SURVIVAL ANALYSIS IN TELEMETRY STUDIES:
THE STAGGERED ENTRY DESIGN

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Abstract: The estimation of survival distributions for animals which are radio-tagged is an important current problem for animal ecologists. Allowance must be made for censoring due to radio failure, radio loss, emigration from the study area and animals surviving past the end of the study period. First we show that the Kaplan-Meier procedure widely used in medical and engineering studies can be applied to this problem. An example using some quail data is given for illustration. As radios malfunction or are lost, new radio-tagged animals have to be added to the study. We show how this modification can easily be incorporated into the basic Kaplan-Meier procedure. Another example using quail data is used to illustrate the extension. We also show how the log rank test commonly used to compare two survival distributions can be generalized to allow for additions. Simple computer programs which can be run on a PC are available from the authors.

Radio-telemetry is becoming an increasingly popular methodology for studying wild animal populations. An animal, captured by trap, dart gun or some other method, is fitted with a small radio transmitter and released. From release, the animal's unique radio signal can be monitored until the animal dies or is censored. An animal is censored if we have lost track of the animal due to radio failure, radio loss, or emigration from the study area.

The most common application of radio-telemetry technology has been to the study of animal movements in relation to daily activity patterns, seasonal changes, habitat types, and interaction with other animals. Time series approaches will become very important to the thorough analyses of these data (see for example Dunn and Gipson, 1977; Pantula and Pollock, 1985).

Biologists have also begun to use radio-tagged animals to study survival. Present techniques for analyzing the data from these studies assume that each survival event (typically an animal surviving a day) is independent and has a constant probability over all animals and all periods (see Trent and Rongstad, 1974; Bart and Robson, 1982; Heisey and Fuller, 1985). These assumptions are often believed to be unrealistic and restrictive. White (1983) has generalized discrete approaches in the framework of band return models (Brownie et al. (1978)). He has developed a flexible computer program, SURVIV, for use with his approach.

Typically an animal's exact survival time (at least to within one or two days) is known unless that survival time is right censored (that is only known to be greater than some value). We suggest an approach based on the continuous survival models allowing right censoring which are widely used

in medical and engineering applications (Kalbfleisch and Prentice, 1980; Cox and Oakes, 1984). Pollock (1984) and Pollock et al. (1987) illustrated the usefulness of this approach and provided examples of the Kaplan-Meier procedure. This procedure does not require specification of a particular parametric continuous distribution such as the exponential or Weibull.

In this paper we first present a simple description of the Kaplan-Meier procedure complete with an example of some quail survival data collected by Curtis. We next show how to generalize the Kaplan-Meier procedure to allow gradual (or staggered entry) into the study. The calculations are again illustrated with an example from the Curtis quail data. Finally, we present the log rank test for comparison of survival distributions (modified for staggered entry of animals) and illustrate it. We also present a discussion of model assumptions and directions for future research.

THE KAPLAN-MEIER OR PRODUCT LIMIT PROCEDURE

The Kaplan-Meier or product limit estimator was developed by Kaplan and Meier (1958) and is discussed in many books on survival analysis. See for example Cox and Oakes (1984, p. 48) or Kalbfleisch and Prentice (1980, p. 13).

The survival function S(t) is the probability of an arbitrary animal in our population surviving t units of time from the beginning of the study. A nonparametric estimator of the survival function can be obtained by just restricting ourselves to the discrete points where deaths occur a_1, a_2, ..., a_g. We define r_1, ..., r_g to be the numbers of animals at risk at these points and d_1, d_2, ..., d_g to be the number of deaths at

the same points. The probability of surviving from 0 to a_1 is then estimated by

$$\hat{S}(a_1) = 1 - d_1/r_1$$

because d_1/r_1 is the estimated proportion dying in that interval. The probability of surviving from a_1 to a_2 is similarly given by

$$1 - d_2/r_2$$

and $\hat{S}(a_2)$ is then given by the product

$$\hat{S}(a_2) = (1 - d_1/r_1)(1 - d_2/r_2)$$

Therefore the estimated survivor function for any arbitrary time t is given by

$$\hat{S}(t) = \prod_{j \mid a_j < t} (1 - d_j/r_j) \tag{1}$$

which is the mathematical way of stating we are considering the product of all j terms for which a_j is less than the time t.

Let us talk further about r_i which is the number at risk at time a_i. In this situation we would start off with a fixed sample of size n. The number at risk at a particular death time a_i will then be n minus the number of deaths before a_j minus the number of animals censored before time a_j.

As an example of the use of this model, we present results from a radio-tagging study on northern bobwhite quail (Colinus virginianus) conducted by one of the authors (P.Curtis) at Fort Bragg, North Carolina. This was a two-year study, but in this section we just consider the data collected in the spring of 1985. This is a small study, the pertinent data on each of the eighteen radio-tagged birds is included in Table 1.

Six birds died and five birds disappeared (were censored) during the study, leaving seven birds which survived for the thirteen weeks of the study.

Let us illustrate how the estimation of $\hat{S}(t)$ was carried out to obtain the last column in Table 1 and Figure 1. The computations involve only the five weeks in which deaths were recorded; therefore $a_1 = 3$, $a_2 = 6$, $a_3 = 8$, $a_4 = 9$ and $a_5 = 10$. Also recall that we have eighteen animals that began the study at time 0. We estimate $\hat{S}(a_1)$ as

$$\hat{S}(a_1) = \hat{S}(3) = 1 - d_1/r_1$$
$$= 1 - 2/18$$
$$= 0.8889,$$

because there are 2 deaths at time 3 and there are 18 animals still at risk (r_1). The next death time a_2 is at 6 weeks $(a_2 = 6)$ and at that time there is one death $(d_2 = 1)$ and 16 animals at risk $(r_2 = 16)$. There are 16 at risk because 2 were lost to death at time 1. Therefore $\hat{S}(a_2)$ is given by

$$\hat{S}(a_2) = \hat{S}(6) = (1 - d_1/r_1)(1 - d_2/r_2)$$
$$= (1 - 2/18)(1 - 1/16)$$
$$= 0.8333$$

Similarly $\hat{S}(a_3)$ is given by

$$\hat{S}(a_3) = \hat{S}(8) = (1 - d_1/r_1)(1 - d_2/r_2)(1 - d_3/r_3)$$
$$= (1 - 2/18)(1 - 1/16)(1 - 1/15)$$
$$= 0.7778,$$

$$\hat{S}(a_4) = \hat{S}(9) = (1 - d_1/r_1)(1 - d_2/r_2)(1 - d_3/r_3)(1 - d_4/r_4)$$
$$= (1 - 2/18)(1 - 1/16)(1 - 1/15)(1 - 1/13)$$
$$= 0.7179,$$

and $\hat{S}(a_5) = \hat{S}(10) = (1 - d_1/r_1)(1 - d_2/r_2)(1 - d_3/r_3)(1 - d_4/r_4)(1 - d_5/r_5)$

$$= (1 - 2/18)(1 - 1/16)(1 - 1/15)(1 - 1/13)(1 - 1/10)$$

$$= 0.6462.$$

Notice that the censored observation at time 8 is still considered at risk until the instant after that time so that $r_3 = 15$, not 14, but then $r_4 = 13$. Notice also that the estimate of the survivor function ($\hat{S}(t)$) is presented for each week in Table 1 but that it only changes at the death times. Thus $\hat{S}(t)$ stays at 1.00 until time 3 where it becomes 0.8889 and stays there until time 6 (the next death time) and so on.

Cox and Oakes (1984 p. 51) also discuss how to estimate the variance of the estimate at an arbitrary time point using Greenwood's formula

$$\text{var } (\hat{S}(t)) = [\hat{S}(t)]^2 \sum_{j|a_j < t} \frac{d_j}{r_j(r_j - d_j)} \tag{2}$$

where the summation is for all death times a_j less than t. They also propose an alternate simpler estimate which is better in the tails of the distribution

$$\text{var } (\hat{S}(t)) = \frac{[\hat{S}(t)]^2 [1 - \hat{S}(t)]}{r(t)} \tag{3}$$

Approximate confidence intervals can be obtained using either of these equations. For example, a 95% confidence interval at $t = t_0$ would be

$$\hat{S}(t_0) \pm 1.96 [\text{var } \hat{S}(t_0)]^{1/2} \tag{4}$$

because of the asymptotic (large sample) normality of the estimates $\hat{S}(t)$. In Table 1 approximate 95% confidence intervals are given at all points using the simpler second variance equation (3). Notice that the confidence intervals get wider and wider as the time increases (Figure 1).

ESTENSION OF THE KAPLAN–MEIER PROCEDURE TO STAGGERED ENTRY OF ANIMALS

Here we extend the concept of the Kaplan–Meier estimates to allow animals to enter at different times and for the time variable to be measured from the point where the first group of animals is tagged. Previously we presented an example of the Kaplan–Meier estimator and showed it is based on the formula (3) which is

$$\hat{S}(t) = \prod_{j \mid a_j < t} (1 - d_j/r_j)$$

for the survival function where r_j is the number at risk and d_j is the number of deaths. Typically we assume r_j is decreasing due to deaths and censoring but there is no reason it has to be. New animals will only be considered in those product terms where they are at risk. The formula for the variance of $\hat{S}(t)$ also allows for new animals to enter during the study. It should be emphasized that any newly tagged animals are assumed to have the same survival function as the previously tagged animals.

Let us illustrate the extension of the Kaplan–Meier estimator to staggered entry of animals by considering some further bobwhite quail data collected by P. Curtis. Here we just consider the radio-tagging carried out in the winter period of 1985–86. The data and estimates are presented in Table 2. In Week 1 there were 20 animals radio-tagged ($r_1 = 20$), no deaths, no censors and one animal was added so that the number of animals radio-tagged in Week 2 is 21 and the survival estimate stays at

1 ($\hat{S}(1)$ = 1.0000). In Week 2 the only change is that another animal is

added so that at Week 3 the number at risk is r_3= 22. In Week 3 the survival

estimate is

$$\hat{S}(3) = 1 - d_3/r_3$$
$$= 1 - 2/22 = 0.9091.$$

In Week 4 the number at risk is 19 (2 deaths and 1 censor in Week 3 gives

22 -2 - 1 = 19) and there were 5 deaths so that

$$\hat{S}(4) = (1 - d_3/r_3)(1 - d_4/r_4)$$
$$= (1 - 2/22)(1 - 5/19) = 0.6699$$

The survival estimates for later times can be obtained similarly. The

approximate 95% confidence limits are also given based on equations (3) and

(4) presented in the previous section. Notice that here the confidence

limits do not necessarily get wider with time because new animals may be

added (Figure 2). At time 13 there is a marked decrease in the confidence

interval width because the number at risk jumps from 10 to 16 and then at

time 14 it jumps again to 22 due to the large numbers of new animals added

at those times.

LOG RANK TEST EXTENSION TO STAGGERED ENTRY OF ANIMALS

Often it is important to compare two estimated survival functions to

see if they could have come from the same underlying true survival curve.

For example, in Table 3 we present some bobwhite quail survival estimates

for Fall 1985 and Fall 1986. We would like to know if the survival patterns

are the same for the two years. Graphical comparison would be possible by

plotting survival functions on the same graph; however, a formal hypothesis

testing procedure is also needed..

There are many possible tests available (Lee, 1980, p. 122) but here we will concentrate attention on the log rank test (Savage, 1956; Kalbfleisch and Prentice, 1980, p. 17; Cox and Oakes, 1984, p. 104). We do this because of the test's simplicity and easy generalization to the case where animals have staggered entry into the study.

To compare two survivor functions let us generalize the formulation we used for defining the Kaplan-Meier estimates. Let a_1, a_2, ..., a_k denote the death times for the sample formed by combining the two samples. Suppose there are d_j deaths and r_j animals at risk at a_j, with d_{0j} and d_{1j} being from samples one and two, respectively. Similarly there are r_{0j} and r_{1j} animals at risk from the two samples.

For each of the k points the data can be represented as a 2 x 2 contingency table. For the jth contingency table we have the formulae for the mean and variance of d_{1j} given by

$$E(d_{1j}) = d_j r_{1j}/r_j$$
$$Var_1(d_{1j}) = r_{0j} r_{1j} d_j (r_j - d_j)/r_j^2 (r_j - 1)$$

An approximate χ^2 test statistic with one degree of freedom can be obtained by combining the results from all the contingency tables (assuming conditional independence and asymptotic normality of the d's) in the following way

$$X^2 = \left(\frac{\sum\limits_{j=1}^{K} d_{1j} - \sum\limits_{j=1}^{K} E(d_{1j})}{\sum\limits_{j=1}^{K} var(d_{1j})} \right)^2$$

Cox and Oakes (1984, p. 105) consider two modifications to the variance of d_{1j} which give rise to X_2^2 and X_3^2 respectively. The first modification is to use

$$var_2(d_{1j}) = \frac{r_{0j} \, r_{1j} \, d_j}{r_j^2}$$

which gives a slightly more conservative test. The second modification is to use

$$\sum_{i=1}^{K} var(d_{1j}) = \left[\frac{1}{\sum_{j=1}^{K}(d_j r_{0j}/r_j)} + \frac{1}{\sum_{j=1}^{K}(d_j r_{1j}/r_j)} \right]^{-1}$$

which gives an even more conservative test. Notice that as mentioned earlier, this test in any of its three versions easily generalizes to the case of staggered entry because r_{0j} and r_{1j}, the number at risk at each time point, can be redefined to include newly tagged animals.

In Table 4 we present the calculations of the log rank test for data given in Table 3 which compares bobwhite quail survival distributions for Fall 1985 and Fall 1986. Notice that we present the number at risk and the number of deaths for every week although deaths do not occur every week. Notice that in those weeks where there are no deaths there are no contributions to the test statistics.

We now present the calculations for the three approximate χ^2 test statistics with one degree of freedom.

$$X_1^2 = \frac{(6 - 3.317)^2}{1.681} = 4.28$$

$$X_2^2 = \frac{(6 - 3.317)^2}{1.729} = 4.16$$

$$X_3^2 = \frac{(6 - 3.317)^2}{[\frac{1}{3.683} + \frac{1}{3.317}]^{-1}} = 4.13$$

where
$$\sum_{i=1}^{K} d_{1j} = 6$$

$$\sum_{i=1}^{K} E(d_{1j}) = 3.317$$

and hence
$$\sum_{i=1}^{K} E(d_{0j}) = 7 - E(d_{1j}) = \sum_{i=1}^{K} d_j r_{0j}/r_j$$
$$= 3.683$$

$$\sum_{i=1}^{K} d_j r_{0j} r_{1j}/r_j^2 = 1.729$$

$$\sum_{i=1}^{K} d_j r_{0j} r_{1j}(r_j - d_j)/r_j^2(r_j - 1) = 1.681$$

Notice that all the chi-square tests are very similar and show slightly increasing conservatism. The approximate p value is 0.04 which indicates there is a significant difference between the two years survival curves at the 5% level.

ASSUMPTIONS

Here we briefly discuss the assumptions of the model as they apply to radio-telemetry data. As this is a new approach, further research on the validity of these assumptions in practice and on model robustness is required.

We have assumed that a random sample of animals of a particular age and sex class has been obtained. Take, for example, a study on winter survival of mallards. If lighter adult males tend to be captured and these have lower survival rates, a negative bias to the survival estimates will result. Of course, this assumption is also crucial to survival estimates obtained from capture-recapture and band return studies (Jolly, 1965; Seber, 1965; Pollock, 1981; Brownie et al., 1978).

This model requires the assumption that survival times are independent for the different animals. Again, this assumption is also required of capture-recapture and band return models. Geese which form tight family groups would be an example where this assumption could fail. Additionally, the death of a large female mammal (e.g., black bear) still nursing her young would not be independent of the fate of those young. We believe that violation of this assumption will not cause bias but it will make our estimates appear more precise than they actually are.

Another assumption which is common to any method involving marked animals is that capturing the animal or having it carry a radio-tag does not influence its future survival. Clearly, failure of this assumption will cause a negative bias on the survival estimates. As radio-tags are becoming more sophisticated and, hence, smaller this is less of a problem. Short term effects could be eliminated by having a conditioning period of say one week after tagging where an animal's survival time is not considered until it has survived that period. (This was done by P. Curtis with the quail data.)

The assumption that the censoring mechanism is random (i.e., is not related to an animal's fate) is extremely important and requires more attention from researchers. Possible violations could result from a predator killing an animal and also destroying the radio or an animal emigrating because it is more (or less) healthy than its companions. Medical studies often suffer a similar "emigration" problem; patients doing poorly (or well) may decide to leave the study. A review of the literature on survival analysis shows that very little has been done regarding alternative, so called, informative censoring models. The authors believe this is mainly due to the difficulty of the problem rather than lack of research. However, "bounds" can be generated for the survival curve by allowing censoring to take two very extreme forms. A lower bound can be obtained by assuming that every censored observation was really a death and an upper bound by assuming that every censored observation was not a death and that the animal survived to the end of the study.

In some cases, it may be reasonable to assume that either emigration or radio failure is zero. Then the likelihood of the censoring time could provide very useful information. For example, in a study of winter survival of waterfowl with reliable radios, the censoring times would primarily reflect emigration. Estimation of this emigration time distribution could be informative to the biologist, especially if it could be related to covariates such as those reflecting weather severity.

One of the most important considerations in application of survival analysis to radio-telemetry data is the definition of a time origin. In medical studies the natural time origin is the time treatment begins. In

radio-telemetry there is no such natural time origin. In studies where all the animals are captured at or near the same time the obvious time origin might be the date when the last animal was captured. It should be kept in mind that survival from the origin could be seriously influenced by seasonal effects. For example, survival for one month from a summer time origin could be quite different than survival for one month from a winter time origin.

In some studies animals may be introduced into the study gradually over a long period of time. This could be due to practical problems of capturing animals all at one time or because the biologist deliberately wants to introduce more animals into the study to increase precision after a lot of the animals have died. We have shown that the Kaplan-Meier estimator of the survivor function and the log rank test for comparing survival curves can be easily generalizedto allow for staggered entry. In this case the time origin will be when the first group of animals is tagged.

A special assumption of the staggered entry design is that newly tagged animals have the same survival function as previously tagged animals. If there were enough animals in both groups contingency table tests of this assumption could be made. In practice, however, the animals will often be added in very small groups thereby prohibiting a quantitative assessment of this assumption..

DISCUSSION

The radio-tagged survival analysis procedure presented above provides a general framework for analyses of these studies. Radio-telemetry is likely

to become an even more common technique as the technology improves and costs are reduced. The large body of statistical research into survival analysis in medicine and engineering should prove a valuable starting point. The techniques we've outlined above, within the constraints of the listed assumptions, provide the researcher with a more realistic and sophisticated analysis than has previously been possible. The Kaplan-Meier procedure is simple and flexible and also easily generalizable to the case of staggered entry of newly tagged animals. Many biologists will find the simple log rank test for comparing survival distributions very valuable. It is also easily adapted to the staggered entry case.

We believe the testing of ecological hypotheses regarding the influence of individual animal covariates on survival using the proportional hazards model (Cox, 1972) is extremely important. In Pollock et al. (1987), this model is illustrated by showing how winter survival of female black ducks is related to their condition index at the start of the winter. The Cox proportional hazards model is described clearly by Cox and Oakes (1984, p. 91). This model can also generalize to the case of staggered entry of animals. We plan to do this and write a computer program for implementation of this model in a future article.

There needs to be more work done to study the efficiency of the Kaplan-Meier estimator (and the power of the log rank test) when there is staggered entry of animals but some preliminary statements can be made based on the analyses presented here. If you look carefully at Table 2 and Figure 2 it is clear that the precision is poor unless the number of animals

tagged at a particular time is above about 20. To get good precision, however, 40-50 animals would need to be tagged at all times as a minimum. (Also if there is a period of interest when mortality is likely to be high the biologist should be prepared to introduce a large number of newly radio-tagged animals at that time.)

In the analysis discussed in this paper we have put most emphasis on the Kaplan-Meier product limit estimator because of its simplicity and generality. An important question is when should one use parametric modelling as compared to non parametric? Previous approaches to analysis of radio-telemetry data (Trent and Rongstadt, 1974; Bart and Robson, 1982; Heisey and Fuller, 1985) could be viewed as very special cases of parametric modelling. Although discrete, their approaches are very similar to fitting an exponential distribution. Miller (1983) has done a comparison of maximum likelihood estimation and the Kaplan-Meier procedure when the underlying distribution is exponential and there is right censoring. As Miller (1983) points out this comparison is biased against the Kaplan-Meier estimator and its efficiency can be low. This is especially troublesome when t is large and Miller (1983) states "Parametric modelling should be considered as a means of increasing the precision in the estimation of small tail probabilities". He further states that it is surprising that so little work has been done on this question considering the importance of survival analysis in many disciplines.

Lagakos (1979) in a review paper on right censoring and survival analysis discusses informative censoring (i.e., where the censoring is related to the fate of the animal). Again it is surprising how little research has been done on this problem. One practical approach discussed in the

assumptions section is to calculate extreme bounds for the estimated survival curve by considering each censored observation to be either a death or a survivor until the end of the study. If there is a lot of censoring early in the study, these bounds can be very wide.

Finally, we wish to discuss the problem where cause of death can be classified into several categories. For example, in some cases the biologists may want to separate hunting deaths from nonhunting deaths. Marginal survival curves can be obtained by treating deaths from any other cause as censored observations. For example, if one were considering the survival curve just related to hunting, then all animals that died of nonhunting causes would be viewed as censored observations. Unfortunately this approach does not take into account that different causes of death may not be independent. A lot of work has been done on competing risks models when there are several possible causes of deaths, but it has been shown that these models are not useful for estimating the dependency. These models are what statisticians refer to as "nonidentifiable". Therefore, the biologist is forced back to using marginal or crude survival curves, but he or she should be aware that these results could be misleading if the different causes of death are not independent. For more information on competing risk models and their problems the reader could refer to Kalbfleisch and Prentice (1980, p. 163).

COMPUTER PROGRAMS

Two computer programs written specifically for the analysis of survival data when all animals enter at the same time are PHGLM (Harrell, 1983) and

LIFETEST (SAS, 1985). Other programs include SURVREG (Preston and Clarkson, 1983) and LIFEREG (SAS, 1985) for use with parametric models when all animals enter at the same time.

A simple computer program which calculates the Kaplan-Meier estimator and the log rank test when there is potentially staggered entry of animals is available from the authors. It will run on IBM compatible personal computers in conjunction with the Lotus Spread Sheet software (Kapor and Sachs, 1983).

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Table 1. Kaplan–Meier survival estimates for bobwhite quail radio-tagged in Spring 1985.

Week (t)	Dates	No. at Risk (r_j)	No. of deaths (d_j)	No. censored	Survival $\hat{S}(t)$	95% Confidence Interval Lower	95% Confidence Interval Upper
1	3/3 –3/9	18	0	0	1.0000	1.0000	1.0000
2	3/10–3/16	18	0	0	1.0000	1.0000	1.0000
3	3/17–3/23	18	2	0	0.8889	0.7520	1.0258
4	3/24–3/30	16	0	0	0.8889	0.7437	1.0341
5	3/31–4/6	16	0	0	0.8889	0.7437	1.0341
6	4/7 –4/13	16	1	0	0.8333	0.6667	1.0000
7	4/14–4/20	15	0	0	0.8333	0.6612	1.0055
8	4/21–4/27	15	1	1	0.7778	0.5922	0.9633
9	4/28–5/4	13	1	2	0.7179	0.5107	0.9252
10	5/5 –5/11	10	1	1	0.6462	0.4079	0.8844
11	5/12–5/18	8	0	0	0.6462	0.3798	0.9125
12	5/19–5/25	8	0	1	0.6462	0.3798	0.9125
13	5/26–6/1	7	0	0	0.6462	0.3614	0.9309

Table 2. Kaplan-Meier survival estimates for bobwhite quail radio-tagged in Winter 1985-86 modified to allow for the staggered entry of new animals.

Week (t)	Dates	No. at Risk (r_j)	No. of deaths (d_j)	No. censored	No. new added	Survival $\hat{S}(t)$	95% Confidence Interval Lower	Upper
1	11/17-11/23	20	0	0	1	1.0000	1.0000	1.0000
2	11/24-11/30	21	0	0	1	1.0000	1.0000	1.0000
3	12/1 -12/7	22	2	1	0	0.9091	0.7946	1.0236
4	12/8 -12/14	19	5	0	0	0.6699	0.4968	0.8429
5	12/15-12/21	14	3	0	0	0.5263	0.3366	0.7161
6	12/22-12/28	11	0	0	0	0.5263	0.3122	0.7404
7	12/29-1/4	11	0	0	0	0.5263	0.3122	0.7404
8	1/5 -1/11	11	2	0	0	0.4306	0.2386	0.6226
9	1/12-1/18	9	1	0	0	0.3828	0.1863	0.5792
10	1/19-1/25	8	0	1	0	0.3828	0.1744	0.5912
11	1/26-2/1	7	0	0	3	0.3828	0.1600	0.6056
12	2/2 -2/8	10	0	0	6	0.3828	0.1964	0.5692
13	2/9 -2/15	16	4	0	10	0.2871	0.1683	0.4059
14	2/16-2/22	22	4	0	5	0.2349	0.1490	0.3207
15	2/23-3/1	23	4	1	6	0.1940	0.1228	0.2652
16	3/2 -3/8	24	4	0	0	0.1617	0.1025	0.2209
17	3/9 -3/15	20	2	0	0	0.1455	0.0866	0.2045

Table 3. Comparison of survival distributions for Fall 1985 and
Fall 1986 of radio-tagged bobwhite quail.

	Fall 1985					Fall 1986				
Week	No. at risk	No. deaths	No. censored	No. added	$\hat{S}$ (t)	No. at risk	No. deaths	No. censored	No. added	$\hat{S}$ (t)
1	7	1	0	0	0.8571	7	0	1	0	1.0000
2	6	0	0	2	0.8571	6	0	0	5	1.0000
3	8	0	0	5	0.8571	11	1	0	0	0.9091
4	13	0	1	6	0.8571	10	0	0	6	0.9091
5	18	0	0	0	0.8571	16	1	0	0	0.8523
6	18	0	0	0	0.8571	15	0	0	0	0.8523
7	18	0	0	0	0.8571	15	1	0	0	0.7955
8	18	0	0	0	0.8571	14	0	0	0	0.7955
9	18	0	0	1	0.8571	14	3	0	0	0.6250

Table 4. Log rank test calculations comparing survival distributions of radio-tagged bobwhite quail for Fall 1985 and Fall 1986 modified to allow for the staggered entry of new animals.

	Fall 1985		Fall 1986		Total		$E(d_{1j})$	$Var_2(d_{1j})$	$Var_1(d_{1j})$
Week	r_{0j}	d_{0j}	r_{1j}	d_{1j}	r_j	d_j	$\dfrac{d_j}{r_j}\cdot r_{1j}$	$\dfrac{d_j r_{1j} r_{0j}}{r_j^{\,2}}$	$\dfrac{d_j r_{1j} r_{0j}(r_j-d_j)}{r_j^{\,2}(r_j-1)}$
1	7	1	7	0	14	1	0.500	0.250	0.250
2	6	0	6	0	12	0	0	0	0
3	8	0	11	1	19	1	0.579	0.244	0.244
4	13	0	10	0	23	0	0	0	0
5	18	0	16	1	34	1	0.471	0.249	0.249
6	18	0	15	0	33	0	0	0	0
7	18	0	15	1	33	1	0.455	0.248	0.248
8	18	0	14	0	32	0	0	0	0
9	18	0	14	3	32	3	1.313	0.738	0.691
Total		1		6		7	3.317	1.729	1.681

FIGURE CAPTIONS

Figure 1.   The Kaplan–Meier survival function for bobwhite quail
            radio-tagged in Spring 1985.

Figure 2.   The Kaplan–Meier survival function, modified for staggered
            entry of animals, for bobwhite quail radio-tagged in
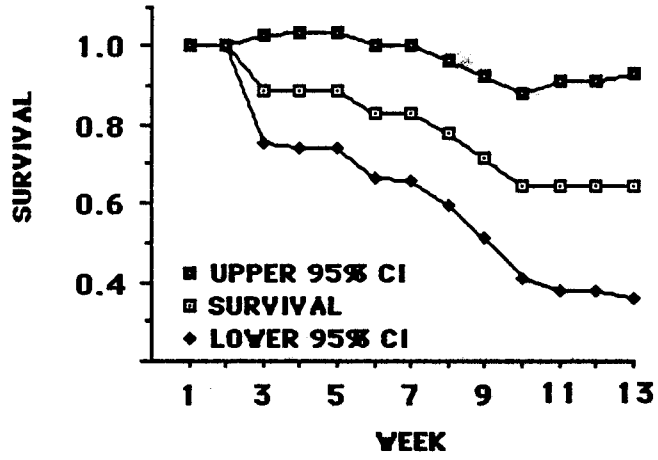   Winter 1985–86.
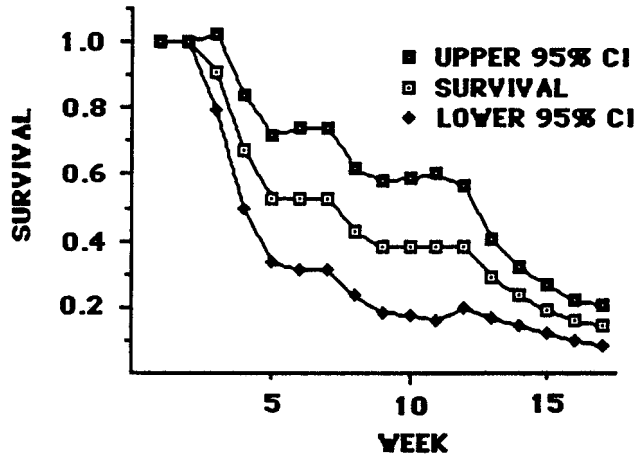
Fig. 1.

Fig. 1.

Fig. 2.

Fig. 2.