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.*

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

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

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

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

□

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

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

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

□

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

$$\mathrm{Var}(X) \coloneqq \mathbb{E}[(X - \mathbb{E}(X))^2]$$

Proposition 1.8.

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

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 $Y \leq X$ and

$$\mathbb{E}(Y) = c \cdot \mathbb{P}(X \geq c) \leq \mathbb{E}(X).$$

□

Theorem 1.11 (Chebychev's Inequality). *If $c > 0$, then*

$$\mathbb{P}(|X - \mu| \geq c) \leq \frac{\mathbb{E}[(X-\mu)^2]}{c^2}$$

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, $\overline{X}_n \xrightarrow{\text{p}} \mu$. That is*

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

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 $n \to \infty$. □

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*

$$\sqrt{n}(\overline{X}_n - \mu) \xrightarrow{\mathscr{D}} \mathcal{N}(0, \sigma^2).$$

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 assume $\mu = 0$ and consider the mgf.

$$M_{\sum X_i/\sqrt{n}}(t) = M_{\sum X_i}\left(\frac{t}{\sqrt{n}}\right) = \left[M_{X_i}\left(\frac{t}{\sqrt{n}}\right)\right]^n.$$

We obtain an approximation though Taylor:

$$M_X\left(\frac{t}{\sqrt{n}}\right) \approx M_X(0) + \frac{t}{\sqrt{n}} M_X'(0) + \frac{t^2}{n} M_X''(0)$$

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 $N(0, \sigma^2)$. $\square$

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*
$$f(X) = \frac{\partial^n F(x)}{\partial x_1 \ldots \partial x_n}.$$
- *If $X_1, X_2, \ldots, X_n$ are independent, then*
$$F_X(x) = F_{X_1}(x_1) \ldots F_{X_n}(x_n).$$
- *If F is differentiable, then*
$$f_X(x) = f_{X_1}(x_1) \ldots f_{X_n}(x_n),$$
and conversely!
- *If $X = (X_1, X_2, \ldots, X_n)$ 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 $S_n := \{1, 2, \ldots, n\}$ 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
$$J_g(y) := \det\left[\frac{\partial g_i}{\partial y_j}\right].$$

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 $Y := h(X)$ 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
$$\Gamma(\lambda) := \int_0^\infty e^{-x} x^{\lambda-1}\, \mathrm{d}x.$$

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

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

Their **correlation** is given by

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

Proposition 2.7. *Properties:*

- $\mathrm{Var}(a+bX) = b^2\,\mathrm{Var}(X)$.
- $\mathrm{Cov}(a+bX, c+dY) = bd\,\mathrm{Cov}(X,Y)$.
- $\mathrm{Var}(X+Y) = \mathrm{Var}(X) + \mathrm{Var}(Y) + 2\,\mathrm{Cov}(X,Y)$.
- *If X and Y are independent, then* $\mathrm{Cov}(X,Y) = 0$. *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* $\mathrm{Cov}(SZ,TZ) = 0$.

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. Point Estimates

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

Definition 3.6. An estimator $\hat{\theta}_n$ of θ is said to be **consistent** if

$$\hat{\theta}_n \xrightarrow{\text{p}} \theta.$$

That is, if for any $\epsilon > 0$,

$$\lim_{n\to\infty} \mathbb{P}(\left|\hat{\theta}_n - \theta\right| > \epsilon) = 0.$$

3.1.1. *Asymptotic Estimation.*

Definition 3.7 (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 3.8. 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 3.9.

- 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 3.10 (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 3.11 (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 3.12 (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 3.13. 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 3.14.

- *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 $\hat{\lambda}_1 := 1/\overline{T}_n$ 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

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

then, is unbiased and has smaller variance.

Remark 3.15. 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 3.16. 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

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

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

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

Remark 3.17. It is often easier to work with the **log likelihood**:

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

Remark 3.18.

- 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:

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

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

3.3. **Likelihood Theory.**

Definition 3.19 (**Score Function**).

$$\dot{\ell}_n(\theta) \coloneqq \frac{\partial}{\partial\theta}\ell_n(\theta) = \sum_{i=1}^{n} \frac{\frac{\partial f}{\partial\theta}(x_i,\theta)}{f(x_i,\theta)} = \sum_{i=1}^{n} \frac{f'(x_i,\theta)}{f(x_i,\theta)}.$$

Remark 3.20. We find the MLE by setting the score function to 0.

Proposition 3.21. *If $f(x,\theta)$ has common support, that is, if $\{x : f(x,\theta) > 0\}$ doe s not depend on θ, then*

$$\mathbb{E}_{\theta_0}\left[\frac{L_n(\theta)}{L_n(\theta_0)}\right] = 1.$$

Equivalently,

$$\mathbb{E}\left[\exp\left(\ell_n(\theta) - \ell_n(\theta_0)\right)\right] = 1.$$

Proposition 3.22. *If the density functions are smooth, then*

(a) $\mathbb{E}_\theta\left[\dot{\ell}_n(\theta)\right] = 0$.

(b) $-\mathbb{E}_\theta\left[\ddot{\ell}_n(\theta)\right] = \mathbb{E}\left[\dot{\ell}_n(\theta)^2\right]$.

Definition 3.23 (**Fisher Information**).

$$I(\theta) \coloneqq \mathbb{E}_\theta[\dot{\ell}(\theta)^2] = \mathbb{E}_\theta[-\ddot{\ell}(\theta)].$$

That is,

$$I(\theta) \coloneqq \mathbb{E}\left[\left(\frac{\partial}{\partial\theta}\log f(X,\theta)\right)^2\right] = -\mathbb{E}\left[\frac{\partial^2}{\partial\theta^2}\log f(X,\theta)\right],$$

where the expectation is taken with respect to $X \sim f(x,\theta)$.

Remark 3.24. Intuitively, the more variation there is in the density functions $f(x,\theta)$ as we vary θ, the more information we can get from data. Fisher information measures the variation in density functions by looking at its derivative.

Theorem 3.25 (Cramér–Rao Inequality)**.** *Let $T(X_n)$ be any unbiased estimator for $g(\theta)$. Then*

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

Remark 3.26. The Cramér–Rao lower bound is attained if and only if

$$\mathrm{Corr}(\hat{\theta}(X), \dot{\ell}(X)) = 1.$$

By Cauchy-Schwarz inequality, this is equivalent to $\hat{\theta}(X)$ and $\dot{\ell}(X)$ being linearly related random variables. That is,

$$\dot{\ell}(\theta) = \alpha(\theta)\hat{\theta}(X) + \beta(\theta)$$

for functions α and β that do not depend on X.

Proposition 3.27. *Under the regularity conditions in the Cramér–Rao inequality, there exists an unbiased estimator $\hat{\theta}(X)$ of θ whose variance attains the Cramér–Rao lower bound if and only if the score can be expressed in the form*

$$\dot{\ell}(\theta) = I(\theta)\left\{\hat{\theta}(X) - \theta\right\},$$

or, equivalently, if and only if the function

$$\frac{\dot{\ell}(\theta)}{I(\theta)} + \theta$$

does not depend on θ.

Theorem 3.28 (Fisher)**.** *Let $X_1, \ldots, X_n \overset{\text{iid}}{\sim} f(x|\theta_0)$, with f satisfying certain smoothness conditions. As $n \to \infty$, we have*

$$\sqrt{nI(\theta_0)} \cdot (\hat{\theta} - \theta_0) \xrightarrow{\mathscr{D}} \mathcal{N}(0, 1)$$

and

$$\sqrt{nI(\hat{\theta})} \cdot (\hat{\theta} - \theta_0) \xrightarrow{\mathscr{D}} \mathcal{N}(0, 1)$$

Remark 3.29. One may also think

$$\hat{\theta} \approx \mathcal{N}\left(\theta_0, \frac{1}{nI(\theta_0)}\right).$$

Proposition 3.30. *Assumptions:*

- *Common support: $\{x : f(x, \theta) > 0\}$ does not depend on x.*
- *Smoothness of densities.*
- *Distinct densities: if $\theta_1 \neq \theta_2$ then $f(x, \theta_1) \neq f(x, \theta_2)$.*

Properties of maximal likelihood estimators under the above assumptions:

- *consistency,*
- *asymptotic normality,*
- *has known and optimal asymptotic variance (**efficiency**). That is, it attains the Cramér–Rao bound.*
- *Invariance in the following sense:*

Theorem 3.31. *If $\hat{\theta}_n$ is an MLE of θ, then $\hat{\tau}_n := g(\hat{\theta}_n)$ is an MLE of $g(\theta)$.*

3.4. **Jensen Inequality.**

Theorem 3.32. *If $f : \mathbb{R} \to \mathbb{R}$ is convex and X is a random variable such that $\mathbb{E}|X| < \infty$, then*

$$f(\mathbb{E}\, X) \leq \mathbb{E}\, f(X).$$

Proof. From the convexity of f we know $f(x) \geq f(y) + f'(y)(x-y)$ for any x and y. Setting $y = \mu =: \mathbb{E}\, X$ gives

$$f(X) \geq f(\mu) + f'(\mu)(X-\mu), \quad \forall x, y.$$

Taking expectation on both sides gives the desired result. □

3.4.1. *Applications of Jensen Inequality.*

- If f is concave, then $f(\mathbb{E}\, X) \geq \mathbb{E}\, f(X)$.
- The convex function $x \mapsto x^2$ and the concave function $x \mapsto \log x$ give two special cases:

$$(\mathbb{E}\, X)^2 \leq \mathbb{E}\, X^2, \quad \log \mathbb{E}\, X \geq \mathbb{E} \log X.$$

- If $x_1, x_2, \ldots, x_n > 0$ and $p_i \geq 0$ such that $\sum p_i = 1$, then

$$\prod x_i^{p_i} \leq \sum p_i x_i.$$

  Remark 3.33. When $p_i = 1/n$, this result reduces to the geometric mean-arithmetic mean equality.

  Proof. Let X be a discrete variable such that $\mathbb{P}(X = x_i) = p_i$. Then

$$\sum p_i \log x_i = \mathbb{E} \log X \leq \log \mathbb{E}\, X \leq \sum p_i x_i.$$

  Taking exp on both sides gives the desired result. □

- **Hölder's inequality**: If $X, Y \geq 0$ are random variables and $p, q > 0$ are such that $1/p + 1/q = 1$, then

$$\mathbb{E}(XY) \leq (\mathbb{E}\, X^p)^{1/p} \cdot (\mathbb{E}\, X^q)^{1/q}.$$

  Proof. If $\mathbb{E}\, X^p = \mathbb{E}\, X^q = 1$, then

$$XY = (X^p)^{1/p}(Y^q)^{1/q} \leq \frac{1}{p}X^p + \frac{1}{q}X^q,$$

  where the last inequality follows from the previous result. Taking expectation on both sides then gives $\mathbb{E}[XY] \leq \mathbb{E}\, X^p\, \mathbb{E}\, Y^q$.

  For the general case, normalize by setting

$$\tilde{X} := \frac{X}{(\mathbb{E}\, X)^{1/p}}, \quad \tilde{Y} := \frac{Y}{(\mathbb{E}\, Y)^{1/q}}.$$

  □

- **Cauchy Inequality**: Taking $p = q = 2$ in Hölder gives

$$\mathbb{E}|XY| \leq \sqrt{\mathbb{E}\, X^2}\sqrt{\mathbb{E}\, Y^2}.$$

- The consistency of likelihood.

3.5. **Multivariate Normal.**

Definition 3.34. The random vector $X = (X_1, X_2, \dots, X_k)$ is said to follow a **multivariate normal distribution** if for each $a \in \mathbb{R}^k$, $a^\intercal x$ is normal. We write

- $\mu = \mathbb{E}\, X \in \mathbb{R}^k$.
- $\Sigma = \mathrm{Var}(X) = \mathbb{E}\left[(X-\mu)(X-\mu)^\intercal\right] \in \mathbb{R}^{2k}$.

Proposition 3.35.

- *If Σ is positive definite, then X has density*
$$f(X) = \frac{1}{(2\pi)^{k/2}\det(\Sigma)} \exp\left(-\frac{1}{2}(X-\mu)^\intercal \Sigma^{-1}(X-\mu)\right).$$
- *If (X_1, X_2) is bivariate normal and* $\mathrm{Cov}(X_1, X_2) = 0$*, then X_1 and X_2 are independent.*
- *If $U \sim N_k(\mu, \Sigma)$, $a \in \mathbb{R}^p$, and B is a $p \times k$ matrix, then*
$$V = a + BU \sim N_p(a + B\mu, B\Sigma B^\intercal).$$

4. Confidence Intervals

Definition 4.1. Suppose $X_1, X_2, \ldots, X_n \overset{\text{iid}}{\sim} F_\theta$. Confidence intervals (CIs) are probabilistic statements on data of the form

$$\mathbb{P}_\theta \left[A(X_1, \ldots, X_n) \le \theta \le B(X_1, \ldots, X_n)\right] = \alpha.$$

The interval

$$[A(X_1, \ldots, X_n), B(X_1, \ldots, X_n)]$$

is called a $\alpha \cdot 100\%$ **confidence interval**.

Remark 4.2. We are typically interested in $\alpha = 0.95$ or $\alpha = 0.99$.

Remark 4.3.

- The probabilistic statement concerns the interval ends, not θ, which is fixed. The interval ends are random variables.
- Interpretation (frequentest): the long run frequency of the CI covering θ is α.

Definition 4.4. The α quantile of $X \sim F$, q_α, is such that

$$\mathbb{P}[X \le q_\alpha] = \alpha.$$

Example 4.5. Let $X_1, X_2, \ldots, X_n \overset{\text{iid}}{\sim} \text{Exponential}(\lambda) = \text{Gamma}(1, \lambda)$. Then $\sum X_i \sim \text{Gamma}(n, \lambda)$ and $\lambda \sum X_i \sim \text{Gamma}(n, 1)$. Note that the distribution of $\lambda \sum X_i$ does not depend on λ. We then have

$$\left[\frac{q_{0.025}}{\sum X_i}, \frac{q_{0.975}}{\sum X_i}\right],$$

where q refers to the quantile of $\text{Gamma}(n, 1)$, is a 95% CI.

Definition 4.6. Let $X_1, X_2, \ldots, X_n \overset{\text{iid}}{\sim} F_\theta$. The function

$$g(X_1, \ldots, X_n, \theta)$$

is called a **pivot** if its distribution does not depend on θ.

Remark 4.7. One may use the distribution of the pivot $g(X_1, \ldots, X_n, \theta) \sim F^*$ to build CIs. Let L and U be the $(1-\alpha)/2$ and $1-(1-\alpha)/2$ quantiles for F^*. Then

$$\alpha = \mathbb{P}\left[L \le g(X_1, \ldots, X_n, \theta) \le U\right] = \mathbb{P}\left[\theta \in S(X_1, \ldots, X_n, L, U)\right]$$

for some set S. If S is an interval, it is a CI.

Theorem 4.8. *Let* $X_1, X_2, \ldots, X_n \overset{\text{iid}}{\sim} \mathcal{N}(\mu, \sigma^2)$. *Let*

$$\overline{X_n} \sim \mathcal{N}\left(\mu, \frac{\sigma^2}{n}\right), \quad S^2 \coloneqq \frac{1}{n-1} \sum (X_i - \overline{X})^2.$$

Then

$$\sqrt{n} \cdot \frac{\overline{X} - \mu}{S} \sim t_{n-1}, \quad (n-1)\frac{S^2}{\sigma^2} \sim \chi^2_{n-1}.$$

Remark 4.9. Thus $\sqrt{n} \cdot \frac{\overline{X}-\mu}{S}$ is a pivot estimator for μ and $(n-1)\frac{S^2}{\sigma^2}$ is a pivot estimator for σ.

Remark 4.10. We may use the central limit theorem and the above results to obtain approximate CIs for large samples. Let $X_1, X_2, \ldots, X_n \overset{\text{iid}}{\sim} F$ with $\mathbb{E}[X] = \mu$ and $\mathrm{Var}[X] = \sigma^2$. The central limit theorem gives

$$\sqrt{n} \cdot \frac{\overline{X} - \mu}{\sigma} \approx \mathcal{N}(0, 1).$$

Thus

$$\left[\overline{X} - q_{\alpha/2}\frac{\sigma}{\sqrt{n}}, \overline{X} + q_{\alpha/2}\frac{\sigma}{\sqrt{n}}\right],$$

where q is the quantiles on $\mathcal{N}(0, 1)$ contains μ with probability α. We can approximate σ using S to obtain the following CI:

$$\left[\overline{X} - q_{\alpha-2}\frac{S}{\sqrt{n}}, \overline{X} + q_{\alpha/2}\frac{S}{\sqrt{n}}\right].$$

Note that we used two approximations: central limit theorem and using S to approximate σ.

Remark 4.11. For a MLE $\hat{\theta}$, we can use the following two results to construct approximate CIs:

$$\sqrt{n}(\hat{\theta} - \theta) \approx \mathcal{N}\left(0, \frac{1}{I(\theta)}\right), \sqrt{nI(\theta)}(\hat{\theta} - \theta) \approx \mathcal{N}(0, 1).$$

Remark 4.12. The above cases fail, however, if either the distribution of the pivot or the variance of the estimators is unknown.

5. The Bootstrap

Definition 5.1. Let $X_1, X_2, \ldots, X_n \overset{\text{iid}}{\sim} F$. The **empirical distribution function (EDF)**, $\hat{F_n}$, is the CDF that puts probability $1/n$ at each X_i.

$$\hat{F_n}(x) \coloneqq \frac{1}{n} \sum \mathbb{1}_{\{X_i \le x\}}.$$

Remark 5.2. Note that $\mathbb{1}_{\{X_i \le x\}} \sim \text{Bernoulli}(F(x))$. This gives the following properties:

Proposition 5.3.

- *$\hat{F}(x)$ is an unbiased estimator for $F(x)$:*
$$\mathbb{E}[\hat{F}(x)] = F(x).$$
- *$\hat{F}(x)$ has variance:*
$$\text{Var}(\hat{F}(x)) = \frac{F(x)(1 - F(x))}{n}.$$
- *By the law of large numbers,*
$$\hat{F}(x) \xrightarrow{\text{p}} F(x).$$
Moreover, $\hat{F_n}(x) \to F(x)$ uniformly. That is:

Theorem 5.4 (Glivenko-Cantelli)**.** *If $X_1, X_2, \ldots, X_n \overset{\text{iid}}{\sim} F$, then as $n \to \infty$ we have*

$$\sup_x \left|\hat{F_n}(x) - F(x)\right| \longrightarrow 0.$$

Remark 5.5. For variable $\theta \coloneqq T(F)$, we can thus construct estimator $\hat{T} \coloneqq T(\hat{F})$.

Example 5.6. For $T = \int x \, \mathrm{d}F(x)$, θ is the mean. For $T = \int (x - \mu)^2 \, \mathrm{d}F(x)$, θ is the variance. For $T = F^{-1}(1/2)$, θ is the median.

Remark 5.7. Let $X_1, X_2, \ldots, X_n \overset{\text{iid}}{\sim} F$ and $T_n \coloneqq g(X_1, \ldots, X_n)$. We want to find $\text{Var}(T_n)$. If it is possible to sample from F, then we may repeated the following procedure

- Take repeated samples of size n.
- Calculate T_n for each sample.

to obtain k samples of T_n, $T_{n,1}, \ldots, T_{n,k}$. We may use

$$\frac{1}{k} \sum \left(T_{n,j} - \overline{T}_n\right)^2$$

as an estimator for the variance of T_n, $\text{Var}_F(T_n)$.

Remark 5.8. If we cannot directly sample from F, we may use $\hat{F}$ as an approximation. That is, given a sample of size n, we sample repeatedly with replacement k samples also of size n from the given sample, and calculate the statistic of interest for each sample to estimate the distribution of T_n. This procedure is called **bootstrapping**, and each sample is called a **bootstrap sample**.

6. Hypothesis Testing

We want to test whether a set of given data is generated by a certain data generating model.

The idea: we use a certain distance between the ecdf and the theoretical cdf in the density space as a test statistic.

Example 6.1. Given $X_i \overset{\text{iid}}{\sim} F$, we want to test if F is the cdf of a normal distribution. Test statistic:

- **Kolmogorov–Smirnov**: $S \coloneqq \sup_x \left|F(x) - \hat{F}(x)\right|$.
- Quantiles: e.g., compare $Q_3 - Q_1$ with $X_{(\lfloor 3N/4 \rfloor)} - X_{(\lfloor N/4 \rfloor)}$.
- **Shapiro-wilk**:
$$W \coloneqq \frac{\left(\sum a_i x_{(i)}\right)^2}{\sum (x - \overline{x})^2}.$$

6.1. Hypothesis Testing for Parametric Models.

Let $X_i \overset{\text{iid}}{\sim} F_\theta$ with $\theta \in \Omega$. The **null hypothesis**:
$$H_0 : \theta \in \Omega_0 \subset \Omega.$$
The **alternative hypothesis**:
$$H_A : \theta \in \Omega_1.$$
We often have $\Omega_1 = \Omega \setminus \Omega_0$.

Remark 6.2. Note a certain asymmetry: we usually know a lot more about H_0 (the "status quo") than H_1.

Definition 6.3. Let S be the set of all possible values for $X = (X_1, \ldots, X_n)$. The values for which we do not reject H_0, S_0, is called the **acceptance region**. The values for which we reject H_0, S_1, is called the **rejection region**. Note that we require $S = S_0 \cup S_1$.

Definition 6.4. $T = T(X)$ is called a **test statistic** if
$$S_1 = \{x : T(x) \in R_1\}$$
for some $R_1 \subset \mathbb{R}$.

Definition 6.5. A **type I error**, or a false positive, is the rejection of the null hypothesis when it is actually true. A **type II error**, or a false negative, is the failure to reject a null hypothesis that is actually false.

Definition 6.6. The function
$$\pi : \Omega \longrightarrow [0, 1], \quad \pi(\theta) \coloneqq \mathbb{P}_\theta(x \in S_1)$$
is called the **power function**.

Remark 6.7. Note we can represent type I errors as $\pi(\theta)$ with $\theta \in \Omega_0$; and type II errors as $1 - \pi(\theta)$ with $\theta \in \Omega_1$. Ideally, we want π to be small on Ω_0 and large on Ω_1. We often find S_1 such that π is low on Ω_0 and hope for the best for Ω_1.

Definition 6.8. The **size** of the test is $\sup_{\theta \in \Omega_0} \pi(\theta)$.

Definition 6.9. A test is a **level α test** if it has size $\leq \alpha$.

Remark 6.10. For convenience of calculating size, we often want either simple H_0 such that $\theta = \theta_0$, or the power function to be constant on Ω_0.

Example 6.11. Let $X_i \overset{\text{iid}}{\sim} F$ such that $\mathbb{E}[X_i] = \mu$ with known variance $\mathrm{Var}[X_i] = \sigma^2$. Let

$$H_0 : \mu = \mu_0, \quad H_A : \mu > \mu_0.$$

Under H_0, the CLT gives

$$T(X) \coloneqq \sqrt{n} \cdot \frac{\overline{X} - \mu_0}{\sigma} \approx \mathcal{N}(0, 1).$$

Then, we may set the rejection region by picking c such that

$$\mathbb{P}_\mu\left(\{T(X) \geq c\}\right) = \alpha.$$

Example 6.12. Same set up as above, with

$$H_0 : \mu = \mu_0, \quad H_A : \mu \neq \mu_0.$$

We may set

$$S_1 \coloneqq \{X : |T(X)| > c_2\}$$

to be such that $\mathbb{P}_\mu(X \in S_1) \approx \alpha$.

Remark 6.13. If σ is unknown, we may use the fact that under H_0,

$$\sqrt{n} \cdot \frac{\overline{X} - \mu_0}{S} \sim t_{n-1}.$$

6.2. ***p*-value.**

Definition 6.14. The ***p*-value** is the smallest level α for which we reject H_0 with the observed data.

Proposition 6.15. *If under H_0, $T \sim F$, then $p = \mathbb{P}(T \geq T_{obs})$. Moreover,* $F(p) \sim \mathrm{Uniform}[0, 1]$.

Appendix A: Common Distributions

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

6.3. **Properties of the uniform distribution.**

Proposition 6.16. *Let* $X_1, \ldots, X_n \overset{\text{iid}}{\sim} \mathcal{N}(\mu, \sigma^2)$.

- $\mathbb{E}[S^2] = \sigma^2$.
- $\overline{X}$ *and* S^2 *are independent.*

6.4. **Properties of the exponential distribution.**

Proposition 6.17.

(i) *The "memoryless" property:*

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

(ii) $\text{Exponential}(\lambda) = \text{Gamma}(1, \lambda)$.

6.5. Properties of the gamma distribution.

Proposition 6.18.

(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*

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

(iii) *If* $X \sim \text{Gamma}(\alpha, \beta)$, *then*

$$\beta X \sim \text{Gamma}\left(\alpha, 1\right).$$

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.

$\square$