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Alternative Models of Recreational Off-Highway Vehicle Site Demand

JEFFREY ENGLIN[1,*], THOMAS HOLMES[2] and REBECCA NIELL[3]
[1]*Department of Resource Economics, University of Nevada, Reno, Nevada, USA;* [2]*Southern Research Station, US Forest Service, Asheville, North Carolina, USA;* [3]*Department of Agricultural and Resource Economics, University of California, Davis, California, USA;* **Author for correspondence (phone +775-7844411; fax +775-7841342; e-mail ENGLIN@UNR.EDU)*

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Abstract. A controversial recreation activity is off-highway vehicle use. Off-highway vehicle use is controversial because it is incompatible with most other activities and is extremely hard on natural eco-systems. This study estimates utility theoretic incomplete demand systems for four off-highway vehicle sites. Since two sets of restrictions are equally consistent with utility theory both are imposed and the best fitting restrictions are identified using Voung's non-nested testing scheme. The demand system is modeled using both Poisson and negative binomial II distributions. Data are provided by a survey conducted at four recreational off-highway vehicle (OHV) sites in western North Carolina.

Key words: incomplete demand system, integrability, off-road vehicle, travel cost

JEL classification: Q26, C35, C51

1. Introduction

Demand systems are becoming increasingly popular for estimating recreational site demand because demand systems address the idea that multiple recreation sites could be substitutes and therefore should be estimated together (Burt and Brewer 1971; Englin et al. 1998). Complete, partial, and incomplete demand systems are three frameworks for estimating systems of demand equations (Epstein 1982). Incomplete demand systems estimate demand as a function of income and the prices of concerned goods without regard to the allocation of remaining income. Incomplete demand systems are favored for modeling consumer choice based on a subset of total goods, because, unlike complete or partial demand systems, incomplete demand systems avoid the restrictive assumptions of aggregation and/or separability.

However, estimation of an incomplete demand system requires implementation of cross-equation parameter restrictions to maintain theoretical

consistency between model estimations and underlying utility theory. Estimation of incomplete demand systems has become possible through the extensive work of LaFrance (1990), LaFrance and Hanemann (1984, 1989), and von Haefen (2002) identifying the necessary restrictions for maintaining integrability of the Slutsky symmetry matrix. LaFrance and LaFrance and Hanemann identified the integrability restrictions necessary for eight specifications of linear, log-linear, and semi-logarithmic demand systems. von Haefen extended the use of incomplete demand systems by adding 16 functional forms including linear, log-linear, and semi-logarithmic forms of expenditure and expenditure share systems.

In this context an important empirical issue is the choice of demand restrictions. This is a specialized form of the general problem of model specification. It is particularly important since the choice of restrictions implies a particular set of underlying behaviors. Since key restrictions in these models take the form of restrictions of own and cross-price effects these restrictions can imply strong restrictions of substitution behavior. Fortunately the data can provide tests of the consistency of restrictions with the data on observed behavior

In this study, two different cross-equation restriction specifications for a semi-logarithmic system of demand equations are estimated for visitation data to four recreational off-highway vehicle (OHV) sites in western North Carolina. The systems are estimated as travel cost models, and count data for OHV site visits are modeled using the Poisson and the negative binomial II distributions. Results from the two different cross-equation restriction specifications are compared using a two-part likelihood ratio non-nested testing procedure (Vuong 1989), and welfare measures are calculated for the ordinary and compensated demands.

The analysis of OHV riding is important because OHV riding is especially hard on the environment. The vehicles used on OHV recreational, whether two or four wheeled, create trails without vegetation, churn stream beds into mud and are uniformly noisy. It is difficult to have OHV use in a multiple use setting. It is also difficult to have OHV use in a setting that does not generate substantial negative externalities for other users. As a result OHV riders must generally be segregated from other recreational activities. Of course, upon reflection, it is also clear that OHV activities require large areas. If one is to ride on a motocross bike for an hour a substantial amount of trail or road will be covered unless it is simply a circular track. Given the impacts and scale of OHV riding it is important to have quality welfare measures so that good public policy can be implemented.

The remainder of this paper is organized as follows. Section 2 reviews the theoretical properties of incomplete demand systems and presents the semi-logarithmic demand equations used in this study. Data are described in

Section 3. Section 4 describes model estimation procedures and presents the likelihood ratio non-nested testing procedure that is used to compare the two different restriction specifications. Section 5 presents the results of model estimations, the non-nested testing procedure, and the welfare estimates. Finally, Section 6 is a discussion of results and conclusions.

2. Background

There are numerous feasible functional forms for modeling incomplete demand systems. The semi-logarithmic functional form is appealing for demand analysis because it restricts demand to be non-negative. This analysis considers one form of a semi-logarithmic demand system for which appropriate econometric cross-equation restrictions have been developed. The model is:

$$x_i = \alpha_i(z) \exp\left(\sum_{j=1}^{n} \beta_{ij}\rho_j + \gamma_i y\right) \tag{1}$$

where x_i is the quantity of visits to site i demanded, $\alpha_i(z)$ is a demand shifter that accounts for influences of a vector (z) of non-income shift variables, β_{ij} is a price coefficient that captures the influence of ρ_j (the cost of visiting site j) on demand for good i, γ_i is the income coefficient, and y is income. Two sets of parameter restrictions satisfy the integrability requirements that are consistent with the Slutsky symmetry condition. The derivations of these restrictions are described in LaFrance (1990) and von Haefen (2002).

2.1. RESTRICTION SET I

The first set of restrictions requires that (i) the demand shifter $\alpha_i(q)$ is positive, (ii) the income parameter is the same across sites, (iii) the cross-price effects are zero, and (iv) own-price coefficients are negative.

$$\alpha_i(q)0, \quad \forall i \tag{2}$$

$$\gamma_i = \gamma_j \tag{3}$$

$$\beta_{ij} = 0, \quad \forall i \neq j \tag{4}$$

$$\beta_{ii} < 0 \tag{5}$$

It is important to realize that, although Eq. (4) restricts the cross-price effects to be zero in the system of ordinary demand functions, the compensated cross-price effects might be non-zero. The expression for the compensated cross-price effect between site j and k for any individual is:

$$S_{jk} = \gamma x_j x_k \tag{6}$$

(Englin et al. 1998). If the parameter estimates representing the income effect (γ) is positive and significantly different than zero, and if individual n visits both sites j and k, then S_{njk} will be positive (i.e., the sites are substitutes). If individual n visits either site j or site k, but not both, then the compensated cross-price effect for individual n is (naturally) zero (i.e., the sites are independent). If none of the recreationists sampled visit both site j and k, then the average value of S_{jk} for the sample will be zero. Of course, Eq. (6) indicates that the compensated cross-price effects are symmetric.

2.2. RESTRICTION SET II

The second set of restrictions that satisfy the integrability requirements is more restrictive because only one site's intercept (demand shifting parameter) is estimated, and other site intercepts are calculated as functions of the first site intercept and own-price coefficients (Eq. 7). As in the first set of restrictions, the income effect is equal across sites. However, the cross-price effects of site k on all other sites are the same and are equal to the own-price coefficient of site k. Because the own price effect is negative (i.e., an increase in price is associated with a decrease in trips to that site), the cross-price effect maintains that recreation sites are complements (an increase in price at one site is associated with a decrease in the number of trips to other sites in the system).

$$\alpha_i = (\beta_{ii}/\beta_{jj})\alpha_j > 0 \tag{7}$$

$$\gamma_i = \gamma_j \tag{8}$$

$$\beta_{ik} = \beta_{jk} = \beta_{kk} \forall k. \tag{9}$$

The empirical fit of these two restriction specifications is compared using recreational site demand data for off-highway vehicle users.

3. Data

Data for these analyses are from a survey of off-highway vehicle (OHV) users that was conducted at four U.S. Forest Service managed OHV recreational sites in western North Carolina (Figure 1). Table I provides an overview of the four OHV sites. Volunteers collected data during the months of July–October 2000 at Badin Lake, Brown Mountain, Upper Tellico, and Wayehutta OHV recreation areas, as users exited the sites. A total of 357 surveys were collected: 97 at Badin Lake, 101 at Brown Mountain, 118 at Upper Tellico and 41 at Wayehutta. The data include numbers of visits to each of the four OHV areas in 1997, 1998, and 1999; site user fees; and demographic characteristics including sex, age, education, income, skill, and zip code.

For these analyses data were limited to surveys completed by individuals 18 years of age and older for which data were complete for at least one of the three focal years. This resulted in a total of 672 observations across the years 1997–1999. One aspect of the data is the issue of recall bias in the 1998 and 1997 trip counts. The fact that OHV trips are fairly work intensive activities lends credence to the likelihood that recall will be accurate, unlike frequent activities like fishing. Other analyses of infrequent activities, like overnight back country hiking (Englin and Shonkwiler (1995)) have used 10 year windows of recall in their analyses. As long as the recall is not systematically too low or too high the recall error will present itself as a form of hetero-

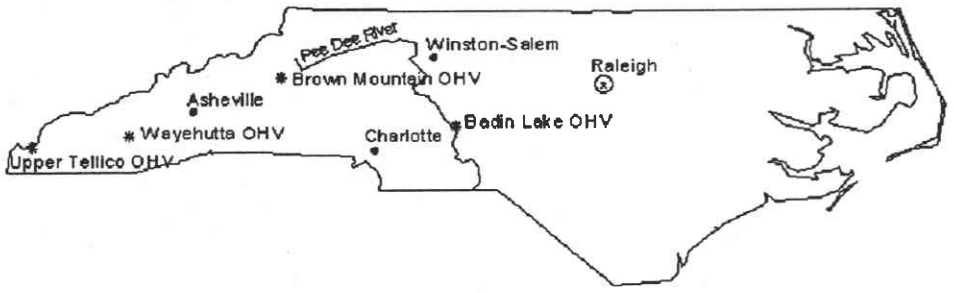

Figure 1. Location of the four United States Forest Service off-highway vehicle sites

Table I. Off-highway vehicle sites

Area	User type	Fee	Trail miles	Season
Upper Tellico	All: ATV, trail bike, 4WD	$5	40	All year
Wayehutta	ATV, trail bike	$5	22	April 1–December 15
Brown Mountain	Mostly ATV & trail bike	$5	33.5	March 16–January 1
Badin Lake	All: ATV, trail bike, 4WD	$5	16	April 1–December 15

scedasticity. As a result of this negative binomial models which can handle heteroscedastic data will be employed as a method of estimation.

A travel cost variable, 2 year variables, and four demographic variables (income, skill, education, and sex) were included in these analyses. COST was calculated as the sum of transit cost, opportunity cost of time, and a site fee which is constant across the sites. Transit cost was calculated as the distance to and from a site (based on the residential zip codes of survey respondents) multiplied by the cost per mile ($0.25/mile). Opportunity cost was calculated as one third of the individual's wage rate (annual income divided by hours worked each year) multiplied by travel time assuming an average speed of 60 mph. Dummy variables were included for 1997 and 1998 to account for annual differences in visitation rates from 1997 to 1999. INCOME was reported in the survey and scaled as $1000's for model estimations. SKILL level is self-reported by the respondents. They rate themselves as fitting into one of three categories: beginner, intermediate, and advanced, and was coded as 1, 2, and 3, respectively for these analyses. EDUCATION is the number of years of school completed by the survey respondent and FEMALE is a dummy variable for females.

4. Model Estimation

Model estimation was conducted in GAUSS, and visits to each OHV site (Badin Lake, Brown Mountains, Upper Tellico, and Wayehutta) were estimated as functions of cost (ρ_i), income (y), year, skill-level, education, and sex. The number of visits to each site was estimated as (i) a Poisson distribution and (ii) a negative binomial II (NB2) distribution with site-specific dispersion parameters (Cameron and Trivedi 1998). The Poisson distribution is commonly used when estimating recreation demand models because it is a non-negative, discrete distribution that allows for zero trips and provides unbiased parameter estimates regardless of the true underlying data generating distribution (Gourieroux et al. 1984). The Poisson log-likelihood function for a single site can be expressed as:

$$l_i = -\exp(x\beta) + x_i \ln(\exp(x\beta) - \ln(x_i!) \tag{10}$$

where $\exp(x\beta) = \lambda$ is the predicted/estimated number of visits for an individual to a site, and x_i is the observed number of visits for an individual to the site. One property of the Poisson distribution is that the variance equals the mean.

Like the Poisson distribution, the negative binomial II (NB2) distribution assumes non-negative integer values. However, NB2 allows the variance to differ from the mean by including a dispersion parameter, thus allowing for

the possibility for over-dispersion in the number of OHV trips. The NB2 log-likelihood for a single observation at a site can be expressed as:

$$l_i = -d^{-1}\ln(1 + d(\exp(x\beta))) + \ln\left(\frac{\Gamma(x_i + d^{-1})}{\Gamma(x_i + 1)\Gamma(d^{-1})}\right) + x_i \ln\left(\frac{d(\exp(x\beta))}{1 + d(\exp(x\beta))}\right) \quad d \geq 0; \quad x_i \geq 0 \tag{11}$$

where d is the site-specific dispersion parameter, $\exp(x\beta) = \lambda$ which is the predicted/estimated number of visits for an individual at the site, x_i is the observed number of visits of an individual to the site, and γ is the gamma function (see Table II).

One limitation of the NB2 distribution is that estimated parameters are not robust to distributional misspecification unless the dispersion parameter d is known with certainty, which it is not (i.e., it must be estimated). To account for potential distributional misspecification, White's (1980) method is employed to correct the variance–covariance matrix and to calculate White's corrected standard errors for both the Poisson and NB2 distributions. White's method provides a robust approach for hypothesis testing.

One other aspect of the distributional specification must be addressed. Parameters of the Poisson and NB2 distributions are *not* corrected for endogenous stratification which might result from over sampling people who often recreate at these sites. The full (rather than truncated) Poisson and NB2 distributions are used because zero trips to a particular site in the system are allowed (i.e., people intercepted at a particular site were asked about their visits to other sites). Finally, it is assumed that visits over consecutive years are independent.

In the log-likelihood functions shown above (Eqs. 10 and 11), λ is calculated using the semi-logarithmic demand equation shown in Eq. 1. However, Eq. 1 is modified as follows to include demographic variables and year:

$$\lambda_{im} = \alpha_i \exp\left(\sum_{j=1}^{n} \beta_{ij}\rho_{jm} + \xi z_m + \gamma y_m\right) \tag{12}$$

where λ_{im} is the estimated number of visits to site i by individual m, ξ is a vector of parameter coefficients for the vector z of demographic variables and year, and all other parameters are the same as described above for Eq. 1.

Four versions of this system of incomplete semi-log demand equations are estimated. These include estimation of both sets of parameter restriction specifications (Eqs. 2–5 and 7–9) using the Poisson and NB2 distributions.

Table II. Parameter estimates

	Restriction set 1		Restriction set 2	
	Poisson	NB2	Poisson	NB2
Badin Lake				
Intercept	0.4610 (1.4734)	0.2052** (0.0915)	0.6513 (0.8745)	0.1861** (0.0751)
Travel cost	−0.0239** (0.0023)	−0.0366** (0.0032)	−0.0022** (0.0004)	−0.0014* (0.0008)
d		1.7943** (0.1726)		2.6539** (0.1317)
Brown Mountain				
Intercept	0.7599 (6.9731)	0.1359** (0.0462)	0.5977	0.1227
Travel cost	−0.0321** (0.0030)	−0.0338** (0.0061)	−0.0021** (0.0006)	−0.0010** (0.0005)
d		1.8428** (0.1429)		2.5091** (0.1337)
Upper Tellico				
Intercept	0.2871 (0.5833)	0.0463** (0.0105)	0.7926	0.2233
Travel cost	−0.0099** (0.0013)	−0.0076** (0.0019)	−0.0027** (0.0006)	−0.0017* (0.0009)
d		1.2591** (0.1086)		1.4531** (0.0978)
Wayehutta				
Intercept	0.8200 (10.7845)	0.0864** (0.0240)	0.4710	0.1382
Travel cost	−0.0392** (0.0082)	−0.0269** (0.0069)	−0.0016** (0.0005)	−0.0011** (0.0005)
d		3.1137** (0.1878)		3.5382** (0.1588)
Demographic variables				
Income ($1000's)	0.0127** (0.0018)	0.0137** (0.0025)	0.0089** (0.0021)	0.0073** (0.0034)
1998	−0.5131** (0.1579)	−0.6744** (0.1499)	−0.5107** (0.1435)	−0.5914** (0.1297)
1997	−0.8004** (0.2342)	−0.9268** (0.1872)	−0.7987** (0.2113)	−0.9552** (0.1660)
Skill	0.5803 (0.3672)	0.8439** (0.1445)	0.5087** (0.1126)	0.8459** (0.1218)
Education	0.0629 (0.1116)	0.1316** (0.0369)	0.0125 (0.0313)	0.0296 (0.0405)
Female	0.6028** (0.2040)	0.3563* (0.2055)	0.5105** (0.1966)	0.5193** (0.2143)
Log-likelihood	−5704.44	−2517.06	−7441.74	−2680.46

* Significant at the 10% level

** Significant at the 5% level

Robust standard errors are calculated using White's method and are used for calculating t-values.

The two sets of parameter restrictions are compared using a two step non-nested likelihood ratio testing procedure (Vuong 1989). This test was run using a Gauss based program written by von Haefen (2002). The first step of Vuong's test determines whether the non-nested models can be distinguished. This is done based on pair-wise comparisons of log-likelihood values for each model specification across individuals. If it is determined that the models can be distinguished, the second step of the Vuong test is implemented to determine which model, if either, is preferred. Detailed descriptions of Vuong's non-nested test are provided in Vuong (1989) and in Englin and Lambert (1995).

5. Results

Table I presents the maximum likelihood estimates for the four models estimated in this study. The site-specific negative binomial dispersion parameters d are significantly different from zero ($P < 0.05$) for both restriction specifications, supporting that the NB2 models are preferred to the Poisson specifications; this is confirmed by comparing the log-likelihood values between the Poisson and NB2 specifications for each set of restrictions. The results of Vuong's two step non-nested likelihood ratio testing procedure show that (i) the NB2 models are distinguishable (test statistic = 390.4; $P < 0.001$) and (ii) the first set of restrictions is the preferred model (test statistic = 8.27; $P < 0.001$). Thus, the best-fit model is the NB2 specification of the semi-log incomplete demand system with the first set of restriction specifications.

Both sets of restrictions and distribution specifications provide same-signed coefficient estimates and similar sets of significant parameters. Travel cost parameters for all sites are negative and significant ($P < 0.05$) in the preferred model and negative and at least marginally significant ($P < 0.10$) in all models. The magnitudes of the own price parameter estimates suggest that demand is more elastic under the second set of restrictions. Income effects are positive and significant ($P < 0.05$) across all models. Individuals made significantly fewer visits to sites in 1997 and 1998 than in 1999 across all models, and the skill of riders had a significant positive effect on site demand in all models except for the Poisson estimation of the first set of restrictions. The coefficient on education was only significant in the preferred model (the NB2 specification of the first set of restrictions) and suggested that individuals with more years of education have a higher demand for the OHV sites. The coefficient on female was positive and significant across all models except the preferred model in which it was only marginally significant ($P < 0.10$). The female result is that women who do participate in OHV riding are more

avid than an otherwise identical male. Since fewer women in the general population OHV ride this suggests that tastes OHV riding amongst women is bimodal, with the women who choose to participate participating more extensively than their male counterparts.

Welfare measures associated with the Poisson and negative binomial models are calculated using the travel cost parameters, such that the per trip consumer surplus for an individual who takes a trip to site i is $1/\beta_i$. Based on the preferred model, consumer surplus for an OHV trip is \$27.32, \$29.59, \$131.58, and \$37.17 for Badin Lake, Brown Mountain, Upper Tellico, and Wayehutta, respectively. Estimated per trip consumer surpluses are slightly different for the Poisson specification of the first set of integrability restrictions, with values of \$41.84, \$31.15, \$101.01, and \$25.51 for Badin Lake, Brown Mountain, Upper Tellico, and Wayehutta, respectively. Interestingly, the ordering of sites by consumer surplus measures is different across the two distributions for the first set of parameter restrictions, but in both cases Upper Tellico OHV area provides the highest level of consumer surplus.

Consumer surplus estimations based on the second set of parameter restrictions are very different and much higher than for the first set of restriction specifications. Per trip consumer surpluses for the Poisson and NB2 specifications, respectively, were of \$454.55 and \$714.29 for Badin Lake, \$476.19 and \$1000.00 for Brown Mountain, \$370.37 and \$588.24 for Upper Tellico, and \$625.00 and \$909.09 for Wayehutta. The ordering of sites based consumer surplus and the second restriction set is both different across distributions and different than for the first restriction set. The compensated variation (CV) can be calculated using a method presented by Englin and Shonkwiler (1995). The formula is:

$$\mathrm{CV} = \frac{1}{\gamma}\ln\left(1 + \lambda\frac{\gamma}{\beta_i}\right) \tag{13}$$

where γ is the income coefficient and β_i is the cost coefficient for site i, these results are found in Table III. Since the income effects are small the difference between the Hicksian and Marshallian measures of welfare are virtually identical.

6. Conclusions

In this study, two sets of integrability restrictions are independently imposed on parameter estimates for a semi-logarithmic incomplete demand system. A non-nested likelihood ratio test revealed that the first set of parameter restrictions (the less restrictive model) fit the empirical data significantly better than the second restriction set ($P < 0.001$). Likewise, model parameters indicated that data are over-dispersed and the negative binomial II

Table III. Per trip consumer surplus and compensating variation estimates

	Restriction set 1				Restriction set 2			
	Poisson		Negative binomial 2		Poisson		Negative binomial 2	
	Consumer surplus	Compensating variation	Consumer surplus	Compensating variation	Consumer surplus	Compensating variation	Consumer surplus	Compensating variation
Badin Lake	$41.84	$41.83	$27.32	$27.32	$454.55	$456.63	$714.29	$712.43
Brown Mountain	$31.15	$31.15	$29.59	$29.58	$476.19	$475.18	$1000.00	$996.37
Upper Tellico	$101.01	$100.95	$131.58	$131.46	$370.37	$369.76	$588.24	$586.98
Wayehutta	$25.51	$25.51	$37.17	$37.17	$625.00	$623.27	$909.09	$906.09

(NB2) distribution fit the data better than the Poisson. However, because the NB2 is less robust to distribution misspecification, it remains unclear which distribution should provide more accurate welfare estimates.

One of the primary goals of demand system estimations is the derivation of welfare measures for different goods or characteristics. In the analyses reported here, welfare estimates varied dramatically depending on the specification of the parameter restrictions. The implication of this analysis is that researchers should not naively apply parameter restrictions when estimating systems of semi-logarithmic incomplete demand equations, but should test alternative sets of utility theoretic restrictions to determine which set best conforms to the data.

References

Burt, O. and D. Brewer (1971), 'Estimation of Net Social Benefits from Outdoor Recreation', *Econometrica* **52**(1971), 32–41.

Cameron, A. C. and P. Trivedi (1998), *Regression Analysis of Count Data*, New York: Cambridge University Press.

Englin, J. and D. Lambert (1995), 'Measuring Angling Quality in Count Data Models of Recreational Fishing: A Non-nested Test of Three Approaches', *Environmental and Resource Economics* **6**, 389–399.

Englin, J., P. Boxall and D. Watson (1998), 'Modeling Recreation Demand in a Poisson System of Equations: An Analysis of the Impact of International Exchange Rates', *American Journal of Agricultural Economics* **80**, 255–263.

Englin, J. and J. S. Shonkwiler (1995), 'Estimating Social Welfare Using Count Data Models: An Application to Long Run Recreation Demand Under Conditions of Endogenous Stratification and Truncation', *Review of Economics and Statistics* **77**, 104–112.

Epstein, L. (1982), 'Integrability of Incomplete Systems of Demand Functions', *Review of Economic Studies* **49**, 411–425.

Gourieroux, C., A. Monfort and A. Trognon (1984), 'Pseudo Maximum Likelihood Methods: Applications to Poisson Models', *Econometrica* **52**, 701–720.

LaFrance, J. (1990), 'Incomplete Demand Systems and Semilogarithmic Demand Models', *Australian Journal of Agricultural Economics* **34**(2), 118–131.

LaFrance, J. T. and W. M. Hanemann (1984), On the Integration of Some Common Demand Systems. Staff Papers in Economics. Staff Paper 83–10.

LaFrance, J. and W. M. Hanemann (1989), 'The Dual Structure of Incomplete Demand Systems', *American Journal of Agricultural Economics* **71**, 262–274.

Haefen, R. H. (2002), 'A Complete Characterization of the Linear, Log-Linear, and Semi-Log Incomplete Demand System Models', *Journal of Agricultural and Resource Economics* **27**(2), 281–319.

Vuong, Q. H. (1989), 'Likelihood Ratio Tests for Model Selection and Non-Nested Hypotheses', *Econometrica* **57**(2), 307–333.

White, H. (1980), 'A Heteroscedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroscedasticity', *Econometrica* **48**, 817–838.

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