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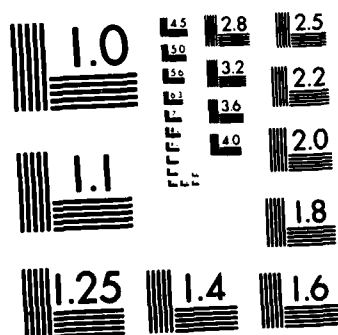
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WEIGHTED DISTRIBUTIONS ARISING OUT OF
METHODS OF ASCERTAINMENT

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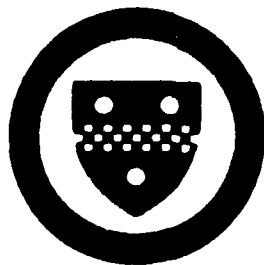
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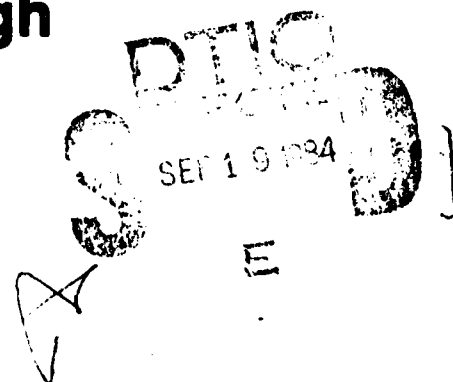
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by

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and

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July 1984

Technical Report No. 84-38

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ABSTRACT

The concept of weighted distributions can be traced to the study of the effects of methods of ascertainment upon the estimation of frequencies by Fisher in 1934, and it was formulated in general terms by the author in a paper presented at the First International Symposium on Classical and Contagious Distributions held in Montreal in 1963. Since then, a number of papers have appeared on the subject. This paper reviews some previous work, points out, through appropriate examples, some situations where weighted distributions arise and discusses the associated methods of statistical analysis.

The importance of specification of the class of underlying probability distributions (or stochastic model) in data analysis based on a detailed knowledge of how data are obtained is emphasized. Failure to take into account the conditions of ascertainment of data can lead to wrong conclusions.



Keywords and Phrases: Damage models, nonresponse, probability sampling, quadrat sampling, size biased sampling, truncation, weighted distributions.

1. IMPORTANCE OF SPECIFICATION

For drawing valid inferences from observed data through statistical methodology it is necessary to identify the proper sample space (all possible outcomes) and specify the class of probability distributions (the model) to which the true distribution of the observations belongs. More precisely, the observed data set x has to be considered as the result of a random experiment, i.e., as a realization of a random variable (rv) X taking values in a space X and subject to a probability distribution P belonging to a specified class \mathcal{P} . Such a knowledge enables us to write down the probability (or probability density) of x for given P , which we write as $\ell(P|x)$. The function $\ell(\cdot|x)$ defined over \mathcal{P} for given x , called the likelihood, together with any apriori information we may have on \mathcal{P} forms the basis of statistical inference. The specification of \mathcal{P} , or the model as it is sometimes called, is thus a datum of the problem of inference. However, not much attention is given to this problem in statistical theory or practice despite the emphasis given to it by the pioneers in statistics like Karl Pearson and R. A. Fisher. Wrong specification may lead to invalid inference, which is sometimes referred to as a third kind of error, the first two being the familiar ones associated with the Neyman-Pearson theory of testing of hypotheses.

It is almost axiomatic to say, although it may need some effort to demonstrate, that inference based on specification \mathcal{P}_1 is possibly more precise than that on \mathcal{P}_2 if $\mathcal{P}_1 \subset \mathcal{P}_2$ provided \mathcal{P}_1 includes the true distribution. It is therefore of considerable value to specify the smallest possible set. (See Althum, 1984; Bishop, Fienberg and Holland, 1975, p. 313). Perhaps past experience can be of help in the choice of such a set. But it should also be possible to start with a wider set and narrow it down by using the observed

data themselves, although the appropriate methodology for this purpose is not fully developed. On the other hand, statisticians seem to be content with studies on robustness, i.e., in determining the widest class P for which a given statistical procedure is valid.

The problem of specification is not a simple one. A detailed knowledge of the procedure actually employed in acquiring data is an essential ingredient in arriving at a proper specification. The situation is more complicated with field observations and nonexperimental data where nature produces events according to a certain model, which are observed and recorded by investigators. There does not always exist a suitable sampling frame necessary for the application of the classical sampling theory. One needs to work with visibility analysis instead. In practice, it is not always possible to observe and record events as they occur. For instance, certain events may not be observable by the method we employ and therefore missed in the record (truncated, censored, and incomplete samples). Or an event may be observable only with a certain probability depending on the nature of the event such as its conspicuousness and the procedure employed to observe it (unequal probability sampling). Or an event may change in a random way by the time or during the process of observation so that what comes on record is a modified event (damage models). Sometimes, events produced under two or more different mechanisms with unspecified relative frequencies get mixed up and brought into the same record (outliers, contaminated samples). In all these cases, the specified class P for the original events (as they occur) may not be appropriate for the events as they are recorded (observed data) unless it is suitably modified.

In a classical paper, Fisher (1934) demonstrated the need for such adjustment in specification depending on the way the data are ascertained. In extending the basic ideas of Fisher, the author (Rao, 1965) introduced the concept of a weighted distribution as a method of adjustment applicable to many situations. In the present paper we discuss, through suitable examples, some procedures for making adjustments in specification based on methods of ascertaining data.

Although I have mentioned only field observations which are collected without the help of a suitable sampling frame, I must emphasize that similar problems of specification arise with data collected in large scale sample surveys and also with data acquired through field and laboratory experiments. Survey practioners are faced with problems of incomplete frame which raise questions of representativeness of a sample for a given population (see Kruskal and Mosteller, 1980 and references therein), nonresponse which raises questions of repeated visits to sampled units, substitution of nonresponding units by others with possibly similar characteristics, and imputation of values (Fienberg and Tanur, 1983; Fienberg and Stasny, 1983; Rubin, 1976, 1980), and nonsampling errors which raise questions about their recognition, detection, measurement and making adjustments in expressing precision of estimates (Mahalanobis 1944; Mosteller, 1978). Similarly in design of experiments, difficulties in random allocation of treatments and choice of controls in field trials, pooling of evidence from different experiments conducted over space and time and missing values (drop outs) introduce additional uncertainties in statistical inference and interpretation of results for practical use or policy purposes (for typical problems and references see Fienberg, Singer and Tanur, 1984; Neyman, 1977).

2. TRUNCATION AND CENSORING

Some events, although they occur, may be unascertainable so that the observed distribution would be truncated to a certain region of the sample space. An example is the frequency of families with both parents heterozygous for albinism but having no albino children. There is no evidence that the parents are heterozygous unless they have an albino child, and the families with such parents and having no albino children get confounded with normal families. The actual frequency of the event 'zero albino children' is, thus, not ascertainable. Adjustment to the probability distribution applicable to observable events in such a case is simple.

In general, if $p(x, \theta)$ is the pdf (probability density function), where θ denotes unknown parameters, and the rv X is truncated to a specified region $T \subset X$, then the pdf of the truncated random variable X^w is

$$p^w(x, \theta) = w(x, T)p(x, \theta) \div u(T, \theta) \quad (2.1)$$

where $w(x, T) = 1$ if $x \in T$ and $= 0$ if $x \notin T$ and $u(T, \theta) = E[w(X, T)]$. If x_1, \dots, x_n are independent observations subject to truncation, then the likelihood is

$$p(x_1, \theta) \dots p(x_n, \theta) \div [u(T, \theta)]^n. \quad (2.2)$$

In some cases we may have independent observations x_1, \dots, x_n arising from a truncated distribution in addition to a number m (and not the actual values) of observations falling outside T . Then the likelihood is

$$\frac{(n + m)!}{m!} p(x_1, \theta) \dots p(x_n, \theta) [1 - u(T, \theta)]^m. \quad (2.3)$$

A more complicated case is the following.

Suppose that we have a measuring device which records the time at which a bulb fails. If we are experimenting with n bulbs in a life testing problem using a measuring device which may itself fail at a random time, then the observations would be of the type

$$x_1, \dots, x_{n_1}, n_2, n_3 \quad (2.4)$$

where x_1, \dots, x_{n_1} are the life times of n_1 bulbs recorded before an unknown time point T at which the measuring device failed, n_2 is the number of bulbs that failed between T and T_0 , the known time at which the experiment was terminated, and n_3 is the number of bulbs still burning after T_0 . Let

$$w_1(T, \theta) = P(x \leq T), \quad w_2(T, \theta) = P(T < x \leq T_0), \quad w_3(T, \theta) = 1 - w_1(T, \theta) - w_2(T, \theta).$$

Then the likelihood based on the data (2.4) is

$$\frac{n!}{n_2!n_3!} p(x_1, \theta) \dots p(x_{n_1}, \theta) [w_2(T, \theta)]^{n_2} [w_3(T, \theta)]^{n_3} \quad (2.5)$$

where T is unknown besides the basic parameters θ . Inference on T and θ based on (2.5) does not seem to have been fully worked out but could be developed on standard lines.

The expressions (2.2), (2.3), and (2.5) are simple examples of weighted distributions, whose general definition is given in Section 3.

3. WEIGHTED DISTRIBUTIONS

In Section 2, we have considered situations where certain events are unobservable. But a more general case is where an event that occurs has a certain probability of being recorded (or included in the sample). Let X be a rv with $p(x, \theta)$ as the pdf, and suppose that when $X = x$ occurs, the probability of recording it is $w(x, \alpha)$ depending on the observed value x and possibly also

on an unknown parameter α . Then the pdf of the resulting rv X^W is

$$p^W(x, \theta, \alpha) = w(x, \alpha) p(x, \theta) \div E[w(X, \alpha)]. \quad (3.1)$$

Although in deriving (3.1), we chose $w(x, \alpha)$ such that $0 \leq w(x, \alpha) \leq 1$, we can define (3.1) for any arbitrary non-negative weight function $w(x, \alpha)$ for which $E[w(X, \alpha)]$ exists. The distribution (3.1) obtained by using any non-negative weight function $w(x, \alpha)$ is called (see Rao, 1965) a weighted version of $p(x, \theta)$. In particular, the weighted distribution

$$p^W(x, \theta) = |x| p(x, \theta) \div E[|x|] \quad (3.2)$$

where $|x|$ is the norm or some measure of size of x is called the size biased distribution. When x is univariate and non-negative, the weighted distribution

$$p^W(x, \theta) = x p(x, \theta) \div E(X) \quad (3.3)$$

is called length (size) biased distribution. For example, if X has the logarithmic series distribution

$$\frac{\alpha^r}{-r \log(1 - \alpha)}, \quad r = 1, 2, \dots \quad (3.4)$$

then the distribution of the size biased variable is

$$\alpha^{r-1}(1 - \alpha), \quad r = 1, 2, \dots \quad (3.5)$$

which shows that $X^W - 1$ has a geometric distribution. A truncated geometric distribution is sometimes found to provide a good fit to an observed distribution of family size (Feller, 1966). But, if the information on family size has been ascertained from school children, then the observations would have a size biased distribution. In such a case a good fit of the geometric distribution to the observed family sizes would indicate that the underlying distribution of family size is, in fact, a logarithmic series.

Table 1 gives a list of some basic distributions and their size biased forms. it is seen that the size biased form belongs to the same family as the original distribution in all cases except the logarithmic series (see Rao, 1965; Patil and Ord, 1975; Janardhan and Rao, 1983 for characterizations and examples of size biased distributions).

Table 1. Certain Basic Distributions and their Size-Biased Forms

Random Variable(rv)	pf(pdf)	Size-biased rv
Binomial, $B(n, p)$	$\binom{n}{x} p^x (1-p)^{n-x}$	$1 + B(n-1, p)$
Negative Binomial, $NB(k, p)$	$\binom{k+x-1}{x} q^x p^k$	$1 + NB(k+1, p)$
Poisson, $Po(\lambda)$	$e^{-\lambda} \lambda^x / x!$	$1 + Po(\lambda)$
Logarithmic series, $L(\alpha)$	$\{-\log(1-\alpha)\}^{-1} \alpha^x / x$	$1 + NB(1, \alpha)$
Hypergeometric, $H(n, M, N)$	$\binom{n}{x} M^x (N-M)^{n-x} / N^n$	$1 + H(n-1, M-1, N-1)$
Bimomial beta, $BB(n, \alpha, \gamma)$	$\binom{n}{x} \beta(\alpha+x, \gamma+n-x) / \beta(\alpha, \gamma)$	$1 + BB(n-1, \alpha, \gamma)$
Negative binomial beta, $NBB(k, \alpha, \gamma)$	$\binom{k+x-1}{x} \beta(\alpha+x, \gamma+k) / \beta(\alpha, \gamma)$	$1 + NBB(k+1, \alpha, \gamma)$
Gamma, $G(\alpha, k)$	$\alpha^k x^{k-1} e^{-\alpha x} / \Gamma(k)$	$G(\alpha, k+1)$
Beta first kind, $B_1(\delta, \gamma)$	$x^{\delta-1} (1-x)^{\gamma-1} / \beta(\delta, \gamma)$	$B_1(\delta+1, \gamma)$
Beta second kind, $B_2(\delta, \gamma)$	$x^{\delta-1} (1+x)^{-\gamma} / \beta(\delta, \gamma-\delta)$	$B_2(\delta+1, \gamma-\delta-1)$
Pearson type V, $Pe(k)$	$x^{-k-1} \exp(-x^{-1}) / \Gamma(k)$	$Pe(k-1)$
Pareto, $Pa(\alpha, \gamma)$	$\gamma \alpha x^{-(\gamma+1)}, x \geq \alpha$	$Pa(\alpha, \gamma-1)$
Lognormal, $LN(\mu, \sigma^2)$	$(2\pi\sigma^2)^{-\frac{1}{2}} x^{-1} \exp - \left(\frac{\log x - \mu}{\sigma \sqrt{2}} \right)^2$	$LN(\mu + \sigma^2, \sigma^2)$

An example of weighted distributions arises in sample surveys when unequal probability sampling or pps (probability proportional to size) sampling is employed. A general version of the sampling scheme involves two rv's X and Y with pdf, $p(x,y,\theta)$ and a weight function $w(y)$ which is a function of y only giving the weighted pdf

$$p^W(x,y,\theta) = w(y)p(x,y,\theta) \div E[w(Y)]. \quad (3.6)$$

In sample surveys we obtain observations on (X^W, Y^W) from the pdf (3.6) and draw inference on the unknown parameter θ .

It is of interest to note that the marginal pdf of X^W is

$$p^W(x,\theta) = w(x,\theta)p(x,\theta) \div E[w(X,\theta)] \quad (3.7)$$

which is a weighted version of $p(x,\theta)$ with the weight function

$$w(x,\theta) = \int p(y|x)w(y)dy \quad (3.8)$$

which may involve the unknown parameter θ .

There is an extensive literature on weighted distributions since the concept was formalized in Rao (1965), which is reviewed with a large number of references in a paper by Patil (1984) with special reference to ecological work. Reference may also be made to two earlier contributions by Patil and Rao (1977, 1978) and Patil and Ord (1975) which contain reviews of previous work and details of some new results.

In the next sections, we consider several examples where weighted distributions are used in the analysis of data.

4. DIFFERENTIAL PRESERVATION OF SKULLS

The following problem arose in the analysis of cranial measurements. A sample of skulls dug out from ancient graves in Jebel Moya, Africa, consisted of some well-preserved skulls and the rest in a broken condition (see Mukherji, Trevor and Rao, 1955). On each well-preserved skull it was possible to take four measurements, C (capacity), L (length), B (breadth), and H (height), while on a broken skull only a subset of L, B, and H and not C could be measured. The observed data, thus, consisted of samples from a four variate population with several observations missing. There were some sets with all the four measurements C, L, B, H, and some with 1 or 2 or 3 of the measurements L, B, and H only. The problem was to estimate the mean values of C, L, B, and H in the original population of skulls from the recovered fragmentary samples. In a number of papers which appeared in the early issues of Biometrika, it was the practice to estimate the unknown population mean value of any characteristic, say C, by taking the mean of all the available measurements on C. An alternative to this, which is often recommended, is to compute maximum likelihood estimates of the unknown mean values, variances, and covariances by writing down the likelihood function based on all the available data assuming a four variate normal distribution for C, L, B and H and using the derived marginal distribution for an incomplete set of measurements. This is based on the assumption that each skull admitting all the four measurements or any subset of the four can be considered as a random sample from the original population of skulls. Is this assumption valid?

It is a common knowledge that a certain proportion of the original skulls gets broken depending on the length of time and depth at which they lay buried. Let $w(c)$ be the probability that a skull of capacity c is not broken

and $p(c, \theta)$ be the pdf of C in the original population. Then the pdf of C measured on well-preserved skulls is

$$w(c)p(c, \theta) \div E[w(C)]. \quad (4.1)$$

If $w(c)$ depends on c , then the observed measurements on C cannot be considered as a random sample of C from the original population. Further, if $w(c)$ is a decreasing function of c , then there will be a larger representation of small skulls among the unbroken skulls, and therefore the mean of the available measurements on C will be an underestimate of the mean capacity of the original population.

Is there any evidence that $w(c)$ depends on c ? To answer this question, the regression of C on L , B , and H (in terms of logarithms) was estimated from the data sets where all the four measurements were available and used to predict the mean capacity of broken skulls by substituting the observed averages \bar{L} , \bar{B} , and \bar{H} of broken skulls in the regression equation. At least in two series of cranial measurements, (see Rao and Shaw, 1948; Rao, 1973, p. 280) it was found that the average measured capacity of unbroken skulls was smaller than the estimated average capacity of broken skulls. This provided some evidence about the differential preservation of skulls with smaller skulls having a higher chance of remaining unbroken.

This finding invalidates the assumption that skulls providing all the four measurements is a random sample from the original population of skulls. The pdf associated with these measurements is more appropriately (4.1) which is a weighted version of the original pdf with an unknown weight function. Presumably, the pdf associated with observations on any subset of L , B , and H will again be a weighted pdf with a weight function depending on the degree of

damage to a skull. The expression for the correct likelihood will then depend on the original pdf and the probabilities of different degrees of damage as assessed by subsets of measurements that can be taken on a skull, which are likely to be unknown. Is there a reasonable solution to the problem of estimation of mean values in a situation like the above?

There are several possibilities of which the following procedure for estimating the mean of C appears to be a natural one. We use the complete sets of measurements, C, L, B and H, on unbroken skulls to compute the regressions of C on different subsets of L, B and H. Using the appropriate regression function, we estimate (predict) the missing value of C for each broken skull. Then an average is taken of all the measured and estimated values of C. Such an average is likely to be a valid estimate of the mean of C. The estimation is based on the assumption that the complete sets of measurements (C, L, B, H) can provide valid estimates of relationships like the regression functions of C on L, B, H and its subsets, although they are biased samples from the original population. Similar methods can be used to estimate the mean values of L, B and H.

Paleontologists compare the characteristics of fossils of long bones and cranial material discovered in different parts of the world to trace the evolutionary history of hominids. Such studies based on physical measurements may be misleading as the discovered fossils may not be representative samples from the original populations due to differential preservation of skeletal material. It is gratifying to note that attempts are being made to compare the fossils in terms of some basic chemical measurements which are not likely to be subject to the phenomenon of differential preservation.

5. ENQUIRY THROUGH AN OFFSPRING

In genetic and socio-psychological studies it is the common practice to locate an abnormal individual and through him or her collect information on the status of brothers and sisters, parents, uncles, and aunts. From such data estimates are made of the incidence of abnormality in families by sex and parity of birth. A family is the basic unit whose characteristics may have a specified distribution. But our method of ascertainment gives unequal probabilities to families depending on the mechanism inherent in the selection of an abnormal family member. Thus, the distribution applicable to observed data on families is a weighted version of the distribution specified for the families. We consider some examples, discuss the nature of the problems involved in each case, and suggest possible solutions.

5.1 TOO MANY MALES?

During the last few years, while lecturing to students and teachers in different parts of the world, I collected data on the numbers of brothers and sisters in the family of each individual in the audience. The results are summarized in Tables 2, 3, and 4. The data from the male respondents given in Tables 2 and 4 show that the ratio of B , the total number of brothers, including the respondents, to $B + S$, the total number of brothers and sisters is much larger than half in each case indicating a preponderance of male children in the families of male members of the audience.

Rao (1977) showed that the appropriate model for the distribution of brothers and sisters of male respondents is size biased binomial so that the probability of r brothers and $(n - r)$ sisters in a family of size n is

$$r \binom{n}{r} \pi^r (1 - \pi)^{n-r} \div E(r) = \binom{n-1}{r-1} \pi^{r-1} (1 - \pi)^{n-r} \quad (5.1.1)$$

where π is the probability of a male child. Under this hypothesis we find that

$$E \left(\frac{B - k}{B + S - k} \right) = \pi \quad (5.1.2)$$

where k is the number of male respondents, so that $(B - k)/(B + S - k)$ is an estimate of π , and

$$\frac{[B - k - (B + S - k)\pi]^2}{(B + S - k)\pi(1 - \pi)} \quad (5.1.3)$$

has an asymptotic chi-square distribution on 1 degree of freedom. Similar results hold for the data from female respondents in Table 3. It is seen from the chi-square values in Tables 2 and 3 that the data collected from the students are consistent with the hypothesis of size biased binomial with $\pi = 1/2$.

The situation is somewhat different in Table 4 relating to data from the professors. The estimated π is more than half in each case and the chi-square values are high. This implies that the weight function appropriate for these data is of a higher order than r , the number of brothers. A possible sociological explanation for this is that a person coming from a family with a larger number of brothers tends to acquire better academic qualifications to

compete for jobs!

The following example on observed sex ratio is illuminating. In a survey of fertility and mortality, Dandekar and Dandekar (1953) gave the distribution of brothers (excluding the informant) and sisters, and sons and daughters as reported by 1115 'male heads,' contacted through households chosen with equal probability for each household. It may be observed that in a survey of this type, a family with r brothers gets a chance nearly proportional to r , and the conditions for a weighted binomial with $w(r) = r$ holds for the number of brothers in a family. Yet we find from Table 5 that the total number of brothers 1325 (excluding the informants) is far in excess of the number of sisters, 1014 giving

$$\chi^2 = \frac{(1325 - 1014)^2}{1325 + 1014} = 41.35$$

which is very high on 1 degree of freedom. Is the theory of size biased binomial wrong?

But it is clear from Table 5 that the disproportionate sex ratio is confined to the age groups above 15-19 years and the same phenomenon seems to occur in the case of sons and daughters. There is perhaps an underreporting of sisters and daughters who are married off due to a superstitious custom of not including them as members of the household. Underreporting of female members is a persistent feature of data on fertility and mortality collected in developing countries.

Table 2. Data on male respondents (students)

(k = number of students, B = total number of brothers including the respondent, S = total number of sisters).

Place and year	k	B	S	$\frac{B}{B+S}$	$\frac{B-k}{B+S-k}$ [*]	χ^2
Bangalore (India, 75)	55	180	127	.586	.496	.02
Delhi (India, 75)	29	92	66	.582	.490	.07
Calcutta (India, 63)	104	414	312	.570	.498	.04
Waltair (India, 69)	39	123	88	.583	.491	.09
Ahmedabad (India, 75)	29	84	49	.632	.523	.35
Tirupati (India, 75)	592	1902	1274	.599	.484	.50
Poona (India, 75)	47	125	65	.658	.545	1.18
Hyderabad (India, 74)	25	72	53	.576	.470	.36
Tehran (Iran, 75)	21	65	40	.619	.500	.19
Isphahan (Iran, 75)	11	45	32	.584	.515	.06
Tokyo (Japan, 75)	50	90	34	.725	.540	.49
Lima (Peru, 82)	38	132	87	.603	.519	.27
Shanghai (China, 82)	74	193	132	.594	.474	.67
Columbus (USA, 75)	29	65	52	.556	.409	2.91
College St. (USA, 76)	63	152	90	.628	.497	.01
Total	1206	3734	2501	.600	.503	0.14

^{*}Estimate of π under size biased binomial distribution

Table 3. Data on female respondents (students)

Place and year	k	B	S	$\frac{S}{B+S}$	$\frac{S-k}{B+S-k}$	χ^2
Lima (Peru, 82)	16	37	48	.565	.464	.36
Los Banos (Philippines, 83)	44	101	139	.579	.485	.18
Manila (Philippines, 83)	84	197	281	.588	.500	.00
Bilbao (Spain, 83)	14	19	35	.576	.525	.10
Total	158	354	503	.587	.493	.11

Table 4. Data on male respondents (professors)

Place and year	k	B	S	$\frac{B}{B+S}$	$\frac{B-k}{B+S-k}$	χ^2
State College (USA, 75)	28	80	37	.690	.584	2.53
Warsaw (Poland, 75)	18	41	21	.660	.525	2.52
Poznan (Poland, 75)	24	50	17	.746	.567	1.88
Pittsburgh (USA, 81)	69	169	77	.687	.565	2.99
Tirupati (India, 76)	50	172	132	.566	.480	.39
Maracaibo (Venezuela, 82)	24	95	56	.629	.559	1.77
Richmond (USA, 81)	26	57	29	.663	.517	.03
Total	239	664	369	.642	.535	3.95

Table 5. Distribution by age of brothers, sisters, sons and daughters
Dandekar and Dandekar (1953)

age-group	brothers	sisters	sons	daughters
0- 4	5	10	357	348
5- 9	27	31	330	354
10-14	63	62	305	226
15-19	87	85	208	190
.....				
20-24	155	100	167	130
25-29	181	130	85	63
30-34	156	130	29	33
35-40	179	123	18	16
40-44	146	105	13	5
rest	336	228	21	10
.....				
total	1325	1014	1533	1375

5.2 ALBINISM

We introduce a general model which would be useful in genetic studies.

Let π_1 and π_2 be the probabilities that a male child and a female child being an albino respectively. Then the probability that a family of n children has r_1 males of whom t_1 are albinos and r_2 females of whom t_2 are

albinos is

$$p(r_1, t_1; r_2, t_2) = \binom{n}{r_1} \left(\frac{1}{2}\right)^n \binom{r_1}{t_1} \pi_1^{t_1} \phi_1^{r_1-t_1} \binom{r_2}{t_2} \pi_2^{t_2} \phi_2^{r_2-t_2} \quad (5.2.1)$$

where $\phi_1 = 1 - \pi_1$ and $\phi_2 = 1 - \pi_2$, and the probability of a child being a male or a female is taken as half.

There are a number of ways in which we can introduce probabilities of selection of affected families. We consider some models which are extensions of those suggested by Fisher (1934) and Haldane (1938).

Introducing α and $\beta = 1 - \alpha$ as relative probabilities of observing a male or a female albino, we may consider a mixture of two size biased distributions.

$$p^W(r_1, t_1; r_2, t_2) = \left(\frac{2\alpha t_1}{n\pi_1} + \frac{2\beta t_2}{n\pi_2} \right) p(r_1, t_1; r_2, t_2) \quad (5.2.2)$$

as the appropriate distribution of the observed vector (r_1, t_1, r_2, t_2) . If we have data on (r_1, t_1, r_2, t_2) from N ascertained families, we can write down the likelihood using the expression (5.2.2) and estimate the parameters α , π_1 and π_2 . Alternatively, we can use the method of moments, using the statistics Σt_1 , Σt_2 , and Σr_1 to estimate the unknown parameters.

If $\pi_1 = \pi_2 = \pi$, the expression (5.2.2) reduces to

$$\frac{2}{n\pi} (\alpha t_1 + \beta t_2) p(r_1, t_1; r_2, t_2) \quad (5.2.3)$$

and the estimates of α and π can be obtained from the equations

$$\begin{aligned}\bar{t}_1 &= \alpha + \frac{\pi}{2k} \Sigma(n_i - 1) \\ \bar{t}_2 &= \beta + \frac{\pi}{2k} \Sigma(n_i - 1)\end{aligned}\quad (5.2.4)$$

where k is the number of families, n_i is the number of children in the i -th family and \bar{t}_1 and \bar{t}_2 are average numbers of male and female albino children in a family.

Another model is as follows. Let ρ_1 and ρ_2 be the probabilities of observing a male and a female albino respectively. Then the probability that a family with n children having t_1 male albinos and $r_1 - t_1$ normal males, and t_2 female albinos and $r_2 - t_2$ normal females, is investigated s_1 times by observing a male albino and s_2 times by observing a female albino is

$$\binom{t_1}{s_1} \rho_1^{s_1} (1 - \rho_1)^{t_1 - s_1} \binom{t_2}{s_2} \rho_2^{s_2} (1 - \rho_2)^{t_2 - s_2} p(r_1, t_1; r_2, t_2). \quad (5.2.5)$$

Since a family is not investigated unless at least one of t_1 and t_2 is different from zero, the effective distribution for the observed data is (5.2.5) normalized by the dividing factor $[1 - (1 - \rho)^n]$ where $\rho = (\rho_1 \pi_1 + \rho_2 \pi_2)/2$. The method of estimation of ρ_1 , ρ_2 , π_1 and π_2 when we have the additional information on the number of times each family is investigated is discussed in detail in Rao (1965).

In case a family is investigated only once although more than one abnormal child in the family is observed the appropriate distribution is

$$[1 - (1 - \rho_1)^{t_1} (1 - \rho_2)^{t_2}] p(r_1, t_1; r_2, t_2) \div [1 - (1 - \rho)^n] \quad (5.2.6)$$

where $\rho = (\pi_1 \rho_1 + \pi_2 \rho_2)/2$. If $\rho_1 = \rho_2 = \rho$ and $\pi_1 = \pi_2 = \pi$, then the

expression (5.2.6) reduces to

$$\frac{[1 - (1 - \rho)^{t_1+t_2}]}{1 - (1 - \pi\rho)^n} \frac{n!}{t_1!(r_1 - t_1)!t_2!(r_2 - t_2)!} \left(\frac{\pi}{2}\right)^{t_1+t_2} \left(\frac{\phi}{2}\right)^{n-t_1-t_2}. \quad (5.2.7)$$

If sex is ignored then (5.2.7) becomes

$$\frac{1 - (1 - \rho)^t}{1 - (1 - \pi\rho)^n} \frac{n!}{t!(n - t)!} \pi^t \phi^{n-t} \quad (5.2.8)$$

where $t = t_1 + t_2$, which is the expression used by Haldane (1938).

We have considered three different models (5.2.2), (5.2.5) and (5.2.6) for the probability of selection of a family. In the case where we have information only on the number r of abnormal children in a family of size n without any sex distinction we may consider a weighted binomial distribution

$$w(r) \binom{n}{r} \pi^r \phi^{n-r} \div E[w(r)] \quad (5.2.9)$$

with three possible alternatives for $w(r)$

$$w(r) = r \quad (5.2.10)$$

$$= r^\alpha, (\alpha \text{ unknown}) \quad (5.2.11)$$

$$= 1 - (1 - \rho)^n, (\rho \text{ unknown}). \quad (5.2.12)$$

The maximum likelihood method of estimating α and π under the model (5.2.9, 5.2.11) is discussed in Rao (1965), and of ρ and π under the model (5.2.9, 5.2.12) in Haldane (1938). To demonstrate the relevance of the weight function (5.2.11), we compare in Table 6 the observed data on frequencies of albino children in families of different sizes with the expected values under the two different weight functions $w(r) = r$ and $w(r) = r^{1/2}$ choosing $\pi = 1/4$. It is seen that the weight function $w(r) = r^{1/2}$ provides a better fit.

Table 6. Observed and expected frequencies of albino children for each family size (n)

(expected (1) for $w_r = r$ and expected (2) for $w_r = r^{1/2}$)

no. of albinos	n = 2		n = 3		n = 4		n = 5	
	obs- erved	expected	obs- erved	expected	obs- erved	expected	obs- erved	expected
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
1	31	30.00 32.37	37	30.93 35.81	22	21.10 26.07	25	19.00 24.93
2	9	10.00 7.63	15	20.63 16.88	21	21.09 18.43	23	25.31 23.50
3			3	3.44 2.30	7	7.03 5.02	10	12.65 9.59
4					0	0.78 0.48	1	2.81 1.85
5							1	0.23 0.13

no. of albinos	n = 6		n = 7		total	
	obs- erved	expected	obs- erved	expected	obs- erved	expected
	(1)	(2)	(1)	(2)	(1)	(2)
1	18	12.58 17.46	16	8.21 11.98	149	121.82 148.62
2	13	20.96 20.58	10	16.37 16.94	96	114.36 103.98
3	18	13.98 11.20	14	13.64 11.53	47	50.74 39.64
4	3	4.66 3.23	5	6.06 4.43	9	14.31 10.00
5	0	0.77 0.48	1	1.51 0.99	1	2.51 1.61
6	1	0.05 0.03	0	0.20 0.12	1	0.25 0.15
			0	0.01 0.01	0	0.01 0.01

For a general discussion of the type of problems discussed in this section, and a few other models for selection probabilities, the reader is referred to Stene (1981) and other references mentioned in that paper. For estimation of α and π in the model (5.2.9, 5.2.11), reference may be made to Rao (1965).

5.3 ALCOHOLISM, FAMILY SIZE AND BIRTH ORDER

Smart (1963, 1969) and Sprott (1969) examined a number of hypotheses on the incidence of alcoholism in Canadian families using the data on family size and birth order of 242 alcoholics admitted to three alcoholism clinics in Ontario. The method of sampling is thus of the type discussed in Sections 5.1 and 5.2.

One of the hypotheses tested was that "larger families contain larger number of alcoholics than expected." The null hypothesis was interpreted to imply that the observations on family size as ascertained arise from the weighed distribution

$$np(n) \div E(n), n = 1, 2, \dots \quad (5.3.1)$$

where $p(n)$, $n = 1, 2, \dots$ is the distribution of family size in the general population, including families with no alcoholics. It may be noted that the distribution (5.3.1) would be applicable if we had observed an individual (alcoholic or not) at random from the general population and ascertained the size of the family to which he or she belonged. It needs some argument to show that the same distribution holds for family size ascertained by observing the alcoholic individuals only. The following justification of (5.3.1) makes use of an interpretation of the null hypothesis that is being tested.

Let π be the probability of an individual becoming an alcoholic and suppose that the probability that a member of a family becomes an alcoholic is independent of whether another member is alcoholic or not. Further let $p(n)$, $n = 1, 2, \dots$, be the probability distribution of family size (whether a family has an alcoholic or not) in the general population. Then the probability that a family is of size n and has r alcoholics is

$$p(n) \binom{n}{r} \pi^r \phi^{n-r}, \quad r = 0, \dots, n; \quad n = 1, 2, \dots \quad (5.3.2)$$

From (5.3.2), it follows that the distribution of family size in the general population given that a family has at least one alcoholic is

$$(1 - \phi^n)p(n) \div 1 - E(\phi^n), \quad n = 1, 2, \dots \quad (5.3.3)$$

If we had chosen households at random and recorded the family sizes in households containing at least one alcoholic, then the null hypothesis on the excess of alcoholics in larger families could be tested by comparing the observed frequencies with the expected frequencies under the model (5.3.3). However, under the sampling scheme adopted, the weighted distribution of (n, r)

$$p^w(n, r) = rp(n) \binom{n}{r} \pi^r \phi^{n-r} \div E(n) \quad (5.3.4)$$

is more appropriate. If we had information both on the family size (n) as well as on the number of alcoholics (r) in the family, we could have compared the observed joint frequencies of (n, r) with those expected under the model (5.3.4).

From (5.3.4), the marginal distribution of n alone is

$$np(n) \div E(n), \quad n = 1, 2, \dots \quad (5.3.5)$$

which is used by Smart and Sprott as a model for the observed frequencies of family sizes. It is shown in (5.3.3) that in the general population, the distribution of family size in families with at least one alcoholic is

$$(1 - \phi^n)p(n) \div 1 - E(\phi^n)$$

which reduces to (5.3.5) if ϕ is close to unity. Or in other words, if the probability of an individual becoming an alcoholic is small, then the distribution of family size as ascertained is close to the distribution of family size in families with at least one alcoholic in the general population. This is not true if ϕ is not close to unity.

Smart and Sprott found that the distribution (5.3.5) did not fit the observed frequencies, which had heavier tails supporting the hypothesis under test.

An alternative to (5.3.4) is obtained by assuming that each alcoholic has a chance θ of being admitted to a clinic independently of other alcoholic family members. In such a case, the probability that a family of size n has r alcoholics and a member has been admitted to a clinic is

$$p(n) \binom{n}{r} \pi^r \phi^{n-r} (1 - (1 - \theta)^r). \quad (5.3.6)$$

The marginal distribution of n with the normalizing factor is then

$$p(n)(1 - (1 - \pi\theta)^n) \div E(1 - (1 - \pi\theta)^n) \quad (5.3.7)$$

$$n = 1, 2, \dots$$

The distribution (5.3.7) involves one unknown parameter $\pi\theta$ which needs to be estimated in fitting to the observed frequencies of family sizes. Some examples of distributions of the type (5.3.7) have been considered by Barraï, Mi, Morton and Yasuda (1965). The distribution (5.3.7) is close to (5.3.5) if $\pi\theta$ is small.

We may also consider a more complicated model by assuming different probabilities π_1 and π_2 for males and females becoming alcoholic and also different probabilities θ_1 and θ_2 for male and female alcoholics being referred to a clinic. In such a case, the probability of inclusion of a family of size n with r_1 males and s_1 male alcoholics, r_2 females and s_2 female alcoholics is

$$p(n) \binom{n}{r} \left(\frac{1}{2}\right)^n \binom{r_1}{s_1} \pi_1^{s_1} \phi_1^{r_1-s_1} \binom{r_2}{s_2} \pi_2^{s_2} \phi_2^{r_2-s_2} (1 - (1 - \theta_1)^{s_1} (1 - \theta_2)^{s_2}) \quad (5.3.8)$$

which gives the marginal distribution of n as

$$p(n)(1 - 2^{-n}(2 - \pi_1\theta_1 - \pi_2\theta_2)^n) \div E(1 - 2^{-n}(2 - \pi_1\theta_1 - \pi_2\theta_2)^n) \quad (5.3.9)$$

which again involves one unknown parameter $(\pi_1\theta_1 + \pi_2\theta_2)/2$. The marginal distribution of r_1 and r_2 obtained from (5.3.8) is

$$p(n) \binom{n}{r_1} \left(\frac{1}{2}\right)^n (1 - (1 - \pi_1\theta_1)^{r_1} (1 - \pi_2\theta_2)^{r_2}) \div E(1 - 2^{-n}(2 - \pi_1\theta_1 - \pi_2\theta_2)^n) \quad (5.3.10)$$

where $n = (r_1 + r_2)$. If $\pi_1\theta_1$ and $\pi_2\theta_2$ are small, then (5.3.10) becomes

$$p(n) \binom{n}{r_1} \left(\frac{1}{2} \right)^n (r_1 \pi_1^{\theta_1} + r_2 \pi_2^{\theta_2}) \div \frac{\pi_1^{\theta_1} + \pi_2^{\theta_2}}{2} E(n). \quad (5.3.11)$$

If we had the joint frequencies of males and females in the observed families of alcoholics, we could have fitted distributions of the type (5.3.10) and (5.3.11) to test the null hypothesis of larger number of alcoholics in larger families.

It is seen from (5.3.10) and (5.3.11) that the distribution of (r_1, r_2) will not be symmetric unless $\pi_1^{\theta_1} = \pi_2^{\theta_2}$. This may result in excess of males or females in observed families. Such an effect (with an excess of males) can be seen in similar data studied by Freire-Mala and Chakraborty (1975) and Rao, Mazumdar, Waller and Li (1973); these authors have not, however, commented on it.

Another hypothesis considered by Smart was that the later born children have a greater tendency to become alcoholic than the earlier born. The method used by Smart may be somewhat confusing to statisticians. Some comments were made by Sprott criticizing Smart's approach. We shall review Smart's analysis in the light of the model (5.3.4). If we assume that birth order has no relationship on becoming an alcoholic, and the probability of an alcoholic being referred to a clinic is independent of the birth order, then the probability that an observed alcoholic belongs to a family with n children, r alcoholics and has given birth order $s \leq n$ is, using the model (5.3.4),

$$\sum_{s=1}^n p(n) \binom{n}{r} \pi^r \phi^{n-r} \div E(n) \quad (5.3.12)$$

$s = 1, \dots, n; r = 1, \dots, n; n = 1, 2, \dots$

Summing up over r , the marginal distribution of (n,s) , the family size and birth order, applicable to the observed distribution, is

$$p(n) \div E(n) \quad (5.3.13)$$

$$s = 1, \dots, n; n = 1, 2, \dots$$

where it may be recalled that $p(n)$, $n = 1, 2, \dots$, is the distribution of family size in the general population. Smart gave the observed bivariate frequencies of (n,s) , and since $p(n)$ was known, the expected values could have been computed and compared with the observed. He did something else.

From (5.3.13), the marginal distribution of birth rank is

$$\sum_{i=r}^{\infty} p(i) \div E(n), r = 1, 2, \dots \quad (5.3.14)$$

Smart's (1963) analysis in his Table 2 is an attempt to compare the observed distribution of birth ranks with the expected under the model (5.3.14) with $p(i)$ itself estimated from data using the model (5.3.1).

A better method is as follows: from (5.3.13) it is seen that for given family size, the expected birth order frequencies are equal as computed by Smart (1963) in Table 1, in which case individual chi squares comparing the expected and observed frequencies for each family size would provide all the information about the hypothesis under test. Such a procedure would be independent of any knowledge of $p(n)$. But it is not clear whether a hypothesis of the type posed by Smart can be tested on the basis of the available data without further information on the other alcoholics in the family, such as their ages, sexes, etc.

Let us consider a portion of Table 1 in Smart (1963) relating to families up to size 4 and birth ranks up to 4.

Table 7. Distribution of birth rank(s) and family size (n)
Smart (1963, Table 1)

birth rank	family size							
	1		2		3		4	
	O	E	O	E	O	E	O	E
1	21	21	22	16	17	13.3	11	11.75
2			10	16	14	13.3	10	11.75
3					9	13.3	13	11.75
4							13	11.75

O = observed, E = Expected

It is seen that for family sizes 2 and 3, the observed frequencies seem to contradict the hypothesis, and for family sizes above 3 (see Smart's Table 1), birth rank does not have any effect. It is interesting to compare the above data with a similar type of data collected by the author on birth rank and family size of the staff members in two departments at the University of Pittsburgh.

Table 8. Distribution of birth rank and family size (up to 4)
among staff members

birth rank	family size			
	1	2	3	4
1	7	14	9	6
2		6	4	2
3			2	0
4				0

It appears that there are too many earlier borns among the staff members indicating that becoming a professor is an affliction of the earlier born! It is clear that the observed data by themselves do not enable any inference to be drawn on the relationship between birth rank of a child and any attribute under consideration.

6. QUADRAT SAMPLING WITH VISIBILITY BIAS

For estimating wildlife population density, quadrat sampling has been found generally preferable. Quadrat sampling is carried out by first selecting at random a number of quadrats of fixed size from the region under study and ascertaining the number of animals in each. The following assumptions are made:

- A_1 : Animals are found in groups within each quadrat and the number of animals X in a group follows a specified distribution.
- A_2 : The number of groups N within a quadrat has a specified distribution.
- A_3 : The number of groups within a quadrat and the numbers of animals within groups are independent.

Let the method of sampling be such that the probability of sighting (or recording) a group of x animals is $w(x)$. If X^w and N^w represent the rv's of the number of animals in a group and number of groups within a quadrat as ascertained, then we have the following results.

$$(i) \quad P(N^W = m | N = n) = \binom{n}{m} w^m (1 - w)^{n-m} \quad (6.1)$$

where

$$w = \sum_1^{\infty} w(x) P(X = x) \quad (6.2)$$

is the visibility factor (or the probability of recording a group).

$$(ii) \quad P(N^W = m) = \sum_{n=m}^{\infty} \binom{n}{m} w^m (1 - w)^{n-m} P(N = n) \quad (6.3)$$

$$(iii) \quad P(N^W = m, X_1^W = x_1, \dots, X_m^W = x_m) \\ = w^{-m} P(N^W = m) \prod_{i=1}^m w(x_i) P(X = x_i) \quad (6.4)$$

(iv) Let $S^W = X_1^W + \dots + X_m^W$. Then

$$P(S^W = y) = \sum_{m=1}^{\infty} P(N^W = m) P(S^W = y | m) \quad (6.5)$$

$$P(S^W = y | m) = \sum_{\sum x_i = y}^{\infty} \frac{w(x_1)}{w} \dots \frac{w(x_m)}{w} P(X_1 = x_1) \dots P(X_m = x_m). \quad (6.6)$$

The formulae listed above are useful in many practical situations. Usually the sighting probability is of the form

$$w(x) = 1 - (1 - \beta)^x. \quad (6.7)$$

For some applications, the reader is referred to papers by Cook and Martin (1974), and Patil and Rao (1977, 1978).

7. WAITING TIME PARADOX

Patil (1984) reported a study conducted in 1966 by the "Institut National de la Statistique et de l'Economie Appliquée" in Morocco to estimate the mean sojourn time of tourists. Two types of surveys were conducted one by contacting tourists residing in hotels and another by contacting tourists at frontier stations while leaving the country. The mean sojourn time as reported by 3,000 tourists in hotels was 17.8 days and by 12,321 tourists at frontier stations was 9.0. Suspected by the officials in the department of planning, the estimate from the hotels was discarded.

It is clear that the observations collected from tourists while leaving the country correspond to the true distribution of sojourn time, so that the observed average 9.0 is a valid estimate of the mean sojourn time. It can be shown that in a steady state of flow of tourists, the sojourn time as reported by those contacted at hotels has a size biased distribution so that the observed average will be an overestimate of the mean sojourn time. If X^W is a size biased random variable, then

$$E(X^W)^{-1} = \mu^{-1} \quad (7.1)$$

where μ is the expected value of X , the original variable. The formula (7.1) shows that the harmonic mean of the size biased observations is a valid estimate of μ . Thus the harmonic mean of the observations from the tourists in hotels would have provided an estimate comparable with the arithmetic mean of the observations from the tourists at the frontier stations.

It is interesting to note that the estimate from hotel residents is nearly

twice the other, a factor which occurs in the waiting time paradox (see Feller, 1968; Patil and Rao, 1977) associated with the exponential distribution. This suggests, but does not confirm, that the sojourn time distribution may be exponential.

Suppose that the tourists at hotels were asked how long they had been staying in the country up to the time of enquiry. In such a case, under the assumption that the pdf of the rv Y , the time a tourist has been in a country up to the time of enquiry, is the same as that of the product $X^W R$ where X^W is the size biased version of X , the sojourn time, and R is an independent rv with a uniform distribution on $[0,1]$. If $F(x)$ is the distribution function of X , then the pdf of Y is

$$\mu^{-1}[1 - F(y)]. \quad (7.2)$$

The parameter μ can be estimated on the basis of observations on Y , provided the functional form of $F(y)$, the distribution function of the sojourn time, is known.

It is interesting to note that the pdf (7.2) is the same as that obtained by Cox (1962) in studying the distribution of failure-time of a component used in different machines from observations on the ages of the components in use at the time of investigation.

8. DAMAGE MODELS

Let N be a rv with probability distribution, p_n , $n = 1, 2, \dots$, and R be a rv such that

$$P(R = r | N = n) = s(r, n). \quad (8.1)$$

Then the marginal distribution of R truncated at zero is

$$p_r^t = (1 - p)^{-1} \sum_{n=r}^{\infty} p_n s(r, n), \quad r = 1, 2, \dots \quad (8.2)$$

where

$$p = \sum_{r=1}^{\infty} p_r s(0, 1). \quad (8.3)$$

The observation r represents the number surviving when the original observation n is subject to a destructive process which reduces n to r with probability $s(r, n)$. Such a situation arises when we consider observations on family size counting only the surviving children (R). The problem is to determine the distribution of N , the original family size, knowing the distribution of R and assuming a suitable survival distribution.

Suppose that $N \sim P(\lambda)$, i.e., distributed as Poisson with parameter λ and let $R \sim B(\cdot, \pi)$, i.e., binomial with parameter π . Then

$$p_r^t = e^{-\lambda \pi} \frac{(\lambda \pi)^r}{r!} \div (1 - e^{-\lambda \pi}), \quad r = 1, 2, \dots \quad (8.4)$$

It is seen that the parameters λ and π get confounded so that knowing the distribution of R , we cannot find the distribution of N . Similar confounding occurs when N follows a binomial, negative binomial or logarithm series distribution. When the survival distribution is binomial, Sprott (1965) gives a general class of distributions which has this property. What additional

information is needed to recover the original distribution? For instance, if we know which of the observations in the sample did not suffer damage, then it is possible to estimate the original distribution as well as the binomial parameter π .

It is interesting to note that observations which do not suffer any damage have the distribution

$$p_r^u = c p_r \pi^r \quad (8.5)$$

which is a weighted distribution. If the original distribution is Poisson, then

$$p_r^u = e^{-\lambda \pi} \frac{(\lambda \pi)^r}{r!} \div (1 - e^{-\lambda \pi}) \quad (8.6)$$

which is same as (8.4). It is shown in Rao and Rubin (1964) that the equality $p_r^u = p_r^*$ characterizes the Poisson distribution.

The damage models of the type described above were introduced in Rao (1965). For theoretical developments on damage models and characterization of probability distributions arising out of their study, the reader is referred to Alzaid, Rao and Shanbhag (1984).

9. NONRESPONSE: THE STORY OF AN EXTINCT RIVER

Sample survey practitioners define nonresponse as a missing observation or nonavailability of measurements on a unit included in a sample. It is clear

that if the missing values can be considered as a random sample from the population under survey then the observed values constitute a representative sample of the whole population (Rubin, 1976). Usually this is not the case and special techniques are developed in sample surveys to cope with such situations.

In general, nonresponse poses serious issues such as the problem of broken skulls not providing direct measurements on capacity (see Section 4 of this paper). More complex cases are as follows.

For instance, we may try to estimate the underground resources in a given region by making borings at a randomly chosen set of points and taking some measurements. But it may so happen that borings cannot be made at some chosen points due to some reasons such as the presence of rocks. The measurements at such points may be of a different type from the rest in which case the observed sample will not be a representative sample from the whole region.

Such a problem arose in an investigation by geologists at the Indian Statistical Institute to estimate the mean direction of flow of an extinct river of geological times in a certain region (see Sengupta, 1966; J. S. Rao and Sengupta, 1966). The geologists collected a series of observations on direction cosines of flow (two dimensional vector data), which seemed ideal for an application of Fisher's (1953) distribution and the associated theory for estimation of the mean direction of flow. Then the question arose as to what the hypothetical population was from which the observations could be considered as a random sample. It appeared that the measurements on direction cosines could not be made at any chosen point, but only at certain points where there was rock formation with some markings known

as "outcrops." The geologists walked along the region under exploration and made measurements wherever they came across outcrops. If the outcrops had been uniformly distributed over space, then it might have been possible to define a population of which the observations made by the geologists could be a representative sample. The locations at which observations were made when plotted on a topographical map of the region showed an unequal distribution of outcrops in different areas of the region indicating the nonrandom nature of the occurrence of outcrops. In such a case the estimate of mean direction assuming that each observation is an independent sample with a common expectation will be biased. In order to minimize the bias in estimation, the following method of estimation was adopted. A square lattice was imposed on the topographical map and the measurements in each grid were replaced by their average. Then a simple average of these averages was taken as an estimate of the mean direction of flow. This estimate differed somewhat from the average of all the measurements and was considered to have less bias.

This study points out the need for a re-examination of data on directions of rock magnetism collected by geologists and analysed by Fisher (1953) who developed a special theory for that purpose. If the outcrops at which measurements of direction are possible are not uniformly distributed over space, then there will be some difficulty in interpreting the observed mean direction as an estimate of some specific parameter.

10. CONCLUSIONS

Some of the broad conclusions that emerge from the discussion of the live examples in the paper are as follows:

Specification or the choice of a model is of great value in data analysis. An appropriate specification for given data can be arrived at on the basis of past experience, information on the stochastic nature of events, a detailed knowledge of how observations are ascertained and recorded, and an exploratory analysis of current data itself using graphical displays, preliminary tests and cross validation studies.

Inaccuracies in specification can lead to wrong inference. It is therefore worthwhile to review the data under different possible specifications (models) to determine how variant the conclusions could be.

What population does an observed sample represent? What is the widest possible universe to which the conclusions drawn from a sample apply? The answers depend on how the observations are ascertained and what the deficiencies in data are in terms of nonresponse, measurement errors, and contamination.

Every data set has its own unique features which may be revealed in initial scrutiny of data and/or during statistical analysis, which may have to be taken into account in interpreting data. Routine data analysis based on text book methods or software packages may be misleading.

Generally in scientific investigations, a specific question cannot be answered without knowing the answers to several other questions. It often pays to analyse the data to throw light on a broader set of relevant and related questions.

What data should be collected to answer specific questions? Lack of information on certain aspects may create undue complications in applying statistical methods and/or restrict the nature of conclusions drawn from available data. Attempts should be made to collect information on concomitant variables to the extent possible, whose use can enhance the precision of estimators of unknown parameters, and provide broader validity to statistical inference.

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