Lecture 5

Survival analysis

Let T be a non-negative random variable that represents the time to event (for example, T might be unemployment time, or time from a treatment to death of a patient, or time from production to failure of a device, etc.). Denote $F(t) = P(T \leq t)$ the c.d.f. of T and assume that F has a p.d.f. f. A central concept of the survival analysis is the *hazard function*

$$\begin{aligned} h(t) &= \lim_{td \to 0} \frac{P(t \leq T \leq t + td | T \geq t)}{td} \\ &= \lim_{td \to 0} \frac{1}{td} \frac{P(t \leq T \leq t + td)}{P(T \geq t)} \\ &= \frac{f(t)}{1 - F(t)}, \end{aligned}$$

That is, $h(t)$ is the instantaneous rate of failure at a time $T = t$ under the condition of survival to the time t (loosely speaking, $h(t)$ is the probability density of failure at time t, given survival to then). If T is a discrete random variable, $h(t) = P(T = t | T > t) = P(T = t)/\{1 - F(t)\}$. Function

$$S(t) = P\left(T > t\right) = 1 - F(t)$$

is called *survivor function* of T. Next, define the *cumulative hazard function* by

$$H(t) = \int_0^t h(s) ds = \int_0^t \frac{f(s)}{1 - F(s)} ds = \int_0^t \frac{F'(s)}{1 - F(s)} ds = -\log\{S(t)\}.$$

Thus the survivor function may be written as $S(t) = \exp\{-H(t)\}$ and $f(t) = h(t)S(t) = h(t)\exp\{-H(t)\}$.

Examples

1. Exponential distribution: $h(t) = \lambda$ and $S(t) = \exp(-\lambda t)$
2. Weibull distribution: $h(t) = \alpha\lambda^\alpha t^{\alpha-1}$ and $S(t) = \exp\{-(\lambda t)^\alpha\}$
3. Log-logistic distribution: $h(t) = \alpha\lambda^\alpha t^{\alpha-1}\{1 + (\lambda t)^\alpha\}^{-1}$ and $S(t) = \{1 + (\lambda t)^\alpha\}^{-1}$

Ideally, we would have independent realisations of T: $t_1, \ldots, t_n$. However, in practice the failure time can not always be observed due to various reasons. This phenomenon is called *censoring*. The simplest form of censoring occurs when T is observed until some pre-determined time c. If $T < c$, we observe the value t_i of T, if $T > c$, we only know that T survived beyond c. This is called *Type I censoring. Type II censoring* arises when n independent variables are observed until there have been r failures, so only $0 < T_{(1)} < \ldots < T_{(r)}$ are observed. This type of censoring has an open-ended random trial time and is therefore impractical and is rarely used. Under *random censoring* the jth subject in the study has a random censoring time C_j drawn from some distribution G, independent of T_j. These are all examples of *right-censoring.* Left-censoring (the time of origin is not known) is less common.

Hence, under censoring one rather deals with $Y_j = \min\{T_j, C_j\}$, while it is known if $Y_j = T_j$. That is, a pair (y_j, δ_j) is observed, where $\delta_j = 1$ if y_j is the survival time and $\delta_j = 0$, if y_j is the censoring time. Note that the assumption of independence of T and C is crucial.

Now assume that T has a continuous distribution F and there are n data points available $(y_1, \delta_1), \ldots, (y_n, \delta_n)$, where $y_i = \min\{t_i, c_i\}$. Assume that $F(x) = F(x; \theta)$ is a some parametric distribution and that censoring variables C_i (independent on T_i) have c.d.f. G and p.d.f. g, which are independent on θ. Hence, the likelihood contribution from y_i is

$$\begin{cases} f(y_i;\theta)\{1 - G(y_i)\}, & \text{if } \delta_i = 1 \\ S(y_i;\theta)g(y_i), & \text{if } \delta_i = 0. \end{cases}$$

Since G and g are independent on θ, the likelihood becomes

$$\mathcal{L}(\theta) = \prod_{i=1}^{n} f(y_i;\theta)^{\delta_i} S(y_i;\theta)^{1-\delta_i},$$

while log-likelihood is

$$\ell(\theta) = \sum_{i=1}^{n} \left[\delta_i \log\{f(y_i;\theta)\} + (1 - \delta_i)\log\{S(y_i;\theta)\}\right].$$

This includes Type I censoring, for which G puts all its probability at c. Note that we can represent the log-likelihood as

$$\ell(\theta) = \sum_{i=1}^{n} [\delta_i \log\{h(y_i;\theta)\} - H(y_i;\theta)].$$

For exponential distribution $h(t;\lambda) = \lambda$ and $H(t;\lambda) = \lambda t$, so that the log-likelihood becomes

$$\ell(\lambda) = \sum_{i=1}^{n} \{\delta_i \log(\lambda) - \lambda y_i\} = \log(\lambda)\sum_{i=1}^{n}\delta_i - \lambda\sum_{i=1}^{n} y_i,$$

implying

$$\hat{\lambda}_{ML} = \frac{\sum_{i=1}^{n} \delta_i}{\sum_{i=1}^{n} y_i}.$$

In particular, if all observations are censored (=no failures), the estimator is zero. To calculate the (asymptotic) variance of a maximum likelihood estimator, we would need to calculate the Fisher Information $E\{J(\lambda)\}$, where $J(\lambda) = -\partial^2 \ell(\lambda)/\partial\lambda^2$ is the observed information. However, this is not possible without some assumption on G. In practice one approximates $\widehat{\text{var}}(\hat{\lambda}) = J(\hat{\lambda})^{-1} = \hat{\lambda}^2 / \sum_{i=1}^{n} \delta_i$. In particular, this can be used to build an approximate confidence interval for λ (using asymptotic normality of maximum likelihood estimators) as

$$\left[\hat{\lambda}(1 - z_{\alpha/2}/\sqrt{r}), \hat{\lambda}(1 + z_{\alpha/2}/\sqrt{r})\right], \quad r = \sum_{i=1}^{n} \delta_i.$$

The assumption of a constant hazard function is often unrealistic (for example, the instantaneous failure rate (hazard) of a technical device usually grows with the time from being put into service). A commonly used parametric distribution for modelling lifetimes with monotone hazard is the Weibull distribution. Values for λ and α can be estimated by the maximum likelihood similarly to the exponential distribution, however, this has to be done numerically.

Kaplan-Meier and Fleming-Harrington estimator

Often it is unclear which parametric model would be appropriate for the data (if any). A standard tool for initial data inspection, for suggesting plausible models and for checking their fit is a nonparametric estimator of the survivor function. If there were no censored observations, then we could estimate $\hat{S}(t) = n^{-1} \sum_{i=1}^{n} \mathbb{I}(T_i > t)$. For censored observations we have the likelihood

$$\mathcal{L}(S) = \prod_{i=1}^{n} f(y_i)^{\delta_i} S(y_i)^{1-\delta_i}.$$

Let $0 \leq \tau_1 < \tau_2 < \ldots$ be the ordered uncensored failure times. Let r_i denote the number of units that are still in risk at τ_i (=not failed yet or censored) and d_i the number of units that fail at τ_i. It can be shown that the function $\mathcal{L}$ is maximized by the piecewise constant function $\hat{S}_{\text{KM}}$ defined by

$$\hat{S}_{\text{KM}}(t) = \prod_{\{j:\tau_j < t\}} \left(1 - \frac{d_j}{r_j}\right).$$

This is the *Kaplan-Meier* estimator for the survivor function S.

A further estimator for S is the *Fleming-Harrington* estimator $\hat{S}_{\text{FH}}$. It is a plug in estimator defined by $\hat{S}_{\text{FH}} = \exp\{-\hat{H}(t)\}$, where $\hat{H}$ is the *Nelson-Aalen* estimator for H.

It is defined by

$$\hat{H}(t) = \sum_{\{j:\tau_j<t\}} \frac{d_j}{r_j}.$$

Observe that

$$\hat{S}_{\mathrm{FH}} = \exp\{-\hat{H}(t)\} = \prod_{\{j:\tau_j<t\}} \exp\left(-\frac{d_j}{r_j}\right).$$

Since $1-x \approx \exp(-x)$ for small x, the estimators $\hat{S}_{\mathrm{FH}}$ and $\hat{S}_{\mathrm{KM}}$ are quite similar, if many items are still at risk.

We next aim for computing **confidence bands** for the true survivor function S. To this end, assume that $\hat{S}$ is an estimator for S (e.g. the Kaplan-Meier or the Fleming-Harrington estimator) and let $\widehat{\mathrm{var}}(\log \hat{S})$ be some estimate for the variance of $\log \hat{S}$. Then, by the delta-method, it follows that $\mathrm{var}\{\hat{S}(t)\} \approx \hat{S}^2\widehat{\mathrm{var}}(\log \hat{S})$ and hence an approximate confidence band can be constructed by

$$\left[\hat{S}(t) - z_{\alpha/2}\hat{S}(t)\sqrt{\widehat{\mathrm{var}}\{\log \hat{S}(t)\}}, \hat{S}(t) + z_{\alpha/2}\hat{S}(t)\sqrt{\widehat{\mathrm{var}}\{\log \hat{S}(t)\}}\right].$$

The main problem with this is the fact that the upper and lower bounds may be larger than 1 and smaller than 0, respectively. As a way out, one considers confidence bands for the statistic $\log\{-\log(\hat{S})\}$ (that has the range $\mathbb{R}$) and obtains again by the delta-method that

$$\mathrm{var}[\log\{-\log \hat{S}(t)\}] \approx \frac{1}{\log^2 \hat{S}(t)}\widehat{\mathrm{var}}(\log \hat{S}).$$

Setting $B^{\pm} = \log(-\log \hat{S}(t)) \pm z_{\alpha/2} \log^{-1} \hat{S}(t)\sqrt{\widehat{\mathrm{var}}(\log \hat{S})}$ an approximate confidence band (contained in $[0,1]$) is given by

$$\left[\exp(-\exp(B^-), \exp(-\exp(B^+)\right].$$

As an example we obtain for the Kaplan-Meier estimator by Greenwood's formula

$$\widehat{\mathrm{var}}\{\log \hat{S}_{\mathrm{KM}}(t)\} \approx \sum_{\{j:\tau_j<t\}} \frac{d_j}{r_j(r_j-d_j)}.$$

Often we wish to decide whether or not two (or more) samples stem from the same survivor function or not. The **log-rank test** is such a simple test procedure. We will now assume that the failure times $\tau_1 < \tau_2 < \cdots < \tau_k$ are realizations of *two* random variables T_1 and T_2 corresponding to two groups of items (patients). For each observed failure time τ_j we consider the contingency table

Groups	failure at time τ_j	items at risk at time τ_j
1	d_{1j}	r_{1j}
2	d_{2j}	r_{2j}
$1+2$	d_j	r_j

Under the null-hypothesis that $T_1 = T_2$ the expected number of failures at time τ_j in group 1 and 2 are hypergeometrically distributed with parameters r_j, r_{1j}, d_j and r_j, r_{2j}, d_j, respectively. Thus, mean and variance of the number of failures in group 1 and 2 can be computed as

$$e_{1j} = \frac{d_j}{r_j} r_{1j} \quad \text{and} \quad e_{2j} = \frac{d_j}{r_j} r_{2j},$$

and

$$v_{1j} = v_{2j} = \frac{d_j r_{1j} r_{2j} (r_j - d_j)}{r_j^2 (r_j - 1)}.$$

Under the null-hypothesis, the statistic

$$\chi^2 = \frac{\left[\sum_{j=1}^k (d_{1j} - e_{1j})\right]^2}{\sum_{j=1}^k v_{1j}}$$

is χ^2-distributed with 1 degree of freedom.

An important question in practice is: Is the assumption of a Weibull distributed survivor time justified? An indication to the answer of this question can be obtained as follows: Observe that under the assumption that T is Weibull distributed one has

$$\log\{-\log S(t)\} = \alpha(\log t + \log \lambda), \quad t > 0.$$

Now let $\hat{S}(t)$ be a nonparametric estimate for S (e.g. the Kaplan-Meier estimator $\hat{S}_{\mathrm{KM}}$). Then the plot $\log\{-\log \hat{S}(t)\}$ against $\log t$ should approximately be a straight line with slope α and intercept $-\log \lambda$.