Chapter 16

Latent Class Analysis of Survey Data Dealing with Academic Dishonesty

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1. Introduction

Academic cheating by college and university students has long been a concern of researchers in higher education. Typically, the prevalence of various types of cheating behavior has been assessed by means of anonymous questionnaires distributed or mailed to students. Recent estimates of cheating during college on at least one surveyed behavior range from 49% for marketing students (Tom & Borin, 1988) to 88% for premedical students (Sierles, Hendrickx, & Circle, 1980), with a variety of intermediate estimates. For example, Eve & Bromley (1981) report 63% for 681 undergraduate students at a state university, McCabe (1992) reports 67% for over 6,000 respondents at 31 „highly selective“ colleges, and Collison (1990a) reports 78% for students at a state university. At the post-graduate level, Sierles, Hendrickx, and Circle (1980) indicate that 58% of medical students have cheated and Zastrow (1970) reports that 40% of graduate students have cheated. Estimates of cheating behavior during a single academic year are somewhat lower. Haines et al. (1986) surveyed 380 undergraduate students at a state university and found that about 54% reported cheating on examinations, quizzes, etc. during an academic year. While these estimates of academic cheating appear relatively high, there is evidence that estimates relying on anonymous questionnaires may result in severe underestimation of sensitive behaviors. Scheers and Dayton (1987) found that underreporting for five academic cheating behaviors ranged from 39% to 83% for 378 Education students when responses to an anonymous questionnaire were compared to estimates obtained using a randomized response technique designed to reduce jeopardy when reporting sensitive behavior.

There is some evidence that individuals are consistent in their cheating behavior over a variety of situations. That, is individuals who cheat in one type of situation, such as copying answers on an examination, may be more likely to cheat in other types of situations, such as falsifying original research reports. Ward and Beck (1990) see the distinction as that: „between those who cheat once compared with those who cheat more than once throughout a series of test-taking situations. Clearly this constitutes a fundamentally different measure of persistence than that of mere differences in frequency of cheating during a single test-taking event (p.337).“ Often, the type of behavior classified as cheating is not consistent in content or seriousness across studies. In some studies, respondents are classified as persistent cheaters if they report any cheating activity at all, including such things as failure to footnote sources in written work (McCabe, 1992; Eve & Bromley, 1981), copying homework (Tom & Borin, 1988) or adding unread items to bibliographies (Eve & Bromley, 1981; Zastrow, 1970).

Some studies have attempted to explicitly identify persistent cheaters. McCabe (1992) mailed a 72-item questionnaire to 500 students at each of 31 institutions and received over 6,000 responses. Approximately 19% were classified as persistent cheaters based on having five or more self-reported cheating incidents in their college career. Using the same frequency definition for persistent cheating, but based only on incidents during the preceding six months, LaBeff, et al. (1990) report that about 7% of students are persistent cheaters. Eve and Bromley (1981) administered anonymous questionnaires in classes that enrolled students from all academic divisions and extended over all four college years (e.g., physical education classes) and report that about 22% of the respondents admitted cheating five or more times during their college careers while 10% admitted cheating five or more times during the last semester. Using a definition of „hard-core cheating“ as having cheated in eight or more courses during a college career, Collison (1990b) reports estimates for persistent cheaters that range from 5% for science majors to 50% for economics majors.

In view of the long-standing research interest around the topic of cheating in academic settings, the present study was undertaken to apply rigorous statistical methods to the identification and characterization of persistent cheaters. In particular, the methods of latent class analysis are especially appropriate since there exists no reliable, external criterion (i.e., „gold standard“) by which persistent cheaters may be identified. In addition, there is the issue of underreporting of cheating behavior when traditional anonymous questionnaires are utilized. This study presents results from a survey in which responses were obtained both from an anonymous questionnaire and from a randomized response technique. A major focus of the analysis concerns the issue of the degree to which these two survey approaches yield equivalent latent structures.

2. Design

Sample. The target population comprised upper-division undergraduates (i.e., students attaining at least Junior, or third-year, status) at a large state university in the United States. Since prior studies have reported considerable diversity of results associated with academic discipline, students were sampled from four colleges: College of Business Management (BMGT), College of Behavioral and Social Sciences (BSOS), College of Education (EDUC) and College of Engineering (ENGI). Simple random samples of 400 students were selected from each College. These samples were randomly and evenly divided between an anonymous questionnaire group and a randomized-response group. Appropriate forms of a cheating behavior questionnaire were mailed directly to sampled students. Postage-paid, return-address envelopes were included with the survey forms, but there was no follow-up of non-respondents.

Questionnaire Forms. The questionnaire contained 10 items dealing with cheating and asked for information concerning sex, grade point average (in categories) and academic major. The first six items concerned the specific cheating behaviors of lying to a teacher to avoid taking an examination, lying to a teacher to avoid turning in a term paper on time, turning in a term paper that was purchased or obtained from someone else, obtaining a copy of an examination before taking it, copying answers from another test taker, and knowledge of bribes given to teachers in exchange for grades. The last four items were attitudinal in nature and are not analyzed in the current study.

The direct-question anonymous questionnaire form (DQ) consisted of an explanatory cover letter and a one-page questionnaire on which students were requested to circle YES or NO responses. The randomized-response questionnaire form (RRT) consisted of an explanatory cover letter, an instruction sheet that served as the basis for the randomizing procedure and a two-page questionnaire containing the same 10 cheating items presented to the DQ group. However, each cheating item was paired with a non-sensitive item based on either the student's day of birth, Social Security number or telephone number. The randomizing procedure involved entries selected from a table of pseudo-random digits supplied to the student on the instruction sheet. In particular, the student was asked to arbitrarily enter the table of random digits and to copy 10 consecutive digits into spaces provided at the bottom of the instruction sheet. The number copied into a particular space was utilized for randomizing the response to the corresponding item on the questionnaire. The randomizing proportion was approximately .7 in favor of selecting the sensitive, cheating items.

Analysis. The basic analytical approach was that of latent class analysis for dichotomous manifest variables. For the i^{th} dichotomous item (1 for „yes“ and 0 for „no“), the conditional probability for the response, 1, in the g^{th}, of G, latent class is π_{i1g} and the proportion of cases in the g^{th} latent class is π_g. When applied to responses from a randomized-response survey, the parameters, π_{i1g}, represent conditional probabilities for the response „Yes“, but what is required are estimates of the conditional probabilities for the associated sensitive behaviors. For the unrelated-question RRT method that was used in this study, the expected proportion of „Yes“ responses to an item for which the sensitive behavior has true probability, π_{iA}, in the population is, say, $\lambda_i = p \cdot \pi_{iA} + (1 - p) \cdot \pi_Y$ where p is the randomizing proportion and π_Y is the proportion of cases with the non-sensitive characteristic, Y. Assuming π_Y known and that $\hat{\lambda}_i$ is the sample proportion of „Yes“ responses, the maximum-likelihood estimator for π_{iA} is given by $\hat{\pi}_{iA} = [\lambda_i - (1-p)\,\pi_Y]\,/\,p$ restricted to the [0,1] interval. Since $\hat{\pi}_{iA}$ is a linear function of $\hat{\lambda}_i$, maximum-likelihood estimates associated with $\hat{\lambda}_i$ -values under a latent class model can be transformed to corresponding maximum-likelihood estimates of $\hat{\pi}_{iA}$. Thus, if a latent-class solution based on RRT data results in estimates, $\hat{\pi}_g$ and $\hat{\pi}_{ilg}$ on the λ-scale, the corresponding MLEs associated with the sensitive behaviors are $\hat{\pi}_g$ and $\hat{\pi}_{A_{i\,lg}} = [\hat{\pi}_{i1g} - (1-p)\pi_Y]\,/\,p$.

3. Results

Response Rates. Response rates ranged from a low of 23% for the RRT form for students in BSOS to a high of 50% for the DQ form for students in EDUC (Table 1). Overall, a total of 326 cases, or 41% of those sampled, was returned for the DQ form. Due to missing responses, analysis was based on 319 complete-data cases. Only 242 cases, or 30% of those sampled, were returned for the RRT form and all of these were complete. The DQ form showed higher response rates in three of the four Colleges with the response rates in ENG being essentially the same for the two forms. Since relatively higher levels of preparation in mathematics are typical of students in colleges such as ENG, in future applications it seems reasonable to consider the use of the random-number, randomized-response technique for similar populations.

		Direct Questionnaire					*RRT Questionnaire*				
Item		BMG	BSOS	EDUC	ENGI	Total	BMG	BSOS	EDUC	ENGI	Total
1	Lied re exam	0.11	0.20	0.08	0.08	0.11	0.23	0.34	0.11	0.00	0.15
2	Lied re paper	0.11	0.17	0.16	0.03	0.12	0.08	0.47	0.33	0.22	0.25
3	Purchase TP	0.04	0.03	0.04	0.03	0.03	0.12	0.06	0.00	0.13	0.07
4	Exam copy	0.07	0.04	0.05	0.00	0.04	0.03	0.12	0.00	0.06	0.05
5	Copy answer	0.29	0.25	0.20	0.11	0.21	0.30	0.35	0.37	0.22	0.30
6	Bribe	0.01	0.03	0.02	0.00	0.02	0.00	0.12	0.00	0.03	0.02
	Number	84	71	99	72	326	67	45	57	73	242
	Percent	42%	36%	50%	36%	41%	34%	23%	29%	37%	30%

Table 1: Estimates of cheating behavior

Traditional Estimates. Table 1 provides a summary of the proportion of students responding „Yes“ to each of the six cheating items for both the DQ and RRT forms of the questionnaire. The estimates shown for the RRT form are MLEs transformed to the π_A-scale. Two conclusions are apparent from Table 1. First, as expected, the estimates from the RRT survey are, on average, higher than those from the DQ survey. And, second, two clusters of cheating items seem to exist for both questionnaire forms. That is, Items 1 and 2, concerning lying to avoid examinations and papers, and Item 5, concerning copying answers on an examination, occur with distinctly greater prevalence than Items 3, 4 and 6 that concern purchasing a term paper, seeing a copy of an examination before taking it and having knowledge of bribes given to change grades. In ensuing analyses, these clusters of items are referred as Set 1, representing minor cheating infractions, and Set 2 representing more serious cheating offenses.

Preliminary Latent Class Analyses. In order to assess the general type of latent structure required for the surveys, separate series of analyses were conducted for data from the DQ and RRT questionnaire forms (Table 2). One-class to three-class models were fit separately to the total samples of respondents, as well to each of the four Colleges, for each form. Although the models are nested, comparisons among models are based on Bozdogan (1987) CAIC values in addition to chi-square difference statistics since comparisons among alternate models involved setting parameters to boundary values (Titterington, Smith & Makov, 1985). Overall, for data from the DQ survey, the results suggest a latent class structure based on two classes, although for three of the four Colleges the independence model could not be rejected. The results for data from the RRT survey, on the other hand, suggest models more complex than independence only for BMGT and BSOS based on chi-square differences with CAIC favoring independence models in all cases. Two factors that may account for this somewhat anomalous outcome are the relatively small total sample size (i.e., 242 cases) and the additional unreliability introduced into the data by the randomization procedure. For the two-class solutions, the latent class may be interpreted as representing „persistent cheater“ and „non-cheater“ classes. For the DQ and RRT forms, respectively, the sizes for the persistent cheater latent class are 13% and 7%. This result is consistent with estimates for persistent cheaters in studies cited earlier that reported values ranging from about 7% to 22%.

			Log-Likelihoods for:			*CAIC for*			*Min*	*Chi-Sq Diffs for:*		*Chi-Sq*
Form	Group	N	1-Class	2-Class	3-Class	1-Class	2-Class	3-Class	CAIC	1 vs 2	2 vs 3	Choice

DQ	TOTAL	319	-514.55	-481.55	-470.05	1076.46	1057.81	1082.17	2	66.01	22.99	3
	BMGT	81	-128.41	-114.93	-113.14	294.58	305.38	339.56	1	26.95	3.58	2
	BSOS	70	-137.55	-129.19	-127.61	311.84	331.86	365.44	1	16.71	3.16	2
	EDUC	97	-166.52	-140.98	-135.66	372.06	360.01	388.39	2	51.09	10.64	2
	ENGI	71	-63.78	-57.47	-56.91	164.40	188.62	224.34	1	12.62	1.13	1
RRT	TOTAL	242	-665.63	-659.08	-654.85	1376.68	1409.01	1445.97	1	13.09	8.47	1
	BMGT	67	-179.75	-172.44	-169.08	395.93	417.75	447.46	1	14.61	6.72	2
	BSOS	45	-143.68	-139.96	-130.48	321.01	347.21	361.90	1	7.45	18.95	3
	EDUC	57	-136.26	-133.39	-129.14	307.82	337.38	364.18	1	5.74	8.49	1
	ENGI	73	-186.51	-183.02	-178.89	410.05	440.11	468.88	1	6.97	8.26	1
									Chi-Sq(7, 0.95) = 14.07			

Table 2: Preliminary latent class models fitted to cheating survey

Modeling Latent Class Structure per Questionnaire Form. A series of constrained models based on a two-class structure was conducted separately for data from the DQ questionnaire form. Since constraints result in greater degrees of freedom for chi-square fit tests and affect the CAIC information measures, it was believed that these analyses might provide additional insight into the data that was not available on the basis of the unconstrained models described above. Results for these models, along with summaries for the unconstrained solutions, are presented in Table 3.

Model	*# LC*	*Description*	*# Par*	*DF*	*Chi-Square*	*p-value*	*CAIC**
1	1	Independence	7	57	112.363	0.0000	-273.2528
2	2	Unrestricted CP	14	50	46.353	0.6210	-291.9065
3	2	Constant CP per class	4	60	186.742	0.0000	-219.1694
4	2	Class 1 CP = by sets; Class 2 CP unrestricted	10	54	50.861	0.5960	-314.4593
5	2	Class 1 CP unrestricted; Class 2 CP = by sets	10	54	72.764	0.0450	-292.5563
6	2	CP = by sets in both classes	6	58	86.447	0.0090	-305.9340
7	3	Unrestricted CP	21	43	20.048	0.9990	-270.8552

Table 3: Latent class models for direct questionnaire survey

Notes: CAIC* = Chi-Square -2DF differs from the usual CAIC by a constant.
Class 1 = „Cheaters“; Class 2 = „Non-Cheaters“

Model 3, that imposes the 10 restrictions, $\pi_{i1g} = \pi_{1g} \ \forall \ i$, posits a constant conditional probability for the „Yes“ response both within the persistent cheater and non-cheater latent classes. However, this is the worst-fitting model among those evaluated. Model 4 imposes four restrictions, $\pi_{111} = \pi_{211} = \pi_{511}$ and $\pi_{311} = \pi_{241} = \pi_{61}$, that equate conditional probabilities for the „Yes“ response within sets of items for persistent cheaters, but places no restrictions on the non-cheater class (i.e., the π_{i2} values are unconstrained). This models fits the data acceptably in terms of its chi-square value and has an CAIC*value that is smaller than that associated with the unrestricted two-class model. Model 6 extends the restrictions from Model 4 to both the persistent cheater and non-cheater latent classes. This model provides marginal fit but has the second smallest CAIC*value. Thus, among these two-class models, application of a minimum CAIC*strategy suggests that Model 4 provided a useful basis for interpreting the cheating data.

Modeling Latent Class Structure Across Questionnaire Forms. Ordinarily, model comparisons across groups, such as those represented by the DQ and RRT samples, can be carried out by combining the data sets, imposing suitable restrictions during estimation and using chi-square or information-theory based procedures to evaluate the impact of the restrictions. In the present case, such comparisons can not be done with conventional latent class analysis programs such as MLLSA or LCAG since the item conditional probabilities, π_{ilg} are on different scales for the two groups. Since comparisons between estimates from the two groups were an important factor in this study, an alternate methodology was implemented. The general approach was to calculate MLEs separately for the DQ and RRT samples, to suitably transform the RRT item conditional probability estimates to the π_A-scale, and to make comparisons between groups utilizing approximate chi-square and z-test procedures.

A partial test for the equivalence of parameter estimates from the two samples was carried out by fitting expected frequencies for the 64 distinct response patterns based on transformed parameter estimates for an unconstrained two-class model from the RRT sample to the observed frequencies from the DQ sample. The resulting likelihood ratio chi-square statistic was 178.85 which, with 63 degrees of freedom, suggests that the two samples are not equivalent in latent structure. In addition, item-by-item and latent class size comparisons were made between estimates for unconstrained two-class models for the two samples (Table 4, upper panel). Estimated standard errors and sampling covariances for the two-class models in the DQ and RRT (suitably transformed) samples were derived from a program based on a Newton-Raphson solution (Dayton & Macready, 1977). Latent class proportions for the persistent cheater class were not reliably different for the RRT and DQ samples ($z = .71$, $p = .4777$) and, for the unconstrained model, the combined estimate for persistent cheaters is about 10%. With respect to individual items, the two „lie“ items (Items 1 and 2) for the non-cheater latent class were reliably higher at conventional levels for the RRT sample compared with the DQ sample. In addition, one-tailed tests suggest higher RRT rates for Item 3 (purchase of a term paper) and Item 5 (copying answers on an exam). However, none of the differences for persistent cheaters was statistically significant. This suggests that the RRT survey may provide more realistic estimates of cheating behavior for students who are not persistent cheaters but that estimates for persistent cheaters are comparable across the survey methods. It should be noted that estimates for the RRT samples are much less stable (i.e., have substantially larger standard errors) than those for the DQ sample and that this occurs for two reasons. First, the number of DQ respondents was larger and, second, the RRT procedure introduces additional variability into the estimates as a result of the randomizing procedure.

Finally, an effort was made to determine whether or not the model identified as a reasonable representation of the latent structure for the DQ sample could be replicated for the RRT sample. This constrained two-class model, denoted Model 4 in Table 3, equated conditional probabilities in sets for the persistent cheater latent class but did not constrain these probabilities for the non-cheater latent class (Table 4, lower panel). Since this model cannot be estimated directly from the observed frequencies for the RRT sample, an approximate chi-square test representing a multivariate analog of the Wald test (Bishop, Fienberg & Holland, 1975; p. 353) was utilized based on estimated parameters and their associated sampling covariance matrix both of which had been converted from the λ-metric to the π_A-metric.

Unconstrained Two-Class Solution

	RRT Sample				DQ Sample					
Parameter Estimates	LC 1 „Cheaters“	SE	LC 2 „Non-Cheaters“	SE	LC 1 „Cheaters“	SE	LC 2 „Non-Cheaters“	SE	RRT vs DQ z-Tests „Cheaters“	RRT vs DQ z-Tests „Non-Cheaters“
LC Proportion	0.07	0.07	0.93		0.13	0.05	0.87			-0.71
Item 1	0.51	0.27	0.13	0.04	0.63	0.16	0.03	0.02	-0.38	2.20
Item 2	0.61	0.27	0.22	0.05	0.66	0.16	0.04	0.02	-0.17	3.45
Item 3	0.16	0.16	0.07	0.03	0.17	0.07	0.01	0.01	-0.02	1.80
Item 4	1.00	0.87	0.00	0.06	0.12	0.06	0.03	0.01	1.01	-0.45
Item 5	0.55	0.24	0.28	0.05	0.40	0.10	0.18	0.03	0.56	1.87
Item 6	0.40	0.27	0.00	0.03	0.08	0.05	0.01	0.01	1.17	-0.17

Constrained Two-Class Solution

	RRT Sample		DQ Sample	
Parameter Estimates	LC 1 „Cheaters“	LC 2 „Non-Cheaters“	LC 1 „Cheaters“	LC 2 „Non-Cheaters“
LC Proportion	*	*	0.17	0.83
Item 1	0.56**	0.13	0.50	0.02
Item 2	0.56**	0.22	0.50	0.03
Item 3	0.33**	0.07	0.11	0.01
Item 4	0.33**	0.00	0.11	0.03
Item 5	0.56**	0.28	0.50	0.17
Item 6	0.33**	0.00	0.11	0.00
	*Cannot be estimated; see text * *Weighted average using 1/SE as weights			

Table 4: Comparison of latent class structures for RRT and DQ samples

These estimates can be represented by a 13x1 parameter vector, **b**, and a 13x13 covariance matrix, **C**, respectively. Given an m-row hypothesis matrix, **H**, where each row of **H** represents a restriction on the parameters, an approximate chi-square statistic is defined as $\chi^2 = (\mathbf{H} \cdot \mathbf{b})' (\mathbf{H} \cdot \mathbf{C} \cdot \mathbf{H}')^{-1} (\mathbf{H} \cdot \mathbf{b})$ with m degrees of freedom. For restrictions corresponding to Model 4 in Table 3, the calculated chi-square value was 1.56 which, with 4 degrees of freedom, does not indicate lack of fit for the RRT sample. The same model applied to the non-cheater class, however, results in a calculated chi-square value of 10.71 that, with 4 degrees of freedom, suggests lack of fit. The values shown for item conditional probabilities in the lower panel of Table 4 for persistent cheaters in the RRT sample were calculated as weighted averages of the estimates from the unconstrained two-class model, the weights being reciprocals of standard errors. Unfortunately, there is no straight-forward method for estimating the latent class proportions for the constrained model and there are indicated by '*' in Table 4.

In summary, there is evidence that similar latent structures exist for the DQ and RRT samples although item conditional probabilities are different. More specifically, these latent structures may be summarized as follows. For the unconstrained two-class model, the proportion of persistent cheaters is consistent for the two samples with a combined estimate of approximately 10% of respondents falling this category. Within the persistent cheating class for both samples, there is evidence that the relatively minor behaviors represented by Items 1, 2 and 5 occur equally often and, similarly, that the relatively more serious infractions represented by Items 3, 4 and 6 occur equally often. For the unconstrained model, there is no evidence for differ rates in the persistent cheater class for the items and this result may generalize to the constrained model. However, for the non-cheater class, the RRT sample shows distinctly higher rates for certain of the cheating items.

Cheating Behavior as a Function of GPA. The relationship between the latent structure of academic cheating behaviors and academic performance as measured by grade point average (GPA) was studied by means of a covariate latent class model applied to data from the DQ sample only. The number of cheating behaviors acknowledged on the survey was scored 0, 1, 2, 3-and-more since there were very few students reporting larger numbers of cheating behaviors (i.e., only two students reported 4 or 5 such behaviors). Figure 1 shows the results of fitting a binomial logistic covariate latent class model, as proposed by Dayton and Macready (1988a,b), to the frequencies of behaviors. The values plotted are the proportions estimated in the persistent cheating class at each level of GPA. In addition, the proportion of students reporting any cheating at each level of GPA and the usual logistic regression of number of cheating behaviors on GPA are shown for reference. The covariate model fits the data relatively well yielding a chi-square fit statistic of 10.46 based on 11 degrees of freedom. The binomial rates of cheating behavior for the persistent cheating and non-cheating classes were .43 and .05, respectively. These are somewhat higher than the weighted averages of the conditional probabilities for these classes in the unconstrained two-class model presented above (i.e., the average rates were .25 and .03, respectively). As expected, the likelihood of academic cheating increases inversely with academic performance as measured by GPA. Notably, in the lowest GPA category used in this survey, the estimate of persistent cheaters is in excess of 50%.

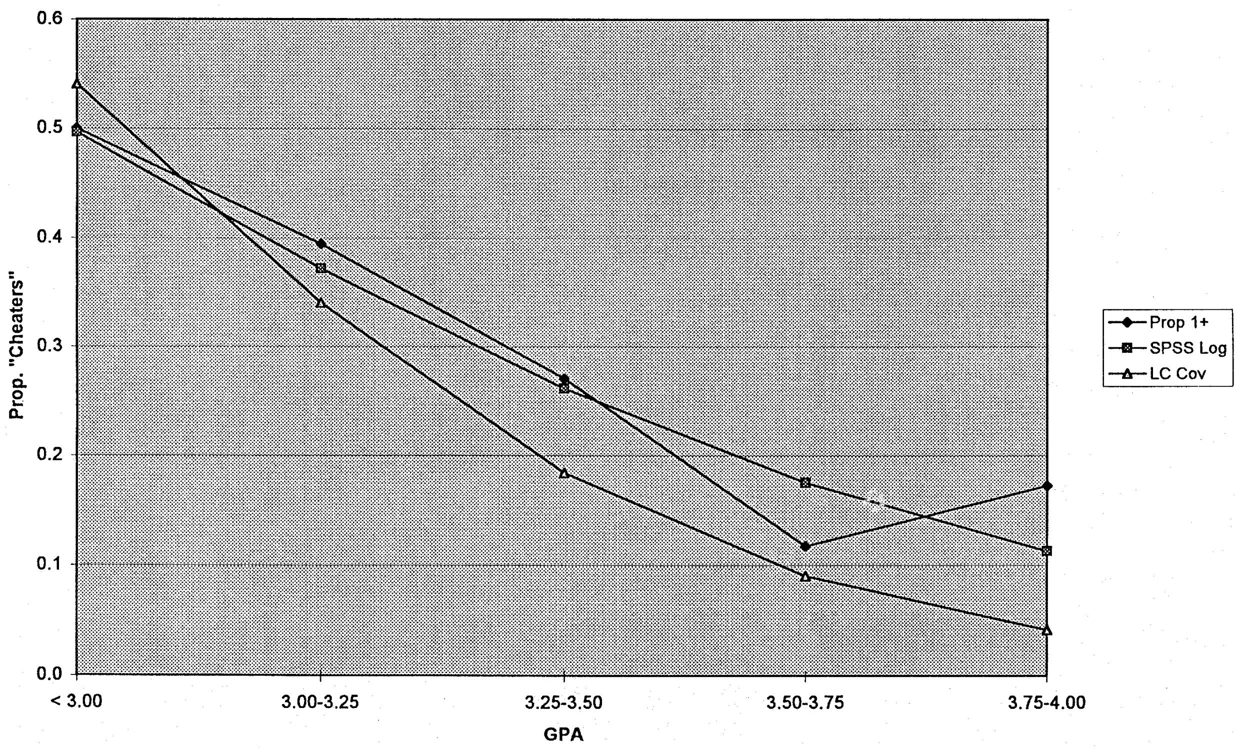

Figure 1: Relation of GPA to cheating

4. Summary

This study focused on the identification and characterization of persistent cheaters in an academic environment. Rather than using arbitrary criteria for identifying persistent cheaters, the present approach utilized the methods of latent class analysis. In addition, results from a sample using a randomized response technique were compared with those from a sample presented with a conventional direct question format.

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