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Valuing Local Public Goods with Advanced Stated Preference Models: Traffic Calming Schemes in Northern England

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SUMMARY

The paper reports the results of three stated preference surveys in urban-rural areas in Northern England. The objective is that of valuing the economic benefits from traffic calming schemes in two areas with different traffic problems from stated preference observations. Both choice-experiments and contingent valuation methods are employed using advanced modelling. Fixed and random coefficient utility models are estimated from responses of the choice-experiments, while double-bound spike models are used for contingent valuation. Welfare estimates from the different methods are compared. The role of accounting for repeated choices is found to be of relevance. Choice modelling is designed to disentangle the values of benefits from 5 major attributes of traffic calming schemes (noise abatement, speed control, community severance, aesthetic layout and tax burden).

Keywords: Local public goods, non-market valuation, stated preference, choice experiments, traffic calming

NON TECHNICAL SUMMARY

The paper reports the results of three stated preference surveys in urban-rural areas in Northern England. The objective is that of valuing the economic benefits from traffic calming schemes in two areas with different traffic problems from stated preference observations.
Because such schemes produce multi-attribute benefits, amongst which the most important is speed-reduction to the imposed speed limits, two stated-preference valuation techniques are used. The first is contingent valuation, which is employed solely to estimate benefits from speed reduction, the second is choice-modelling with which we estimate also benefits from community severance, noise emission and aesthetics of the scheme. Advanced econometric modelling (spike models and mixed logits) is employed to analyse both types of data. Welfare estimates from the different methods are compared and found to be in keeping with theoretical expectations. The role of accounting for repeated choices is found to be of relevance: unsurprisingly variation of taste across people is found to be substantially higher than between choices by the same respondent. This finding has implication for conventional analysis based on fixed parameter logit models.

CONTENTS

1 Introduction

Stated preference (SP) methods are an important tool for valuing public goods. Ever since the seminal work by Paul Samuelson, applied public economists have been aware of the difficulties involved in public good valuation, and of the relative implications in normative public policy. In the last two decades a great effort has been devoted to the development of methodologies capable of delivering reliable estimates of the benefits people enjoy from the provision of this special category of goods. Recent advances and lines of future research in the field of SP for non-market valuation have been summarised in a recent workshop (SEC, 2001).

SP valuation methods have evolved greatly in the last 20-30 years, and although contingent valuation and — more recently — experimental choice modelling (henceforth experiments) have been under severe methodological and theoretical scrutiny, the debate around the suitability of SP estimates to guide policy action is ongoing (Randall, 1998), dividing economists into the 'skeptical' and those who are 'supporters'. In this research note we present and compare estimates of the same set of economic benefits from a local public good: a traffic calming scheme, using SP techniques and focussing on the role of modelling unobserved heterogeneity.

This interdisciplinary study was conducted as a joint effort between civil engineers and applied economists, and funded by the U.K. EPSRC (Engineering and Physical Science Research Council).

Three separate SP surveys were administered. Their main aim was to estimate the benefits enjoyed by local communities in the presence of traffic calming schemes (TCS), under different ex-ante traffic conditions.

Two populations were studied: A and B. Population A was the collection of households (HHs) residing in three rural towns in Northern England (Haydon Bridge, Seaton Sluice and Rowlands Gill), all of which are suffering from the same problem: traffic from main trunk roads passing through the centre of the settlement.

Population B was made-up of HHs residing in two rural towns: Sherburn-in-Elmet and Great Ayton. These show traffic problems of a different nature as the trunk roads mainly affected the periphery of each settlement, hence negative externalities for the resident HHs might be hypothesiezed to be lower. These two rural towns, rather than being simply crossed by a trunk road, have each developed around a major junction or a river crossing. Due to the different nature of the traffic problems experienced by the two populations, preference for TCS 'deliverables' are expected to be different. Because choice models allow the estimation of multi-attribute packages of public goods, the choice of a different setting was motivated by the need to verify whether or not the benefit estimates for the same bundle of public good attributes changed according to expectations from economic theory.

The benefits that the two populations would enjoy from TC were investigated by designing and administering three separate survey instruments:

- *Survey 1.* The first survey, administered to population A, was a 'conventional' contingent valuation survey, aimed at valuing an important component of traffic calming: effective speed reduction (ESR). This is the speed below which 85% of traffic travels along a certain stretch of road.
- *Survey 2.* The second survey, also administered to population A, was developed around a choice experiment investigating three additional deliverables of traffic calming. Noise abatement, reduced community severance and the aesthetic improvement of the traffic calming layout. This survey is referred to as CE1 in the rest of the paper.
- *Survey 3.* The third survey was equivalent to survey 2 and investigated the same set of TC attributes, but in a different population, population B. This survey is referred to as CE2 in the rest of the paper.

Estimates of willingness to pay for ESR from population A can be derived from the data collected from both CE1 and CE2. The estimated willingness to pay is expected to be similar from the contingent valuation questions and the choice experiments, up to sampling noise. In statistical terms this implies that the analysis should fail to reject the null of no difference between the two value estimates. Similarly, validity of multi-attribute choice experiment value estimates, implies that these should be sensitive to different degrees of traffic disturbances suffered by the population of interest. For example, population A suffers an average noise emission higher than population B, which in turns experiences a longer average time to cross the road than A. Estimated values should reflect these differences, and hence show a higher value for noise abatement in population A, and a lower value for reduced crossing time for B.

Recent applications of simulated-likelihood estimation of random utility models (Layton, 2000) have shown how welfare measures from discrete choice models can be sensitive to the treatment of unobserved heterogeneity. This can be present in significant residual form even when conditional heterogeneity is accounted for (Scarpa *et al.*, 2001). For this reason we employ random parameter logit (or mixed) logit models (McFadden and Train, 2000) to estimate a number of plausible specifications and assess the effects that various ways of accounting for unobserved heterogeneity have on welfare estimates.

We find that some basic expectations are met, but the statistical significance of this concordance with *a-priori* inequalities is sensitive to econometric specifications, such as whether or not parameters are assumed to be fixed or random, and — in the latter case — to the distribution assumed to the taste parameters.

The rest of the paper is divided into three sections. In section 2 we illustrate the methods employed for the development of the survey instruments, their administration, and the econometric analyses of the three sets of discrete-choice data obtained. In section 3 the results are reported and discussed. Conclusions are drawn in section 4, and some thoughts are advanced for future directions of work.

2 Theory and methods

Local public goods are defined by the geographical limits of the benefits they create. As such, their non-excludability and non-rivalry in consumption extend only over a geographically limited area. Traffic calming is one such public good, in that there is a limit to the number of households that will enjoy the benefits from the deliverables arising from a traffic calming initiative.

The flow of traffic along trunk roads, especially in relation to commuters and commercial vehicles, generates a series of negative effects which are endured by local residents. TCSs can be designed to address — in a selective fashion — a number of these negative effects, thereby achieving their mitigation and producing benefits to local residents. In as much as traffic speed is reduced, and in the absence of alternative routes, the outcome of traffic calming may well be a net cost for some through-route road users. This study, however, is aimed exclusively at the characterization and valuation of the benefits to local residents, ignoring the costs that such a public programme would generate for non-resident road-users.

2.1 The identification of programme attributes.

Multi-attribute valuation through choice-modelling is normally conducted by means of a set of choice experiments. In these experiments respondents are asked to choose between bundled goods whose attributes have been varied systematically according to an experimental design. By choosing the favourite bundle, or ranking them according to preference, respondents implicitly trade-off attribute levels, hence disclosing their preference structure. From these observations, one can identify marginal rates of substitution between choice attributes. Clearly, in the case of valuation studies one attribute must be the cost of the chosen alternative.

In choice experiments, only a limited number of attributes can be practically accommodated within the cognitive tasks performed by a random sample of individuals. From the civil engineering view point, traffic calming schemes can be designed to achieve a number of goals. The most prominent of these is speed reduction. This attribute was given a particular emphasis in the study, and investigated in detail by a specific contingent valuation survey. However, a number of other 'deliverables' can be generated by TCS. These deliverables were identified by the civil engineers involved in the project and by means of focus groups conducted with representatives of populations A and B. Thus, three more attributes were identified as important: *noise abatement*, reduction in *community severance* and the *aesthetic appeal* of the traffic calming measures themselves.

The final objective of the choice experiment valuation study was therefore that of valuing the benefits associated with traffic calming schemes, with a particular focus on:

(a) cost of the scheme to the household,
(b) effective speed reduction,
(c) noise abatement,
(d) aesthetic appeal of the engineering solutions, and
(e) reduction in community severance.

2.2 Surveys and data analyses.

Given the importance of the issue of estimate validation in the SP literature, and the impossibility of conducting a social-experiment with real payments, an attempt to cross-validate the estimates was made by designing an independent contingent valuation survey exclusively aimed at valuing the most prominent of the traffic schemes deliverables, i.e. ESR. This survey was administered to an independent random sample of households drawn from the same target population as one of the choice experiments (CE1), population A, which was also the population most exposed to high speed.

Both types of data — from choice experiment and contingent valuation — produced an estimate of WTP for ESR, hence provided a comparison, similar to those conducted in other studies (Adamowicz *et al.*, 1998).

2.2.1 The contingent valuation survey.

The salient features of the contingent valuation study are summarised as follows. The survey was conducted over the phone. The random sampling framework was the listing in the '192 people identification' software. The survey instrument and the bid design were tested in two pilot studies. The payment vehicle was an increase in local taxes, while the public good market was that of a local referendum. A follow-up question was also administered and debriefing questions helped identifying zero-bidding behaviour, while in both the first and second questions a discrete-choice elicitation format was used. The bid design was up-dated after the collection of the first 150 responses on the basis of the expected Bayesian posterior probability.

Some expectations were held for the WTP estimates from the CV survey. In particular, under the assumption of procedural invariance of the particular SP approach we expect that:

(1) They compare well in magnitude with those obtained from data collected in study CE1, as the same population was sampled.
(2) Secondly, that they relate to the magnitude of the decrease in speed brought about by the implementation of ESR in different areas. Thus, people who live in areas where the scheme will achieve larger speed reduction from those recorded in the absence of the TCS are expected to be willing to pay more. For example, if the TCS reduces the current speed by 15 miles/hour, this is expected to be more valuable than if it is reduced it only by 10 mph.

2.2.2 Contingent valuation data analysis.

Parametric WTP estimates were obtained by maximizing the likelihood of the sample responses using the following probability specification (Ayala and An, 1996):

$$Pr(WTP < x) = \begin{cases} 0, & \text{if } x < 0; \\ p, & \text{if } x = 0; \\ p + (1-p)F(x), & \text{if } x > 0. \end{cases}$$

Where $F(x)$ is the normal cdf and $x = (\ln(t) - \mu)/\sigma$, and the location and scale parameter are to be estimated from the data, along with the proportion of zero bidders p.

In the estimation we assumed that both the first and the follow-up responses be distributed according to the same underlying WTP distribution and used the interval-data likelihood.

With the above probability specification, there is a positive probability mass — or ‘spike’ (Kriström) — at zero, which is measured by p. Under the above hypotheses, the estimators for median $\hat{M}(WTP)$ and mean $\hat{E}(WTP)$ can be expressed as close-form solutions of the ML parameters estimates $\hat{\mu}, \hat{\sigma}$ and $\hat{p}$, as follows (Reiser and Shecther, 1999):

$$\hat{E}(WTP) = (1 - \hat{p}) \exp(\hat{\mu} + 0.5\hat{\sigma}^2).$$

$$\hat{M}(WTP) = \begin{cases} \exp\left[\hat{\mu} + \hat{\sigma}\Phi^{-1}\left(\frac{0.5-\hat{p}}{1-\hat{p}}\right)\right], & \text{if } \hat{p} < 0.5; \\ 0, & \text{if } \hat{p} \geq 0.5. \end{cases}$$

Approximate confidence intervals around maximum-likelihood estimates for $\hat{M}(WTP)$ and mean $\hat{E}(WTP)$ can be obtained by resampling from the asymptotic sampling distribution of the parameter estimates, which is multivariate normal, using the Krinsky and Robb approach.

It should be noted that, although we interpret the contingent valuation responses using a variation function approach based on the expenditure function, the scale and location parameters can be used to derive their counterparts (constant and slope) under an indirect utility function interpretation.

However, we need not assume an identical functional form to compare estimates of benefits from two different samples and SP methods, as here we do not use joint estimation.

2.2.3 The choice experiment surveys.

Two choice experiment studies were conducted at five sites. The first (CE1) in three towns crossed by trunk roads (population A); with the second (CE2) survey administered in two towns that were less affected by trunk road traffic (population B). In the absence of a TCS, population A suffers more severely from negative traffic externalities.

In particular, given the different nature of the traffic and the layout of the settlements and after examining the physical data measured on-site, we expected the following differences from the benefit estimates for residents in the two areas:

(1) ESR should be valued more by population A than by B, as the current average speed is higher in the first context, and hence the speed decrease to achieve compliance at 30 mph is also higher (more good delivered);
(2) Noise reduction should be valued more by population A than by B, as through traffic at high speed was noisier (background dim noise levels where 7dB higher in A than in B);
(3) Reduction in community severance (waiting time for crossing at main points) should be valued more in A than in B, as in the presence of junction and slower traffic the main road is easier to cross;

(4) No particular expectation was held for the value of aesthetic improvement of the TCS layout, which is a matter of taste of local residents. However, it can be argued that economic theory suggests this to be a 'luxury' good and be positively linked to HH income. Since this is highly correlated with property values, one could reasonably expect that locations with high property values show a higher WTP for aesthetic improvement. In our case this translates into higher expected values for population A, as here property values are on average higher than in population B.

Both surveys were conducted by a market research firm, whose professional enumerators interviewed respondents in person and in their own homes, asking to talk — whenever possible — to the person in charge of paying the local council taxes. The two choice experiment studies were designed to investigate the same set of attributes, but the preference elicitation mechanism was slightly different in the two studies. In population A respondents were asked to identify the favourite choice out three alternative: two profiles and the zero option. In population B respondents were asked to rank four alternatives: the same two profiles as in population A, plus a third randomly selected non-dominated profile and the zero option.

For a start we wanted to reduce the cognitive task of choosing amongst many alternatives. We therefore only employed choice situations with only either three or four options. In each choice situation the preservation of the status quo — for which we had physical measurements — was one of the possible alternatives. As a result the respondent had to compare either 2 or 3 attribute profiles in each choice task. Each attribute profile is a potentially deliverable bundle in the engineering implementation of the TCS. Eight choice tasks were carried out by each respondent, who was remunerated with a 3 or 5 pound voucher redeemable at outlets of a major retail chain. To enable them to assess different degrees of aesthetic solutions in TCS, respondents were shown photos from existing TCSs. To enable them to experience decibel sound measurements, they were exposed to recorded sounds of traffic at 60, 70 and 80 decibels, which were the levels employed in the profiles.

The core of the experiment was a basic fully factorial experimental design for pair-wise undominated choices. An algorithm was written in GAUSS to produce all of the possible combinations of pair-wise cases, and for the subsequent elimination of all the cases in which the choice was dominated by one particular profile. As a consequence we do not expect the model to display the usual high fit and parameter stability that is normally (artificially) achieved by more efficient designs, such as the partial factorial, D-optimal, etc. However, the data collected in this fashion support more complex model specifications (i.e. quadratic utility with interactions between taste parameters) than those obtained with factorial main-effect designs, although these more complex specifications are not explored in the present paper.

This basic design was employed to generate the choices for CE1 that were administered to a sample of 413 respondents drawn from the same population as the one for the contingent valuation study (i.e. the rural towns bisected by trunk roads). So, respondents in this context were asked to select a preferred choice amongst a set which included a 'zero-option' and two traffic-calming schemes, each of which was represented by a profile of five attributes expressed at a certain level (see Table 1).

The design employed for CE2 and administered to the sample (N= 407) of households drawn from population B was slightly more complicated, as it involved a ranking exercise amongst four choices, one of which was the 'zero-option'. The other three were made up by the same basic experimental design as for CE1, with an additional non-dominated choice randomly drawn from the available set. In this study, however, to maintain homogeneity with the approach used in CE1, we only employ the 'preferred' choice, and ignore ranking information.

2.2.4 CE data analyses.

The discrete choices collected in the two choice experiments can be analysed in many different ways.

In the selected set of models, we adopted the classic random utility specification, with Gumbel distributed error terms in each alternative-specific indirect utility, so that the unobservable stochastic component associated with the utility difference is logistically distributed.

Indirect utility is a linear additive index of the taste parameters multiplied by the attributes levels, and no individual-specific interaction variables are employed. This leads to the well-known choice probability specification for the favourite choice j^* from j choices:

$$P(U_{j^*}) = \frac{\exp(\theta' \mathbf{x}_{j^*})}{\sum_j^J \exp(\theta' \mathbf{x}_j)} \tag{1}$$

Given the recent interest in random parameter logit models (mixed logit) for multinomial discrete-choices, we present some comparisons between conventional fixed parameter (conditional) logit (standard logit) and a variety of mixed logit models with different choices of mixing distributions.

the marginal probability needs to be integrated over all the possible values of the varying parameter, and weighted by its density conditional on the density parameters ω:

$$P(U_{j^*} \mid \omega) = \int_{\underline{\tilde{\theta}}}^{\bar{\tilde{\theta}}} P(U_{j^*} \mid \tilde{\theta}) f(\tilde{\theta} \mid \omega) d\tilde{\theta}. \quad (2)$$

In mixed logit taste-heterogeneity is generic, and is represented by a parametric randomness of taste parameters, conditional on the values of the distribution parameters $\tilde{\theta}$. The intensities of taste are varying according to a distribution that can be chosen to be a parametric function and therefore the task becomes that of obtaining estimates $\hat{\omega}$ of the parameters of the assumed distribution. For example, for a distribution with two-parameter, say with location μ and scale σ, one would have to estimate the vector $\hat{\omega} = \{\hat{\mu}, \hat{\sigma}\}$.

Taste variation may occur across all observed choices in the sample. In this case the source of unobserved heterogneity is choice-specific, regardless of the individual (Revelt and Train, 1998). In this case the marginal probability of observing choice j^* needs to be integrated over all the possible values of the varying parameter, and weighted by its density conditional on the density parameters ω:

$$P(U_{j^*} \mid \omega) = \int_{\underline{\tilde{\theta}}}^{\bar{\tilde{\theta}}} P(U_{j^*} \mid \tilde{\theta}) f(\tilde{\theta} \mid \omega) d\tilde{\theta}. \quad (3)$$

However, with few exceptions (for example, Scarpa *et al.*, 2001), in most choice modelling studies respondents are asked to repeat choice tasks, so as to economize on sample sizes. In these instances taste variation may be thought as being prevalently driven by taste-differences across individual respondents, and being invariant across the sequence of repeated choices by the same respondent (Brownstone and Train, 1999). In this case the object of integration is the joint probability of observing a sequence of k choices by the n^{th} respondent, and the marginal becomes:

$$P(U_1, U_2, \ldots, U_{k(n)} \mid \omega) = \int_{\underline{\tilde{\theta}}}^{\bar{\tilde{\theta}}} \prod_{k(n)} P(U_{j^*} \mid \tilde{\theta}) f(\tilde{\theta} \mid \omega) d\tilde{\theta}. \quad (4)$$

Mixing across all recorded choices ignores the dependence in the set of repeated choices by the same respondent. This is consistent with the unobservable component of utility varying across choices independently of who makes them.

Mixing across choices from different respondents, instead, accounts for the fact that during each of the repeated choices the respondent holds the same set of taste parameters, that is, taste parameters are respondent-specific, although a stochastic component is still associated with the utility of each alternative.

The issue of whether the latter is nested in the former, so that comparisons of the model fit are possible through comparisons of the simulated likelihood at a maximum, is still unclear.

Estimates for the scale and location parameters of the distributions of tastes are obtained by maximization of the simulated likelihood of the sample. The presence of some parameter varying according to a given distribution, the marginal probability of each choice needs to be approximated via simulation by the average of logit probabilities each computed at a random draw of taste parameters from the postulated distribution with parameters ω:

$$SP_{j|\omega} = \frac{1}{R}\sum_{r}^{R} SP_i(\theta_i^{r|R} \mid \omega) \qquad r = 1, 2, \ldots, R. \tag{5}$$

The simulation approximates the integral in eq. (3) or that in eq.(4) according to whether the estimation is for each observed choice (non-panel) or each sequence of observed choices by the same respondent (panel). The estimates we present were derived using 200 well equi-dispersed random draws based on Halton sequences (Train, 1999), which is sufficient to obtain the desirable properties of the log-likelihood function, and the tolerance for the convergence is 10^{-5} of the improvement of the log-likelihood gradient [2].

2.2.5 Mixing specifications.

The taste parameter is held fixed *a-priori*, while taste parameters for the other attributes of the TCS are assumed to be random.

The *a-priori* assumptions on the mixing distribution vary in two different ways:

- in the first case (assumption 1) we assume all the varying parameters to be normal. This specification is employed for the non-panel as well as for the panel estimates;
- in the second case (assumption 2), which is estimated only for the panel case, we impose restrictions on the signs of some parameters, in particular

[2] All the estimation were conducted using the GAUSS code available from http://elsa.berkeley.edu/~train. The authors wish to thank Prof. Train for making this source code available to the public.

on 'noise' and 'aesthetic improvements' (beauty). The first is constrained to be negative, while the second to be positive. The constraint is implemented by using a log-normal distribution, so that it be limited to the appropriate orthant by an adequate change of sign of the vector of data.

Assuming a log-normal distribution implies that if $\hat{\mu}$ and $\hat{\sigma}$ are the location and scale parameter estimates for the distribution of the 'log-noise' taste, then the median, mean and standard deviation of the 'noise' parameter are $\exp(\hat{\mu}), \exp(\hat{\mu} + 0.5\hat{\sigma}^2)$ and $\exp(\hat{\mu} + 0.5\hat{\sigma}^2)\sqrt{\exp(\hat{\sigma}^2 - 1)}$, respectively. However, with this being a public good study, and because of the relevance of the median voter attitude to determine political outcomes, we focus our attention on median estimation. So, the reported estimates are computed at the median value of the estimated taste parameter distribution. That is, for the taste parameter whose distribution is normal we will have:

$$WTP_k = MRS_{k,£} = U'_k/U'_£ = \beta_k/\beta_£ = \hat{\mu}_k/\hat{\mu}_£.$$

Where β is the taste parameter, and k and £ index the generic TCS attribute and the cost to the household respectively.

While for one whose distribution is assumed to be log-normal we will have:

$$WTP_k = MRS_{k,£} = U'_k/U'_£ = \beta_k/\beta_£ = \exp(\hat{\mu}_k)/\hat{\mu}_£.$$

The choice of the median avoids the problem of excessively large $\hat{WTP}$ which is often encountered when using the first central moments in log-normal specifications of taste parameters.

As the small sample properties of ratios of simulated maximum likelihood estimates are unknown, interval estimates around median WTP are approximated by randomly drawing from an approximation of the asymptotic sampling distribution of the simulated maximum likelihood estimates, which is a variant of Krinsky and Robb parametric bootstrapping for non-linear functions of parameter estimates.

3 Econometric issues and results.

3.1 Contingent valuation estimates.

The results from the contingent valuation data are reported in table 2. A more detailed study of the different estimates obtained in the three towns are reported in Scarpa *et al.* (forthcoming), while here we only report the pooled estimates. The estimated fraction of zero bidders is just over 20% and this is taken to represent the fraction of HHs not willing to pay for ESR to the customary 30 mph limit. However, the median WTP estimate is nearly £8 per year in increased council taxes, while the mean WTP estimate is just over £20. This skewed distribution of WTP is implicit in the log-normal assumption, but it fits the nature of the data well. In fact, as it also emerged from the focus groups, those categories of people that feel threatened by speed (the elderly and families with small children) are indeed willing to pay high amounts.

3.2 Fixed parameters conditional logit.

The conventional fixed parameter logit estimates for the two populations are reported at the top of tables 3 and 5 respectively (Standard Logit A and Standard Logit B). The tables also report the estimated benefit estimates and their 90% confidence intervals. It can be noted that:

(1) Noise reduction is estimated to be worth around £2 a decibel by population A, and only approximately £1 by population B. The approximate confidence intervals of the two estimates do not overlap, — although this is not a proper test of the null of no difference — it suggests that this hypothesis might be rejected at conventional values of α and this is in keeping with the *a-priory* expectation.

(2) Speed reduction is estimated to be worth exactly the same to both populations in these models, that is about £0.22 per mile reduction of excessive speed (over 30 mph). This estimate for population A, shows an approximated asymptotic 90% confidence interval of $0.19 \div 0.58$. Assuming linearity of this value over the interval of speed reduction needed to bring down speed from the currently observed 45 mph to the ESR limit of 30 mph (which was the scenario hypothesized in the contingent valuation survey), the estimated median (and mean) WTP 90% confidence interval is £2.85 ÷ £8.7. Comparing this interval with that obtained from the contingent valuation study we find that they substantially overlap (over the interval £6.72 ÷ £8.7). The data seem not to reject the hypotheses of convergence between estimates across the two independent samples from different SP approaches.

(3) The point estimates for the impact of an 'aesthetically improved' engineering solution are around £1.5 in population A and £2.5 in population B, but the approaximated confidence intervals overlap, suggesting that the null of no difference may well not be rejected, at least in these models. So the evidence in the data does not resolve this issue, for which we only had a weak *a-priori* expectation about population A being willing to pay more than population B.

(4) Finally, reduction of waiting time for crossing from the status-quo to less than a minute, which was taken as a measure of community severance,

appears to be valued much more (over £7) by population A, than (about £1.5) by population B. Yet again, the noise around these two point estimates is of such a magnitude that the null of no difference may well fail to be rejected by the data.

3.3 Mixing across choices with normality

The second specification employed is a mixed logit assuming independence across choices made by the same respondent, where all the unobserved heterogeneity is normal. These models are labelled 'Mixed Logit A1' and 'Mixed Logit B1' and they are reported in the intermediate section of tables 3 and 5.

The following observations are noteworthy:

(1) This kind of mixed logit specification improves the log-likelihood in both samples, but more dramatically so in the sample from population A. A likelihood ratio test conducted under the null that all four scale parameters be zero rejects the null of fixity of the taste parameters in both models, hence providing evidence of significant unobserved heterogeneity.

(2) In the model estimated from the sample drawn from population A (Mixed Logit A1) we notice that accounting for unobserved heterogeneity through mixing induces a marked increase in the estimated value for the 'aesthetics' of the TCS, which jumps to over £8, while the estimates for speed reduction also doubles to £0.50 per mph reduction. The location parameter estimate $\hat{\mu}$ for waiting time is now negative, showing that the majority of people are unwilling to pay for this kind of improvement.

For population B, the estimates also change somewhat, but to a lesser degree. The only attribute showing a change in benefits is that for waiting time. Speed reduction seems to be valued less, but the approaximated c.i. are still quite large to be taken as indicative of a clear-cut evidence.

3.4 Mixing across households with normality

The fixity of tastes across choices made by the same individual household can be accounted for by employing the panel version of the random parameter logit. We do so in models 'Mixed Logit Panel A1' and 'Mixed Logit Panel B1'. These estimates are reported in the upper part of tables 4 and 6.

With respect to these estimates, the following observations can be made:

(1) In the first instance we notice that the fit of the panel model is dramatically better. The log-likelihood values at a maximum for population A and

B are respectively 86.57% and 56.14% of the values observed in the FPL models, and 87.22% and 56.33% of the RPL models 'non-panel'. This indicates that accounting for stability of taste across repeated choices matters, and panel models should be considered whenever adequate and possible. This findings is of major relevance when one considers that most of the published work relied on estimation ignoring dependence across choices by the same respondent.

(2) As regards the estimates for the attribute values, it can be seen that benefits from noise reduction are now significantly higher in population A than in B. This is in keeping with our expectations, and was not evident in previous models.

(3) Speed reduction benefits are now more 'noisy'. This is in accordance with the statements collected during the post-interview debriefing during which many respondents who were drivers, or who drove as part of their jobs, declared that they consider TCS imposing speed reductions as a 'nuisance'. The randomness in the panel version seems to pick up this bi-polarity of behaviour, as represented by the large scale parameter estimate.

(4) Similar 'noisier' behaviour seems to be present for taste intensities pertaining to attributes such as 'aesthetic improvement'. This attribute appears to be more important for population A, as it is now estimated at £14 with a significant difference from population B for whom the estimate is now centred on £6.5.

(5) Finally, in both samples this model produces an inversion of signs in the location parameters for the attribute of reduction of waiting time at crossing. We have no explanation at the moment for this behaviour and speculations are welcome.

3.5 Mixing across households with normality and log-normality

Benefits from local goods often have a distance-decay effect. In another study on this issue we found these to be present in the sample of responses drawn from population A (Garrod *et al.* forthcoming). Another way to model these spatially related effects is to allow for a skewed distribution of taste parameters, by specifying randomness with a skewed and sign-constrained distribution.

In models 'Mixed Logit Panel A2' and 'Mixed Logit Panel B2' we assume that the taste parameter for 'noise' and 'aesthetic appeal' (beauty for short) be log-normally distributed respectively on the negative and positive orthant, while we maintain the normal distribution for 'speed' and 'waiting time' reductions to be normal. As we have seen we have evidence of residents being both in favour and against these TCS deliverables, and normality accommodates both of these behaviours by allowing positive and negative signs.

The estimates of these models, which are still panel models, are reported in the bottom part of tables 4 and 6.

With respect to the estimates in 'Mixed Logit Panel A2' and 'Mixed Logit Panel B2' the following observations can be made:

(1) For a start, the log-likelihood values seem to indicate that this change in the assumptions of randomness, fit the sample from population A better than those of normality in model 'Mixed Logit Panel A1', while the opposite is true for the model estimated from the sample of population B, where the likelihood in fact increases. We speculate that this might be in keeping with the different layout of the residents with respect to the trunk road, which is the major source of nuisance.
(2) By evaluating the benefits at the median in the log-normal parameter distribution we notice that in either sample the estimates for 'noise' and 'beauty' are slightly lower than they counterparts in 'Mixed Logit Panel A1' and 'Mixed Logit Panel B1'. Benefits from a reduction of speed and waiting time are estimated to be more or less the same as before.

4 Discussion and conclusions.

In the present paper we have presented and discussed estimates of benefits for a particular category of local public goods: traffic calming schemes (TCS) in affected towns. Both contingent valuation and multi-attribute choice-modelling surveys were developed for two populations of residents with different traffic problems. The study was designed to allow some comparisons of estimated benefit for effective speed reduction across two major stated-preference methods: contingent valuation and choice-experiments. The data cannot reject the null that estimated benefits be the same. This is encouraging, as the two survey instruments were quite different.

Expectations about the size of benefits across the two populations were postulated on the basis of the physical measurements of traffic effects for speed, noise emissions and waiting time for crossing which characterize the current status-quo. With a varying degree, most of the expected differences in the sizes of benefits for the other deliverables of TCS were confirmed by the benefit estimates, although these are sensitive to the particular kind of random utility specification employed.

Both these results can be interpreted as evidence in favour of procedural invariance between direct and indirect stated preference methods, and represent an encouraging result for practitioners.

Unobserved taste heterogeneity was accounted for by using mixed logit models. Mixing which accounts for taste stability across repeated choices seems to fit the pattern of observed choices markedly better than when mixing is across single choices. So, taste variation across individuals seems to be stronger than within repeated choices made by the same individual. This finding is relevant for most previously published choice modelling studies, where maximum likelihood estimates of taste parameters are derived by fixed logit models, and under the clearly erroneous assumption of independence across choices by the same individual.

Because of the policy relevance of information characterizing these two distinct sources of heterogeneity, we advocate that further research in mixed logit estimation should address this issue. Possibly by allowing simultaneous estimation of both the individual-specific and the choice-specific sources of unobserved heterogeneity.

In conclusion we believe that these results add to findings reported elsewhere (Layton, 2000) in indicating how relevant mixed-logit modelling can be in the derivation of improved estimates of welfare measures.

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5 Tables

Table 1
Attributes and levels.

Attribute	Unit of Measurement	Levels
Noise	Decibels	60, 70, 80
Speed	in ESR miles/hour	20, 30, 40
Beauty	Basic/Improved	0, 1
Cost	Pounds	£10, £20, £30
Severance	Long/short Waiting time for crossing	$0 = L < 1', 1 = 1' < L < 3'$

Table 2
Estimates of CVM spike model interval data.

N = 682, $\mathcal{L} = -1,001.38$		
Variable	Estimate	st.err./C.I.
μ	2.460	0.015
σ	1.267	0.065
p	0.205	0.069
$\hat{M}(WTP)$	7.71	[6.72 , 8.80]
$\hat{E}(WTP)$	20.77	[17.51 , 24.78]

Table 3
CE1: estimates of mixing distributions' parameters.

Standard Logit A. $\mathcal{L}$ = -3,530.38.					
Parameter	Estimate	St.Err.	p-values of z	$W\hat{T}P_k$	90% C.I. $W\hat{T}P_k$
cost	-0.020	0.003	0.000		
Noise	-0.039	0.003	0.000	1.96	[1.56 , 2.57]
Speed	-0.009	0.003	0.007	0.42	[0.19 , 0.62]
Beauty	0.010	0.026	0.707	0.49	[-2.63 , 1.77]
Wait	-0.075	0.048	0.119	3.75	[0.14 , 8.73]
Mixed Logit 1. $\mathcal{L}$ = -3,504.03.					
$cost$	-0.033	0.005	0.000		
$\hat{\mu}_{nse}$	-0.059	0.007	0.000	1.79	[1.45 , 2.31]
$\hat{\sigma}_{nse}$	0.089	0.022	0.000		
$\hat{\mu}_{spd}$	-0.018	0.005	0.001	0.56	[0.35 , 0.75]
$\hat{\sigma}_{spd}$	0.077	0.014	0.000		
$\hat{\mu}_{bty}$	0.268	0.078	0.001	8.20	[4.44 , 12.61]
$\hat{\sigma}_{bty}$	0.689	0.390	0.078		
$\hat{\mu}_{wait}$	0.041	0.073	0.574	-1.26	[-4.83 , 2.66]
$\hat{\sigma}_{wait}$	0.416	0.426	0.328		
Mixing across 3,304 **choices**.					

Table 4
CE1: estimates of mixing distributions' parameters.

Parameter	Estimate	St.Err.	p-values of z	$\hat{WTP}_k$	90% C.I. $\hat{WTP}_k$
Mixed Logit Panel A1. $\mathcal{L}$ = -2,249.54.					
$cost$	-0.047	0.005	0.000		
$\hat{\mu}_{nse}$	-0.105	0.008	0.000	2.26	[1.88 , 2.79]
$\hat{\sigma}_{nse}$	0.098	0.010	0.000		
$\hat{\mu}_{spd}$	-0.015	0.018	0.400	0.33	[-0.32 , 0.96]
$\hat{\sigma}_{spd}$	0.251	0.017	0.000		
$\hat{\mu}_{bty}$	0.651	0.107	0.000	13.99	[10.21 , 18.48]
$\hat{\sigma}_{bty}$	1.141	0.138	0.000		
$\hat{\mu}_{wait}$	0.354	0.096	0.000	-7.62	[-11.15 , 4.33]
$\hat{\sigma}_{wait}$	0.881	0.133	0.000		
Mixed Logit Panel A2. $\mathcal{L}$ = -2,228.21.					
$cost$	-0.051	0.005	0.000		
$\hat{\mu}_{\ln(nse)}$	-2.664	0.119	0.000	1.38	[1.11 , 1.75
$\hat{\sigma}_{\ln(nse)}$	1.178	0.106	0.000		
$\hat{\mu}_{spd}$	-0.007	0.014	0.594	0.14	[-0.32 , 0.58]
$\hat{\sigma}_{spd}$	0.260	0.017	0.000		
$\hat{\mu}_{\ln(bty)}$	-1.003	0.262	0.000	7.26	[4.74 , 11.16]
$\hat{\sigma}_{\ln(bty)}$	1.317	0.168	0.000		
$\hat{\mu}_{wait}$	0.341	0.099	0.001	-6.75	[-10.12 , 3.65]
$\hat{\sigma}_{wait}$	0.941	0.130	0.000		
Mixing across 413 **households**.					

Table 5
CE2: estimates of mixing distributions' parameters.

Standard Logit B. $\mathcal{L}$ = -4,007.58.					
Parameter	Estimate	St.Err.	p-values of z	$W\hat{T}P_k$	90% C.I. $W\hat{T}P_k$
cost	0.058	0.003	0.000		
Noise	-0.056	0.003	0.000	0.97	[0.85 , 1.10]
Speed	-0.013	0.004	0.001	0.22	[0.12 , 0.32]
Beauty	0.143	0.052	0.006	2.48	[0.95 , 4.15]
Wait	-0.084	0.051	0.101	1.46	[0.02 , 3.04]
Mixed Logit B. $\mathcal{L}$ = -3,992.82.					
$cost$	-0.060	0.005	0.000		
$\hat{\mu}_{nse}$	-0.065	0.006	0.000	1.09	[0.94 , 1.26]
$\hat{\sigma}_{nse}$	-0.025	0.036	0.495		
$\hat{\mu}_{spd}$	-0.007	0.006	0.181	0.12	[-0.03 , 0.26]
$\hat{\sigma}_{spd}$	0.087	0.013	0.000		
$\hat{\mu}_{bty}$	0.200	0.081	0.014	3.34	[1.09 , 5.68]
$\hat{\sigma}_{bty}$	0.550	0.451	0.223		
$\hat{\mu}_{wait}$	0.097	0.084	0.252	-1.62	[-3.97 , 0.70]
$\hat{\sigma}_{wait}$	0.787	0.356	0.027		
Mixing across 3,256 **choices**.					

Table 6
CE2: estimates of mixing distributions' parameters.

Parameter	Estimate	St.Err.	p-values of z	$\hat{WTP}_k$	90% C.I. $\hat{WTP}_k$
Mixed Logit Panel B1. $\mathcal{L}$ = -3,056.18.					
$cost$	-0.053	0.004	0.000		
$\hat{\mu}_{nse}$	-0.082	0.006	0.000	1.57	[1.31 , 1.88]
$\hat{\sigma}_{nse}$	0.083	0.007	0.000		
$\hat{\mu}_{spd}$	0.047	0.014	0.001	-0.89	[-1.39 , -0.43]
$\hat{\sigma}_{spd}$	-0.232	0.014	0.000		
$\hat{\mu}_{bty}$	0.345	0.079	0.000	6.57	[4.15 , 8.99]
$\hat{\sigma}_{bty}$	0.691	0.118	0.000		
$\hat{\mu}_{wait}$	0.285	0.077	0.000	-5.43	[-7.76 , -3.06]
$\hat{\sigma}_{wait}$	0.668	0.119	0.000		
Mixed Logit Panel B2. $\mathcal{L}$ = -3,059.34.					
$cost$	-0.052	0.004	0.000		
$\hat{\mu}_{\ln(nse)}$	-2.921	0.116	0.000	1.04	[0.84 , 1.32]
$\hat{\sigma}_{\ln(nse)}$	1.013	0.102	0.000		
$\hat{\mu}_{spd}$	0.046	0.014	0.001	-0.90	[-1.42 , -0.44]
$\hat{\sigma}_{spd}$	0.233	0.014	0.000		
$\hat{\mu}_{\ln(bty)}$	-1.280	0.317	0.000	5.39	[3.23 , 8.95]
$\hat{\sigma}_{\ln(bty)}$	0.854	0.270	0.002		
$\hat{\mu}_{wait}$	0.290	0.075	0.000	-5.62	[-8.02 , -3.32]
$\hat{\sigma}_{wait}$	0.632	0.107	0.000		
Mixing across 407 **households**.					

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