

1 Basic Statistical Concepts

Consider a poll with two answers, A and B, regarding political parties. Let:

- N : total number of voters,
- M : number of voters supporting A,
- n : size of the poll,
- X_1, X_2, \dots, X_n : responses,
- Each $X_i \in \{0, 1\}$ if $X_i = 1$ supports A.

Additionally, assume:

- We select n individuals from N at random and record their truthful reply,
- Every person asked replies (no selection bias),
- People can be asked repeatedly.

The aim of the poll is to estimate the fraction of party A supporters, say θ .

Definition 1 (Estimator). *An intuitive estimator is:*

$$\hat{\theta} = \frac{1}{n} \sum_{i=1}^n X_i$$

This estimator will be analyzed in the following sections to determine whether it is unbiased, consistent, and optimal.

2 Statistical Models

Let (X, \mathcal{F}) be a measurable space, i.e., a set X with a sigma-algebra \mathcal{F} , in which our statistical observations take values.

Definition 2 (Statistical Model). *Let (X, \mathcal{F}) be some sample space. We call the parameter space Θ . A statistical model is a family of probability measures $\{P_\theta\}_{\theta \in \Theta}$.*

Remark 1. *Often (X, \mathcal{F}) is a product space. For example, if $X_i \in \{0, 1\}$, each P_θ is a product distribution, i.e., X_1, X_2, \dots, X_n are independent and identically distributed (iid). Then we say $\{P_\theta : \theta \in \Theta\}$ is an iid statistical model.*

Remark 2. *If every person could only be asked once, we would have P_θ as a hypergeometric distribution, which converges to the Bernoulli model as $N, M \rightarrow \infty$.*

3 Parameter Estimation

Assume $(\Omega, \mathcal{F}, P_\theta)$ is the setting of parametric statistics. Assume Θ is measurable.

Definition 3 (Estimator). *An estimator for θ is any measurable function $\hat{\theta} : X \rightarrow \Theta$, i.e., any function that, based on some data X , outputs a guess $\hat{\theta}(X)$ for θ .*

4 Unbiased and Consistent Estimators

4.1 Unbiased Estimator

Definition 4 (Unbiased Estimator). *Let $(\Omega, \mathcal{F}, P_\theta)$ be a measurable space. An estimator $\hat{\theta}$ is called unbiased if:*

$$\mathbb{E}[\hat{\theta}] = \theta \quad \forall \theta \in \Theta$$

where \mathbb{E}_{P_θ} denotes expectation under the law P_θ . In more explicit terms, unbiasedness means no systematic error.

Proof. For the Bernoulli model, we compute:

$$\mathbb{E}[\hat{\theta}_n] = \mathbb{E}\left[\frac{1}{n} \sum_{i=1}^n X_i\right] = \frac{1}{n} \sum_{i=1}^n \mathbb{E}[X_i] = \frac{1}{n} \sum_{i=1}^n \theta = \theta$$

Thus, $\hat{\theta}_n$ is an unbiased estimator of θ . □

4.2 Consistent Estimator

Definition 5 (Consistent Estimator). *Let $\{P_{\theta,n} : n \geq 1\}$ be a sequence of statistical models on the same parameter space. Let $\hat{\theta}_n$ be a sequence of estimators. The sequence $\hat{\theta}_n$ is called consistent if for every $\theta \in \Theta$:*

$$\hat{\theta}_n \rightarrow \theta \quad \text{in probability as } n \rightarrow \infty$$

or equivalently:

$$P_\theta \left(\lim_{n \rightarrow \infty} \hat{\theta}_n = \theta \right) = 1$$

Proof. For the Bernoulli model:

$$\hat{\theta}_n = \frac{1}{n} \sum_{i=1}^n X_i$$

We know $\mathbb{E}[\hat{\theta}_n] = \theta$ and $\text{Var}(\hat{\theta}_n) = \frac{\theta(1-\theta)}{n}$. Using Chebyshev's inequality, for any $\epsilon > 0$:

$$P \left(|\hat{\theta}_n - \theta| > \epsilon \right) \leq \frac{\text{Var}(\hat{\theta}_n)}{\epsilon^2} = \frac{\theta(1-\theta)}{n\epsilon^2}$$

As $n \rightarrow \infty$, this probability tends to 0, proving that $\hat{\theta}_n$ is consistent. □

5 Maximum Likelihood Estimation (MLE)

Definition 6 (Maximum Likelihood Estimator). *The maximum likelihood estimator (MLE) is the parameter that maximizes the likelihood function:*

$$L(\theta) = \prod_{i=1}^n P_\theta(X_i)$$

5.1 Proof: MLE for Bernoulli Model

Proof. For the Bernoulli model, $P_\theta(X_i) = \theta^{X_i}(1-\theta)^{1-X_i}$, so the likelihood function is:

$$L(\theta) = \prod_{i=1}^n \theta^{X_i}(1-\theta)^{1-X_i} = \theta^{\sum X_i} (1-\theta)^{n-\sum X_i}$$

Taking the logarithm:

$$\log L(\theta) = \sum X_i \log \theta + (n - \sum X_i) \log(1 - \theta)$$

Setting the derivative with respect to θ equal to 0 gives:

$$\frac{d}{d\theta} \log L(\theta) = \frac{\sum X_i}{\theta} - \frac{n - \sum X_i}{1 - \theta} = 0$$

Solving for θ , we get:

$$\hat{\theta}_n = \frac{1}{n} \sum_{i=1}^n X_i$$

which is the MLE. □

6 Bayesian Methods

Definition 7 (Posterior Distribution in Bayesian Inference). *In Bayesian statistics, a key element is the prior distribution, denoted by $\pi(\theta)$, which reflects our beliefs about the parameter θ before observing data. The posterior distribution is given by:*

$$\pi(\theta|X) \propto P_\theta(X)\pi(\theta)$$

6.1 Example: Posterior for Bernoulli Model

Example 1. Suppose we have a Beta prior for θ , $\pi(\theta) \sim \text{Beta}(\alpha, \beta)$, and observe X_1, \dots, X_n as Bernoulli trials. The likelihood is:

$$P(X|\theta) = \theta^{\sum X_i} (1 - \theta)^{n - \sum X_i}$$

The posterior is proportional to the product of the prior and likelihood:

$$\pi(\theta|X) \propto \theta^{\sum X_i + \alpha - 1} (1 - \theta)^{n - \sum X_i + \beta - 1}$$

Thus, $\pi(\theta|X) \sim \text{Beta}(\sum X_i + \alpha, n - \sum X_i + \beta)$.

Notes on Bayes and Posterior

Posterior = prior \times likelihood

Normalizing Constant

$$\int \text{Posterior } dx = 1$$

So,

$$\int \text{Posterior } dx = 1$$

Prior \rightarrow Posterior via Bayes.

Let \mathcal{F}_0 be a σ -algebra on Ω and suppose $(\Omega, \mathcal{F}_0, P_\theta)$ is a dominated statistical model with densities $p(x|\theta)$. Assume

$$x, \theta \in \Omega \Rightarrow p(x|\theta)$$

is jointly measurable with respect to $\mathcal{F}_0 \times \mathcal{F}_1$.

Let π be a prior distribution on Ω with density $\pi(\theta)$ with respect to measure ν . Define posterior density

$$\pi(\theta|x) = \frac{p(x|\theta)\pi(\theta)}{\int p(x|\theta)\pi(\theta) d\theta}$$

The corresponding probability measure is called the **posterior distribution**.

Think of $p(x|\theta)$ as a Lebesgue measure. Let ν be a Lebesgue density.

Exception: If $\Omega = \{0, 1\}$, then we take ν to be the counting measure.

From the posterior, we can derive several estimators. For example, $E[\theta|X = x]$ is convex:

$$\int \theta p(x|\theta) d\theta = E[\theta|X = x]$$

Example: Binomial model $X|\theta \sim \text{Binomial}(n, \theta)$ with prior $\theta \sim \text{Unif}(0, 1)$.

For a uniform prior, we know the MAP and MLE.

Posterior mean:

$$\theta_{\text{MAP}} = \frac{k+1}{n+2}$$

In the case of coin flips, $X \sim \text{Binomial}(n, \theta)$, where k is the number of heads, we conclude $\theta|X \sim \text{Beta}(k+1, n-k+1)$.

$$\theta|X \sim \text{Beta}(k+1, n-k+1)$$

Conjugate Bayes Models: Let $P_\theta \in \mathcal{P}$ be a statistical model. Then some family of priors is called **conjugate** if

$$P_\theta \in \mathcal{P} \Rightarrow \theta|X \in \mathcal{P}$$

for all $X \in \mathcal{X}$, where \mathcal{X} is the sample space.

$$\theta|X \sim \text{Beta}(a, b), \quad X \sim \text{Bernoulli}(p)$$

Loss Functions and Risk

Loss Function: A function $L : \Theta \times \mathcal{X} \rightarrow [0, \infty)$ is a basis function if for every $\theta \in \Theta$, $L(\theta, \cdot)$ is measurable.

Given an estimator δ , the expected loss is

$$R(\theta, \delta) = E_\theta[L(\theta, \delta)]$$

Mean Squared Error (MSE):

$$L(x, y) = (x - y)^2 \Rightarrow R(\theta, \delta) = E_\theta[(\delta - \theta)^2]$$

Bias-Variance Decomposition:

$$L(x, y) = (x - y)^2$$

Proof: Let $\delta(x) = E[\theta|X = x]$.

$$R(\theta, \delta) = E_\theta[(\delta(X) - \theta)^2]$$

Bias-variance decomposition:

$$E[(\delta(X) - \theta)^2] = \text{Var}(\delta(X)) + (\text{Bias})^2$$

Minimax and Bayes Risk

Minimax Risk: Given an estimator δ in a model $P_\theta \in \mathcal{P}$, the maximal risk of it is

$$\sup_{\theta \in \Theta} R(\theta, \delta)$$

The minimax of a model P_θ is given as $\inf_\delta \sup_\theta R(\theta, \delta)$, where the inf is over all estimators.

An estimator is called minimax if

$$\sup_\theta R(\theta, \delta) = \inf_\delta \sup_\theta R(\theta, \delta)$$

Bayes Risk: Given an estimator δ and prior π on Θ , the Bayes risk of δ is defined as

$$R_\pi(\delta) = \int R(\theta, \delta) d\pi(\theta)$$

The posterior risk of an estimator $\delta(X)$ is defined by

$$R(\delta|X = x) = E[L(\theta, \delta(X))|X = x]$$

Suppose δ^* is an estimator that minimizes the posterior risk, $\delta^*(x) = E[\theta|X = x]$. Then it also minimizes the Bayes risk.

If $L(x, y) = (x - y)^2$, the Bayes optimal estimator $\delta(x)$ is the posterior mean.

We want to construct $C(x)$ s.t. $P_\theta(\theta \in C(x)) \geq 1 - \alpha, \forall \theta \in [0, 1]$

$$x^{(1)} \quad (\quad) \quad C(x^{(1)})$$

$$x^{(k)} \quad (\quad) \quad C(x^{(k)})$$

$$\theta \rightarrow \quad \rightarrow \quad \rightarrow \quad \text{contains true param } 3/4 \text{ times}$$

Example cont.:

Best guess: $C(x) = \left[\frac{\bar{X}_n - a}{n}, \frac{\bar{X}_n + b}{n} \right]$

$$P_\theta^n(\theta \in C(x)) = P_\theta^n \left(\frac{\bar{X}_n}{n} - \theta \in [-b, a] \right)$$

$$= F_\theta^n(a) - F_\theta^n(-b) + \rho_n$$

where $F_\theta^n : \mathbb{R} \rightarrow [0, 1]$, $F_\theta^n(t) = P_\theta^n \left(\frac{\bar{X}_n - \theta}{n} \leq t \right)$ is the CDF of $\frac{\bar{X}_n - \theta}{n}$ under P_θ and $\rho_n = P_\theta^n \left(\frac{\bar{X}_n}{n} - \theta = -b \right)$.

How to choose a and b:

$$\text{CDF} \quad \text{CDF} \quad \leftarrow \quad -b \quad a \rightarrow t$$

We'd like to choose $a = (F_\theta^n)^{-1} \left(1 - \frac{\alpha}{2} \right)$ and $b = (F_\theta^n)^{-1} \left(\frac{\alpha}{2} \right)$, where

$$(F_\theta^n)^{-1}(p) := \inf \{ t \in \mathbb{R} : F_\theta^n(t) \geq p \} \quad (\text{Quantile Function})$$

Let's use a normal approximation, for $\sigma^2 = \theta(1 - \theta)$:

$$\sqrt{n} \left(\frac{\bar{X}_n}{n} - \theta \right) = \frac{1}{\sqrt{n}} \sum_{k=1}^n \frac{X_k - \theta}{\sigma} \xrightarrow{d} \mathcal{N}(0, 1) \quad [\text{CLT}]$$

$$X_k \sim \text{Ber}(\theta)$$

Then it follows that

$$F_\theta^n(a_n) = P_\theta^n \left(\frac{\bar{X}_n}{n} - \theta \leq a_n \right)$$

$$= P_\theta^n \left(\frac{\sqrt{n}}{\sigma} \left(\frac{\bar{X}_n - \theta}{n} \right) \leq \sqrt{n} a_n \right)$$

$$= \Phi \left(\frac{\sqrt{n}}{\sigma} a_n \right),$$

where the convergence is valid if $a_n := \text{const.} \cdot \frac{1}{\sqrt{n}}$.

Now, let us choose

$$a := \frac{\sigma}{\sqrt{n}} z_{1 - \frac{\alpha}{2}}$$

where $z_{1 - \frac{\alpha}{2}} = \Phi^{-1} \left(1 - \frac{\alpha}{2} \right)$ is the $1 - \frac{\alpha}{2}$ quantile of $\mathcal{N}(0, 1)$ and $b = a$. Then

$$C(x) = \left[\frac{\bar{X}_n}{n} - \frac{\sigma}{\sqrt{n}} z_{1 - \frac{\alpha}{2}}, \frac{\bar{X}_n}{n} + \frac{\sigma}{\sqrt{n}} z_{1 - \frac{\alpha}{2}} \right]$$

It follows

$$P_\theta^n(\theta \in C(x)) = F_\theta^n(a_n) - F_\theta^n(b) + \rho_n = 1 - \frac{\alpha}{2} + o(1) + o(1)$$

$$= 1 - \alpha + o(1) \text{ as } n \rightarrow \infty$$

\Rightarrow Asymptotically valid confidence set

One more problem: σ depends on θ

- Upper bound: $\sup_{\theta \in [0, 1]} \theta(1 - \theta) = \frac{1}{4}$ (maximized at $\theta = \frac{1}{2}$)
- Empirical Variance: $\hat{\sigma}^2 = \frac{1}{n} \sum_{i=1}^n (X_i - \frac{1}{n} \sum_{i=1}^n X_i)^2$

$$\frac{\hat{\sigma}^2}{\sigma^2} \xrightarrow{P_\theta} 1$$

Slutsky's Theorem:

$$X_n \xrightarrow{d} X, \quad Y_n \xrightarrow{d} \text{const.} \Rightarrow X_n Y_n \xrightarrow{d} CX$$

Exercise: Use this to deduce that $a_n = \frac{\hat{\sigma}}{\sqrt{n}} z_{1-\frac{\alpha}{2}}$ is also valid

Remark:**Hypothesis Testing**

Definition: Let $(P_\theta : \theta \in \Theta)$ be a statistical model and let $\Theta = \Theta_0 \cup \Theta_1$ be a partition. Then:

- A statistical test is a measurable function of the data $\varphi : (\mathcal{X}, \mathcal{F}) \rightarrow [0, 1]$
- If $\forall x \in \mathcal{X}, \varphi(x) \in \{0, 1\}$, then φ is a non-randomized test
- Else φ is randomized

Definitions:

- $H_0 : \theta \in \Theta_0$ is called the null hypothesis
- $H_1 : \theta \in \Theta_1$ is called the alternative hypothesis
- The map $\theta \rightarrow \beta_\varphi(\theta) = P_\theta[\varphi = 1]$ is called the power function of a test φ

$$\begin{array}{ccccc} 1 & \beta_\varphi(\theta) & 0 & \Theta_0 & \Theta_1 & \Theta \end{array}$$

- For $\theta \in \Theta_0$, $\beta_\varphi(\theta)$ is the type-I-error under θ [Wrongly rejecting the null]
- For $\theta \in \Theta_1$, $1 - \beta_\varphi(\theta)$ is the type-II-error

Note:

$$1 - P_\theta(\varphi = 1) = P_\theta(\varphi = 0) = P_\theta(\text{wrongly accepting the null})$$

Definition: [Level]

$\varphi : \mathcal{X} \rightarrow [0, 1]$ has level $\alpha \in [0, 1]$ if

$$\sup_{\theta \in \Theta_0} \beta_\varphi(\theta) \leq \alpha$$

Definition: [Uniformly most powerful test]

Given a level $\alpha \in (0, 1)$, $\varphi : \mathcal{X} \rightarrow [0, 1]$ is called UMP if for every other test φ' of level α and all $\theta \in \Theta_1$,

$$\beta_\varphi(\theta) \geq \beta_{\varphi'}(\theta)$$

$$\begin{array}{ccccc} 1 & \alpha & 0 & \beta_\varphi(\theta) & \beta_{\varphi'}(\theta) & \Theta_0 & \Theta_1 \end{array}$$

Remark:

In general, it is very hard to find UMP tests. But: for simple hypotheses, i.e. $\Theta_0 = \{\theta_0\}, \Theta_1 = \{\theta_1\}$, it is possible. Here, likelihood ratio tests are UMP.

Theorem: [Neyman-Pearson Lemma]

Let $\Theta_0 = \{\theta_0\}, \Theta_1 = \{\theta_1\}$ be simple:

1. **Existence:** There exists a test φ and a constant $k \in [0, \infty)$, s.t. $P_{\theta_0}(\varphi = 1) = \alpha$, of the form

$$\varphi(x) = \begin{cases} 1, & \text{if } \frac{p_{\theta_1}(x)}{p_{\theta_0}(x)} > k \\ 0, & \text{if } \frac{p_{\theta_1}(x)}{p_{\theta_0}(x)} < k \end{cases} \quad (*)$$

Here $p_{\theta_1}, p_{\theta_0}$ are densities w.r.t. some dominated measure μ , e.g. $\mu = p_{\theta_0} + p_{\theta_1}$. Finite Θ implies measure is always dominated (likelihood always exists).

2. **Sufficiency:** If φ satisfies $P_{\theta_0}(\varphi = 1) = \alpha$ and $(*)$ then φ is a UMP level α test.
3. **Necessity:** If φ_k is UMP for level α , then it must be of the form $(*)$, and it also satisfies $P_{\theta_0}(\varphi_k = 1) = \alpha$, or else it must satisfy $P_{\theta_1}(\varphi_k = 1) = 1$.

Proof:

1. Define $r(x) = \frac{p_{\theta_1}(x)}{p_{\theta_0}(x)} \in [0, \infty) \cup \{\pm\infty\}$. Let F_0 be the CDF of $r(x)$ under P_{θ_0} .

$$F_0(t) = P_{\theta_0}(r(x) \leq t)$$

Then define also $\alpha(t) = 1 - F_0(t) = P_{\theta_0}(r(x) > t)$

- α is right-continuous:

$$\lim_{\epsilon \rightarrow 0} \alpha(t + \epsilon) = \lim_{\epsilon \rightarrow 0} P_{\theta_0}(r(x) > t + \epsilon) = P_{\theta_0}(r(x) > t) = \alpha(t)$$

- α is non-increasing
- α has left limits

$$\lim_{\epsilon \rightarrow 0} \alpha(t - \epsilon) = P_{\theta_0}(r(x) > t - \epsilon) = \alpha(t^-)$$

α is **cadlag**:

- Continuous from the right
- Limit from the left

There exists some $k \in [0, \infty)$ s.t. $\alpha \leq \alpha(k^-)$ and $\alpha \geq \alpha(k)$

We define our test

$$\varphi(x) = \begin{cases} 1 & \text{if } r(x) > k \\ \gamma & \text{if } r(x) = k \quad [\text{reject null w.p. } \gamma] \\ 0 & \text{if } r(x) < k \end{cases}$$

We set

$$\gamma = \frac{\alpha - \alpha(k)}{\alpha(k^-) - \alpha(k)}$$

The level of φ is

$$\begin{aligned} E_{\theta_0}[\varphi(x)] &= P_{\theta_0}(\varphi(x) = 1) \\ &= P_{\theta_0}(r(x) > k) + P_{\theta_0}(r(x) = k) \cdot \gamma \\ &= \alpha(k) + [\alpha(k^-) - \alpha(k)] \cdot \frac{\alpha - \alpha(k)}{\alpha(k^-) - \alpha(k)} = \alpha \\ &\quad (\text{randomizing the test}) \end{aligned}$$

Lecture 6

Neyman-Pearson

Power of a test:

$$E_{\theta_1}[\varphi] = P_{\theta_1}(\varphi = 1)$$

Likelihood ratio test:

$$\frac{p_{\theta_1}(x)}{p_{\theta_0}(x)} = r(x)$$

LR test

$$\varphi(x) = \begin{cases} 1 & \text{if } r(x) > k \\ \gamma & \text{if } r(x) = k \\ 0 & \text{if } r(x) < k \end{cases}$$

for some $k \in [0, \infty)$, $\gamma \in [0, 1]$.

Note: LR tests are UMP for simple hypothesis testing:

- Given some α , if LR satisfies $E_{\theta_0}[\varphi] = \alpha$, it represents a Type I error.
- φ minimizes the Type II error

$$E_{\theta_1}[\varphi] \geq E_{\theta_1}[\varphi'] \quad \forall \varphi'$$

Cont. of proof (part of UMP)

Let φ' be another level α test, $E_{\theta_0}[\varphi'] \leq \alpha$.

Goal: $E_{\theta_1}[\varphi] \geq E_{\theta_1}[\varphi']$. Let μ be the dominating measure.

Consider

$$\int (\varphi(x) - \varphi'(x))(p_{\theta_1}(x) - kp_{\theta_0}(x)) d\mu(x) = 0$$

Claim: $p \geq 0$.

Observe:

- If $p_{\theta_1}(x) - kp_{\theta_0}(x) > 0 \Rightarrow \frac{p_{\theta_1}(x)}{p_{\theta_0}(x)} > k \Rightarrow \varphi(x) = 1$.
- If $p_{\theta_1}(x) - kp_{\theta_0}(x) < 0 \Rightarrow \varphi(x) = 0$.
- If $p_{\theta_1}(x) - kp_{\theta_0}(x) = 0 \Rightarrow \text{integrand} = 0$.

$$\Rightarrow p = 0$$

$$\Rightarrow \int (\varphi - \varphi') p_{\theta_1} d\mu = \int (\varphi - \varphi') p_{\theta_0} d\mu = k [E_{\theta_0}[\varphi] - E_{\theta_0}[\varphi']] \geq 0$$

$$\Rightarrow E_{\theta_1}[\varphi] \geq E_{\theta_1}[\varphi']$$

Part (3) UMP \Rightarrow (LR): Take φ^* a UMP test, $E_{\theta_0}[\varphi^*] = \alpha$, and let φ be the LR test with $E_{\theta_0}[\varphi] = \alpha$ with (*).

Goal: $\varphi = \varphi^*$ a.e. except on $\{r(x) = k\}$.

Define

$$x^+ = \{x : \varphi(x) > \varphi^*(x)\}$$

$$x^- = \{x : \varphi(x) < \varphi^*(x)\}$$

$$x^0 = \{x : \varphi(x) = \varphi^*(x)\}$$

$$\tilde{x} = (x^+ \cup x^-) \cap \{x : p_{\theta_1}(x) \neq kp_{\theta_0}(x)\}$$

It suffices to show $\mu(\tilde{x}) = 0$.

Like before, we have

$$(\varphi - \varphi^*)(p_{\theta_1} - kp_{\theta_0}) > 0 \text{ on } \tilde{x}$$

Thus if $\mu(\tilde{x}) > 0$,

$$\begin{aligned} \int_{\mathcal{X}} (\varphi - \varphi^*)(p_{\theta_1} - kp_{\theta_0}) d\mu &\geq 0 \\ \int_{\tilde{x}} (\varphi - \varphi^*)(p_{\theta_1} - kp_{\theta_0}) d\mu &\geq 0 \end{aligned}$$

But also

$$E_{\theta_1}[\varphi] - E_{\theta_1}[\varphi^*] > k [E_{\theta_0}[\varphi] - E_{\theta_0}[\varphi^*]] \geq 0$$

\Rightarrow Cannot be φ^* is UMP.

Example (Gaussian Location Model)

$$X_1, \dots, X_n \stackrel{\text{iid}}{\sim} \mathcal{N}(\mu, \sigma^2)$$

$$H_0 : \mu = \mu_0, \quad H_1 : \mu = \mu_1, \quad \mu_0 < \mu_1$$

Then:

$$\begin{aligned} \frac{p_1(X_1, \dots, X_n)}{p_0(X_1, \dots, X_n)} &= \exp \left(-\frac{1}{2\sigma^2} \sum_{i=1}^n (X_i - \mu_1)^2 + \frac{1}{2\sigma^2} \sum_{i=1}^n (X_i - \mu_0)^2 \right) \\ &= \exp \left(-\frac{1}{2\sigma^2} \sum_{i=1}^n (\mu_1^2 - \mu_0^2) - \frac{2(\mu_1 - \mu_0)}{\sigma^2} \sum_{i=1}^n X_i \right) \\ &= \exp \left(-\frac{n}{2\sigma^2} (\mu_1^2 - \mu_0^2) - \frac{2(\mu_1 - \mu_0)}{\sigma^2} \sum_{i=1}^n X_i \right) \geq K_\alpha \\ &\Rightarrow \frac{1}{n} \sum_{i=1}^n X_i \geq K_\alpha, \text{ some } K_\alpha \in \mathbb{R} \end{aligned}$$

To determine K_α :

$$\begin{aligned} \bar{X}_n &:= \frac{1}{n} \sum_{i=1}^n X_i \stackrel{H_0}{\sim} \mathcal{N}(\mu_0, \sigma^2/n) \\ \Rightarrow \mathbb{L} &= P_{H_0}(\bar{X}_n \geq K_\alpha) = 1 - P_{H_0}(\bar{X}_n < K_\alpha) \\ &= 1 - \Phi \left(\frac{\sqrt{n}}{\sigma} (K_\alpha - \mu_0) \right) \quad (\text{CDF for } \mathcal{N}(0, 1)) \\ \Rightarrow \text{solving for } K_\alpha &\text{ gives } K_\alpha = \mu_0 + \frac{\sigma}{\sqrt{n}} \Phi^{-1}(1 - \alpha), \\ \varphi(X_1, \dots, X_n) &= \begin{cases} 1 & \text{if } \bar{X}_n \geq \mu_0 + \frac{\sigma}{\sqrt{n}} \Phi^{-1}(1 - \alpha) \\ 0 & \text{else} \end{cases} \end{aligned}$$

Corollary

Consider simple hypothesis testing. Let φ be UMP, for level α . Then,

$$\alpha = E_{H_0}[\varphi] = E_{\theta_0}[\varphi] \leq E_{\theta_1}[\varphi]$$

Suppose $E_{\theta_1}[\varphi] = E_{\theta_1}[\varphi_0]$ then φ_0 is also UMP, $\Rightarrow \varphi_0$ is an LR test.

$$\varphi_0 = \begin{cases} 1 & \text{if } \frac{p_{\theta_1}}{p_{\theta_0}} \geq K \quad \text{a.s., some } K \\ 0 & \text{if } \frac{p_{\theta_1}}{p_{\theta_0}} < K \end{cases}$$

Also since $\varphi_0 \in \{\varphi, \beta\}$ we conclude that $p_{\theta_1} = K p_{\theta_0}$ a.s.

But

$$L = \int p_{\theta_0} d\mu = K \int p_{\theta_0} d\mu = 1 \Rightarrow K = 1$$

Correspondence theorem

$$\text{Tests} \longleftrightarrow \text{Confidence regions } C(x)$$

$$\Pr_{\theta}(\theta \in C(x)) \geq 1 - \alpha$$

$$\text{If } \Pr_{\theta}(\phi_{\theta} = 1) = \alpha$$

Theorem: Let $(P_{\theta} : \theta \in \Theta)$ be a statistical model, $\alpha \in (0, 1)$.

(i) Let $C = C(X)$ be a level- α confidence set, then

$$\phi_{\theta_0}(x) = 1 \{\theta_0 \notin C(x)\}$$

is a level- α test of $\theta = \theta_0$ vs. $\theta \neq \theta_0$.

(ii) Suppose $\{\phi_{\theta_0} : \theta_0 \in \Theta\}$ is a family of level- α tests, then

$$C(X) = \{\theta \in \Theta : \phi_{\theta}(X) = 0\}$$

is a $(1 - \alpha)$ confidence set.

Proof:

$$(i) \quad \Pr_{\theta_0}(\phi_{\theta_0} = 1) = \Pr_{\theta_0}(\theta_0 \notin C(X)) = \alpha$$

$$(ii) \quad \Pr_{\theta}(\theta \notin C(X)) = \Pr_{\theta}(\theta \notin \{\tilde{\theta} \in \Theta : \phi_{\tilde{\theta}}(X) = 0\}) = \Pr_{\theta}(\phi_{\theta}(X) = 1) \leq \alpha$$

UMPT Tests in Models with Monotone Likelihoods

Proposition: Let $\Theta \subseteq \mathbb{R}$. Consider testing $H_0 : \theta \leq \theta_0$ vs. $H_1 : \theta > \theta_0$, for some $\theta_0 \in \mathbb{R}$.

Assume there exists some test statistic $T : X \rightarrow \mathbb{R}$ and a function $h : \mathbb{R} \times \Theta \times \Theta$ such that

$$\frac{P_{\theta}(X)}{P_{\tilde{\theta}}(X)} = h(T(X), \theta, \tilde{\theta})$$

and for all $\theta \geq \tilde{\theta}$, $t \mapsto h(t, \theta, \tilde{\theta})$ is monotone increasing.

The simplest model for the relationship between Y_i and X_i assumes a linear relationship:

$$Y_i = aX_i + b + \varepsilon_i$$

for $i = 1, \dots, n$, where ε_i is centered, i.e., $E(\varepsilon) = 0$ and $\text{Var}(\varepsilon) = \sigma^2$. Suppose $\varepsilon \sim N(0, \sigma^2)$ with σ known.

The statistical model is given by

$$(\mathbb{R}, B(\mathbb{R}), (\bigotimes_{i=1}^n N(ax_i + b, \sigma^2))_{(a,b) \in \mathbb{R}^2})$$

The likelihood within the statistical model is

$$L((a, b)|y) = \prod_{i=1}^n (2\pi\sigma^2)^{-1/2} \exp\left(-\frac{1}{2\sigma^2} \sum_{i=1}^n (y_i - ax_i - b)^2\right)$$

The MLE satisfies the optimization problem

$$(\hat{a}, \hat{b}) = \arg \min_{(a,b) \in \mathbb{R}^2} \sum_{i=1}^n (y_i - (ax_i + b))^2$$

Provided that $x_i \neq x_j$ for $i \neq j$, the least squares problem has a solution with minimum given by (Gauss, 1801):

$$(\hat{a}, \hat{b}) = \left(\frac{\frac{1}{n-1} \sum_{i=1}^n (x_i - \bar{x})(y_i - \bar{y})}{\frac{1}{n-1} \sum_{i=1}^n (x_i - \bar{x})^2}, \bar{y} - \hat{a}\bar{x} \right)$$

Definition 8 (Linear Model). A random vector $Y = (Y_1, \dots, Y_n)^T \in \mathbb{R}^n$ stems from a linear model if there exists a parameter vector $\beta \in \mathbb{R}^p$, a matrix $X \in \mathbb{R}^{n \times p}$, and a random vector $\varepsilon \in \mathbb{R}^n$ such that

$$Y = X\beta + \varepsilon$$

1. A linear model is called regular if

- (a) $p \leq n$ (parameter size is smaller than sample size),
- (b) X has full rank. $\text{rank}(X) = p \leq n$ (design with full rank)
- (c) $E(\varepsilon) = 0$ (noise is controlled)
- (d) The covariance matrix is positive definite, $\Sigma = (\text{Cov}(\varepsilon_i, \varepsilon_j))_{i,j \in [n]}$

2. A linear model is called ordinary if $\Sigma = \sigma^2 E_n$ (and is usually the noise is Gaussian)

Remark 3. 1. There are several synonyms

- (a) Y a dependent variable, response, regressand
- (b) X , a independent variable, predictor, design matrix, regressor
- (c) ε Error, perturbation, reression function

2. The matrix Σ is symmetric and diagonalizable, i.e. $\Sigma = UDU^T$ for some diagonal matrix, $D = \text{diag}(\lambda_1, \dots, \lambda_n) \in \mathbb{R}^{n \times n}$

3. Positive semi-definite, i.e. $\lambda_i \geq 0$

$$\begin{aligned} \langle \Sigma u, u \rangle &= \langle E[(\varepsilon - E[\varepsilon])(\varepsilon - E[\varepsilon])^T] u, u \rangle \\ &= E[(\varepsilon - E[\varepsilon])^2] \geq 0, u \in \mathbb{R}^n \end{aligned}$$

item If Σ is positive definite ($\lambda_i > 0$) for $i = 1, \dots, n$, then there exists the inverse $\Sigma^{-1} = UD^{-1}U^T$ and $\Sigma^{-1/2} = UD^{-1/2}U^T$.

4. If X is not deterministic, we speak of random design.

In the regular linear model, $\hat{\beta}$ is called weighted least squares estimate, (LSE). if

$$\|\sigma^{-1/2}(Y - X\hat{\beta})\|^2 = \inf_{\beta \in \mathbb{R}^n} \|\sigma^{-1/2}(Y - X\beta)\|^2 = \inf_{\beta \in \mathbb{R}^n} \|\sigma^{-1/2}Y - X_{\Sigma}\beta\|^2$$

where $X_{\Sigma} = \Sigma^{-1/2}X$. $X_{\Sigma}\hat{\beta}$ is the point within the subspace,

$$U = \{X_{\Sigma}\beta \mid \beta \in \mathbb{R}^n\} \subseteq \mathbb{R}^n$$

with the smallest distance to the vector $\Sigma^{-1/2}Y$. Thus, $X_{\Sigma}\hat{\beta} = \Pi_U(\Sigma^{-1/2}Y)$ where Π_U is the orthogonal projection onto U . $\Pi_U u = u$ for all $u \in U$. $\langle \Pi_U v - v, u \rangle = 0$ for all $u \in U$ and $v \in \mathbb{R}^n$. Provided that $(X_{\Sigma}^T X_{\Sigma})^{-1}$ exists, we can confirm by direct computation that the projection satisfies

$$\Pi_U = X_{\Sigma}(X_{\Sigma}^T X_{\Sigma})^{-1} X_{\Sigma}^T$$

For $u = X_{\Sigma}\beta$ we have,

$$X_{\Sigma}(X_{\Sigma}^T X_{\Sigma})^{-1} X_{\Sigma}^T X_{\Sigma}\beta = X_{\Sigma}\beta = u$$

By symmetry,

$$\langle \Pi_U v - v, u \rangle = \langle v, \Pi_U u \rangle - \langle v, u \rangle = \langle v, u \rangle - \langle v, u \rangle = 0$$

for all $u \in U$.

Lemma 1. *Representation for the LSE Consider a regular linear model, then the LSE exists uniquely, and is given by*

$$\hat{\beta} = (X_{\Sigma}^T X_{\Sigma})^{-1} X_{\Sigma}^T \Sigma^{-1/2} Y = X_{\Sigma}^+ \Sigma^{-1/2} Y$$

Proof. $\ker(X_{\Sigma}^T X_{\Sigma})$ is invertible. Suppose that $X_{\Sigma}^T X_{\Sigma} v = 0$ ($v \in \ker(X_{\Sigma}^T X_{\Sigma})$)

$$0 = v^T X_{\Sigma}^T X_{\Sigma} v = (X_{\Sigma}^T v)^T X_{\Sigma} v = \langle X_{\Sigma} v, X_{\Sigma} v \rangle = \|X_{\Sigma} v\|^2 = \|\Sigma^{-1/2} X v\|^2 \implies \|X v\|^2 = 0 \implies v = 0$$

So then

$$\begin{aligned} X_{\Sigma}\hat{\beta} &= \Pi_U \Sigma^{-1/2} Y = X_{\Sigma}(X_{\Sigma}^T X_{\Sigma})^{-1} X_{\Sigma}^T \Sigma^{-1/2} Y \\ X_{\Sigma}^T X_{\Sigma}\hat{\beta} &= X_{\Sigma}^T X_{\Sigma}(X_{\Sigma}^T X_{\Sigma})^{-1} X_{\Sigma}^T \Sigma^{-1/2} Y \\ &\implies \hat{\beta} = (X_{\Sigma}^T X_{\Sigma})^{-1} X_{\Sigma}^T \Sigma^{-1/2} Y \end{aligned}$$

□

Remark 4. 1. If $p > n$, then $(X_{\Sigma}^T X_{\Sigma})^{-1}$ does not exist and the LSE is not unique.

$$\left\{ \beta \cdot \|\Sigma^{-1/2} Y - X_{\Sigma}\beta\|^2 = 0 \right\}$$

is a $p - n$ dim subspace and each solution interpolates the data

Theorem 1. *Optimality of the LSE, Gauss-Markov Theorem Consider an ordinary linear model for $\sigma > 0$, then*

1. The least squares estimator $\hat{\beta} = (X^T X)^{-1} X^T Y$ is linear and the unbiased parameter for the parameter β .
2. For the desired parameter $\alpha = \langle \beta, v \rangle$ for $v \in \mathbb{R}$, the estimator $\hat{\alpha} = \langle \hat{\beta}, v \rangle$ is the best linear unbiased estimator (BLUE), meaning that $\hat{\alpha}$ has the optimal value within the class of linear unbiased estimators for α
3. $\hat{\sigma}^2 = \frac{\|Y - X\hat{\beta}\|^2}{n-p}$ is an unbiased estimator of σ^2

Proof.

$$\hat{\beta}(y + \tilde{y}) = \hat{\beta}(y) + \hat{\beta}(\tilde{y}) \text{ for } y, \tilde{y} \in \mathbb{R}^n$$

$$E[\hat{\beta}] = (X^T X)^{-1} X^T E[Y] \tag{1}$$

$$= (X^T X)^{-1} X^T E[X\beta + \varepsilon] \tag{2}$$

$$= (X^T X)^{-1} (X^T X)\beta \tag{3}$$

$$= \beta \tag{4}$$

Suppose that $\tilde{\alpha}$ is some other linear unbiased estimator of α . Since the estimator is linear, there exists some element w such that $\tilde{\alpha} = \langle y, w \rangle$

$$\langle \beta, v \rangle = \alpha = E[\tilde{\alpha}] = E[\langle y, w \rangle] = \langle X\beta, w \rangle = \langle \beta, X^T w \rangle$$

This implies that $v = X^T w$, therefore we have,

$$\text{Var} = \text{Var}(\langle x\beta, w \rangle + \langle \varepsilon, w \rangle) \quad (5)$$

$$= \text{Var}(\langle \varepsilon, w \rangle) + E \left[\left(\sum_{i=1}^n \varepsilon_i w_i \right)^2 \right] \quad (6)$$

$$= \sigma^2 \sum_{i=1}^p w_i^2 = \sigma^2 \|w\|^2 \quad (7)$$

$$\text{Var}(\hat{\alpha}) = E[\langle \hat{\beta} - \beta, v \rangle^2] \quad (8)$$

$$= E[\langle (X^T X)^{-1} X^T \beta + (X^T X)^{-1} X^T \varepsilon - \beta, v \rangle^2] \quad (9)$$

$$= E[\langle (X^T X)^{-1} X^T \varepsilon, v \rangle^2] \quad (10)$$

$$= \sigma^2 \|X(X^T X)^{-1} v\|^2 = \sigma^2 \|X(X^T X)^{-1} X^T w\|^2 \quad (11)$$

$$= \sigma^2 \|\Pi_u w\|^2 \quad (12)$$

Thus, $\text{Var}(\hat{\alpha}) \leq \text{Var}\tilde{\alpha}$ □

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Recall linear model

$$Y = X\beta + \varepsilon$$

where $\text{cov}(\varepsilon) = \Sigma$.

OLD: $\hat{\beta} = (X_\Sigma^T X_\Sigma)^{-1} X_\Sigma^T \Sigma^{-1/2} Y$.

$X\hat{\beta}$ = Projection of $\Sigma^{-1/2} Y$ onto $\text{span} \{X_{\varepsilon,1}, \dots, X_{\varepsilon,p}\}$

Theorem 2 (Gauss-Markov). 1. $\hat{\beta}_{OLS}$ is the best linear unbiased est (BLUE)

2. $\alpha_i = \langle \beta, v \rangle$ is BLUE.

3. $\hat{\sigma}^2 = \frac{\|Y - X\hat{\beta}\|^2}{n-p}$ is unbiased est for $\sigma^2 > 0$

$$\begin{pmatrix} y_1 \\ y_2 \\ \vdots \\ y_n \end{pmatrix} = \begin{pmatrix} x_1 \\ x_2 \\ \vdots \\ x_n \end{pmatrix} \begin{pmatrix} \beta_1 \\ \beta_2 \\ \vdots \\ \beta_p \end{pmatrix}^T + \begin{pmatrix} \varepsilon_1 \\ \varepsilon_2 \\ \vdots \\ \varepsilon_n \end{pmatrix} \quad \text{Where our data is } (Y_i, X_i)_{i=1}^n \in (\mathbb{R} \times \mathbb{R}^p)^{\otimes p}$$

Remark 5. Is this an iid model? Depends!

1. Typically ε_i are iid.

2. If X_i are random then "random design".

3. If X_i are iid, then linear model is iid model.

4. If X_i are deterministic, then not iid model.

$$\beta \mapsto \|Y - X\hat{\beta}\|.$$

Proof. This is a continuation of point 3 in our theorem above.

We already introduced $\Pi_U = X(X^T X)^{-1} X^T$ projection onto col space U of X . Thus $I_n - \Pi_U$ is another projection operator, onto U^\perp (orthogonal complement),

$$U^\perp = \{z \in \mathbb{R}^n \mid \langle z, X_k \rangle \forall k = 1, \dots, p\}.$$

Choose a basis e_1, \dots, e_{n-p} , orthonormal, of U^\perp , then

$$(I_n - \Pi_U)z = \Pi_{U^\perp} z = \sum_{k=1}^{n-p} \langle z, e_k \rangle e_k.$$

$$\|Y - X\hat{\beta}\| = \|Y - \underbrace{X(X^T X)^{-1} X^T Y}_{\Pi_U}\|^2 \quad (13)$$

$$= \|(I_n - \Pi_U)Y\|^2 \quad (14)$$

$$= \|(I_n - \Pi_U)(X\beta + \varepsilon)\|^2 \quad (15)$$

$$= \|(I_n - \Pi_U)\varepsilon\|^2 \quad (16)$$

$$= \sum_{i=1}^{n-p} \langle \varepsilon, e_i \rangle^2 \quad (17)$$

$$(18)$$

Hence,

$$E[\|Y - X\hat{\beta}\|^2] = \sum_{i=1}^{n-p} E[\langle \varepsilon, e_i \rangle^2] = n - p \implies E[\hat{\sigma}] = n - p$$

□

Remark 6. Recall the $N(\mu, \sigma^2)$ model, where the MLE is

$$\hat{\mu} = \frac{1}{n} \sum_{i=1}^n X_i, \quad \hat{\sigma}^2 = \frac{1}{n} \sum_{i=1}^n (X_i - \hat{\mu})^2.$$

The unbiased estimator for σ^2 was $\frac{1}{n-1} \sum_{i=1}^n (X_i - \hat{\mu})^2$. This is related to the $n - p$ factor in point 3.

Remark 7. 1. If linearity is dropped, there exists better estimators than $\hat{\beta}_{OLS}$. For example a constant estimator, $\hat{\beta} = \beta^*$

2. The MSE of $\hat{\beta}_{OLS}$ is

$$E[\|\hat{\beta}_{OLS} - \beta\|^2] = E\left[\sum_{i=1}^p \langle \hat{\beta}_{OLS} - \beta, \underbrace{e_i}_{\text{ONB of } \mathbb{R}^n} \rangle^2\right] = \sum_{i=1}^p \text{Var}_\beta(\langle \hat{\beta}_{OLS}, e_i \rangle) = \sum_{i=1}^p \sigma^2 \|X(X^T X)^{-1} e_i\|^2$$

We say X satisfies orthogonal design if

$$X^T X = nI_p$$

"The different covariants are uncorrelated." $(X^T X)_{ij} = \langle X_i, X_j \rangle = n\delta_{ij}$ For orthogonal design,

$$E_\beta[\|\hat{\beta}_{OLS} - \beta\|^2] = \frac{1}{n^2} \sigma^2 \sum_{i=1}^p \underbrace{\|Xe_i\|^2}_n = \frac{\sigma^2 p}{n}.$$

and this is equal to noise level times the number of parameters, divided by the number of data points.

Theorem 3 (Bayes in Linear Models). Consider a linear model $Y = X\beta + \varepsilon$, and $\varepsilon \sim N(0, \sigma^2 I_n)$ with $\sigma > 0$ known and $\beta \sim N(m, M)$ where $m \in \mathbb{R}^p, M \in \mathbb{R}^{p \times p}$ positive semi definite. Then, the posterior $\Pi(\beta|Y, X)$ is given by

$$\Pi(\beta|Y, X) = N(\mu_{past}, \Sigma_{past}) \text{ for}$$

$$\mu_{past} = \sigma_{past}^{-2} X^T y + M^{-1} m \quad \Sigma_{past} = (\sigma_{past}^{-2} X^T X + M^{-1})^{-1}$$

Remark 8. Σ_{past} independent of Y . For " $M^{-2} \rightarrow 0$ ", then " $\mu_{past} \rightarrow \hat{\beta}_{OLS}$ "

Proof.

$$L(X, Y, \beta) \pi(\beta) \propto \exp \left(-\frac{1}{2\sigma^2} \|Y - X\beta\|^2 - \frac{1}{2} (\beta - m)^T M^{-1} (\beta - m) \right)$$

We want this to be proportional to $\exp \left(-\frac{1}{2} (\beta - \mu_{\text{past}})^T \sigma_{\text{past}}^{-1} (\beta - \mu_{\text{past}}) \right)$.

Now,

$$\exp \left(-\frac{1}{2} (\beta - \mu_{\text{past}})^T \sigma_{\text{past}}^{-1} (\beta - \mu_{\text{past}}) \right) \propto \exp \left(-\frac{1}{\sigma^2} \beta^T X^T X \beta - \frac{1}{2} \beta^T M^{-1} \beta + \frac{1}{\sigma^2} \beta^T X^T Y + \beta^T M^{-1} m \right)$$

and this is equal to

$$\exp \left(-\frac{1}{2} \beta^T \left(\frac{1}{\sigma^2} X^T X + M^{-1} \right) \beta + \beta^T \left(\frac{1}{\sigma^2} X^T Y + M^{-1} m \right) \right)$$

and this is

$$\propto \exp \left(-\frac{1}{2} (\beta - \mu_{\text{past}})^T \sigma_{\text{past}}^{-1} (\beta - \mu_{\text{past}}) \right)$$

□

Corollary 1. For $\ell = \|\cdot\|^2$, the Bayes estimator is $\hat{\beta}_{\Pi} = \mu_{\text{past}}$

Proposition 1. Consider the previous setting (from the theorem), with $m = 0$, and $M = \tau^2 I_p$ (centered, isotropic, normal prior). The, $\mu_{\text{past}} = \hat{\beta}_{\Pi}$ minimizes

$$\beta \mapsto \|Y - X\beta\|_{\mathbb{R}^n}^2 + \underbrace{\frac{\sigma^2}{\tau^2} \|\beta\|_{\mathbb{R}^p}^2}_{\text{"penalty" or "regularization"}}$$

Proof.

□

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Proof:

Take gradient of $\mathcal{J}(\beta)$ w.r.t. β :

$$\nabla_{\beta} \mathcal{J}(\beta) = 2\mathbf{X}^{\top} (\mathbf{Y} - \mathbf{X}\beta) + \frac{2\sigma^2}{\tau^2} \beta$$

Set = 0:

$$\Rightarrow \nabla_{\beta} \mathcal{J}(\beta) = 2(\mathbf{X}^{\top} \mathbf{X} + \frac{\sigma^2}{\tau^2} \mathbf{I}) \beta - 2\mathbf{X}^{\top} \mathbf{Y} = 0$$

$$\Rightarrow \beta = (\mathbf{X}^{\top} \mathbf{X} + \frac{\sigma^2}{\tau^2} \mathbf{I})^{-1} \mathbf{X}^{\top} \mathbf{Y}$$

Posterior mean:

$$\begin{aligned} \mu_{\text{post}} &= \Sigma_{\text{post}}^{-1} (\mathbf{X}^{\top} \mathbf{Y} + \mathbf{M}_0^{-1} \mu_0) \\ &= (\sigma^{-2} \mathbf{X}^{\top} \mathbf{X} + \tau^{-2} \mathbf{I}_p)^{-1} \sigma^{-2} \mathbf{X}^{\top} \mathbf{Y} \\ &= (\mathbf{X}^{\top} \mathbf{X} + \frac{\sigma^2}{\tau^2} \mathbf{I})^{-1} \mathbf{X}^{\top} \mathbf{Y} \end{aligned}$$

Remark:

β is defined even if $\text{rank}(\mathbf{X}) < p$, in particular even for $n < p$.

Definition:

$$\hat{\beta}_{\text{ridge}} = \underset{\beta \in \mathbb{R}^p}{\operatorname{argmin}} \|\mathbf{Y} - \mathbf{X}\beta\|^2 + \lambda \|\beta\|^2,$$

is called a **Ridge Regression** estimator. Here, $\lambda > 0$ is called a regularization parameter. $\hat{\beta}_{\text{ridge}}$ is always uniquely defined.

For $\mathbf{Y} = \mathbf{X}\beta + \varepsilon$, UM:

$$\hat{\beta}_{\text{ridge}} = (\mathbf{X}^\top \mathbf{X} + \lambda \mathbf{I}_p)^{-1} \mathbf{X}^\top \mathbf{Y}.$$

Estimator independent of σ^2 .

Proposition:

MSE of $\hat{\beta}_{\text{ridge}}$.

Consider a linear model with $\varepsilon \sim N(0, \sigma^2 \mathbf{I}_n)$, $\sigma^2 > 0$ known, and $\mathbf{X}^\top \mathbf{X} = n \mathbf{I}_p$ (orthonormal design).

Let $\mathcal{J} := \langle \beta, \mathbf{v} \rangle$ for $\mathbf{v} \in \mathbb{R}^p$, and:

$$\delta_{\text{ridge}} = \langle \hat{\beta}_{\text{ridge}}, \mathbf{v} \rangle.$$

Then:

1.

$$\mathbb{E}_{\beta}[(\delta_{\text{ridge}} - \mathcal{J})^2] = (1 + \lambda)^{-2} \langle \beta, \mathbf{v} \rangle^2 + \frac{\sigma^2}{n} \|\mathbf{v}\|^2 (1 + \lambda)^{-2}.$$

2.

$$\mathbb{E}_{\beta}[\|\hat{\beta}_{\text{ridge}} - \beta\|^2] = (1 + \lambda)^{-2} \|\beta\|^2 + \frac{p\sigma^2}{n} \frac{1}{(1 + \lambda)^2}.$$

We have:

$$\begin{aligned} \hat{\beta}_{\text{ridge}} &= (\mathbf{X}^\top \mathbf{X} + \lambda \mathbf{I}_p)^{-1} \mathbf{X}^\top \mathbf{Y} \\ &= (n \mathbf{I}_p + \lambda \mathbf{I}_p)^{-1} \mathbf{X}^\top \mathbf{Y} \\ &= \frac{1}{(1 + \frac{\lambda}{n})} (\mathbf{X}^\top \mathbf{X} \beta + \mathbf{X}^\top \varepsilon), \end{aligned}$$

where $\varepsilon \sim N(0, \sigma^2 \mathbf{I}_n)$.

$$\begin{aligned} &= \frac{1}{1 + \frac{\lambda}{n}} (\mathbf{X}^\top \mathbf{X} \beta + \mathbf{X}^\top \varepsilon), \\ &= \frac{1}{1 + \frac{\lambda}{n}} \beta + \frac{1}{1 + \frac{\lambda}{n}} \mathbf{X}^\top \varepsilon. \end{aligned}$$

Bias-Variance Decomposition:

$$\begin{aligned} \mathbb{E}[(\hat{\beta}_{\text{ridge}} - \mathcal{J})^2] &= (\mathbb{E}[\hat{\beta}_{\text{ridge}}] - \mathcal{J})^2 + \operatorname{Var}(\hat{\beta}_{\text{ridge}}). \\ &= ((1 + \frac{\lambda}{n})^{-1} \langle \beta, \mathbf{v} \rangle)^2 + \frac{\lambda^2}{(1 + \lambda)^2} \operatorname{Var}(\mathbf{X}^\top \varepsilon, \mathbf{v}). \end{aligned}$$

Observe:

$$(1 + \frac{\lambda}{n})^{-1} = \frac{1}{(1 + \frac{\lambda}{n})}.$$

Also:

$$\operatorname{Var}(\mathbf{X}^\top \varepsilon, \mathbf{v}) = \mathbf{v}^\top \mathbf{X} \operatorname{Cov}(\varepsilon) \mathbf{X}^\top \mathbf{v} = \sigma^2 \|\mathbf{v}\|^2.$$

Corollary:

Under the same assumptions:

$$\begin{aligned} \mathbb{E}[\|\hat{\beta}_{\text{ridge}} - \beta\|^2] &= \mathbb{E} \left[\sum_{k=1}^p (\langle \beta, \mathbf{e}_k \rangle - \beta_k)^2 \right] \\ &= \frac{1}{(1 + \frac{\lambda}{n})^2} \|\beta\|^2 + \frac{p\sigma^2}{n(1 + \frac{\lambda}{n})^2}. \end{aligned}$$

Remark:

For small $\|\beta\|$, Ridge \rightarrow OLS. The optimal choice of λ depends on $\|\beta\|$.

1.7 Confidence Sets & Tests in Linear Model:

The estimators we studied are independent of σ^2 , but uncertainty quantification will depend on σ^2 !

Assume $\varepsilon \sim N(0, \sigma^2 \mathbf{I}_n)$ throughout.

Easy Case:

For $\sigma^2 > 0$ known:

$$\hat{\beta}_{\text{OLS}} \sim N(\beta, (\mathbf{X}^\top \mathbf{X})^{-1}).$$

Indeed:

$$\text{Cov}((\mathbf{X}^\top \mathbf{X})^{-1} \mathbf{X}^\top \varepsilon) = (\mathbf{X}^\top \mathbf{X})^{-1}.$$

And for $\mathcal{J} = \langle \beta, \nu \rangle$,

$$\hat{\mathcal{J}} = \langle \hat{\beta}_{\text{OLS}}, \nu \rangle \sim N(\mathcal{J}, \sigma^2 \nu^\top (\mathbf{X}^\top \mathbf{X})^{-1} \nu).$$

Then a 95% confidence set for \mathcal{J} is:

$$I_{95\%}(\mathcal{J}) = \left[\hat{\mathcal{J}} \pm 1.96 \sqrt{\nu^\top (\mathbf{X}^\top \mathbf{X})^{-1} \nu} \right].$$

Notes on t - and F -distributions:

BUT: Normally, σ^2 is unknown. Replace σ by its estimator $\hat{\sigma}$. We need the t - and F -distributions.

Definitions:

Definition (t-distribution): The t -distribution with $n \geq 1$ degrees of freedom on \mathbb{R} has density:

$$f_n(x) = C_n \left(1 + \frac{x^2}{n} \right)^{-\frac{n+1}{2}},$$

where C_n is the normalizing constant.

Note: For $n = 1$:

$$f_1(x) = C_1 \frac{1}{1 + x^2},$$

which corresponds to the **Cauchy distribution**.

Definition (F-distribution): The F -distribution with $(m, n) \in \mathbb{N}^2$ degrees of freedom has density:

$$f_{m,n}(x) = C_{m,n} \frac{x^{\frac{m}{2}-1}}{(mx + n)^{\frac{m+n}{2}}}, \quad x \in (0, \infty),$$

where $C_{m,n}$ is the normalizing constant.

Why is this useful?

Lemma: Let $X_1, \dots, X_n, Y_1, \dots, Y_n$ be i.i.d. $N(0, \Delta)$ random variables. Then:

1.

$$T_n := \frac{X_n}{\sqrt{\frac{1}{n} \sum_{i=1}^n Y_i^2}} \sim t_n.$$

2.

$$F_{m,n} := \frac{\frac{1}{m} \sum_{i=1}^m X_i^2}{\frac{1}{n} \sum_{j=1}^n Y_j^2} \sim F_{m,n}.$$

Remarks:

1. The t -distribution arises when considering the "empirical mean" and "empirical variance."
2. For $n \rightarrow \infty$, $T_n \xrightarrow{d} N(0, 1)$.

Proof:**(b) Observe:**

$$T_n^2 = F_{1,n}.$$

By a change of measure ($y \mapsto y^2$ in $(0, \infty)$):

$$f_{F_{m,n}}(x) = f_{F_{m,n}}(x^2)2x, \quad x > 0.$$

Since t is symmetric around 0, we obtain for all $x \in \mathbb{R}$:

$$f_{T_n}(x) = f_{F_{m,n}}(x^2)|x| = C_n \left(1 + \frac{x^2}{n}\right)^{-\frac{n+1}{2}}.$$

It remains to show the claim for $F_{m,n}$.

Let:

$$X = \sum_{i=1}^m X_i^2, \quad Y = \sum_{j=1}^n Y_j^2.$$

Then:

$$X \sim \chi_m^2, \quad Y \sim \chi_n^2,$$

where the density of χ_m^2 is:

$$f(x) \propto x^{m/2-1} e^{-x/2}, \quad x > 0.$$

Derivation:

Writing $W = \frac{X}{Y}$, we have:

$$\mathbb{P}\left(\frac{X}{Y} < z\right) = \int_0^\infty \int_0^{zy} 1 f_X(x) f_Y(y) dx dy.$$

Substituting $x = wy$, we get:

$$\begin{aligned} &= \int_0^\infty \int_0^z 1 f_X(wy) f_Y(y) y dw dy \\ &= \int_0^\infty f_X(zy) f_Y(y) y dy \\ &\propto \int_0^\infty (zy)^{\frac{m}{2}-1} y^{\frac{n}{2}-1} e^{-(z+y)/2} dy. \end{aligned}$$

Change of Variable:

Let $a = \frac{z}{z+1}y$, then:

$$\begin{aligned} &\propto \int_0^\infty \left(\frac{z}{z+1}\right)^{\frac{m}{2}} a^{\frac{m}{2}-1} e^{-\frac{z}{z+1}a} \frac{1}{z+1} da \\ &\propto z^{\frac{m}{2}-1} (z+1)^{-\frac{m+n}{2}} \int_0^\infty a^{\frac{m}{2}-1} e^{-a} da. \end{aligned}$$

It follows:

$$\begin{aligned} \frac{\partial}{\partial z} \mathbb{P}\left(\frac{X}{Y} < z\right) &= f_{X,Y}(z) = \int_0^\infty f_X(zy) f_Y(y) \frac{1}{y} dy \\ &\propto z^{\frac{m}{2}-1} (z+1)^{-\frac{m+n}{2}}. \end{aligned}$$

Change of Variable:

Let $F = \frac{X}{Y}$, given $f_F(z) = \frac{m}{n} f_{X,Y} \left(\frac{m}{n} z \right) = f_{m,n}(z)$.

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t -distribution $\cdot t_n(x) \propto \left(\frac{n^2}{n} + 1 \right)^{-(n+1)/2}$

$$F \cdot f_{m,n}(x) \propto \frac{x^{m/2-1}}{(n+mx)^{(m+n)/2}}$$

In the linear model $Y = X\beta + \varepsilon$, $\varepsilon \sim N(0, \sigma^2 I_n)$,

$$\hat{\beta} = (X^T X)^{-1} X^T Y \sim N(0, \sigma^2 A)$$

where

$$\sigma^2 = \hat{\sigma}^2 = \left\| \frac{Y - X\hat{\beta}}{n-p} \right\|^2 \sim \sigma^2 \frac{\chi^2(n-p)}{n-p}$$

We now have

$$t_n = \frac{N(0,1)}{\chi^2(n)n} \quad f_{m,n} = \frac{n\chi^2(m)}{m\chi^2(n)}$$

Lemma 2. Let $\xi \sim N(0, I_n)$, a random variable in \mathbb{R}^n , and let $R \in \mathbb{R}^{n \times n}$ be an orthogonal projection ($R = R^2, R = R^T$), with $\text{rank}(R) = r \leq n$.

1. $\xi^T R \xi = \|R\xi\|^2 \sim \chi^2(r)$.
2. If $B \in \mathbb{R}^{p \times n}$ is such that $BR = 0$, then $B\xi$ is independent from $R\xi$
3. If $S \in \mathbb{R}^{n \times n}$ is another orthogonal projection, $\text{rank}(S) = s \leq n$ and $RS = 0$, then

$$\frac{s}{r} \frac{\xi^T R \xi}{\xi^T S \xi} \sim F(r, s)$$

Proof. 1. Since R is an orthogonal projection, there exists an orthogonal matrix $T^T = T^{-1}$ such that

$$R = T \begin{pmatrix} I_r & \mathbf{0} \\ \mathbf{0} & \mathbf{0} \end{pmatrix} T^T = T D_r T^T.$$

Then we have $T^T \sim N(0, T^T T) = N(0, I_n)$.

$$\xi^T R \xi = \xi^T (T D_r T^T) \xi = (T^T \xi)^T D_r T^T \xi = \sum_{i=1}^n (T^T \xi)_i^2 \sim \chi^2(r).$$

2. Let $A_1 = B\xi$, $A_2 = R\xi$, then

$$\text{Cov}(A_1, A_2) = \text{Cov}(B\xi, R\xi) = B \text{Cov}(\xi, \xi) R^T = B R^T = B R = 0$$

3. By (2), we know $S\xi$ and $R\xi$ are independent. By (1), $\xi^T S \xi \sim \chi^2(s)$, $\xi^T R \xi \sim \chi^2(r)$. The claim follows from the definition of $F(r, s)$. □

Theorem 4. Linear Model Confidence Sets -unknown σ^2 Assume regular linear model, $Y = X\beta + \varepsilon$, $\text{rank}(X) = p \leq n$, $\varepsilon \sim N(0, \sigma^2 I_n)$. Let $\alpha \in (0, 1)$

1. Let $q_{F_{p,n-p}, 1-\alpha}$ be the $1-\alpha$ quantile of $F_{p,n-p}$ distribution. Then $C(Y, X) = \left\{ \beta \in \mathbb{R}^p \mid \frac{\|X(\beta - \hat{\beta}_{OLS})\|^2}{p\hat{\sigma}^2} \leq q_{F_{p,n-p}, 1-\alpha} \right\}$ is a level $1-\alpha$ confidence set.
2. Let $\alpha = \langle \beta, v \rangle$, for some $v \in \mathbb{R}^p$. Then a $1-\alpha$ confidence set is

$$C = C(Y, X) = \left\{ \alpha \in \mathbb{R} \mid \left| \frac{\alpha - \hat{\alpha}}{\hat{\sigma} \sqrt{v^T (X^T X)^{-1} v}} \right| < q \right\}$$

where $\hat{\alpha} = \langle \hat{\beta}_{OLS}, v \rangle$ and q is the $1-\alpha/2$ quantile of t_n .

Proof. 1. We know $X\hat{\beta}_{OLS} = \Pi_U Y = \Pi_U X\beta + \Pi_U \varepsilon = X\beta + \Pi_U \varepsilon$ Moreover,

$$\hat{\sigma}^2 = \frac{\|X(\beta - \hat{\beta}_{OLS})\|^2}{n-p} = \frac{\|(I_n - \Pi_U)Y\|^2}{n-p} = \frac{\|\Pi_{U^\perp} Y\|^2}{n-p} = \frac{\|\Pi_{U^\perp} \varepsilon\|^2}{n-p}$$

This implies

$$\frac{\|X\beta - X\hat{\beta}_{OLS}\|^2}{p\hat{\sigma}^2} = \frac{(n-p)\|\Pi_U \varepsilon\|^2}{p\|\Pi_{U^\perp} \varepsilon\|^2} \sim \frac{(n-p)\sigma^2\chi^2(p)}{p\sigma^2\chi^2(n-p)} \sim F(p, n-p).$$

2. We know

$$\hat{\sigma} = \langle \hat{\beta}_{OLS}, v \rangle = v^T \hat{\beta}_{OLS} \sim v^T N(\beta, (X^T X)^{-1} \sigma^2) = N(\alpha, v^T (X^T X)^{-1} v \sigma^2)$$

And this implies

$$\frac{\alpha - \hat{\alpha}}{\sigma \sqrt{v^T (X^T X)^{-1} v}} \sim N(0, 1).$$

Finally, also, as in (1), $\hat{\sigma}^2 \sim \sigma^2 \chi^2(n-p)$. This implies

$$\frac{\alpha - \hat{\alpha}}{\hat{\sigma} \sqrt{v^T (X^T X)^{-1} v}} \sim t_{n-p}.$$

□

9.1 The t - and F -test

Remark 9 (Method (t-test)). In a regular linear model with $\varepsilon \sim N(0, \sigma I_n)$, consider $H_0 : \gamma = \gamma_0$ vs $H_1 : \gamma \neq \gamma_0$ ($\gamma = \langle \beta, v \rangle$). The two sided t -test is

$$\varphi_{\alpha_0}(Y, X) = \mathbf{1}(\{|T_{\alpha_0, n-p}(Y, X)| > q\}),$$

where

$$T_{\alpha_0, n-p} = \frac{\alpha_0 - \hat{\alpha}}{\hat{\sigma} \sqrt{v^T (X^T X)^{-1} v}}$$

and q is the $1 - \alpha/2$ -quantile of t_{n-p} .

Remark 10 (Method (F-test)). Same setting as before for t -test, $H_0 : \beta = \beta_0$ vs $H_1 : \beta \neq \beta_0$ since $\beta_0 \in \mathbb{R}^p$. Then the F -test is

$$\varphi_{\beta_0}(Y, X) = \mathbf{1}(|F_{\beta_0, n-p}(Y, X)| > q)$$

where

$$F_{\beta_0, n-p}(Y, X) = \frac{\|X(\beta - \hat{\beta}_{OLS})\|^2}{p\hat{\sigma}^2}$$

and $q = (1 - \alpha)$ -quantile of $F_{p, n-p}$.

9.2 General linear hypothesis testing problems

Definition 9. A linear hypothesis testing pb. is of the form $H_0 : K\beta = d$ vs $H_1 : K\beta \neq d$, where $K \in \mathbb{R}^{r \times p}$ with $\text{rank}(K) = r \leq p$, $d \in \mathbb{R}^r$. In other words "r linear constraints on β "

K is called the "contrast matrix"

Theorem 5. Assume regular linear model, with $\varepsilon \sim N(0, \sigma^2 I_n)$, and consider $H_0 : K\beta = d$ vs. $K\beta \neq d$.

Defin residual sum of squares as $RSS = \|Y - X\hat{\beta}_{OLS}\|^2$ and $RSS_{H_0} = \|Y - X\hat{\beta}_{H_0}\|^2$ and $\hat{\beta}_{H_0}$ over $\{\beta : K\beta = d\}$.

Proof. 1.

$$\hat{\beta}_{H_0} = \hat{\beta}_{OLS} - (X^T X)^{-1} K^T (K(X^T X)^{-1} K^T)^{-1} (K\hat{\beta}_{OLS} - d)$$

$$2. RSS_{H_0} - RSS = (K\hat{\beta}_{OLS} - d)(K(X^T X)^{-1} K^T)^{-1} (K\hat{\beta}_{OLS} - d), \quad \frac{RSS_{H_0} - RSS}{\sigma^2} \sim \chi^2(r)$$

3. Define

$$F = \frac{n-p}{r} = \frac{RSS_{H_0} - RSS}{RSS} = \frac{RSS_{H_0} - RSS}{r\hat{\sigma}^2} \sim F_{r, n-p}$$

under H_0 .

□

10 Lecture 12

ANOVA

linear model, factor/category, F -test for equality of means, $Y_{i,j} = \mu_i + \varepsilon_{ij}$ for $i = 1, \dots, k$ and $j = 1, \dots, n_i$.

First a note, $X^T X = \|X\|_{\mathbb{R}^n}^2$

Theorem 6. In the ANOVA model with $\varepsilon_{ij} \sim N(0, \sigma^2)$:

1. The OLS estimate is

$$\hat{\mu} = (\bar{y}_{1,\cdot}, \dots, \bar{y}_{k,\cdot}) \quad \text{Recall } \bar{y}_{i,\cdot} = \frac{1}{n_i} \sum_{j=1}^{n_i} y_{ij}$$

2.

$$\frac{SSW}{\sigma^2} = \frac{1}{\sigma^2} \sum_i \sum_j (y_{ij} - \bar{y}_{i,\cdot})^2 \sim \chi^2(n - k)$$

3. Under $H_0: \mu_0 = \mu_1 = \dots = \mu_k$, $\frac{SSB}{\sigma^2} = \frac{1}{\sigma^2} \sum_i n_i (\bar{y}_{i,\cdot} - \bar{y})^2 \sim \chi^2(k - 1)$

4. SSW and SSB are independent and under H_0 ,

$$\frac{n - k}{k - 1} \frac{SSB}{SSW} \sim F(k - 1, n - k)$$

Proof. (a) We have $\hat{\mu} = (X^T X)^{-1} X^T Y$, with

$$(X^T X)^{-1} = \begin{pmatrix} \frac{1}{n_1} & \dots & 0 \\ 0 & \dots & 0 \\ 0 & \dots & \frac{1}{n_k} \end{pmatrix}$$

and this implies that

$$\hat{\mu} = \begin{pmatrix} \frac{1}{n_1} & \dots & 0 \\ 0 & \dots & 0 \\ 0 & \dots & \frac{1}{n_k} \end{pmatrix} \begin{pmatrix} \mathbb{1} & \dots & 0 \\ 0 & \dots & 0 \\ 0 & \dots & \mathbb{1} \end{pmatrix} Y = \begin{pmatrix} \frac{1}{n_1} & \dots & 0 \\ 0 & \dots & 0 \\ 0 & \dots & \frac{1}{n_k} \end{pmatrix} \begin{pmatrix} \sum_j Y_{1,j} \\ \vdots \\ \sum_j Y_{k,j} \end{pmatrix} = \begin{pmatrix} \bar{Y}_{1,\cdot} \\ \vdots \\ \bar{Y}_{k,\cdot} \end{pmatrix}$$

(b)

$$SSW = \|Y - X\hat{\mu}\|_{\mathbb{R}^n}^2 = \sum_i \sum_j (y_{ij} - \bar{y}_{i,\cdot})^2 = RSS \implies \frac{SSW}{\sigma^2} \sim \chi^2(n - k).$$

(c) We know that $SSW = RSS$ and we know that $SSW + SSB = SST = \sum_{i,j} (y_{ij} - \bar{y}_{i,\cdot})^2$.

We also know $\frac{RSS_{H_0} - RSS}{\sigma^2} \sim \chi^2(k - 1)$ from before, it suffices to show $SST = RSS_{H_0}$,

$$RSS_{H_0} = \min_{\mu \in \mathbb{R}} \|Y - \mu\|_{\mathbb{R}^n}^2 = \|Y - \bar{Y}_{\cdot,\cdot}\|_{\mathbb{R}^n}^2 = SST.$$

(d) Follows from general lin hypotheses testing theorem, Theorem 2.2.30 in Methoden der Statistik book.

□

10.1 Exponential Families

General Model($P_\theta : \theta \in \Theta$) \supseteq Exp. families \supseteq Linear Model

Regularity Assumptions:

Let $(P_\theta : \theta \in \Theta)$ be a statistical model

1. Dominated, there exists μ such that $P_\theta \ll \mu$ for all $\theta \in \Theta$
2. $\Theta \in \mathbb{R}^p$ is an open set $p \geq 1$.
3. Likelihood $p_\theta(x) > 0$ for all $\theta \in \Theta, x \in X$, in particular $\log p_\theta(x)$ is well defined.

Definition 10. *Score* The score vector is $U_\theta(x) = \nabla_\theta \log p_\theta(x) = \begin{pmatrix} \frac{\partial}{\partial \theta_1} \log p_\theta(x) \\ \vdots \\ \frac{\partial}{\partial \theta_p} \log p_\theta(x) \end{pmatrix}$ whenever it exists

Definition 11. *Fisher Information* For $\theta \in \Theta$, the FI, whenever it exists, is $I(\theta) = E(U_\theta(x)U_\theta(x)^T) \in \mathbb{R}^{p \times p}$

More Regularity Assumptions

1. $p_\theta(x)$ is twice differentiable, in particular, $U_\theta(x)$ is well defined.
2. $E_\theta[||U_\theta(x)||_{\mathbb{R}^p}^2] < \infty$ for all $\theta \in \Theta$, so $I(\theta)$ is well defined.
- 3.

$$\int h(x) \nabla_\theta p_\theta(x) \mu(dx) = \nabla_\theta \int h(x) \mu(dx)$$

for relevant $h(x)$.

Lemma 3. ($P_\theta \in \Theta$) regular model (as above),

1. $E_\theta(U_\theta(x)) = 0$
2. $I(\theta) = \text{Cov}(U_\theta)$

Proof.

$$E_\theta[U_\theta(x)] = \int_X \frac{\nabla_\theta p_\theta(x)}{p_\theta(x)} p_\theta(x) d\mu(x) = \int_X \nabla_\theta p_\theta(x) d\mu(x) = \nabla_\theta [\theta \mapsto 1] = 0$$

□

Definition 12. *Uniform minimum variance unbiased estimators (UMVUE)* Let $p(\theta) \in \mathbb{R}$ be some quantity of interest, $T: X \rightarrow \mathbb{R}$ is a UMVUE if $E_\theta[T(x)] = g(\theta)$ and for all other unbiased estimators, $S: X \rightarrow \mathbb{R}$, $\text{Var}_\theta(T) \leq \text{Var}_\theta(S)$ for all $\theta \in \Theta$.

Remark 11. 1. UMVUE are the best possible among unbiased estimators.

2. Compare to Gauss Markov, $\hat{\beta}_{OLS}$ is UMVUE
- 3.

$$E_\theta[||T(x) - \rho(\theta)||^2] = \text{Bias}^2 + \text{Var}(T) = \text{Var}(T) \leq E_\theta[||S(x) - \rho(\theta)||^2]$$

Theorem 7. Let $(P_\theta | \theta \in \Theta)$ be regular. Let $\rho: \Theta \rightarrow \mathbb{R}$ continuous differentiable, then for any unbiased estimator T of S , $E_\theta[T] = S(\theta)$,

$$\text{Var}_\theta \geq \nabla_\theta \rho(\theta)^T I(\theta)^{-1} \nabla_\theta \rho(\theta)$$

Remark 12. If $\Theta = \mathbb{R}$, $\rho(\theta) = \theta$, $\text{Var}_\theta(T) \geq I(\theta)^{-1}$

Proof. Let us assume $\Theta \subseteq \mathbb{R}$,

$$\text{Cov}_\theta(U_\theta, T) = E_\theta[U_\theta T] - E_\theta[U_\theta]E_\theta[T] = E_\theta[U_\theta T]$$

More over, by Cauchy Swartz,

$$\text{Cov}_\theta(U_\theta T) \leq \text{Var}_\theta(U_\theta)^{1/2} \text{Var}_\theta(T)^{1/2} I(\theta)^{1/2} \text{Var}_\theta(T)^{1/2}$$

But then

$$\begin{aligned} E_\theta[U_\theta T] &= \int_X \nabla_\theta \log p_\theta(x) T(x) p_\theta(x) d\mu(x) \\ &= \int_X T(x) \nabla_\theta p_\theta(x) d\mu(x) \\ &= \nabla_\theta \int_X T(x) p_\theta(x) d\mu(x) \\ &= E_\theta[T] = \rho'(\theta) \end{aligned}$$

Thus, $\text{Var}_\theta(T) \geq I(\theta)^{-1} \rho'(\theta)^2$

□

Another regularity condition, $I(\theta)$ is invertible.