STAT24410 NOTES

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Contents

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1. Probability

1.1. CDF.

1.1.1. *Properties of CDF.*

- Nondecreasing.
- Right continuous.
- $\lim_{x\to-\infty} F(x) = 0$, $\lim_{x\to\infty} F(x) = 1$.

1.1.2. *Inverse of CDF.*

$$F^{-}(x) \coloneqq \inf\{u : x \le F(u)\}.$$

Proposition 1.1. *Let F be the cdf of X. If F is continuous and strictly increasing, then $Y \coloneqq F(X) \sim \mathrm{Uniform}[0,1]$.*

Proof. For any $y \in [0,1]$,

$$\mathbb{P}(F(X) \le y) = F(F^{-1}(y)) = y.$$

□

Proposition 1.2. *Let $U \sim \mathrm{Uniform}[0,1]$ and X be the cdf of X. Then $F^{-1}(U) \overset{\mathcal{D}}{=} X$.*

Proof. For any $x \in [0,1]$,

$$\mathbb{P}(F^{-1}(U) \le x) = \mathbb{P}(U \le F(x)) = F(x).$$

□

Remark 1.3. This is useful for simulation.

1.2. Transformations.

For $Y \coloneqq h(X)$, if h is one-to-one and differentiable, then

$$f_Y(y) = f_X(h^{-1}(y)) \cdot \left|\frac{\mathrm{d}h^{-1}(y)}{\mathrm{d}y}\right|.$$

1.3. Expectation.

For an r.v. X. We define

$$X^+ = \max\{X, 0\}, \quad X^- = \max\{-X, 0\}.$$

Note that $X \equiv X^+ - X^-$.

Since X^+ is nonnegative,

$$\mathbb{E}(X^+) \coloneqq \int_0^\infty x \,\mathrm{d}F(x)$$

in the Riemann–Stieltjes sense, and similarly X^-.

Definition 1.4. X has expected value if at least one of $\mathbb{E}(X^+)$ and $\mathbb{E}(X^-)$ is finite, and when it does

$$\mathbb{E}(X) \coloneqq \mathbb{E}(X^+) - \mathbb{E}(X^-).$$

Definition 1.5. We say Y **stochastically dominates** X, $Y \succeq X$, if

$$\mathbb{P}(X > t) \le \mathbb{P}(Y > t), \quad \forall t.$$

Proposition 1.6.

- *Linearity.*
- *If*
$$\int_{\mathbb{R}} |x| f(x)\,\mathrm{d}x < \infty$$
then
$$\mathbb{E}(X) = \int_{\mathbb{R}} x f(x)\,\mathrm{d}x.$$
- *If X is stochastically dominated by Y then $\mathbb{E}(X) \le \mathbb{E}(Y)$.*
- *If X and Y are independent, then $\mathbb{E}(XY) = \mathbb{E}(X)\,\mathbb{E}(Y)$.*

Definition 1.7. The **variance** of X is given by
$$\operatorname{Var}(X) \coloneqq \mathbb{E}[(X - \mathbb{E}(X))^2]$$

Proposition 1.8.

- $\operatorname{Var}(X) = \mathbb{E}(X^2) - (\mathbb{E}(X))^2$.
- $\operatorname{Var}(cX) = c^2 \operatorname{Var}(X)$.
- *If X and Y are independent, then* $\operatorname{Var}(X+Y) = \operatorname{Var}(X) + \operatorname{Var}(Y)$.

Proposition 1.9. *If $X \ge 0$ and there exists an at most countable subset $S = \{x_1, x_2, \dots\}$ of isolated points such that F_X is continuously differentiable on $[0,\infty) \setminus S$, then*
$$\mathbb{E}(X) = \sum_{x \in S} x\mathbb{P}(X = x) + \int_0^\infty x F_X'(x)\,\mathrm{d}x.$$

1.4. **Probability Inequalities.**

Theorem 1.10 (Markov's Inequality)**.** *If $X \ge 0$ and $c > 0$, then*
$$\mathbb{P}(X \ge c) \le \frac{\mathbb{E}(X)}{c}.$$
(Equality is attained when $\mathbb{P}(X = 0$ or $X = c) = 1$.)

Proof. Construct
$$Y \coloneqq \begin{cases} c, & x \ge 0 \\ 0, & X < c. \end{cases}$$
Then $Y \le X$ and
$$\mathbb{E}(Y) = c \cdot \mathbb{P}(X \ge c) \le \mathbb{E}(X).$$
□

Theorem 1.11 (Chebychev's Inequality)**.** *If $c > 0$, then*
$$\mathbb{P}(|X - \mu| \ge c) \le \frac{\mathbb{E}[(X-\mu)^2]}{c^2}$$
for any μ.

Proof. Apply Markov's inequality to $(X - \mu)^2$. □

Theorem 1.12 (Chernoff's Inequaality)**.** *If $c \in \mathbb{R}$ and $t > 0$, then*

$$\mathbb{P}(X \geq c) \leq e^{-tc}\, \mathbb{E}(e^{tX})$$

and

$$\mathbb{P}(X \leq c) \leq e^{tc}\, \mathbb{E}(e^{-tX}).$$

Proof. Apply Markov's inequality to e^{tX} and e^{-tX}. □

Theorem 1.13 (Weak Law of Large Numbers)**.** *Let $X_1, X_2, \ldots$ be i.i.d. with finite expectation μ and variance σ^2. Then as n goes to infinity,*

$$\mathbb{P}\left[\left|\overline{X_n} - \mu\right| > \epsilon\right] \longrightarrow 0.$$

That is, $\overline{X_n} \xrightarrow{\text{p}} \mu$.

Proof. Note that $\mathbb{E}(\overline{X_n}) = \mu$ and $\mathrm{Var}(\overline{X_n}) = \sigma^2/n$. Chebyshev's gives

$$\mathbb{P}\left(\left|\overline{X_n} - \mu\right| < \epsilon\right) \leq \frac{\sigma^2}{n \cdot \epsilon^2} \longrightarrow 0$$

as $n \longrightarrow \infty$. □

Proposition 1.14 (Large Deviations)**.** *Let $X_1, X_2, \ldots$ be i.i.d. with finite expectation μ and variance σ^2. Let $c > \mu$. Then*

$$\lim_{n\to\infty} \frac{1}{n} \log \mathbb{P}(\overline{X_n} > c) = -\sup_t [tc - \kappa(t)],$$

where $\kappa(t) = \log \mathbb{E}(e^{tX})$.

We do not yet have the tools to prove that this is the limit, but we will use Chernoff's inequality to obtain a bound:

Proof. From Chernoff's inequality, for any t we have

$$\mathbb{P}(\overline{X_n} \geq c) = \mathbb{P}\left(\sum X_i \geq c \cdot n\right) \leq e^{-tnc}\, \mathbb{E}\left[e^{t(\sum Xi)}\right] = e^{-tnc + n\kappa(t)},$$

where $\kappa(t) = \log \mathbb{E}(e^{tX})$. Thus we have

$$\frac{1}{n} \log \mathbb{P}(\overline{X_n} \geq c) \leq -\sup_t [tc - \kappa(t)].$$

□

Remark 1.15.

- $\mathbb{E}[e^{tX}]$ is the **moment generating function**.
- $\kappa(t)$ is the **cumulant generating function**.
- $\sup_t [tc - \kappa(t)]$ is the **Legendre Transform**.

Definition 1.16. X_n **converges in distribution** to X, $X_n \xrightarrow{\mathscr{D}} X$, if

$$F_{X_n}(x) \longrightarrow F_X(x), \quad \forall x \in \mathbb{R}.$$

Definition 1.17. The **moment generating function** of X is

$$\begin{aligned} M : \mathbb{R} &\longrightarrow [0, \infty] \\ t &\longmapsto \mathbb{E}[e^{tX}]. \end{aligned}$$

Proposition 1.18. *Properties of the moment generating function:*

- $\mathbb{E}[X^m] = M_X^{(n)}(0)$ *when Fubini grants*
$$\mathbb{E}\left[e^{tX}\right] = \mathbb{E}\left[\sum_{n=0}^{\infty}\frac{(tX)^n}{n!}\right] = \sum_{n=0}^{\infty}\frac{t^n\,\mathbb{E}(X^n)}{n!}.$$
- $M_{cX}(t) = M_X(ct)$.
- *If* X *and* Y *are independent, then*
$$M_{X+Y}(t) = M_X(t) + M_Y(t).$$
- *If* $X_1, X_2, \ldots$ *are i.i.d., then*
$$M_{\sum X_i} = \prod M_{X_i}.$$
- $X_n \xrightarrow{\mathscr{D}} X$ *if and only if* $M_{X_n} \to M_X$ *in a neighborhood of* 0.

Theorem 1.19 (Central Limit Theorem)**.** *If* $X_1, X_2, \ldots$ *are i.i.d.,* $\mathbb{E}(X_i) = \mu$*, and* $\mathrm{Var}(X_i) = \sigma^2$*, then*
$$\frac{\sum X_i}{\sqrt{n}} \xrightarrow{\mathscr{D}} \mathcal{N}(\mu, \sigma^2).$$
Or, equivalently,
$$\sqrt{n}\cdot \overline{X}_n \xrightarrow{\mathscr{D}} \mathcal{N}(\mu, \sigma^2).$$

The following proof works only when we have enough regularity; it is meant to provide a certain intuition (the general proof needs complex analysis):

Proof. We consider the mgf.
$$M_{\sum X_i/\sqrt{n}}(t) = M_{\sum X_i}\left(\frac{t}{\sqrt{n}}\right) = \left[M_{X_i}(\frac{t}{\sqrt{n}})\right]^n.$$
We obtain an approximation though Taylor:
$$M_X(\frac{t}{\sqrt{n}}) \approx M_X(0) + \frac{t}{\sqrt{n}}M_X'(0) + \frac{t^2}{n}M_X''(0)$$
Noting that $M_X'(0) = \mathbb{E}[X] = 0$ and $M_X''(0) = \mathbb{E}[X^2] = \sigma^2$, we have
$$M_{\sum X_i/\sqrt{n}}(t) \approx \left[1 + \frac{t^2\sigma^2}{n}\right]^n \longrightarrow e^{t^2\sigma^2}.$$
The last term is precisely the mgf of $N(0, \sigma^2)$. □

2. Joint Distribution

2.1. Random Vectors and Joint Distributions.

Proposition 2.1.

- $$F(x) = \int_{-\infty}^{x_1} \cdots \int_{-\infty}^{x_n} f(x)\, \mathrm{d}x.$$
- *If F is continuous and differentiable, then X has density*
$$f(X) = \frac{\partial^n F(x)}{\partial x_1 \dots \partial x_n}.$$
- *If $X_1, X_2, \dots, X_n$ are independent, then*
$$F_X(x) = F_{X_1}(x_1) \dots F_{X_n}(x_n).$$
- *If F is differentiable, then*
$$f_X(x) = f_{X_1}(x_1) \dots f_{X_n}(x_n),$$
and conversely!
- *If $X = (X_1, X_2, \dots, X_n)$ has density f_X, then X_I has density*
$$f_I(x_I) = \int_{\mathbb{R}^{n-|I|}} f\left(x_I, x_{S_n \setminus I}\right)\, \mathrm{d}x_{S_n \setminus I},$$
where $S_n \coloneqq \{1, 2, \dots, n\}$ are all the indices. Think "integrating out" the other variables.

2.2. Transformations.

Definition 2.2. The **Jacobian** of $g : G \to H \subset \mathbb{R}^n$, where G and H are open, is given by
$$J_g(y) \coloneqq \det \left[\frac{\partial g_i}{\partial y_j}\right].$$

If $X : \Omega \to H \subset \mathbb{R}^n$ and $h : H \to G \subset \mathbb{R}^n$, where H and F are open, are such that h is one-to-one and differentiable and $h^{-1} : G \to H$ is differentiable. Then $Y \coloneqq h(X)$ has density
$$f_Y(y) = \begin{cases} f_X(h^{-1}(y)) \cdot \left|Jh^{-1}(y)\right|, & y \in G \\ 0, & y \notin G. \end{cases}$$

Definition 2.3. The Gamma function is given by
$$\Gamma(\lambda) \coloneqq \int_0^\infty e^{-x} x^{\lambda - 1}\, \mathrm{d}x.$$

Proposition 2.4. *Properties:*

- $\Gamma(1) = 1$.
- $\Gamma(1/2) = \sqrt{\pi}$.
- $\Gamma(x+1) = x\Gamma(x)$.
- $\Gamma(n) = (n-1)!$ *for any* $n \in \mathbb{N}$.

2.3. **Conditional distribution.** The continuous case:

Definition 2.5. We define the **conditional density** as

$$f_{X|Y}(x|y) \coloneqq \frac{f_{X|Y}(x,y)}{f_Y(y)},$$

2.4. **Covariance and Correlation.**

Definition 2.6. The **covariance** of random variables X and Y is

$$\operatorname{Cov}(X,Y) = \mathbb{E}\left((X-\mu_X)\cdot(Y-\mu_Y)\right).$$

Their **correlation** is given by

$$\operatorname{Corr}(X,Y) = \frac{\operatorname{Cov}(X,Y)}{\sigma_X\sigma_Y}.$$

Proposition 2.7. *Properties:*

- $\operatorname{Var}(a+bX) = b^2\operatorname{Var}(X)$.
- $\operatorname{Cov}(a+bX, c+dY) = bd\operatorname{Cov}(X,Y)$.
- $\operatorname{Var}(X+Y) = \operatorname{Var}(X) + \operatorname{Var}(Y) + 2\operatorname{Cov}(X,Y)$.
- *If X and Y are independent, then* $\operatorname{Cov}(X,Y) = 0$. *But the converse is not true. For example, if $Z \sim N(0,1)$, and S and T are random signs (1 or -1), then* $\operatorname{Cov}(SZ,TZ) = 0$.

Theorem 2.8.

- *If (X,Y) has density f, then $X|Y$ has density*
$$\frac{f(x,y)}{f_Y(y)}.$$
- *If (X,Y) has a pmf, then $X|Y$ is discrete with pmf*
$$\frac{p(x,y)}{p_Y(y)}.$$

Note that $E(X|Y=y)$ is a number, and $\mathbb{E}(X|Y)$ is a random variable.

Proposition 2.9.

(i) If X and Y are independent, then
$$\mathbb{E}(X|Y) = \mathbb{E}(X) \quad \textit{with probability 1}.$$
(ii) Law of total expectation / Tower law:
$$\mathbb{E}[\mathbb{E}[X|Y]] = \mathbb{E}[X].$$
(iii)
$$\mathbb{E}[g(X)h(Y)|Y] = h(Y)\,\mathbb{E}(g(X)|Y) \quad \textit{with probability 1}.$$
And
$$\mathbb{E}[X|T(Y)] = \mathbb{E}[\mathbb{E}[X|T(Y)|Y]] \quad \textit{with probability 1}.$$

(iv) *Law of total variations*

$$\mathrm{Var}(Y) = \mathbb{E}[\mathrm{Var}(Y|X)] + \mathrm{Var}[\mathbb{E}(Y|X)],$$

where

$$\mathrm{Var}(Y|X) \coloneqq \mathbb{E}(Y^2|X) - (\mathbb{E}(Y|X))^2.$$

2.5. **Rejection Sampling.** If for some constant c we have

$$h(x) \geq c \cdot f(x), \quad \forall x,$$

then we can obtain a sample from distribution with density f using samples from distribution with density h using **rejection sampling**:

(1) Sample Y from g and U from Uniform$(0, 1)$, with Y and U independent.
(2) Set $X \coloneqq Y$ if

$$U \leq \frac{c \cdot f(Y)}{h(Y)}$$

and return to (1) otherwise.

Remark 2.10.

- Think sampling on the area under f (as a subset of the area under g).
- Rejection sampling can also be used if

$$f(x) = \frac{g(x)}{N},$$

where N is an unknown constant (e.g., an integral with numerical approximations but no closed form solutions). We need only find h such that

$$h(x) \geq cN \cdot g(x).$$

Think

$$h(x) \gg g(x).$$

3. Statistical Inference

Example 3.1. Modeling lifetime $T : \Omega \to [0, \infty)$.

Definition 3.2.

- The **survival** function is defined as
$$\begin{aligned} S : [0,\infty) &\longrightarrow [0,1] \\ x &\longmapsto \mathbb{P}(T > x) = 1 - F_Y(x). \end{aligned}$$
- The **failure rate** function is defined as
$$h(x) \coloneqq \frac{f(x)}{S(x)}.$$

Remark 3.3.
$$\mathbb{P}(T \leq x+\Delta x | T > x) = \frac{\mathbb{P}[x < T \leq x + \Delta x]}{\mathbb{P}[T > x]} = \frac{F(x+\Delta x) - F(x)}{S(x)} \approx \Delta x \cdot \frac{f(x)}{S(x)} = \Delta x \cdot h(x).$$
Think of an increasing failure rate as "aging."

Given h we can recover f:
$$h(x) = \frac{f(x)}{1 - F(x)} = -\frac{\partial}{\partial x} \log(1 - F(x)).$$
So,
$$\log(1 - F(x)) = -\int_0^x h(t)\mathrm{d}t + C.$$
Since $F(0) = 0$ we know $C = 0$. We have
$$s(x) = \exp\left(-\int_0^x h\right)$$
and
$$f(x) = h(x)\exp\left(-\int_0^x h\right).$$

Example 3.4.

- If $h(x) = \lambda$ is a constant function, we have $T \sim \text{Exponential}(\lambda)$:
$$f(x) = \lambda \exp\left(-\int_0^x \lambda \mathrm{d}t\right) = \lambda \exp(-\lambda x), \quad \forall x > 0.$$
- If $h(x) = \alpha + \beta x$ with $\alpha, \beta > 0$, then T follows the Gompertz distribution.
- If $h(x) = \lambda\beta x^{\beta - 1}$, then T follows the Weibull distribution.

3.1. **Estimating parameters.**

We next assume $T_1, T_2, \ldots \overset{\text{iid}}{\sim} \text{Exponential}(\lambda)$ and estimate λ.

Remark 3.5. Metrics to evaluate an estimator:

- Bias: $\mathbb{E}(\hat{\lambda}) - \lambda$.
- Variance: $\text{Var}[\hat{\lambda}]$.
- Mean Squared Error: $\text{MSE}[\hat{\lambda}] = \mathbb{E}[(\hat{\lambda} - \lambda)^2] = \text{Bias}^2 + \text{Variance}$.

3.1.1. *Asymptotic Estimation.*

Definition 3.6 (Method of Moments)**.** Let $X_1, X_2, \ldots, X_n \overset{\text{iid}}{\sim} F$ with n parameters. To estimate the parameters, we equate n (usually the first n) theoretical moments to the n corresponding sample moments:

$$\mathbb{E}[X^k] = \frac{1}{n}\sum X_i^k, \quad 1 \le k \le n.$$

Example 3.7. Consider $T_n \overset{\text{iid}}{\sim} \text{Exponential}(\lambda)$.

- Since $\mathbb{E}(\overline{T}_n) = 1/\lambda$, we may use $\hat{\lambda} \coloneqq 1/\overline{T}_n$ as an estimator for λ.
- Since
$$\mathbb{E}\left[\sum T_i^2/n\right] = \frac{2}{\lambda^2},$$
we may also use
$$\hat{\lambda}_2 = \sqrt{\frac{2n}{\sum T_i^2}}$$
as an estimator.

Example 3.8.

- Consider $X_1, X_2, \ldots, X_n \overset{\text{iid}}{\sim} \text{Uniform}[0, \theta]$. We have $\mathbb{E}[X] = \theta/2$.
$$\hat{\theta} \coloneqq 2\hat{X}.$$
- Consider $X_1, X_2, \ldots, X_n \overset{\text{iid}}{\sim} \text{Gamma}(\alpha, \beta)$. We have $\mathbb{E}[X] = \alpha/\beta$ and $\mathbb{E}[X^2] = \alpha/\beta^2 + (\alpha/\beta)^2$. Thus we solve
$$\frac{\hat{\alpha}}{\hat{\beta}} = \overline{X}, \quad \frac{\hat{\alpha}}{\hat{\beta}^2} + \frac{\hat{\alpha}^2}{\hat{\beta}^2} = \frac{\sum X_i^2}{n}.$$

The following theorems help us characterize these estimators.

Theorem 3.9 (Continuous mapping theorem)**.**

(i) if $X_n \xrightarrow{\text{p}} X$ and g is continuous, then $g(X_n) \xrightarrow{\text{p}} g(X)$.
(ii) If $X_n \xrightarrow{\mathscr{D}} X$ and g is continuous, then $g(X_n) \xrightarrow{\mathscr{D}} g(X)$.

Lemma 3.10 (Slutsky)**.** *If $X_n \xrightarrow{\mathscr{D}} X$ and $Y_n \xrightarrow{\text{p}} c$, where c is a constant, then*

$$X_n + Y_n \xrightarrow{\mathscr{D}} X + c, \quad X_n Y_n \xrightarrow{\mathscr{D}} cX.$$

Theorem 3.11 (Delta Method)**.** *If X_n is such that*

$$\sqrt{n}(X_n - \theta) \xrightarrow{\mathscr{D}} \mathcal{N}(0, \sigma^2)$$

and g is continuously differentiable, then

$$\sqrt{n}(g(X_n) - g(\theta)) \xrightarrow{\mathscr{D}} \mathcal{N}(0, \sigma^2[g'(\theta)]^2).$$

Remark 3.12. Intuition: We can write

$$\sqrt{n}(g(X_n) - g(\theta)) = g'(\tilde{\theta}_n)\sqrt{n}(X_n - \theta), \quad \tilde{\theta}_n \in (x_n, \theta).$$

We know that $g'(\tilde{\theta}_n) \xrightarrow{\text{p}} g'(\theta)$ and $\sqrt{n}(X_n - \theta) \xrightarrow{\mathscr{D}} \mathcal{N}(0, \sigma^2)$, so Slutsky's gives the desired result.

We can now characterize estimators obtained from the method of moments:

Proposition 3.13.

- *Non-uniqueness: we can obtain multiple estimators.*
- *Consistency: Law of Large Numbers gives*

$$\overline{X} \xrightarrow{\text{p}} \mathbb{E}[X],$$

  and the continuous mapping theorem then gives consistency (under certain conditions).
- *Asymptotic normality: the Delta Method gives normality (under certain conditions).*

3.1.2. *Estimators for Smaller n.* We can obtain the exact distribution of $\overline{T}_n$. Since

$$T \overset{\text{iid}}{\sim} \text{Exponential}(\lambda) = \text{Gamma}(1, \lambda),$$

we know by the properties of gamma distributions that

$$\sum T_i \sim \text{Gamma}(n, \lambda).$$

Again by properties of gamma distributions, we know that the estimator $\hat{\lambda}_1 \coloneqq 1/\overline{T}_n$ is biased for small n:

$$\mathbb{E}[\hat{\lambda}_1] = n \cdot \mathbb{E}\left[\frac{1}{\sum T_i}\right] = \frac{n\lambda}{n-1}.$$

The estimator

$$\hat{\lambda}_3 \coloneqq \frac{n-1}{n}\hat{\lambda}_1,$$

then, is unbiased and has smaller variance.

Remark 3.14. This is a rare case. Oftentimes, we have instead a tread off between bias and variance.

3.2. **Maximum Likelihood Estimator.** The above may be summed up as the following steps:

- Estimators
- Evaluations
- Distribution for estimators (which allows for the construction of probabilistic statements)

Maximum Likelihood estimator accomplishes all the above in a streamlined fashion.

Definition 3.15. Let $X_1, X_2, \ldots, X_n \overset{\text{iid}}{\sim} F_\theta$, where $\theta \in \mathbb{R}^k$ is a parameter for the distribution. Let $f(x, \theta)$[1] be the density or pmf of F_θ. The **Likelihood** of θ given observations $X_1, X_2, \ldots, X_n$ is

$$L(\theta) = L_n(\theta) \coloneqq \prod_{i=1}^{n} f(X_i, \theta).$$

The **maximum likelihood estimator** is the value at which L obtains its maximum:

$$\hat{\theta} = \hat{\theta}_n \coloneqq \arg\max_{\theta} L(\theta).$$

Remark 3.16. It is often easier to work with the **log likelihood**:

$$\ell(\theta) = \ell_n(\theta) \coloneqq \log L(\theta).$$

Remark 3.17.

- Might be non-unique. Consider $X_1, X_2, \ldots, X_n \overset{\text{iid}}{\sim} \text{Uniform}(\theta, \theta + 1)$.
- Might not exist. Consider $X_1, X_2, \ldots, X_n$ iid with density

  $$f(x, \mu, \sigma^2) = \frac{1}{2}\left[\frac{1}{\sqrt{2\pi}} e^{-x^2/2} + \frac{1}{\sigma\sqrt{2\pi}} e^{-\frac{(x-\mu)^2}{2\sigma^2}}\right].$$

  Think $X \sim \mathcal{N}(0, 1)$ with probability $1/2$ and $X \sim \mathcal{N}(\mu, \sigma^2)$ with probability $1/2$. The likelihood function is unbounded:

  $$\max_{\mu, \sigma^2} L(\mu, \sigma^2) \geq \max_{\sigma} L(X_1, \sigma^2) \geq \frac{1}{2^n}\left[\frac{1}{\sqrt{2\pi}\sigma}\right] \prod_{k=1}^{n} e^{-X_1^2/2}.$$

Theorem 3.18 (Cramér–Rao Inequality)**.** *Let $T(X_n)$ be any unbiased estimator for $g(\theta)$. Then*

$$\text{Var}[T(X_n)] \geq \frac{[g'(\theta)]^2}{nI(\theta)}.$$

Proposition 3.19. *Properties of maximal likelihood estimators:*

- *consistency,*
- *asymptotic normality,*
- *has known and optimal asymptotic variance (**efficiency**). That is, it attains the Cramér–Rao bound.*
- *Invariance: if $\hat{\theta}$ is the estimator of θ, then for any function f, $f(\hat{\theta})$ is the estimator of $f(\theta)$.*

[1]Some also write $f_\theta(x)$ or $f(x|\theta)$.

Appendix A: Common Distributions

3.3. **Exponential.** $X \sim \text{Exponential}(\lambda)$, $\lambda > 0$.

- Support: $[0, \infty)$
- pdf: $\lambda e^{\lambda x}$
- cdf: $1 - e^{\lambda x}$

Definition 3.20. If $X \sim \text{Gamma}(\alpha, \beta)$ and has a density, then

$$f(x) = \frac{\beta^\alpha}{\Gamma(\alpha)} x^{\alpha-1} e^{-\beta x}, \quad x > 0.$$

Proposition 3.21.

- $$\mathbb{E}(X) = \frac{\alpha}{\beta}, \quad \text{Var}(X) = \frac{\alpha}{\beta^2}.$$

Distribution	**Support**	**PMF**	**Mean**	**Variance**
Binomial(n, p)	$\{0, 1, 2, \dots, n\}$	$\binom{n}{x} p^x (1-p)^{n-x}$	np	$np(1-p)$
Geometric(p)	$\{1, 2, 3, \dots\}$	$(1-p)^{x-1} p$	$\frac{1}{p}$	$\frac{1-p}{p^2}$
Poisson(λ)	$\{0, 1, 2, \dots\}$	$\frac{\lambda^x e^{-\lambda}}{x!}$	λ	λ

Table 1. Key Properties of Discrete Distributions

Distribution	**Support**	**PDF**	**Mean**	**Variance**
Uniform(a, b)	$[a, b]$	$\frac{1}{b-a}$	$\frac{a+b}{2}$	$\frac{(b-a)^2}{12}$
$\mathcal{N}(\mu, \sigma^2)$	$(-\infty, \infty)$	$\frac{1}{\sqrt{2\pi\sigma^2}} e^{-\frac{(x-\mu)^2}{2\sigma^2}}$	μ	σ^2
Exponential(λ)	$[0, \infty)$	$\lambda e^{-\lambda x}$	$\frac{1}{\lambda}$	$\frac{1}{\lambda^2}$
Gamma(α, β)	$(0, \infty)$	$\frac{\beta^\alpha x^{\alpha-1} e^{-\beta x}}{\Gamma(\alpha)}$	$\frac{\alpha}{\beta}$	$\frac{\alpha}{\beta^2}$
Beta(α, β)	$(0, 1)$	$\frac{x^{\alpha-1}(1-x)^{\beta-1}}{B(\alpha, \beta)}$	$\frac{\alpha}{\alpha+\beta}$	$\frac{\alpha\beta}{(\alpha+\beta)^2(\alpha+\beta+1)}$

Table 2. Key Properties of Continuous Distributions

Proposition 3.22. *Properties of Exponential distribution:*

(i) The "memoryless" property:

$$\mathbb{P}(T \leq x + y | T > x) = \mathbb{P}(T \leq y).$$

Proposition 3.23. *Properties of Gamma distribution:*

(i) If $X_i \overset{\text{iid}}{\sim} \text{Gamma}(\alpha_i, \beta)$ *for* $i = 1, 2, \ldots, N$*, then*

$$\sum X_i \sim \text{Gamma}\left(\sum \alpha_i, \beta\right).$$

(ii) If $X \sim \text{Gamma}(\alpha, \beta)$ *and* $\alpha > 1$*, then*

$$\mathbb{E}\left[1/X\right] = \frac{\beta}{\alpha - 1}.$$

Proof.

(i) Note that

$$\mathbb{E}\left[e^{tX_i}\right] = \left(1 - \frac{t}{\beta}\right)^{-\alpha_i}, \quad \forall t < \beta.$$

We then have

$$M_{\sum X_i}(t) = \prod M_{X_i}(t) = \left(1 - \frac{t}{\beta}\right)^{-\sum \alpha_i}.$$

(ii) We have

$$\mathbb{E}[1/X] = \int_0^\infty \frac{1}{x} \cdot \frac{\beta^\alpha}{\Gamma(\alpha)} \cdot x^{\alpha-1} e^{-\beta x} \, \mathrm{d}x,$$

which we can integrate by reducing to the Γ function.

$\square$