered data. There is a significant increase in fit moving from model 1 (log-lik = -5227, adj.pseudo R-sq = 0.36) to Model 2 (log-lik = -5184, adj.pseudo R-sq =0.36) and to Model 3 (log-lik = -5118, adj.pseudo R-sq =0.37).

Taste heterogeneity across respondents

Models 4 and 5 are mixed logit models built on model 3 in which the coefficient for price is assumed log-normal, while coefficients for the presence of the policies of interest are assumed normally distributed. The only difference is that model 5 allows for correlation across non-monetary attributes, producing a substantial and jointly significant improvement in statistical fit (from a log-lik = -4301, adj.pseudo R-sq = 0.47 to a log-lik = -4052, adj.pseudo R-sq = 0.5). Model 5 also produces the sharpest estimates for both ρ_t and θ_k. The pattern of ρ_t is reproduced in Figure 1, with a cubic curve interpolating the values of Model 5, which has the best fit. This result indicates that the scale increases gradually from the first to the 11th rank-order task, and then declines quite rapidly for ranking tasks 14-16. The discontinuity at the 6th rank-order is not significantly different from zero. The latter result is consistent resonates with Caussade et al. (2005). Once taste heterogeneity is addressed, the scale effect for choice tasks with 5 alternatives is no longer significant, while the estimated values for θ_3 and θ_4 still indicate significantly higher variances in utility evaluations after the first best/worst round.

Validation regressions

From each of the 107 respondents we estimated four sets of means for the conditional distributions of marginal WTPs. We expect the 4 means of each individual to be correlated, so we use both fixed and random effects regressions to explain the pattern of variation. We explore the role of socio-economic covariates for which we have information as determinants of estimated value. Estimates of fixed and random effects regressions are reported in Table 4. Having higher education (HIGH_EDU) has a significant and positive effect of about 20 Eurocents. Being/having been a livestock keeper or having a relative who is one (BREEDER) has a very strong positive and significant effect of about 50 Eurocents. Expenditure linked to the day out also is significant and positive (EXPEND), as is whether a person regularly practices sports (SPORT_PR). On the other hand, being married (MARRIED) has a negative effect on mean WTP, and having seen similar sites in another region of the Alps called Valdaosta

(VALDOSTA) also has a significant negative effect, perhaps due to substitution effects, or a “seen one, seen them all” effect. Finally, the effect of income is described using a dummy variable that separates those willing to report personal incomes from those who did not (INC_MISS). Compared to those who did not, those who did (91 out of 107) exhibit significantly lower mean WTP. If high income respondents were those most reluctant to disclose income levels in surveys, the estimated income effect would be as expected. Finally, being interviewed at the end of the visit while waiting for the shuttle bus at the car parking lot (BUS_STOP) has a positive effect, but it is significant only in the random effects model. Taken together, these variables do a reasonable job of explaining the variation in conditional means of marginal WTP, with an R-squared of 33.35% in the fixed effects model.

VI. Conclusions

In many operating contexts in-depth multi-attribute stated preference surveys are the only practical way to derive good quality estimates for multi-attribute goods. This is the case in our research context where we interviewed mountain visitors in a relatively large area about the external effects of alpine grazing. In these situations coupling rank-ordered data with efficient experimental design seems to be an effective way to obtain high quality information. Prior work has criticized rank-ordered data because different error scales are associated with different choice task contexts (Hausman and Ruud 1987), which confounds estimation of taste-intensities in random utility models. Past research has associated estimates of Gumbel error scale with choice complexity, with complex choice sets associated with higher utility variance (smaller scale). Recent research results suggest that asking respondents to choose alternatives using an iterated best/worst selection rule may facilitate respondents’ choices. We find support for this in so far as the first rounds of best/worst choices are associated with lower variance in utility even though the choices are made from a larger set of alternatives. Cognitive facilitation associated with the best-worst approach seems to over-ride the complexity associated with a large number of alternatives from which to choose. This makes the iterated best/worst approach of interest in this area of stated choice practice.

We also found that utility variance decreases with the sequence of ranks in the panel or tasks by respondents up to the 10^{th}-11^{th} task, and then increases. We note that these forms of scale variation are all intra-personal, and hence interpersonal scale variation remains unaddressed. This is clearly beyond the scope of this paper, and we leave it as an objective for further methodological developments. Similarly, we leave as a subject of further analysis the investigation of whether the preferences of each rank are sufficiently consistent with each other.

In terms of the implied marginal WTPs estimated at the respondent level we found that they respond in a plausible fashion to variation of covariates, which suggests that the best-worst approach produces valid WTP estimate for policy analysis. From the viewpoint of local policy making we found that typical dairy transformation is the attribute associated with highest marginal WTP (around 5 euros), followed closely by the alpine landscape component (around 4.80 euros), then by conservation of historical buildings (around 4.40 euros). The least WTP was found for biodiversity (around 3.95 euro), taking the overall WTP for a policy improving all four attributes to the maximum at an average value of 18 euros. As this was a hypothetical stated choice exercise, one needs to correct for hypothetical bias. Murphy et al. (2003) in a meta-analysis study find the median ratio of hypothetical to real valuations to be 1.35, with considerable skewness to the right. In the absence of locally developed calibration functions between hypothetical and real choices one can take a conservative stand and assume the real average WTP to be no more than 2/3 of the stated hypothetical WTP. Hence, an average access fee of eight to twelve euros per person would appear to be acceptable to most visitors represented by our sample for a policy providing the highest improvements. Local politicians may therefore be advised that a visitors' access fee to support the continuation of alpine transhumance is a viable proposition.

VII. References

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VIII. Tables and figures

Table 1. Experimental design

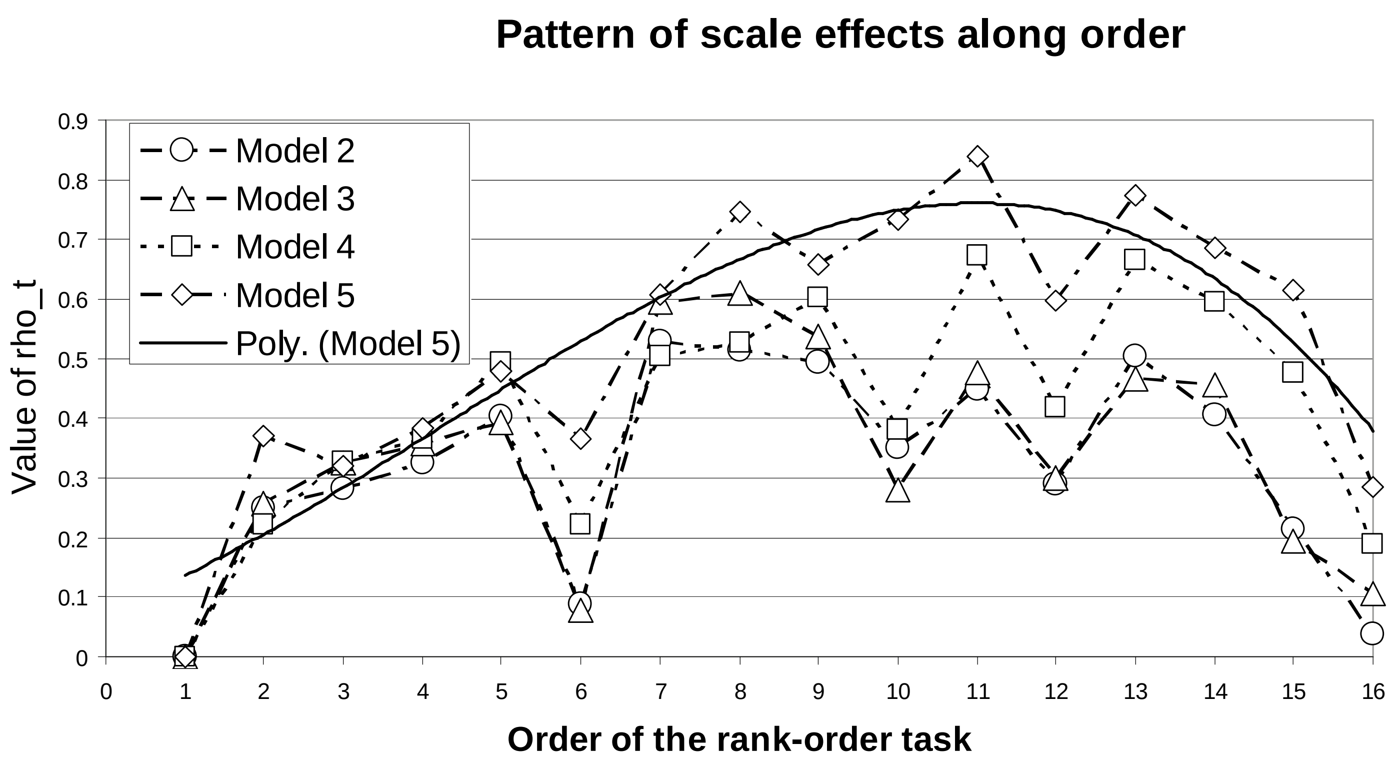
		Order in ranking-task sequence															
		1	2	3	4	5	6	7	8	9	10	11	12	13	14	15	16
Alternative 1	PR	3	3	3	2	1	0	2	2	0	1	2	3	0	3	0	1
	AL	1	1	0	0	1	0	0	1	1	0	1	0	1	1	0	0
	BD	1	0	0	1	0	1	1	0	1	1	0	1	1	0	0	0
	ST	1	1	0	1	0	0	1	1	0	0	0	1	1	0	0	1
	TR	0	1	1	1	1	0	0	1	1	0	0	1	0	0	1	0
Alternative 2	PR	1	2	2	1	3	3	0	3	2	2	0	1	3	2	2	3
	AL	0	0	0	0	0	1	1	0	0	1	0	1	1	1	1	1
	BD	0	0	1	0	1	1	0	0	0	1	1	0	0	1	1	1
	ST	1	0	1	0	0	1	1	0	0	1	0	1	0	1	0	1
	TR	1	0	1	1	0	1	1	0	0	1	1	0	0	0	0	1
Alternative 3	PR	2	0	1	3	2	1	3	1	3	0	1	2	2	0	3	0
	AL	1	1	1	1	1	0	0	1	0	0	1	0	0	0	0	1
	BD	0	1	1	1	1	0	0	1	1	0	1	0	0	0	1	0
	ST	0	0	0	0	1	1	0	0	1	1	1	0	1	1	1	0
	TR	0	1	0	0	1	0	0	1	0	0	0	1	1	1	1	1
Alternative 4	Pr	0	1	0	0	0	2	1	0	1	3	3	0	1	1	1	2
	AL	0	0	1	1	0	1	1	0	1	1	0	1	0	0	1	0
	BD	1	1	0	0	0	0	1	1	0	0	0	1	1	1	0	1
	ST	0	1	1	1	1	0	0	1	1	0	1	0	0	0	1	0
	TR	1	0	0	0	0	1	1	0	1	1	1	0	1	1	0	0

Table 2. Mapping between best/worst choice and the structure of the exploded logit

Instance	Choice	Rank	Alternatives in Exploded logit Choice set	Alternatives in the selection context	Scale Coefficient
1	Best 1	1	5	5	θ_5
	Worst 1	5	2	4	θ_2
2	Best 2	2	4	3	θ_4
	Worst 2	4	2	2	θ_2
Residual alternative		3	3	1	θ_3

Table 3: Estimation results

	Model 1		Model 2		Model 3		Model 4					Model 5				
Coefficient	Coeff.	|t-val.|	Coeff.	|t-val.|	Coeff.	|t-val.|	Coeff.	|t-val.|		Coeff.	|t-val.|	Coeff.	|t-val.|		Coeff.	|t-val.|
β_{pr} Access Fee	-0.123	(25.7)	-0.087	(10.0)	-0.089	(8.3)	-2.351	(14.3)	σ_{pr}	1.448	(11.2)	-1.821	(13.1)	σ_{pr}	1.392	(13.0)
β_{al} Landscape	0.362	(20.8)	0.262	(9.5)	0.260	(8.0)	0.387	(6.3)	σ_{al}	0.333	(5.0)	0.420	(7.0)	σ_{al}	0.269	(7.2)
β_{bd} Biodiversity	0.276	(16.7)	0.203	(9.6)	0.226	(8.5)	0.309	(6.0)	σ_{bd}	0.252	(4.8)	0.341	(6.3)	σ_{bd}	0.055	(0.9)
β_{st} Historical	0.329	(19.9)	0.235	(9.5)	0.283	(8.6)	0.331	(5.7)	σ_{st}	0.439	(4.3)	0.378	(6.7)	σ_{st}	0.362	(6.6)
β_{tr} Dairy Trans.	0.371	(21.2)	0.274	(9.9)	0.303	(8.7)	0.422	(6.6)	σ_{tr}	0.043	(0.9)	0.434	(6.0)	σ_{tr}	0.226	(3.3)
β_{sq} Status quo	-4.311	(47.5)	-3.028	(11.2)	-3.356	(10.3)	-4.767	(8.2)				-5.010	(9.0)			
β_1 Left option			0.073	(2.0)	0.077	(2.2)	0.116	(2.7)				0.113	(2.6)	$\eta_{bd,al}$	0.300	(6.0)
β_2 Middle-Left			0.110	(2.8)	0.058	(1.5)	0.121	(2.6)				0.109	(2.5)	$\eta_{st,al}$	0.165	(3.3)
β_3 Middle-Right			0.076	(2.0)	0.050	(1.3)	0.093	(2.4)				0.084	(2.2)	$\eta_{st,bd}$	0.230	(6.3)
ρ_2 2nd in panel			0.249	(2.2)	0.257	(2.2)	0.221	(2.2)				0.369	(3.7)	$\eta_{tr,al}$	0.065	(2.0)
ρ_3 3rd in panel			0.283	(2.2)	0.326	(2.5)	0.326	(2.2)				0.320	(2.4)	$\eta_{tr,bd}$	0.316	(5.0)
ρ_4 "			0.324	(2.4)	0.358	(2.8)	0.366	(2.9)				0.384	(3.4)	$\eta_{tr,st}$	-0.093	(1.8)
ρ_5 "			0.402	(3.4)	0.394	(3.4)	0.494	(4.4)				0.479	(4.7)			
ρ_6 "			0.088	(0.7)	0.078	(0.6)	0.221	(2.0)				0.366	(3.1)			
ρ_7 "			0.530	(4.3)	0.594	(5.0)	0.503	(4.8)				0.608	(5.4)			
ρ_8 "			0.514	(4.0)	0.609	(4.8)	0.526	(4.5)				0.746	(7.5)			
ρ_9 "			0.494	(4.3)	0.536	(4.7)	0.602	(6.2)				0.657	(6.1)			
ρ_{10} "			0.352	(2.9)	0.280	(2.4)	0.381	(3.1)				0.734	(5.4)			
ρ_{11} "			0.448	(3.6)	0.475	(3.9)	0.674	(4.9)				0.841	(5.1)			
ρ_{12} "			0.290	(2.3)	0.300	(2.3)	0.418	(3.9)				0.597	(4.9)			
ρ_{13} "			0.505	(4.2)	0.467	(4.0)	0.665	(5.8)				0.774	(6.5)			
ρ_{14} "			0.407	(3.3)	0.455	(3.6)	0.596	(4.4)				0.685	(5.5)			
ρ_{15} "			0.214	(1.5)	0.195	(1.4)	0.477	(4.0)				0.614	(5.0)			
ρ_{16} "			0.038	(0.3)	0.107	(0.9)	0.189	(1.6)				0.285	(2.4)			
θ_3 1st-2nd worst					-0.393	(6.2)	-0.447	(3.1)				-0.622	(3.7)			
θ_4 2nd best					-0.200	(2.9)	-0.217	(1.6)				-0.253	(2.1)			
θ_5 1st best					0.238	(2.9)	0.149	(1.5)				0.080	(0.9)			
N. Param.	6		24		27		32					38				
-ln(*L*)	5227.27		5184.96		5118.2		4301.49					4052.23				
Adj-Pseudo- R^2	0.36153		0.36447		0.37224		0.47131					0.50096				



Pattern of scale effects along order
Model 2
Model 3
Model 4
Model 5
Poly. (Model 5)
Value of rho_t
Order of the rank-order task
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
0.1
0
0 1 2 3 4 5 6 7 8 9 10 11 12 13 14 15 16

Figure 1

Table 4: Validation regressions

	Classical OLS Fixed effects		Random effects	
Name	Coeff.	|t-values|	Coeff.	|t-values|
AL	2.336	(1.1)	-0.075	(0.5)
BD	0.754	(0.5)	-0.519	(3.2)
ST	0.377	(0.3)	-0.312	(1.9)
HIGH_EDU	0.188	(2.3)	0.228	(3.3)
BREEDER	0.505	(3.6)	0.582	(4.7)
EXPEND	0.005	(3.2)	0.008	(5.6)
SPORT_PR	0.295	(2.1)	0.127	(1.0)
MARRIED	-0.331	(2.4)	-0.606	(4.9)
VALDOSTA	-0.405	(2.4)	-0.316	(2.3)
INC_MISS	-1.025	(5.8)	-0.918	(5.9)
BUS_STOP	0.074	(0.4)	0.371	(2.5)
Constant			2.484	(8.2)
Log-likelihood:		-572.807		
Adjusted rho-square:		0.333546		
F		2.83		
Fixed vs. Random Effects (Hausman)			43.5	

Example of Choice card:

Please make each choice as if the alternatives were the only one available.
Please, examine the 5 different alternatives and in the bottom row of each card indicate

(1) the <u>best</u> alternative;	**(mark with 1)**
(2) the <u>worst</u> alternative;	**(mark with 5)**

Then repeat the exercise and show

(3) the <u>best</u> alternative among the remaining three;	**(mark with 3)**
(4) the <u>worst</u> alternative among the remaining two.	**(mark with 4)**

Card D.1

	Altern. A	**Altern. B**	**Altern. C**	**Altern. D**	**Altern. E**
Access fee €	**12**	**5**	**8**	**2**	**0**
Landscape	**<u>Very</u>** tidy	Quite tidy	**<u>Very</u>** tidy	Quite tidy	**ABANDONMENT**
Biodiversity	**<u>High</u>**	Medium	Medium	**<u>High</u>**	
Historical-cultural function	**<u>Visitable</u>**	**<u>Visitable</u>**	Not accessible	Not accessible	
Milk processing	At the valley	**<u>In “Malga”</u>**	At the valley	**<u>In “Malga”</u>**	

IX. Endnotes

[1] In the province of Trento, for instance, there is a total of 613 mountain pastures registered that are estimated to cover an area of 33,000 hectares, yet recent surveys show that only about 260 pastures are actually grazed.