ESTIMATING RESIDENTS' WTP FOR EFFECTIVE SPEED RESTRICTION WITH REFERENDUM-CV

Riccardo Scarpa

DEAR, Universita` della Tuscia, Viterbo, Italy.

Ken Willis

CREAM and Department of Town and Country Planning,

University of Newcastle uponTyne, U.K

Guy Garrods

CREAM and Department of Agricultural Economics and Food Marketing,

University of Newcastle uponTyne, U.K

Version of 26/01/2000

ESTIMATING RESIDENTS' WTP FOR EFFECTIVE SPEED RESTRICTION

WITH REFERENDUM-CV

ABSTRACT

We present an empirical estimation of the distribution of WTP for effective speed restriction via implementation of local traffic calming schemes. Random samples are drawn from the populations of households of three centres intersected by main trunk roads with varying through traffic conditions. We retrieve the underlying WTP distributions from discrete-choice responses to site-specific referendum contingent valuation studies accounting for zero-bidders. We then test the hypothesis of different distributions across villages. The statistical analysis is first conducted by means of a parametric specification and then by a totally non-parametric one. Stated welfare changes for effective speed reduction are found to be sizeable and the parameters of the random utility models are plausibly related to differences in objective speed measures across centres. The results appear to encourage the use of the referencum-CV method in the estimation of local public goods.

JEL Classification: D12,D62,C42,C35,C14

ESTIMATING WTP FOR EFFECTIVE ENFORCEMENT OF SPEED REDUCTION FROM DICHOTOMOUS-CHOICE CV: THE CASE OF RURAL TRUNK ROADS

1. Introduction

High speed is probably perceived as the most undesirable externality of traffic along roads crossing residential areas. This is particularly true in rural villages crossed by long distance through routes, such as trunk roads. Here, the local authorities often do not take action to enforce speed restrictions effectively, because this is costly and the magnitude of benefits to the local residents are not well known. The fraction of traffic which is most inclined to exceed speed limits is that destined to further destinations along the through road, so the construction of a by-pass diverting fast traffic may suffice in reaching satisfactory speed behaviour within these developed areas. However, by-pass construction is a particularly expensive solution, and it is often impracticable.

In the UK, traffic engineering standards assess that effective speed restriction (henceforth ESR) is achieved when 85 percent of the vehicles cruise at a speed up to or below the speed limit. This degree of compliance is neither easy nor cheap to achieve. Traffic calming schemes are amongst the various measures that local authorities can apply to effectively enforce speed limits in residential areas in a relatively inexpensive fashion. Amongst the "losers" of such a local policy are people who use the through road for commercial transport on wheels and for commuting to and from other destinations. These lose out because their journey lengthens by their being held-up when driving through these centres. The potential beneficiaries are the local residents, who will enjoy increased safety and possibly decreased disturbance. But not all of the residents may find this policy beneficial. Some may oppose it, while other may feel indifferent towards it. Hence the economic analyst ought to account for this fraction. Our survey probes respondents households for all three types of behaviour.

The costs of implementing traffic calming schemes for effective speed restriction are easy enough to define. On the other hand, the estimation of the economic benefits enjoyed by the resident population as a consequence of a decreased average speed is a more challenging task. One major difficulty is related to the public nature of this good and the inherent absence of a private market for increased road safety. As is well known, the individual propensity to reveal reservation prices for public goods is quite low for lack of proper incentives for truth-telling.

One avenue to estimate the benefits of avoiding speed-related traffic accidents could be that of employing an actuarial approach. Another approach could involve hedonic techniques on residential property values. However, both these approaches have some well known shortcomings.

For example, collecting site-specific statistics of the type required for actuarial analysis is time consuming. Meanwhile injuries and even deaths may occur. Moreover, safety from speed-related accidents may well be worth to people more than the equivalent loss of earnings. When the statistics on the risk of injury are transferred from other sites, the transfer estimate often relies on assumptions that many may find implausible.

Hedonic approaches suffer from the lack of a reliable data on both market transactions and of property exposure to speed-related risks.

As an alternative, in this paper we present results from a referendum contingent valuation study. At the moment, this is the most commonly employed method for public good valuation, and it is particularly well investigated in the environmental economics literature (Bateman and Willis 1999). The objective is to assess the properties of the distribution of willingness-to-pay in the relevant population from statements collected in a referendum contingent valuation study. As a means of validation of the estimation method, we assess the consistency of the resulting estimates with *a-priori* objective information on recorded speed at each site.

In the statistical analysis of the results we make an attempt at taking stock of the recent modeling advances for discrete choice CV as they appear in the relevant literature. In particular, following Kriström 1997 and Ayala and An 1996a, we model the distribution of willingness-to-pay (henceforth WTP) amounts with a random utility model with positive probability mass at zero to account for those who are indifferent to the proposed public good. This fraction of the population is clearly not in the market for the public good. To validate the results that we obtained with this method with some *a-priori* expectations, we then proceed by formally testing various null hypotheses about the differences between the estimated WTP distributions at different sites and provide point and interval estimates of expected WTP.

Welfare estimates are obtained in three ways. First by accounting for respondents who declared to be indifferent to the proposed policy package (zero-bidders) in a "full information" model; then by ignoring such information with a "partial information" one; finally by means of the more robust KTM non-parametric estimation. The differences in the resulting estimates are contrasted and discussed.

The parametric estimates of expected WTP are computed by truncating the integral of the expectation at the upper bid amount (Duffield and Patterson 1991). This limits the effect of "fat-tails" which characterize the log-logistic distribution, hence producing a conservative estimate on the support of the data. We find *E*(WTP) can be placed with a 95 percent confidence interval between £12.53/year and £22.10/year across the three sites, while the estimated fraction of population outside the relevant market for ESR varies between 13 percent - recorded in the site with highest recorded actual average speed - and 28 percent, in the site with the lowest.

The remainder of the paper is organized as follows. In section 2 we outline the theory of benefit estimation with discrete choice contingent valuation data with a follow-up in the case of speed reduction. In section 3 we describe the methods employed for the data collection and the survey administration along with those for the econometric analyses. The results obtained are discussed in section 4, while our conclusions are drawn in section 5.

2. Theory

2.1 The nature of the problem

Speed reduction of through traffic in residential areas can be seen as a non-excludable and non-rival positive externality for local residents: a local public good. Microeconomic theory suggests that public goods are to be suboptimally supplied and allocated by competitive markets. Optimal supply can be achieved by local government intervention only with knowledge of the costs and benefits associated with the provision schedule. In the case of a traffic calming scheme the costs are relatively easy to determine, however, benefit estimation poses a challenge to the economic analyst for the lack of observable transactions related to this phenomenon.

In as much as speed reduction is effectively achieved by provisions enforced by local authorities and funded via taxes paid by local residents, the benefits associated with it are to be interpreted as provided by a local political market. Since no alternative market exists, these benefits remain unpriced and their marginal value unknown, unless transactions in simulated political markets are observed. Voting on referenda are one mechanism through which private preference can be revealed for collective goods (Deacon and Shapiro 1975, Mitchell and Carson 1989). Yes and no responses to referendum questions can be used to retrieve the salient features of utility changes (Hanemann, 1984, 1989; Hanemann and Kanninen 1999).

The introduction of ESR via locally funded traffic calming schemes, is correctly measured in welfare terms by the compensating variation C. That is, the amount that implicitly equalises the following utility levels:

$$U(\mathbf{x},0,m) = U(\mathbf{x},1,m\text{-}C), \qquad (1.)$$

where $U(\cdot)$ is the household's indirect utility function, $\mathbf{x}$ is the current consumption level of all other private and public goods, m is income, 0 and 1 indicate respectively the absence and presence of ESR. It is the money amount which makes a given household indifferent between enjoying ESR at the cost of C and not enjoying it and saving the amount C, other consumption levels being equal.

One problem with locally provided collective goods is that some members of the relevant population may well not be in the market for that good at all. In the case of ESR, for example, those households resident in sites sufficiently away from the main road, or who do not cross or use the main road frequently, may well receive no benefits. For those households not in the market $C = 0$. Further, it is plausible to assume C to be non-negative for ESR across the population (ESR is not a "bad"), this translates in the need for an econometric specification allowing for $C \in \Re^{+}$.

Under the above conditions it is necessary to employ a distribution which allows for a positive probability mass at WTP = zero. This class of models, in the context of discrete-choice CV, are called spike-models (Kriström 1997, Hanemann and Kanninen

1999), where the word "spike" refers to the parametric probability estimate of observing a WTP = zero in the population.

So the objective of the investigation is to estimate the distribution of WTP as a proxy for C, starting from sample responses to a simulated political market, that is, from discrete Yes-No responses to survey questions. This estimate should identify the fraction of population indifferent to the proposed change (implementation of ESR) as well as the distribution of values of C across those for whom $C > 0$.

In multi-site analyses, a further interesting question to ask is whether or not the distributions estimated from samples drawn from different sites are significantly different, and whether these differences can be explained by observed differences in traffic features across sites.

2.2 Linking referendum responses to utility changes and the role of follow-ups

In the typical referendum CV question the respondent is asked whether or not she would vote in favour of a government programme that would bring about a change in the provision of a given public good which involved a personal cost of t, in terms of increased taxes or expenditures.

The respondent will reply with a "Yes" only if her utility level in the presence of the proposed change and cost exceeds the utility level in its absence. In other words:

$$\text{Yes} \rightarrow U(\mathbf{x},1,m\text{-}t) > U(\mathbf{x},0,m) \qquad (2.)$$

However, although discrete choice responses require low cognitive effort and may be consistent with real referendum formats (Carson *et al.* 1999), they are also very sample inefficient. In fact, an observed positive response only reveals that $C > t$. To increase the sample efficiency of this method Hanemann *et al.* (1991) have proposed to follow-up the first question with a reiteration. If the first amount encounters a positive response, then the follow-up question is reiterated at a higher amount *th*, with $th > t$. If, instead, the first amount t is rejected, the follow-up employs a lower bid amount $tl < t$. This approach relies on the assumption that both first and second response are driven by the same distribution of C. In this case the following implications hold:

$$\text{Yes-Yes} \rightarrow C > th \,, \tag{3.}$$

$$\text{Yes-No} \rightarrow th > C > t, \tag{4.}$$

$$\text{No-Yes} \rightarrow t > C > tl, \tag{5.}$$

$$\text{No - No} \rightarrow tl > C \,, \tag{6.}$$

with strict equalities that are undefined. Thus, in the presence of zero-bidders (those who are not in the market for the proposed change) amongst the "No - No" respondents there will be those for which $C = 0$, which need to be identified by a de-briefing question. As it is evident from the conditions laid out above, the use of a follow-up restricts the interval within which the real measure is contained, hence making estimation of the underlying distribution more efficient at any given sample size. Further bounding (a second follow-up) have been shown to bring about only minor efficiency gains under conventional specifications (Scarpa and Bateman, forthcoming) while they encounter a higher risk of respondent tiredness and reiteration bias.

3. Data and methods

3.1 The Survey

The CV survey was conducted in three separate villages in the North East of England: Haydon Bridge, Seaton Sluice and Rowlands Gill. These are all crossed by fairly busy trunk roads with a sustained through traffic. Actual speed conditions were measured on site and detected to be below the 85 percent compliance threshold, which is deemed to be the definition of effective speed limit enforcement. In particular, the measured 85th percentiles were 42 mph in Seaton Sluice, 40 mph in Rowlands Gill and 35 mph in Haydon Bridge.

The HHs in the sample were randomly drawn from the residential telephone listings of each village. Interviews were conducted by phone in the period between March and May 1999 and at times in which the head of the HH was likely to be found at home. Whenever possible, interviewers aimed at surveying the member of the HH in charge of council taxes payments. The vehicle of hypothetical payment was an increase in the yearly HH council tax for the duration of the traffic-calming scheme ensuring ESR. About a quarter of the selected sample declined to conduct the interview.

Three focus groups were conducted and the outcomes enabled the testing of the wording of the questionnaire and the identification of the general sensitivity of how the public regards traffic and speed reduction. They also allowed us to identify the elements of the initial bid vector, which were then up-dated after the first 300 responses had been collected (Table 1). Since both parametric and non-parametric distribution estimation were intended, and these imply different prescriptions for efficiency in bid-design, the bid up-date had to accommodate two needs. Parametric estimation of measures of central tendency for WTP makes little use of bid amounts placed away from the mean WTP, such as those in the tails, (Kanninen 1993) while non-parametric estimation requires a good investigation of the behaviour along the whole investigated bid range (McFadden 1994). As a compromise we proceeded by increasing the probes on the estimated percentiles around the estimated mean and reduced those in intermediate ones, while maintaining some probes on the extreme percentiles.

The relevant content of the questionnaire is presented in the appendix to this paper. The observed pattern of responses is presented in Table 2.

3.2 Parametric Estimation

Unconditional probability estimates of positive responses to referendum questions are sufficient to identify the relevant parameter of the WTP distribution in the population (McFadden, 1994). We hence concentrate on unconditional estimation.

Parametric estimation of a WTP distribution allowing for a positive probability mass at zero has been proposed by Kriström (1997) and requires the decomposition of the probability of positive response into a mixture of a cumulative distribution function over the plausible WTP range and a parameter called "spike", denoted here by ρ. This represents the probability of a HH producing a "No-No" response as a consequence of not being in the market for the public good of interest.

Following An and Ayala 1996a, we define the probability as follows:

$$M(WTP \leq x;\theta,0) = M(x;\theta,0) = \begin{cases} 0, & if \;\; x<0 \\ \rho, & if \;\; x=0 \\ \rho+(1-\rho)H(x,\theta), & if \;\; x>0 \end{cases} \qquad (7.)$$

Where $H(x,\mathbf{q})$ is a cumulative distribution function of the probability of positive response equivalent to that of WTP, with $H(0,\mathbf{q}) = 0$, while the probability of WTP = 0 is independently estimated by ρ.

The identification of a HH as being into the market for the proposed public good is equivalent to the HH having a positive value for the perspective good, in our case ESR. These "zero bidders" were identified in our study by asking whether or not the HH is willing to pay any amount for ESR to those respondents who answer "No" to both the first and the follow-up bid-amount.

No HH in the sample showed negative values for ESR. Negative values are also implausible for residents, who are unlikely to lose out in a scenario with reduced speed of through traffic. Hence a cumulative distribution function spanning the non-negative orthant is adequate to model probability at various bid amounts. In our econometric analysis we adopted the log-normal distribution.

Following An and Ayala 1996a, the spike parameter, that is the probability of being a zero-bidding HH, can be computed using the information of self-revealed zero-bidders in the sample in a "full information" model, or ignoring this information, the "partial information" model.

3.2.1.*Full information model (FIM)*

This model gives rise to the following log-likelihood function:

$$\ln L(\rho,\theta)=\sum_{i=1}^{N} \{(I_i^1 I_i^2)\ln[(1-\rho)(1-H_i(t^h;\theta))]+I_i^1(1-I_i^2)\ln[(1-\rho)(H_i(t^h;\theta)-H_i(t;\theta))]+(1-I_i^1)I_i^2$$

$$\ln[(1-\rho)(H_i(t;\theta)-H_i(t^l;\theta))]+[(1-I_i^1)(1-I_i^2)-\mathrm{ZB}_i]\ln[(1-\rho)H_i(t^l;\theta)]+\mathrm{ZB}_i\ln(\rho)\}, \qquad (8.)$$

for i= 1, 2, … , N , where $H(x,\theta) = \Phi(\alpha + \beta \ln(t))$, $\Phi(\cdot)$ is the standard normal cdf, $\theta = \{\alpha,\beta\}$, I^1 and I^2 are the indicator functions for a first and second "Yes" response respectively, ZB is the "zero-bidding" indicator function, and t indicates the first bid amount while t^h and t^l indicate the high and low follow-up bids respectively.

3.2.2.*Partial information model (PIM)*

This model gives rise to the following log-likelihood function:

$$\ln L(\rho,\theta) = \sum_{i=1}^{N} (I_i^1 I_i^2)\ln[(1-\rho)(1-H_i(t^h;\theta))] + I_i^1(1- I_i^2)\ln[(1-\rho)(H_i(t^h,\theta)-H_i(t;\theta))]$$

$$+(1- I_i^1)I_i^2\ln[(1-\rho)(H_i(t;\theta)-H_i(t^l;\theta))] + (1- I_i^1)(1- I_i^2)\ln[\rho + (1-\rho)H_i(t^l;\theta)] \quad (9.)$$

Note that in this model the spike parameter draws from the entire pool of "No-No" responses, while in the previous model only self-declared zero bidders are assigned to the computation of the spike. As a result the "full information" model is more constrained than the "partial information" one, and it achieves a maximum at lower values.

3.2.3. *Parametric expected WTP estimates*

The parameter estimates from both of the above models allow the estimation of expected WTP as an integral truncated at a given upper amount t^{max}, using the following formula:

$$E[WTP(t^{max})] = \int_0^{t^{max}} (1-\rho)\Phi[\alpha + \beta\, ln(t)]dt \quad (10.)$$

These estimates will clearly be sensitive to the choice of t^{max}, so that higher t^{max} will produce higher values. As a matter of general practice t^{max} = maximum bid amount, which in our case is £50. The sensitivity of the expectation to higher upper limits of integration may also be investigated.

3.3 Parametric Hypothesis Testing

A number of restrictions can be tested from the pooled sample of responses collected in the three villages to test the existence of significant differences across sites. Given the simplest linear-in-parameter specification of the indirect utility difference (Hanemann 1984) $v = \alpha + \beta\ln(t)$, site-specific dummies can be tested for both the effect on the constant α and the slope β given a baseline site. The slope parameter β can be interpreted in this context as the negative of the marginal utility of money and the relative site-specific dummies can therefore be interpreted as site effects on the marginal utility of money. The constant parameter can instead be interpreted as the mean effect in utility change, therefore the constant site specific dummies represent site effects on this mean. Sites with higher observed speed are expected to show higher site-specific utility effects.

We choose to obtain parameter estimates via maximization of the sample likelihood. An adequate specification test to assess the significance of these site-specific dummies in this context is the likelihood ratio test. The statistic of relevance is known to be asymptotically distributed χ^2_k, where k are the degrees of freedom, represented by the number of parameter restrictions.

To assess site-specific effects we proceed by testing parameter restrictions for constant dummies, slope dummies individually and then for the joint addition of both types of dummies.

Let us identify the constant (α) and slope (β) site specific dummies with the subscripts HB and SS for Haydon Bridge and Seaton Sluice respectively. α_{HB} and $\beta_{HB} = 1$ if the response was recorded at a HH resident in Haydon Bridge, 0 otherwise. Similarly for α_{SS} and β_{SS} that correspond to the Seaton Sluice dummy. Hence dummies pick up differences with respect to the baseline case of Rowlands Gill.

The following null hypotheses are of interest:

1) H_0^1 : $\{\alpha_{HB} , \alpha_{SS} \} = \mathbf{0}$, H_A^1 : at least one element of $\{\alpha_{HB} , \alpha_{SS} \} \neq \mathbf{0}$.

This test is implemented by using Model II as the unresticted model, where

$v = \alpha + \beta\ln(t) + \alpha_{HB} + \alpha_{SS}$.

Rejection of the null implies that mean effects in utilty changes exist in at least one of the two sites from the Rowlands Gill estimated baseline.

2) H_0^2 : $\{\beta_{HB}, \beta_{SS}\} = \mathbf{0}$, H_A^2 : at least one element of $\{\beta_{HB}, \beta_{SS}\} \neq \mathbf{0}$.

This test is implemented by using Model IV as the unresticted model, where

$v = \alpha + \beta\ln(t) + \beta_{HB}\ln(t_{HB}) + \beta_{SS}\ln(t_{SS})$.

Rejection of the null implies that marginal utility of log-money is different at least in one of the two sites from that estimated for Rowland Gill. This can be due to differences in disposable income across residents in the various villages.

3) H_0^3 : $\{\alpha_{HB}, \alpha_{SS}, \beta_{HB}, \beta_{SS}\} = \mathbf{0}$, H_A^3 : at least one element of $\{\alpha_{HB}, \alpha_{SS}, \beta_{HB}, \beta_{SS}\} \neq \mathbf{0}$

This test is implemented by using Model III as the unresticted model, where

$v = \alpha + \beta\ln(t) + \alpha_{HB} + \alpha_{SS} + \beta_{HB}\ln(t_{HB}) + \beta_{SS}\ln(t_{SS})$.

Rejection of the null implies that at least one of these effects is significantly different from that estimated for Rowland Gill. Throughout tests 1) - 3) the restricted model is represented by the simple constant-slope specification.

Individual restrictions on constant and slope site-specific dummies can also be informative and are here tested using the unrestricted Model III, down to Models II and IV, respectively for constant or slope dummies. These give rise to the following tests.

4) H_0^4 : $\{\alpha_{HB}, \alpha_{SS}\} = \mathbf{0}$, H_A^4 : at least one element of $\{\alpha_{HB}, \alpha_{SS}\}$ is significantly different from zero along with $\{\alpha, \beta, \beta_{HB}, \beta_{SS}\}$ in explaining the probability of positive response.

This test is implemented by using Model IV as the restricted model.

Rejection of the null implies that at least one of the site-specific constant effects is significantly different from that estimated for Rowland Gill when slope differences are accounted for.

5) H_0^5 : $\{\beta_{HB},\beta_{SS}\}$ = $\mathbf{0}$, H_A^5 : at least one element of $\{\beta_{HB},\beta_{SS}\}$ $\neq$ is significantly different from zero along with $\{\alpha,\ \beta,\ \alpha_{HB},\ \alpha_{SS}\}$ in explaining the probability of positive response.

This test is implemented by using Model II as the restricted model.

Rejection of the null implies that at least one of the site-specific slope effects is significantly different from that estimated for Rowland Gill when constant differences are accounted for.

In tests 4) and 5) the unrestricted model is Model III.

The above five tests are conducted for both the FIM as well as the PIM specifications so as to assess the invariance of the conclusions with regard to these alternative specifications.

3.4 Nonparametric estimation with Kaplan-Meier-Turnbull (KMT)

Parametric estimation provides a powerful means to relate economic theory to observed data. Often the estimated parameters have a clear economic interpretation in terms of marginal effects or elasticities, hence easing econometric analysis. Unfortunately the statistical identification of theoretically meaningful parameters often comes at a high cost. The assumptions required for parametric estimation are often theoretically unsubstantiated and inherently unverifiable. In the context of commonly used parametric maximum likelihood estimators, model misspecification leads to biased estimates. For this reason more robust non-parametric estimators have been developed and applied to CV discrete responses with follow-ups. Robust estimation allows the identification of robust estimates of probability of positive response in the population over the investigated bid range, and hence it delivers robust welfare estimates.

In the context of CV nonparametric interval and censored data, probability estimation can be achieved by using the so called Kaplan-Meier-Turnbull (KMT) estimator. Seminal papers proposing this estimator are Kaplan and Meier (1958) and Turnbull (1974, 1976). Haab and McConnell (1997) investigate the properties of this estimator in the specific context of CV studies. An and Ayala (1996b) provide a generalised algorithm to compute these estimates in the context of CV follow-up data.

The KMT probability estimator produces a monotonically non-decreasing step function over the investigated bid values. In our case, the point probability of positive response is estimated at the vector {1,2,5,10,15,20,30,40,50}. Point estimates of expected WTP can be obtained by discrete integration under the step function probability estimates. These are often referred to as lower-bound estimates for two reasons. Firstly because the imposed monotonicity creates a downward bias in the probability estimates, secondly because integration can be conducted only within the support of the bid range, up to the maximum bid, as extrapolation is infeasible due to the non-parametric nature of the estimates. Close-form expressions for the variance of the expected WTP are available for single bound discrete choice responses (Haab and McConnell, 1997). Confidence intervals for expected WTP derived from interval-data from follow-ups, can instead be approximated by means of naïve bootsrap techniques (Efron, 1981).

4. Results and discussion

4.1 Pooled Sample Parametric Estimates

The pooled sample gave rise to the full (FIM) and partial information model (PIM) parameter estimates presented in Table 3 and 4. With the exception of the slope dummy for Seaton Sluice, all the parameter estimates are individually quite significant on the basis of the estimated asymptotic standard errors. In the FIM the negative signs on α_{HB} and α_{SS} in Model II would seem to indicate that on average the utility change for achieving ESR is superior in Rowlands Gill compared with the other two sites. However, when the slope dummies are included (Model III), accounting for differences in income as reflected in changes in the marginal utility of money, the net effect on Seaton Sluice (α_{SS}) becomes positive and significant while that in Haydon Bridge is much more negative. This is in accordance with the gradient of observed speeds of through traffic: in sites where ESR reduces speed most the effect in estimated utility difference is higher. Constant site-specific dummies reflect changes with respect to Rowlands Gill, which is the site with intermediate observed speed (85^{th} percentile = 40mph). An observed positive change for α_{SS} is consistent with the fact that in Seaton Sluice ESR at 30 mph will reduce it from the observed 42 mph, hence producing a higher utility than in Rowlands Gill where the reduction is only from 40mph. Similarly, a negative effect in the Haydon Bridge constant dummy is consistent with a lower improvement with respect to Rowlands Gill, since the observed 85^{th} percentile in this site was 35 mph. Similar results are born out by the PIM estimates.

The slope site-specific dummies in Model III are concordant in both FIM and PIM specification in indicating that in Haydon Bridge money is more valued than in Rowlands

Gill, while the coefficient for Seaton Sluice is negative in FIM and positive in PIM, but statistically significant in neither.

4.2 Parametric Specification Tests

The hypotheses tests outlined in paragraph 3.3 were conducted at the 5 percent significance level and results are shown in Table 5. All hypotheses of restrictions were rejected in the PIM, indicating that for this specification relevant site specific dummies significantly help improve the fit to the observed pattern of responses. With the exception of H_0^1 and H_0^2 all the hypotheses were rejected in the FIM specification too.

These results seem to indicate that site-specific model estimations may be warranted, so as to disentangle the various effects at the village level and to obtain more reliable welfare estimates associated with the proposed public good.

4.3 Site-specific parametric estimates

A test of whether site-specific parametric spike models fit significantly better than a unique model estimated from the pooled sample can be conducted by means of a likelihood ratio test. The unrestricted log-likelihood is represented by the sum of those estimated for each site-specific model , $\ln L_{\mathrm{HB}} + \ln L_{\mathrm{SS}} + \ln L_{\mathrm{RG}}$, while the restricted one is represented by the relevant pooled sample Model I. The degrees of freedom are 6, the number of parameter restrictions.

The estimates necessary for conducting such a test are reported in Tables 3 and 4. In the case of the FIM specification we obtain a test statistic of 38.342 with a *p*-value of nearly one, strongly rejecting the restrictions implicit in the pooled model in favour of the site-specific ones. The analogue test for the PIM gives a *p*-value of 0.964, rejecting the restrictions at minimum confidence level of 3.6 percent.

Given these rejections of the pooled model we obtain point estimates of expected WTP by computing the integral of the expectation at four different truncation points. These are presented in the bottom rows of Table 3 and 4. The superscript 1 refers to a truncation point of £50, the maximum amount in the bid range; the superscripts 2, 3 and 4 refer respectively to extrapolations at £80, £100 and £120. It can be seen that the resulting estimated expectations are quite sensitive to choice of truncaton point due to the typical "fat tails" of the log-normal distribution. For each of the estimates are reported the corresponding 95 percent confidence interval, obtained by parametric boostrapping 10 thousand times the asymptotic distribution of the ML estimates (Krisnky and Robb, 1986, Cooper 1994). Expected WTP estimates from FIM and PIM models are very similar. However, while for Haydon Bridge and Seaton Sluice these amount to approximately £15/year (± ≈£2.70), in Rowlands Gill they amount to approximately £19/year (± ≈£2.76).

4.4 Nonparametric estimates

It is of some interest to compare the parametric estimates of E(WTP) values with those obtained using the KMT estimator. Point and interval estimates of the probability of positive response along with the respective expected values are presented in Table 7 for the pooled model and in Tables 8 to 10 for the individual sites. The 95 percent interval estimates approximations are obtained via naïve bootstrap (Efron 1981), by resampling with restitution the empirical distribution 10,000 times. We also report the median of the simulated WTP distributions.

All the non-parametric estimates are inferior to the parametric ones as they are conservative lower-bound measures. However, the estimated distributions of WTP in Haydon Bridge and Seaton Sluice still produce very similar expected WTP values

(£13.50/year and £13.80/year), while that for Rowlands Gill produces the highest expected WTP at £17.46/year.

5. Conclusions

Effective speed reduction is a sensitive issue in many small urban centres crossed by through roads with heavy commercial or commuting traffic. By-pass constructions are expensive enterprises and traffic-calming may be in inexpensive alternative to achieve effective speed restriction. However, benefits from these policies are of difficult estimation, as they provide local public goods without a proper market.

In this study we apply the referendum contingent valuation method to estimate the benefits of effective speed restriction in three peripheral centres crossed by trough roads in North-East England. Households were sampled from local residents and the interview was conducted by phone. Respondents were faced with a referendum scenario where traffic calming could be voted in subject to their willingness to pay an increased level of annual local taxes for the duration of the scheme. The survey instrument employed addressed the issue of indifference and negative WTP for the proposed programme, and found the latter not to be represented in the sample, nor in the focus groups or the pilot study. The sample fraction self-declaring to be indifferent to the policy varied between 28 and 13 percent.

The nearly seven hundred observed responses were analysed using parametric and more robust non-parametric methods allowing for estimation of zero-bidding (spike). The parametric random utility difference analysis shows that site-effects on utility differences are consistent with objective speed measurements, while those in marginal utility of log-income are consistent with differences in property values across the three villages. Specification testing indicates that individual site models perform better than pooled sample ones. Estimates of welfare measures derived from site-specific parametric models are concordant with those derived from the more robust non-parametric Kaplan-Meier-Turnbull estimator. These benefit estimates can be used in cost-benefit analysis as lower bounds on real benefits.

The estimation of the parameters of the indirect utility function from the pooled model are consistent with expectations with respect to the absolute values of speed reductions needed to meet the 85 percent standard of effective speed restriction. However, probably given to lower marginal utility of money associated with higher average wealth, the resident in Rowlands Gill have an estimated WTP per household which is about £4 higher than in Seaton Sluice and Haydon Bridge.

Altogether we find that this application of referendum CV provides estimates that appear to make economic sense, in terms of both absolute magnitude and of relative difference across sites.

6. References

An, M. Y. and Ayala R. A., 1996a. A mixture model of willingness to pay distributions, Mimeograph, available by FTP at “pub/man/papers/npecross.ps”.

An, M. Y. and Ayala, R. A., 1996b. Nonparametric estimation of a survivor function with across-interval-censored data, Mimeograph, available by FTP at “pub/man/papers/npecross.ps” or at “http://econwpa.wustl.edu”.

Bateman I. J. and Willis K. G., 1999. Valuing Environmental Preferences: Theory and Practice of the Contingent Valuation Method in the US, EC and Developing Countries, a cura di, Oxford University Press.

Carson, R. I., Mitchell R. C., 1989. Using Surveys to Value Public Good : The Contingent Valuation Method. Resources for the Future, Washington DC.

Carson Richard T., Groves Theodore and Machina Mark J. 1999. Incentive and Informational Properties of Preference Questions, Paper presented at the IX Conference of the European Association of Environmental and Resource Economists, Oslo 25-27 June, 1999.

Cooper, J. C., 1994. A comparison of approaches to calculating confidence intervals for benefit measures from dichotomous choice Contingent Valuation surveys, Land Economics **70**:111-22.

Deacon, R. and Shapiro P., 1975. Private preference for collective goods revealed through voting on referenda, The American Economic Review, December, **65**:943-955

Efron, B., 1981. Nonparametric standard errors and confidence intervals, Canadian Journal of Statistics, **9**:139-172.

Haab, T. C. and McConnell, K. E., 1997. Referendum Models and Negative WTP: Alternative Solutions, Journal of Environmental Economics and Management, **32**:251-270.

Hanemann, W. M. and Kanninen, B., 1999. The Statistical Analysis of Discrete Response CV Data, in Bateman and Willis, 1999, pp. 302-441.

Hanemann, W. M., 1984. Welfare Evaluations in Contingent Valuations Experiments with Discrete Responses, American Journal of Agricultural Economics, **66**:332-341.

Hanemann, W. M., 1989. Welfare Evaluations in Contingent Valuations Experiments with Discrete Response data: a reply, American Journal of Agricultural Economics, **71**:1057-1061.

Hanemann, W. M., Loomis J. and Kanninen, B., 1991. Statistical Efficiency of Double Bounded Dichotomous Choice Contingent Valuation, American Journal of Agricultural Economics, **73**:1255-63.

Kanninen, B. J., 1993. Optimal Experimental Design for Double-Bounded Dichotomous Choice Contingent Valuation, May, Land Economics, **69**:138-46.

Kaplan, E. and Meier, P., 1958. Nonparametric Estimation from Incomplete Observations, Journal of the American Statistical Association, **53**:457-81.

Krinsky, I. and Robb, A., 1986. Approximating the statistical properties of elasticities, Review of Economics and Statistics, **68**:715-719.

Kriström, B., 1995. Spike Models in Contingent Valuation: Theory and Illustrations. Invited paper to the First Toulouse Conference on Environmental and Resource Economics, Toulouse, France, March 30-31, 1995.

Kriström, B., 1997. Spike Models in Contingent Valuation. American Journal of Agricultural Economics, Aug, **79**:1013-1023.

McFadden D., 1994, Contingent Valuation and Social Choice, American Journal of Agricultural Economics, November, **76**:689-708.

Scarpa, R., and Bateman I., 1999. Efficiency Gains Afforded by Improved Bid Design Versus Follow-Up Valuation Questions in Discrete Choice CV Studies. Forthcoming in Land Economics, Issue of May, 2000.

Turnbull, B. W., 1974. Nonparametric estimation of a survivorship function with doubly censored data, Journal of the American Statistical Association, **69**:169-173.

Turnbull, B. W., 1976. The Empirical Distribution Function with Arbitrarily Grouped, Censored and Truncated Data, Journal of the Royal Statistical Society, Series B, **38**:290-295.

7. Tables

Table 1.

VALUE OF BID AMOUNTS

t^l	t	t^h
1	2	5
2	5	10
5	10	15
10	15	20
15	20	30
20	30	40
30	40	50

Table 2.

SUMMARY OF CV RESPONSES

First Bid	Total OBEs.	YY	YN	NY	NN	Zero Bidders
			Haydon Bridge			
2	48	25	6	1	16	9
5	11	6	1	1	3	2
10	60	12	11	7	30	18
15	7	3	0	0	4	2
20	59	12	6	6	35	18
30	8	1	0	1	6	4
40	44	8	4	4	28	13
			Seaton Sluice			
2	47	21	14	0	12	4
5	4	0	0	0	4	1
10	50	13	10	4	23	8
15	2	0	0	0	2	2
20	46	8	11	2	25	7
30	4	1	0	0	3	1
40	51	7	8	4	32	6
			Rowlands Gill			
2	37	21	7	1	8	6
5	16	7	6	2	1	0
10	46	14	16	1	15	13
15	10	4	1	0	5	2
20	45	10	13	2	20	11
30	11	2	1	1	7	3
40	36	10	6	2	18	10

Table 3.

FULL INFORMATION PROBIT MODELS FOR

POOLED SAMPLE *N* =682

	Model I	Model II	Model III	Model IV
Model *LL*	-1070.47	-1061.64	-1052.63	-1093.39
ρ	0.2053 (0.0155)	0.2053 (0.0155)	0.2053 (0.0155)	0.2053 (0.0155)
α	1.5452 (0.099)	1.8274 (0.1258)	2.0075 (0.1491)	1.4885 (0.0993)
β	-0.6023 (0.0343)	-0.6106 (0.0346)	-0.6793 (0.0461)	-0.6066 (0.0401)
α_{HB}		-0.2575 (0.1262)	-1.5391 (0.159)	
α_{SS}		-0.5075 (0.1208)	0.0768 (0.0057)	
β_{HB}			-0.5133 (0.141)	0.0198 (0.0016)
β_{SS}			-0.0005 (0.0038)	-0.0061 (0.0034)

Table 4.

PARTIAL INFORMATION PROBIT MODELS

FOR POOLED SAMPLE *N* =682

	Model I	Model II	Model III	Model IV
Model *LL*	-831.931	-829.019	-812.739	-830.11
ρ	0.2315 (0.0285)	0.2212 (0.0297)	0.1649 (0.0308)	0.2238 (0.0305)
α	2.1826 (0.1867)	2.3133 (0.2001)	2.3934 (0.2312)	2.0322 (0.2106)
β	-0.7998 (0.0564)	-0.7871 (0.0578)	-0.8332 (0.0696)	-0.7212 (0.0712)
α_{HB}		-0.2640 (0.1366)	-1.9667 (0.1892)	
α_{SS}		-0.3057 (0.1375)	0.0923 (0.0076)	
β_{HB}			-0.3448 (0.1595)	-0.0053 (0.0043)
β_{SS}			0.0009 (0.0043)	-0.0071 (0.0039)

Table 5.

RESULTS OF TESTS OF HYPOTHESES

	H_o^1	Deg.Fr. 2	H_o^2	Deg.Fr. 2	H_o^3	Deg.Fr. 4
	FIM	PIM	FIM	PIM	FIM	PIM
Restricted	-831.93	-1070.47	-831.93	-1070.47	-831.93	-1070.47
Unrestricted	-829.02	-1061.64	-830.11	-1062.94	-812.74	-1052.63
L-Ratio Stat.	5.82	17.67	3.64	15.07	38.38	35.68
P-value	0.95	1.00	0.84	1.00	1.00	1.00
Conclusion	*Not Rej.*	Str. Rej.	*Not Rej.*	Str. Rej.	Str. Rej.	Str. Rej.

	H_o^4	Deg.Fr. 2	H_o^5	Deg.Fr. 2
	FIM	PIM	FIM	PIM
Restricted	-830.11	-1062.94	-829.02	-1061.64
Unrestricted	-812.74	-1052.63	-812.74	-1052.63
L-Ratio Stat.	34.75	20.61	32.56	18.01
P-value	1.00	1.00	1.00	1.00
Conclusion	Str. Rej.	Str. Rej.	Str. Rej.	Str. Rej.

Table 6.

SITE-SPECIFIC PROBIT MODELS AND EXPECTED WTP ESTIMATES

	Full information			Partial Information		
	Haydon B.	Seaton S.	Rowlands G.	Haydon B.	Seaton S.	Rowlands G.
N	237	224	221	237	224	221
Mean ln$L$	-1.52362	-1.53015	-1.57216	-1.13442	-1.22259	-1.27815
Model ln$L$	-361.098	-342.754	-347.447	-268.858	-273.86	-282.471
ρ	0.2785 (0.0291)	0.1295 (0.0224)	0.2036 (0.0271)	0.2785 (0.0291)	0.251 (0.0569)	0.1441 (0.0396)
α	1.513 (0.1774)	1.1593 (0.1533)	2.1222 (0.1958)	1.513 (0.1774)	1.9441 (0.3346)	2.381 (0.294)
β	-0.5875 (0.0623)	-0.5438 (0.0534)	-0.723 (0.0667)	-0.5875 (0.0623)	-0.7531 (0.0975)	-0.8362 (0.0919)
$E(\mathrm{WTP})^1$	14.99 (12.53-17.73)	14.81 (12.33-17.49)	19.22 (16.58-22.10)	14.90 (12.50-17.44)	14.76 (12.36-17.33)	19.38 (16.88-22.02)
$E(\mathrm{WTP})^2$	18.79 (15.06-23.12)	18.34 (14.69-22.47)	23.71 (19.76-28.27)	17.47 (14.17-21.39)	17.40 (14.12-21.35)	22.88 (19.25-27.04)
$E(\mathrm{WTP})^3$	20.66 (16.21-26.06)	20.06 (15.74-25.12)	25.77 (21.12-31.42)	18.52 (14.69-23.26)	18.52 (14.77-23.34)	24.32 (20.08-29.49)
$E(\mathrm{WTP})^4$	22.19 (17.10-28.57)	21.48 (16.60-27.45)	27.38 (22.06-34.08)	19.28 (15.09-24.76)	19.35 (15.20-24.97)	25.36 (20.62-31.40)

The superscript 1 refers to a truncation point of £50, the maximum amount in the bid range; the superscripts 2, 3 and 4 refer respectively to extrapolations at £80, £100 and £120.

Table 7.

KMT NON-PARAMETRIC ESTIMATES FOR POOLED SAMPLE

t	Point Estimate	2.5º perc.	Median	97.5º perc.
0	1	1	1	1
1	0.7730	0.7195	0.7731	0.8245
2	0.7604	0.7075	0.7609	0.8117
5	0.6041	0.5579	0.6042	0.6494
10	0.5289	0.4818	0.5288	0.5754
15	0.3818	0.3385	0.3815	0.4241
20	0.3424	0.3001	0.3418	0.3854
30	0.2292	0.1910	0.2291	0.2683
40	0.1857	0.1466	0.1856	0.2258
50	0.108	0.0727	0.1077	0.1446
E(WTP)	14.84	13.39	14.83	16.30

Table 8.

KMT NON-PARAMETRIC ESTIMATES FOR HAYDON BRIDGE

t	Point Estimate	2.5º perc.	Median	97.5º perc.
0	1	1	1	1
1	0.6988	0.6059	0.6990	0.7913
2	0.6799	0.5891	0.6797	0.7690
5	0.5679	0.4915	0.5688	0.6455
10	0.4769	0.4012	0.4772	0.5548
15	0.3443	0.2758	0.3438	0.4148
20	0.2852	0.2161	0.2848	0.3557
30	0.2148	0.1537	0.2145	0.2781
40	0.1642	0.1026	0.1632	0.2275
50	0.1095	0.0502	0.1077	0.1696
E(WTP)	13.50	11.11	13.48	16.00

Table 9.

KMT NON-PARAMETRIC ESTIMATES FOR SEATON SLUICE

A	Point Estimate	2.5º perc.	Median	97.5º perc.
0	1	1	1	1
1	0.755	0.6584	0.7558	0.8502
2	0.755	0.6582	0.7555	0.8500
5	0.5407	0.4575	0.5411	0.6236
10	0.4884	0.4047	0.4892	0.5758
15	0.3642	0.2917	0.3637	0.4395
20	0.3404	0.2672	0.3406	0.4179
30	0.2163	0.1521	0.2164	0.2856
40	0.173	0.1076	0.1720	0.2423
50	0.0807	0.0305	0.0795	0.1374
E(WTP)	13.80	11.37	13.77	16.46

Table 10.

KMT NON-PARAMETRIC ESTIMATES FOR ROWLANDS GILL

t	Point Estimate	2.5º perc.	Median	97.5º perc.
0	1	1	1	1
1	0.8624	0.7859	0.8621	0.9361
2	0.8452	0.7663	0.8450	0.9203
5	0.7059	0.6286	0.7061	0.7820
10	0.6243	0.5423	0.6243	0.7034
15	0.4424	0.3670	0.4425	0.5213
20	0.4098	0.3319	0.4095	0.4873
30	0.2613	0.1906	0.2610	0.3380
40	0.2240	0.1502	0.2221	0.2979
50	0.1400	0.0703	0.1387	0.2139
E(WTP)	17.46	14.79	17.39	20.30

8. Appendix

Traffic regulations impose speed limits that are not always enforced successfully. In (NAME OF THE VILLAGE) our researchers have found out that 85% of the vehicles driving through the village on the main road (that's eight or nine vehicles in every ten) travel at a speed of (SPEED MEASURED AT THE VILLAGE) mph or faster, compared with the speed limit of 30 mph.

Now, suppose that the council were considering introducing a traffic calming scheme in your village. It is certain that this scheme will reduce the speed of traffic through the main road so that 85% of vehicles would be driving through the village at a speed of 30 mph or slower.

Unfortunately, the cost of this scheme is not covered by the local council budget. The only way to implement the speed reduction scheme in (NAME OF THE VILLAGE) is that each household resident in (NAME OF THE VILLAGE) be paying an additional fee to the council for this new safety scheme. This fee would be on top of the normal council tax.

We are now going to mention some money amounts, these can sound ridicolously high or low to you, please do not take these proposed amounts as an indication of value, because there are not. We are just interested in your honest answer.

Suppose, that the council wants to know the residents' opinion about this public programme and in order to do so it asks residents to vote for or against the realization of such a scheme in a local referendum.

This programme will cost your household additional yearly fees. Please, before answering, consider that there are other things you can buy with this money. Would you vote in favour of the scheme if it would cost you £(bid 1) every year in additional fees.? (if you are not paying local taxes because you are a pensioner, this fee would still apply to your household).

Yes **No**

If yes

Suppose now that the programme will cost your household £ (bid2) every year in additional local fees. Would you still vote in favour of the programme?

If No

Suppose now that the scheme it will personally cost your household £(bid2) every year in additional local fees. Would you vote in favour of the programme at this lower amount?

If two Nos.

Would you be WTP any amount of money at all for such a speed reduction programme through additional local fees?

Yes **No**

If No again.

With which of the two following statements would you most agree with?

a) "The reduction of traffic speed is of no value to my household and I am therefore willing to pay nothing for the proposed traffic calming scheme".

b) "I actively oppose the realization of traffic calming schemes and my household is willing to pay *not to have it implemented in my village.*"