

Advanced Theory of Statistics

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1 Probability Theory

1.1 Measure space, measurable function, and integration

Definition 1: A collection of subsets of Ω, \mathcal{F} , is a σ -field (or σ -algebra) if (i) The empty set $\emptyset \in \mathcal{F}$; (ii) If $A \in \mathcal{F}$, then the complement $A^c \in \mathcal{F}$; (iii) If $A_i \in \mathcal{F}, i = 1, 2, \dots$, then their union $\cup A_i \in \mathcal{F}$. (Ω, \mathcal{F}) is a measurable space if \mathcal{F} is a σ -field on Ω .

Example 1: \mathcal{C} = a collection of subsets of interest. $\sigma(\mathcal{C})$ = the smallest σ -field containing \mathcal{C} (the σ -field generated by \mathcal{C}). $\sigma(\mathcal{C}) = \mathcal{C}$ if \mathcal{C} itself is a σ -field. $\sigma(\{A\}) = \{\emptyset, A, A^c, \Omega\}$.

Example 2 (Borel σ -field): \mathbb{R}^k : the k -dimensional Euclidean space ($\mathbb{R}^1 = \mathbb{R}$ is the real line). \mathcal{O} = all open sets, \mathcal{C} = all closed sets. $\mathcal{B}^k = \sigma(\mathcal{O}) = \sigma(\mathcal{C})$: the Borel σ -field on \mathbb{R}^k . $C \in \mathcal{B}^k, \mathcal{B}_C = \{C \cap B : B \in \mathcal{B}^k\}$ is the Borel σ -field on C .

Definition 2: Let (Ω, \mathcal{F}) be a measurable space. A set function ν defined on \mathcal{F} is a measure if (i) $0 \leq \nu(A) \leq \infty$ for any $A \in \mathcal{F}$; (ii) $\nu(\emptyset) = 0$; (iii) If $A_i \in \mathcal{F}, i = 1, 2, \dots$, and A_i 's are disjoint, i.e. $A_i \cap A_j = \emptyset$ for any $i \neq j$, then $\nu(\cup_{i=1}^{\infty} A_i) = \sum_{i=1}^{\infty} \nu(A_i)$. $(\Omega, \mathcal{F}, \nu)$ is a measure if ν is a measure on \mathcal{F} in (Ω, \mathcal{F}) .

Convention 1: For any $x \in \mathbb{R}$, $\infty + x = \infty$, $x\infty = \infty$ if $x > 0$, $x\infty = -\infty$ if $x < 0$. $0\infty = 0$, $\infty + \infty = \infty$, $\infty^a = \infty$ for any $a > 0$. $\infty - \infty$ or ∞/∞ is not defined.

Example 3 (Important examples of measures): (a) Let $x \in \Omega$ be a fixed point and $\delta_x(A) = \begin{cases} c & x \in A \\ 0 & x \notin A \end{cases}$. This is called a point mass at x . (b) Let \mathcal{F} = all subsets of Ω and $\nu(A)$ = the number of elements in $A \in \mathcal{F}$ ($\nu(A) = \infty$ if A contains infinitely many elements). Then ν is a measure on \mathcal{F} and is called the counting measure. (c) There is a unique measure m on $(\mathbb{R}, \mathcal{B})$, that satisfies $m([a, b]) = b - a$ for every finite interval $[a, b]$, $-\infty < a \leq b < \infty$. This is called the Lebesgue measure.

Proposition 1 (Properties of measures): Let $(\Omega, \mathcal{F}, \nu)$ be a measure space. (1) Monotonicity: If $A \subset B$, then $\nu(A) \leq \nu(B)$. (2) Subadditivity: For any sequence A_1, A_2, \dots , $\nu(\cup_{i=1}^{\infty} A_i) \leq \sum_{i=1}^{\infty} \nu(A_i)$. (3) Continuity: If $A_1 \subset A_2 \subset A_3 \subset \dots$ (or $A_1 \supset A_2 \supset A_3 \supset \dots$ and $\nu(A_1) < \infty$), then $\nu(\lim_{n \rightarrow \infty} A_n) = \lim_{n \rightarrow \infty} \nu(A_n)$ where $\lim_{n \rightarrow \infty} A_n = \cup_{i=1}^{\infty} A_i$ (or $= \cap_{i=1}^{\infty} A_i$).

Definition 3: Let P be a probability measure on $(\mathbb{R}, \mathcal{B})$. The cumulative distribution function (c.d.f.) of P is defined to be $F(x) = P((-\infty, x])$, $x \in \mathbb{R}$.

Proposition 2 (Properties of c.d.f.'s): (i) Let F be a c.d.f. on \mathbb{R} . (a) $F(-\infty) = \lim_{x \rightarrow -\infty} F(x) = 0$; (b) $F(\infty) = \lim_{x \rightarrow \infty} F(x) = 1$; (c) F is nondecreasing, i.e. $F(x) \leq F(y)$ if $x \leq y$; (d) F is right continuous, i.e. $\lim_{y \rightarrow x+0} F(y) = F(x)$. (ii) Suppose a real-valued function F on \mathbb{R} satisfies (a)-(d) in part (i). Then F is the c.d.f. of a unique probability measure on $(\mathbb{R}, \mathcal{B})$.

Definition 4 (Product space): $\mathcal{I} = \{1, \dots, k\}$, k is finite or ∞ . $\Gamma_i, i \in \mathcal{I}$, are some sets. $\prod_{i \in \mathcal{I}} \Gamma_i = \Gamma_1 \times \dots \times \Gamma_k = \{(a_1, \dots, a_k) : a_i \in \Gamma_i, i \in \mathcal{I}\}$. Let $(\Omega_i, \mathcal{F}_i), i \in \mathcal{I}$ be measurable spaces. $\sigma(\prod_{i \in \mathcal{I}} \mathcal{F}_i)$ is called the product σ -field on the product space $\prod_{i \in \mathcal{I}} \Omega_i$. $(\prod_{i \in \mathcal{I}} \Omega_i, \sigma(\prod_{i \in \mathcal{I}} \mathcal{F}_i))$ is denoted by $\prod_{i \in \mathcal{I}} (\Omega_i, \mathcal{F}_i)$.

Definition 5 (σ -finite): A measure ν on (Ω, \mathcal{F}) is said to be σ -finite iff there exists a sequence $\{A_1, A_2, \dots\}$ such that $\cup A_i = \Omega$ and $\nu(A_i) < \infty$ for all i . Any finite measure is clearly σ -finite. The Lebesgue measure on \mathcal{F} is σ -finite.

Proposition 3 (Product measure theorem): Let $(\Omega_i, \mathcal{F}_i, \nu_i), i = 1, \dots, k$, be measure spaces with σ -finite measures. There exists a unique σ -finite measure on σ -field $\sigma(\mathcal{F}_1 \times \dots \times \mathcal{F}_k)$, called the product measure and denoted by $\nu_1 \times \dots \times \nu_k$, such that $\nu_1 \times \dots \times \nu_k(A_1 \times \dots \times A_k) = \nu_1(A_1) \dots \nu_k(A_k)$ for all $A_i \in \mathcal{F}_i, i = 1, \dots, k$.

Definition 6 (Measurable function): Let (Ω, \mathcal{F}) and (Λ, \mathcal{G}) be measurable spaces. Let f be a function from Ω to Λ . f is called a measurable function from (Ω, \mathcal{F}) to (Λ, \mathcal{G}) iff $f^{-1}(\mathcal{G}) \subset \mathcal{F}$.

Definition 7 (Integration): (a) The integral of a nonnegative simple function ϕ w.r.t. ν is defined as $\int \phi d\nu = \sum_{i=1}^k a_i \nu(A_i)$. (b) Let f be a nonnegative Borel function and let \mathcal{S}_f be the collection of all nonnegative simple functions satisfying $\phi(\omega) \leq f(\omega)$ for any $\omega \in \Omega$. The integral of f w.r.t. ν is defined as $\int f d\nu = \sup\{\int \phi d\nu : \phi \in \mathcal{S}_f\}$ (Hence, for any Borel function $f \geq 0$, there exists a sequence of simple functions ϕ_1, ϕ_2, \dots such that $0 \leq \phi_i \leq f$ for all i and $\lim_{n \rightarrow \infty} \int \phi_n d\nu = \int f d\nu$). (c) Let f be a Borel function, $f_+(\omega) = \max\{f(\omega), 0\}$ be the positive part of f , and $f_-(\omega) = \max\{-f(\omega), 0\}$ be the negative part of f . We say that $\int f d\nu$ exists if and only if at least one of $\int f_+ d\nu$ and $\int f_- d\nu$ is finite, in which case $\int f d\nu = \int f_+ d\nu - \int f_- d\nu$. (d) When both $\int f_+ d\nu$ and $\int f_- d\nu$ are finite, we say that f is integrable. Let A be a measurable set and I_A be its indicator function. The integral of f over A is defined as $\int_A f d\nu = \int I_A f d\nu$.

Example 4 (Extended set): For convenience, we define the integral of a measurable f from $(\Omega, \mathcal{F}, \nu)$ to $(\bar{\mathbb{R}}, \bar{\mathcal{B}})$, where $\bar{\mathbb{R}} = \mathbb{R} \cup \{-\infty, +\infty\}, \bar{\mathcal{B}} = \sigma(\mathcal{B} \cup \{\infty, -\infty\})$. Let $A_+ = \{f = \infty\}$ and $A_- = \{f = -\infty\}$. If $\nu(A_+) = 0$, we define $\int f_+ d\nu$ to be $\int I_{A_+} f_+ d\nu$; otherwise $\int f_+ d\nu = \infty$. $\int f_- d\nu$ is similarly defined. If at least one of $\int f_+ d\nu$ and $\int f_- d\nu$ is finite, then $\int f d\nu = \int f_+ d\nu - \int f_- d\nu$ is well defined.

1.2 Integration theory and Radon-Nikodym derivative

Proposition 1: $(\Omega, \mathcal{F}, \nu)$ be a measure space and f and g be Borel functions. (i) If $f \leq g$ a.e., then $\int f d\nu \leq \int g d\nu$, provided that the integrals exist. (ii) If $f \geq 0$ a.e. and $\int f d\nu = 0$, then $f = 0$ a.e.

Theorem 1: Let f_1, f_2, \dots be a sequence of Borel functions on $(\Omega, \mathcal{F}, \nu)$. (i) Fatou's lemma: If $f_n \geq 0$, then $\int \liminf_n f_n d\nu \leq \liminf_n \int f_n d\nu$. (ii) Dominated convergence theorem: If $\lim_{n \rightarrow \infty} f_n = f$ a.e. and $|f_n| \leq g$ a.e. for integrable g , then $\int \lim_{n \rightarrow \infty} f_n d\nu = \lim_{n \rightarrow \infty} \int f_n d\nu$. (iii) Monotone convergence theorem: If $0 \leq f_1 \leq f_2 \leq \dots$ and $\lim_{n \rightarrow \infty} f_n = f$ a.e., then $\int \lim_{n \rightarrow \infty} f_n d\nu = \lim_{n \rightarrow \infty} \int f_n d\nu$.

Example 1 (Interchange of differentiation and integration): Let $(\Omega, \mathcal{F}, \nu)$ be a measure space and, for any fixed $\theta \in \mathbb{R}$, let $f(\omega, \theta)$ be a Borel function on Ω . Suppose that $\partial f(\omega, \theta)/\partial \theta$ exists a.e. for $\theta \in (a, b) \subset \mathbb{R}$ and that $|\partial f(\omega, \theta)/\partial \theta| \leq g(\omega)$ a.e., where g is an integrable function on Ω . Then for each $\theta \in (a, b)$, $\partial f(\omega, \theta)/\partial \theta$ is integrable and, by Theorem 1(ii), $\frac{d}{d\theta} \int f(\omega, \theta) d\nu = \int \frac{\partial f(\omega, \theta)}{\partial \theta} d\nu$.

Theorem 2 (Change of variables): Let f be measurable from $(\Omega, \mathcal{F}, \nu)$ to (Λ, \mathcal{G}) and g be Borel on (Λ, \mathcal{G}) . Then $\int_\Omega g \circ f d\nu = \int_\Lambda g d(\nu \circ f^{-1})$, i.e., if either integral exists, then so does the other, and the two are the same.

Theorem 3 (Fubini's theorem): Let ν_i be a σ -finite measure on $(\Omega_i, \mathcal{F}_i), i = 1, 2$, and f be a Borel function on $\prod_{i=1}^2 (\Omega_i, \mathcal{F}_i)$ with $f \geq 0$ or $\int |f| d\nu_1 \times \nu_2 < \infty$. Then $g(\omega_2) = \int_{\Omega_1} f(\omega_1, \omega_2) d\nu_1$ exists a.e. ν_2 and defines a Borel function on Ω_2 whose integral w.r.t. ν_2 exists, and $\int_{\Omega \times \Omega} f(\omega_1, \omega_2) d\nu_1 \times \nu_2 = \int_{\Omega_2} [\int_{\Omega_1} f(\omega_1, \omega_2) d\nu_1] d\nu_2$.

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Definition 1 (Absolutely continuous): Let λ and ν be two measures on a measurable space $(\Omega, \mathcal{F}, \nu)$. We say λ is absolutely continuous w.r.t. ν and write $\lambda \ll \nu$ iff $\nu(A) = 0$ implies $\lambda(A) = 0$.

Theorem 4 (Radon-Nikodym theorem): Let ν and λ be two measure on (Ω, \mathcal{F}) and ν be σ -finite. If $\lambda \ll \nu$, then there exists a nonnegative Borel function f on Ω such that $\lambda(A) = \int_A f d\nu, A \in \mathcal{F}$. Furthermore, f is unique a.e. ν , i.e. if $\lambda(A) = \int_A g d\nu$ for any $A \in \mathcal{F}$, then $f = g$ a.e. ν .

Example 2: A continuous c.d.f. may not have a p.d.f. w.r.t. Lebesgue measure. A necessary and sufficient condition for a c.d.f. F having a p.d.f. w.r.t. Lebesgue measure is that F is absolute continuous in the sense that for any $\epsilon > 0$, there exists a $\delta > 0$ such that for each finite collection of disjoint bounded open intervals (a_i, b_i) , $\sum(b_i - a_i) < \delta$ implies $\sum[F(b_i) - F(a_i)] < \epsilon$.

Proposition 2 (Calculus with Radon-Nikodym derivatives): Let ν be a σ -finite measure on a measure space (Ω, \mathcal{F}) . (i) If λ is a measure, $\lambda \ll \nu$, and $f \geq 0$, then $\int f d\lambda = \int f \frac{d\lambda}{d\nu} d\nu$. (ii) If $\lambda_i, i = 1, 2$, are measures and $\lambda_i \ll \nu$, then $\lambda_1 + \lambda_2 \ll \nu$ and $\frac{d(\lambda_1 + \lambda_2)}{d\nu} = \frac{d\lambda_1}{d\nu} + \frac{d\lambda_2}{d\nu}$ a.e. ν . (iii) If τ is a measure, λ is a σ -finite measure, and $\tau \ll \lambda \ll \nu$, then $\frac{d\tau}{d\nu} = \frac{d\tau}{d\lambda} \frac{d\lambda}{d\nu}$ a.e. ν . In particular, if $\lambda \ll \nu$ and $\nu \ll \lambda$ (in which case λ and ν are equivalent), then $\frac{d\lambda}{d\nu} = \left(\frac{d\nu}{d\lambda}\right)^{-1}$ a.e. ν or λ . (iv) Let $(\Omega_i, \mathcal{F}_i, \nu_i)$ be a measure space and ν_i be σ -finite, $i = 1, 2$. Let λ_i be a σ -finite measure on (Ω, \mathcal{F}_i) and $\lambda_i \ll \nu_i, i = 1, 2$. Then $\lambda_1 \times \lambda_2 \ll \nu_1 \times \nu_2$ and $\frac{d(\lambda_1 \times \lambda_2)}{d(\nu_1 \times \nu_2)}(\omega_1, \omega_2) = \frac{d\lambda_1}{d\nu_1}(\omega_1) \frac{d\lambda_2}{d\nu_2}(\omega_2)$ a.e. $\nu_1 \times \nu_2$.

1.3 Densities, moments, inequalities, and generating functions

Example 1: Let X be a random variable on (Ω, \mathcal{F}, P) whose c.d.f. F_X has a Lebesgue p.d.f. f_x and $F_x(c) < 1$, where c is a fixed constant. Let $Y = \min\{X, c\}$. Note that $Y^{-1}((-\infty, X]) = \Omega$ if $x \geq c$ and $Y^{-1}((-\infty, x]) = X^{-1}((-\infty, x])$ if $x < c$. Hence Y is a random variable and the c.d.f. of

$$Y \text{ is } F_Y(x) = \begin{cases} 1 & x \geq c \\ F_X(x) & x < c \end{cases}. \text{ This c.d.f. is discontinuous at } c, \text{ since } F_x(c) < 1. \text{ Thus, it does}$$

not have a Lebesgue p.d.f. It is not discrete either. Does P_Y , the probability measure corresponding to F_y , have a p.d.f. w.r.t. some measure? Consider the point mass probability measure on $(\mathbb{R}, \mathcal{B})$:

$$\delta_c(A) = \begin{cases} 1 & c \in A \\ 0 & c \notin A \end{cases}, A \in \mathcal{B}. \text{ Then } P_Y \ll m + \delta_c, \text{ and the p.d.f. of } P_Y \text{ is } f_Y(x) = \frac{dP_Y}{d(m + \delta_c)}(x) =$$

$$\begin{cases} 0 & x > c \\ 1 - F_X(c) & x = c \\ f_X(x) & x < c \end{cases}. \text{ To show this, it suffices to show that } \int_{(-\infty, x]} f_Y(t) d(m + \delta_c) = P_Y((-\infty, x])$$

for any $x \in \mathcal{B}$.

Proposition 1 (Transformation): Let X be a random k -vector with a Lebesgue p.d.f. f_X and let $Y = g(X)$, where g is a Borel function from $(\mathbb{R}^k, \mathcal{B}^k)$ to $(\mathbb{R}^l, \mathcal{B}^l)$. Let A_1, \dots, A_m be disjoint sets in \mathcal{B}^k such that $\mathcal{B}^k - (A_1 \cup \dots \cup A_m)$ has Lebesgue measure 0 and g on A_j is one-to-one with a nonvanishing Jacobian, i.e., the determinant $\text{Det}(\partial g(x)/\partial x) \neq 0$ on $A_j, j = 1, \dots, m$. Then Y has the following Lebesgue p.d.f.: $f_Y(x) = \sum_{j=1}^m |\text{Det}(\partial h_j(x)/\partial x)| f_X(h_j(x))$, where h_j is the inverse function of g on $A_j, j = 1, \dots, m$.

Example 2 (F-distribution): Let X_1 and X_2 be independent random variables having the chi-

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square distributions $\chi_{n_1}^2$ and $\chi_{n_2}^2$, respectively. One can show that the p.d.f. of $Y = (X_1/n_1)/(X_2/n_2)$ is the p.d.f. of the F-distribution F_{n_1, n_2} .

Example 3 (t-distribution): Let U_1 be a random variable having the standard normal distribution $N(0, 1)$ and U_2 a random variable having the chi-square distribution χ_n^2 . One can show that if U_1 and U_2 are independent, then the distribution of $T = U_1/\sqrt{U_2/n}$ is the t-distribution t_n .

Example 4 (Noncentral chi-square distribution): Let X_1, \dots, X_n be independent random variables and $X_i \sim N(\mu_i, \sigma^2)$. The distribution of $Y = (X_1^2 + \dots + X_n^2)/\sigma^2$ is called the noncentral chi-square distribution and denoted by $\chi_n^2(\delta)$, where $\delta = (\mu_1^2 + \dots + \mu_n^2)/\sigma^2$ is the noncentrality parameter. If Y_1, \dots, Y_k are independent random variables and Y_i has the noncentral independent chi-square distribution $\chi_{n_i}^2(\delta_i)$, $i = 1, \dots, k$, then $Y = Y_1 + \dots + Y_k$ has the noncentral chi-square distribution $\chi_{n_1 + \dots + n_k}^2(\delta_1 + \dots + \delta_k)$.

Definition 1 (Moments): If $\mathbb{E}X^k$ is finite, where k is a positive integer, $\mathbb{E}X^k$ is called the k -th moment of X or P_X . If $\mathbb{E}|X|^a < \infty$ for some real number a , $\mathbb{E}|X|^a$ is called the a -th absolute moment of X or P_X . If $\mu = \mathbb{E}X$, $\mathbb{E}(X - \mu)^k$ is called the k -th central moment of X or P_X . $\text{Var}(X) = \mathbb{E}(X - \mathbb{E}X)^2$ is called the variance of X or P_X . For random matrix $M = (M_{ij})$, $\mathbb{E}M = (\mathbb{E}M_{ij})$. For random vector X , $\text{Var}(X) = \mathbb{E}(X - \mathbb{E}X)(X - \mathbb{E}X)^T$ is its covariance matrix, whose (i, j) -th element, $i \neq j$, is called the covariance of X_i and X_j and denoted by $\text{Cov}(X_i, X_j)$. If $\text{Cov}(X_i, X_j) = 0$, then X_i and X_j are said to be uncorrelated. Independence implies uncorelation, not converse. If X is random and c is fixed, then $\mathbb{E}(c^T X) = c^T \mathbb{E}(X)$ and $\text{Var}(c^T X) = c^T \text{Var}(X)c$.

Definition 2 (Moment generating and characteristic functions): Let X be a random k -vector. (i) The moment generating function (m.g.f.) of X or P_X is defined as $\psi_X(t) = \mathbb{E}e^{t^T X}$, $t \in \mathbb{R}^k$. (ii) The characteristic function (ch.f.) of X or P_X is defined as $\phi_X(t) = \mathbb{E}e^{it^T X} = \mathbb{E}[\cos(t^T X)] + i\mathbb{E}[\sin(t^T X)]$, $t \in \mathbb{R}^k$.

Proposition 2 (Properties of m.g.f. and ch.f.): If the m.g.f. is finite in a neighborhood of $0 \in \mathbb{R}^k$, then (i) moments of X of any order are finite; (ii) $\phi_X(t)$ can be obtained by replacing t in $\psi_X(t)$ by it . If $Y = A^T X + c$, where A is a $k \times m$ matrix and $c \in \mathbb{R}^m$, then $\psi_Y(u) = e^{c^T u} \psi_X(Au)$ and $\phi_Y(u) = e^{ic^T u} \phi_X(Au)$, $u \in \mathbb{R}^m$. For independent X_1, \dots, X_k , $\psi_{\sum_i X_i}(t) = \prod_i \psi_{X_i}(t)$ and $\phi_{\sum_i X_i}(t) = \prod_i \phi_{X_i}(t)$, $t \in \mathbb{R}^k$. For $X = (X_1, \dots, X_k)$ with m.g.f. ψ_X finite in a neighborhood of 0, $\frac{\partial \psi_X(t)}{\partial t}|_{t=0} = \mathbb{E}X$, $\frac{\partial^2 \psi_X(t)}{\partial t \partial t^T}|_{t=0} = \mathbb{E}(XX^T)$. If $\mathbb{E}|X_1^{r_1} \dots X_k^{r_k}| < \infty$ for nonnegative integers r_1, \dots, r_k , then $\frac{\partial \phi_X(t)}{\partial t}|_{t=0} = i\mathbb{E}X$, $\frac{\partial^2 \phi_X(t)}{\partial t \partial t^T}|_{t=0} = -\mathbb{E}(XX^T)$.

Theorem 1 (Uniqueness): Let X and Y be random k -vectors. (i) If $\phi_X(t) = \phi_Y(t)$ for all $t \in \mathbb{R}^k$, then $P_X = P_Y$; (2) If $\psi_X(t) = \psi_Y(t) < \infty$ for all t in a neighborhood of 0, then $P_X = P_Y$.

1.4 Conditional expectation and independence

Definition 1: Let X be an integrable random variable on (Ω, \mathcal{F}, P) . (i) The conditional expectation of X given \mathcal{A} (a sub- σ -field of \mathcal{F}), denoted by $\mathbb{E}(X|\mathcal{A})$, is the a.s. unique random variable satisfying the following two conditions: (a) $\mathbb{E}(X|\mathcal{A})$ is measurable from (Ω, \mathcal{A}) to $(\mathbb{R}, \mathcal{B})$; (b) $\int_A \mathbb{E}(X|\mathcal{A}) dP = \int_A X dP$ for any $A \in \mathcal{A}$. (ii) The conditional probability of $B \in \mathcal{F}$ given \mathcal{A} is defined to be $P(B|\mathcal{A}) = \mathbb{E}(I_B|\mathcal{A})$. (iii) Let Y be measurable from (Ω, \mathcal{F}, P) to (Λ, \mathcal{G}) . The conditional expectation of X given Y is defined to be $\mathbb{E}(X|Y) = \mathbb{E}[X|\sigma(Y)]$.

Theorem 1: Let Y be measurable from (Ω, \mathcal{F}) to (Λ, \mathcal{G}) and Z a function from (Ω, \mathcal{F}) to \mathbb{R}^k . Then Z is measurable from $(\Omega, \sigma(Y))$ to $(\mathbb{R}^k, \mathcal{B}^k)$ iff there is a measurable function h from (Λ, \mathcal{G}) such that $Z = h \circ Y$.

Example 1: Let X be an integrable random variable on (Ω, \mathcal{F}, P) , A_1, A_2, \dots be disjoint events on (Ω, \mathcal{F}, P) such that $\cup A_i = \Omega$ and $P(A_i) > 0$ for all i , and let a_1, a_2, \dots be distinct real numbers. Define $Y = a_1 I_{A_1} + a_2 I_{A_2} + \dots$. We can show that $\mathbb{E}(X|Y) = \sum_{i=1}^{\infty} \frac{\int_{A_i} X dP}{P(A_i)} I_{A_i}$.

Proposition 1: Let X be a random n -vector and Y a random m -vector. Suppose that (X, Y) has a joint p.d.f. $f(x, y)$ w.r.t. $\nu \times \lambda$, where ν and λ are σ -finite measures on $(\mathbb{R}^n, \mathcal{B}^n)$ and $(\mathbb{R}^m, \mathcal{B}^m)$, respectively. Let $g(x, y)$ be a Borel function on \mathbb{R}^{n+m} for which $\mathbb{E}|g(X, Y)| < \infty$. Then $\mathbb{E}[g(X, Y)|Y] = \frac{\int g(x, Y)f(x, Y)d\nu(x)}{\int f(x, Y)d\nu(x)}$ a.s.

Definition 2 (Conditional p.d.f.): Let (X, Y) be a random vector with a joint p.d.f. $f(x, y)$ w.r.t. $\nu \times \lambda$. The conditional p.d.f. of X given $Y = y$ is defined to be $f_{X|Y}(x|y)/f_Y(y)$ where $f_Y(y) = \int f(x, y)d\nu(x)$ is the marginal p.d.f. of Y w.r.t. λ .

Proposition 2: Let X, Y, X_1, X_2, \dots be integrable random variables on (Ω, \mathcal{F}, P) and \mathcal{A} be a sub- σ -field of \mathcal{F} . (i) If $X = c$ a.s., $c \in \mathbb{R}$, then $\mathbb{E}(X|\mathcal{A}) = c$ a.s. (ii) If $X \leq Y$ a.s., then $\mathbb{E}(X|\mathcal{A}) \leq \mathbb{E}(Y|\mathcal{A})$ a.s. (iii) If $a, b \in \mathbb{R}$, then $\mathbb{E}(aX + bY|\mathcal{A}) = a\mathbb{E}(X|\mathcal{A}) + b\mathbb{E}(Y|\mathcal{A})$ a.s. (iv) $\mathbb{E}[\mathbb{E}(X|\mathcal{A})] = \mathbb{E}X$. (v) $\mathbb{E}[\mathbb{E}(X|\mathcal{A})|\mathcal{A}_0] = \mathbb{E}(X|\mathcal{A}_0) = \mathbb{E}[\mathbb{E}(X|\mathcal{A}_0)|\mathcal{A}]$ a.s., where \mathcal{A}_0 is a sub- σ -field of \mathcal{A} . (vi) If $\sigma(Y) \subset \mathcal{A}$ and $\mathbb{E}|XY| < \infty$, then $\mathbb{E}(XY|\mathcal{A}) = Y\mathbb{E}(X|\mathcal{A})$ a.s. (vii) If X and Y are independent and $\mathbb{E}|g(X, Y)| < \infty$ for a Borel function g , then $\mathbb{E}[g(X, Y)|Y = y] = \mathbb{E}[g(X, y)]$ a.s. P_Y . (viii) If $\mathbb{E}X^2 < \infty$, then $[\mathbb{E}(X|\mathcal{A})]^2 \leq \mathbb{E}(X^2|\mathcal{A})$ a.s. (ix) Fatou's lemma: If $X_n \geq 0$ for any n , then $\mathbb{E}(\liminf_n X_n|\mathcal{A}) \leq \liminf_n \mathbb{E}(X_n|\mathcal{A})$ a.s. (x) Dominated convergence theorem: If $|X_n| \leq Y$ for any n and $X_n \rightarrow_{\text{a.s.}} X$, then $\mathbb{E}(X_n|\mathcal{A}) \rightarrow_{\text{a.s.}} \mathbb{E}(X|\mathcal{A})$.

Definition 3 (Independence): Let (Ω, \mathcal{F}, P) be a probability space. (i) Let \mathcal{C} be a collection of subsets in \mathcal{F} . Events in \mathcal{C} are said to be independent iff for any positive integer n and distinct events $A_1, \dots, A_n \in \mathcal{C}$, $P(A_1 \cap A_2 \cap \dots \cap A_n) = P(A_1)P(A_2) \dots P(A_n)$. (ii) Collections $\mathcal{C}_i \subset \mathcal{F}, i \in \mathcal{I}$ are said to be independent iff events in any collection of the form $\{A_i \in \mathcal{C}_i : i \in \mathcal{J}\}$ are independent. (iii) Random elements $X_i, i \in \mathcal{I}$, are said to be independent iff $\sigma(X_i), i \in \mathcal{I}$ are independent.

Theorem 2: Let $\mathcal{C}_i, i \in \mathcal{I}$ be independent collections of events. If each \mathcal{C}_i is a π -system, then $\sigma(\mathcal{C}_i), i \in \mathcal{I}$ are independent.

Proposition 2: Let X be a random variable with $\mathbb{E}|X| < \infty$ and let Y_i be random k_i vectors, $i = 1, 2$. Suppose that (X, Y_1) and Y_2 are independent. Then $\mathbb{E}[X|(Y_1, Y_2)] = \mathbb{E}(X|Y_1)$ a.s.

Definition 4 (Conditional independence): Let X, Y, Z be random vectors. We say that given Z , X and Y are conditionally independent iff $P(A|X, Z) = P(A|Z)$ a.s. for any $A \in \sigma(Y)$.

1.5 Convergence modes and relationships

Definition 1 (Convergence modes): Let X, X_1, X_2, \dots be a random k -vectors defined on a probability space. (i) We say that the sequence $\{X_n\}$ converges to X almost surely and write $X_n \rightarrow_{\text{a.s.}} X$ iff $\lim_{n \rightarrow \infty} X_n = X$ a.s. (ii) We say that $\{X_n\}$ converges to X in probability and write $X_n \rightarrow_p X$ iff for every fixed $\epsilon > 0$, $\lim_{n \rightarrow \infty} P(\|X_n - X\| > \epsilon) = 0$. (iii) We say that $\{X_n\}$ converges to X in L_r (or in r th moment) with a fixed $r > 0$ and write $X_n \rightarrow_{L_r} X$ iff $\lim_{n \rightarrow \infty} \mathbb{E}\|X_n - X\|_r^r = 0$. (iv)

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Let $F, F_n, n = 1, 2, \dots$ be c.d.f.'s on \mathbb{R}^k and $P, P_n, n = 1, 2, \dots$ be their corresponding probability measures. We say that $\{F_n\}$ converges to F weakly (or $\{P_n\}$ converges to P weakly) and write $F_n \rightarrow_w F$ (or $P_n \rightarrow_w P$) iff, for each continuity point x of F , $\lim_{n \rightarrow \infty} F_n(x) = F(x)$. We say that $\{X_n\}$ converges to X in distribution (or in law) and write $X_n \rightarrow_d X$ iff $F_{X_n} \rightarrow_w F_X$.

Proposition 1: If $F_n \rightarrow_w F$ and F is continuous on \mathbb{R}^k , then $\lim_{n \rightarrow \infty} \sup_{x \in \mathbb{R}^k} |F_n(x) - F(x)| = 0$.

Theorem 1: For random k -vectors X, X_1, X_2, \dots on a probability space, $X_n \rightarrow_{a.s.} X$ iff for every $\epsilon > 0$, $\lim_{n \rightarrow \infty} P(\cup_{m=n}^{\infty} \{\|X_m - X\| > \epsilon\}) = 0$.

Theorem 2 (Borel-Cantelli lemma): Let A_n be a sequence of events in a probability space and $\limsup_n A_n = \cap_{n=1}^{\infty} \cup_{m=n}^{\infty} A_m$. (i) If $\sum_{n=1}^{\infty} P(A_n) < \infty$, then $P(\liminf_n A_n) = 0$. (ii) If A_1, A_2, \dots are pairwise independent and $\sum_{n=1}^{\infty} P(A_n) = \infty$, then $P(\limsup_n A_n) = 1$.

Definition 2: Let X_1, X_2, \dots be random vectors and Y_1, Y_2, \dots be random variables defined on a common probability space. (i) $X_n = O(Y_n)$ a.s. iff $P(\|X_n\| = O(|Y_n|)) = 1$. (ii) $X_n = o(Y_n)$ a.s. iff $X_n/Y_n \rightarrow_{a.s.} 0$. (iii) $X_n = O_p(Y_n)$ iff, for any $\epsilon > 0$, there is a constant $C_\epsilon > 0$ such that $\sup_n P(\|X_n\| \geq C_\epsilon |Y_n|) < \epsilon$. (iv) $X_n = o_p(Y_n)$ iff $X_n/Y_n \rightarrow_p 0$.

Theorem 3: (i) If $X_n \rightarrow_{a.s.} X$, then $X_n \rightarrow_p X$. (The converse is not true). (ii) If $X_n \rightarrow_{L_r} X$ for an $r > 0$, then $X_n \rightarrow_p X$. (The converse is not true). (iii) If $X_n \rightarrow_p X$, then $X_n \rightarrow_d X$. (The converse is not true). (iv) (Skorohod's theorem). If $X_n \rightarrow_d X$, then there are random vectors Y, Y_1, Y_2, \dots defined on a common probability space such that $P_Y = P_X, P_{Y_n} = P_{X_n}, n = 1, 2, \dots$ and $Y_n \rightarrow_{a.s.} Y$. (v) If, for every $\epsilon > 0$, $\sum_{n=1}^{\infty} P(\|X_n - X\| \geq \epsilon) < \infty$, then $X_n \rightarrow_{a.s.} X$. (vi) If $X_n \rightarrow_p X$, then there is a subsequence such that $X_{n_j} \rightarrow_{a.s.} X$ as $j \rightarrow \infty$. (vii) If $X_n \rightarrow_d X$ and $P(X = c) = 1$, where $c \in \mathbb{R}^k$ is a constant vector, then $X_n \rightarrow_p c$. (viii) Suppose that $X_n \rightarrow_d X$. Then for any $r > 0$, $\lim_{n \rightarrow \infty} \mathbb{E}\|X_n\|_r^r = \mathbb{E}\|X\|_r^r < \infty$ if $\{\|X_n\|_r^r\}$ is uniformly integrable in the sense that $\lim_{t \rightarrow \infty} \sup_n \mathbb{E}(\|X_n\|_r^r I_{\{\|X_n\|_r > t\}}) = 0$.

Proposition 2 (Sufficient conditions for uniform integrability): $\sup_n \mathbb{E}\|X_n\|_r^{r+\delta} < \infty$ for a $\delta > 0$.

Proposition 3 (Properties of the quotient random variables): (i) Suppose X, X_1, X_2, \dots are positive random variables. Then $X_n \rightarrow_{a.s.} X$ iff for every $\epsilon > 0$, $\lim_{n \rightarrow \infty} P(\sup_{k \geq n} \frac{X_k}{X} > 1 + \epsilon) = 0$, and $\lim_{n \rightarrow \infty} P(\sup_{k \geq n} \frac{X}{X_k} > 1 + \epsilon) = 0$. (ii) Suppose X, X_1, X_2, \dots are positive random variables. If $\sum_{n=1}^{\infty} P(X_n/X > 1 + \epsilon) < \infty$ and $\sum_{n=1}^{\infty} P(X/X_n > 1 + \epsilon) < \infty$, then $X_n \rightarrow_{a.s.} X$.

1.6 Uniform integrability and weak convergence

Definition 1 (Tightness): A sequence $\{P_n\}$ of probability measure on $(\mathbb{R}^k, \mathcal{B}^k)$ is tight if for every $\epsilon > 0$, there is a compact set $C \subset \mathbb{R}^k$ such that $\inf_n P_n(C) > 1 - \epsilon$. If $\{X_n\}$ is a sequence of random k -vectors, then the tightness of $\{P_{X_n}\}$ is the same as the boundedness of $\{\|X_n\|\}$ in probability.

Proposition 1: Let $\{P_n\}$ be a sequence of probability measures on $(\mathbb{R}^k, \mathcal{B}^k)$. (i) Tightness of $\{P_n\}$ is a necessary and sufficient condition that for every subsequence $\{P_n\}$ there exists a further subsequence $\{P_{n_j}\} \subset \{P_n\}$ and a probability measure P on $(\mathbb{R}^k, \mathcal{B}^k)$ such that $P_{n_j} \rightarrow_w P$ as $j \rightarrow \infty$. (ii) If $\{P_n\}$ is tight and if each subsequence that converges weakly at all converges to the same probability measure P , then $P_n \rightarrow_w P$.

Theorem 1 (Useful sufficient and necessary conditions for convergence in distribution): Let X, X_1, X_2, \dots be random k -vectors. (i) $X_n \rightarrow_d X$ is equivalent to any one of the following conditions:

(a) $\mathbb{E}[h(X_n)] \rightarrow \mathbb{E}[h(X)]$ for every bounded continuous function h ; (b) $\limsup_n P_{X_n}(C) \leq P_X(C)$ for any closed set $C \subset \mathbb{R}^k$; (c) $\liminf_n P_{X_n}(O) \geq P_X(O)$ for any open set $O \subset \mathbb{R}^k$. (ii) Lévy-Cramér continuity theorem. Let $\phi_X, \phi_{X_1}, \phi_{X_2}, \dots$ be the ch.f.'s of X, X_1, X_2, \dots , respectively. $X_n \rightarrow_d X$ iff $\lim_{n \rightarrow \infty} \phi_{X_n}(t) = \phi_X(t)$ for all $t \in \mathbb{R}^k$. (iii) Cramér-Wold device. $X_n \rightarrow_d X$ iff $c^T X_n \rightarrow_d c^T X$ for every $c \in \mathbb{R}^k$.

Example 1: Let X_1, \dots, X_n be independent random variables having a common c.d.f. and $T_n = X_1 + \dots + X_n, n = 1, 2, \dots$. Suppose that $\mathbb{E}|X_1| < \infty$. It follows from a result in calculus that the ch.f. of X_1 satisfies $\phi_{X_1}(t) = \phi_{X_1}(0) + \sqrt{-1}\mu t + o(|t|)$ as $|t| \rightarrow 0$, where $\mu = \mathbb{E}X_1$. Then, the ch.f. of T_n/n is $\phi_{T_n/n}(t) = [\phi_{X_1}(\frac{t}{n})]^n = [1 + \frac{\sqrt{-1}\mu t}{n} + o(\frac{t}{n})]^n \rightarrow e^{\sqrt{-1}\mu t}$ for any $t \in \mathbb{R}$ as $n \rightarrow \infty$. $e^{\sqrt{-1}\mu t}$ is the ch.f. of the point mass probability measure at μ . Thus $T_n/n \rightarrow_d \mu$ and $T_n/n \rightarrow_p \mu$.

Proposition 2 (Scheffé's theorem): Let $\{f_n\}$ be a sequence of p.d.f.'s on \mathbb{R}^k w.r.t. ν . Suppose that $\lim_{n \rightarrow \infty} f_n(x) = f(x)$ a.e. and $f(x)$ is a p.d.f. w.r.t. ν . Then $\lim_{n \rightarrow \infty} \int |f_n(x) - f(x)| d\nu = 0$.

1.7 Convergence of transformations and law of large numbers

Theorem 1 (Continuous mapping theorem): Let X, X_1, X_2, \dots be random k -vectors defined on a probability space and g be a measure function from $(\mathbb{R}^k, \mathcal{B}^k)$ to $(\mathbb{R}^l, \mathcal{B}^l)$. Suppose that g is continuous a.s. P_X . Then (i) $X_n \rightarrow_{a.s.} X$ implies $g(X_n) \rightarrow_{a.s.} g(X)$; (ii) $X_n \rightarrow_p X$ implies $g(X_n) \rightarrow_p g(X)$; (iii) $X_n \rightarrow_d X$ implies $g(X_n) \rightarrow_d g(X)$.

Theorem 2 (Slutsky's theorem): Let $X, X_1, X_2, \dots, Y_1, Y_2, \dots$ be random variables on a probability space. Suppose that $X_n \rightarrow_d X$ and $Y_n \rightarrow_p c$, where c is a constant. Then (i) $X_n + Y_n \rightarrow_d X + c$; (ii) $Y_n X_n \rightarrow_d cX$; (iii) $X_n/Y_n \rightarrow_d X/c$ if $c \neq 0$.

Theorem 3: Let X_1, X_2, \dots and $Y = (Y_1, \dots, Y_k)$ be random k -vectors satisfying $a_n(X_n - c) \rightarrow_d Y$, where $c \in \mathbb{R}^k$ and $\{a_n\}$ is a sequence of positive numbers with $\lim_{n \rightarrow \infty} a_n = \infty$. Let g be a function from $\mathbb{R}^k \rightarrow \mathbb{R}$. (i) If g is differentiable at c , then $a_n[g(X_n) - g(c)] \rightarrow_d [\nabla g(c)^T]Y$, where $\nabla g(x)$ denotes the k -vector of partial derivatives of g at x . (ii) Suppose that g has continuous partial derivatives of order $m > 1$ in a neighborhood of c , with all the partial derivatives of order $j, 1 \leq j \leq m-1$, vanishing at c , but with the m th-order partial derivatives not all vanishing at c . Then $a_n^m[g(X_n) - g(c)] \rightarrow_d \frac{1}{m!} \sum_{i_1=1}^k \dots \sum_{i_m=1}^k \frac{\partial^m g}{\partial x_{i_1} \dots \partial x_{i_m}}|_{x=c} Y_{i_1} \dots Y_{i_m}$.

Theorem 4 (The δ -method): If Y has the $\mathcal{N}_k(0, \Sigma)$ distribution, then $a_n[g(X_n) - g(c)] \rightarrow_d \mathcal{N}(0, [\nabla g(c)]^T \Sigma \nabla g(c))$.

Theorem 5: Let X_1, X_2, \dots be i.i.d. random variables. (i) The WLLN. A necessary and sufficient condition for the existence of a sequence of real numbers $\{a_n\}$ for which $\frac{1}{n} \sum_{i=1}^n X_i - a_n \rightarrow_p 0$ is that $nP(|X_1| > n) \rightarrow 0$, in which case we may take $a_n = \mathbb{E}(X_1 1_{\{|X_1| \leq n\}})$. (ii) The SLLN. A necessary and sufficient condition for the existence of a constant c for which $\frac{1}{n} \sum_{i=1}^n X_i \rightarrow_{a.s.} c$ is that $\mathbb{E}|X_1| < \infty$, in which case $c = \mathbb{E}X_1$ and $\frac{1}{n} \sum_{i=1}^n c_i(X_i - \mathbb{E}X_1) \rightarrow_{a.s.} 0$ for any bounded sequence of real numbers $\{c_i\}$.

Theorem 6: Let X_1, X_2, \dots be independent random variables with finite expectations. (i) The SLLN. If there is a constant $p \in [1, 2]$ such that $\sum_{i=1}^{\infty} \frac{\mathbb{E}|X_i|^p}{i^p} < \infty$, then $\frac{1}{n} \sum_{i=1}^n (X_i - \mathbb{E}X_i) \rightarrow_{a.s.} 0$. (ii) The WLLN. If there is a constant $p \in [1, 2]$ such that $\lim_{n \rightarrow \infty} \frac{1}{n^p} \sum_{i=1}^n \mathbb{E}|X_i|^p = 0$, then $\frac{1}{n} \sum_{i=1}^n (X_i - \mathbb{E}X_i) \rightarrow_p 0$.

1.8 The central limit theorem

Theorem 1 (Lindeberg's CLT): Let $\{X_{nj}, j = 1, \dots, k_n\}$ be independent random variables with $k_n \rightarrow \infty$ as $n \rightarrow \infty$ and $0 < \sigma_n^2 = \text{var}(\sum_{j=1}^{k_n} X_{nj}) < \infty, n = 1, 2, \dots$. If $\frac{1}{\sigma_n^2} \sum_{j=1}^{k_n} \mathbb{E}[(X_{nj} - \mathbb{E}X_{nj})^2 I_{\{|X_{nj} - \mathbb{E}X_{nj}| > \epsilon \sigma_n\}}] \rightarrow 0$ for any $\epsilon > 0$, then $\frac{1}{\sigma_n} \sum_{j=1}^{k_n} (X_{nj} - \mathbb{E}X_{nj}) \rightarrow_d \mathcal{N}(0, 1)$.

Theorem 2 (Multivariate CLT): For i.i.d. random k -vectors X_1, \dots, X_n with a finite $\Sigma = \text{var}(X_1)$, $\frac{1}{\sqrt{n}} \sum_{i=1}^n (X_i - \mathbb{E}X_1) \rightarrow_d \mathcal{N}_k(0, \Sigma)$.

Theorem 3 (Berry-Esséen bound): For i.i.d. $\{X_n\}$ and $W_n = \sqrt{n}(\bar{X} - \mu)/\sigma$, $\sup_t |F_{W_n}(t) - \phi(t)| \leq \frac{33}{4} \frac{\mathbb{E}|X_1 - \mu|^3}{\sigma^3 \sqrt{n}}, n = 1, 2, \dots$. Thus, the convergence speed of F_{W_n} to ϕ is of the order $n^{-1/2}$.

2 Fundamentals of Statistics

2.1 Models, data, statistics, and sampling distributions

Definition 1: A set of probability measures P_θ on (Ω, \mathcal{F}) indexed by a parameter $\theta \in \Theta$ is said to be a parametric family or follow a parametric model iff $\Theta \subset \mathbb{R}^d$ for some fixed positive integer d and each P_θ is a known probability measure when θ is known. The set Θ is called the parameter space and d is called its dimension. $\mathcal{P} = \{P_\theta : \theta \in \Theta\}$ is identifiable iff $\theta_1 \neq \theta_2$ and $\theta_i \in \Theta$ imply $P_{\theta_1} \neq P_{\theta_2}$, which may be achieved through reparameterization.

Definition 2 (Dominated family): A family of populations \mathcal{P} is dominated by ν (a σ -finite measure) if $P \ll \nu$ for all $P \in \mathcal{P}$, in which case \mathcal{P} can be identified by the family of densities $\{\frac{dP}{d\nu} : P \in \mathcal{P}\}$ or $\{\frac{dP_\theta}{d\nu} : \theta \in \Theta\}$.

Definition 3 (Exponential families): A parametric family $\{P_\theta : \theta \in \Theta\}$ dominated by a σ -finite measure ν on (Ω, \mathcal{F}) is called an exponential family iff $\frac{dP_\theta}{d\nu}(\omega) = \exp\{[\eta(\theta)]^T T(\omega) - \xi(\theta)\} h(\omega), \omega \in \Omega$ where $\xi(\theta) = \log\{\int_\Omega \exp\{[\eta(\theta)]^T T(\omega)\} h(\omega) d\nu(\omega)\}$. In an exponential family, consider the parameter $\eta = \eta(\theta)$ and $f_\eta(\omega) = \exp\{\eta^T T(\omega) - \zeta(\eta)\} h(\omega), \omega \in \Omega$. This is called the canonical form for the family, and $\Xi = \{\eta : \zeta(\eta) \text{ is defined}\}$ is called the natural parameter space. An exponential family in canonical form is a natural exponential family. If there is an open set contained in the natural parameter space of an exponential family, then the family is said to be of full rank.

Theorem 1: Let \mathcal{P} be a natural exponential family. (i) Let $T = (Y, U)$ and $\eta = (\theta, \phi)$, Y and θ have the same dimension. Then, Y has the p.d.f. $f_\eta(y) = \exp\{\theta^T y - \zeta(\eta)\}$. In particular, T has a p.d.f. in a natural exponential family. Furthermore, the conditional distribution of Y given $U = u$ has the p.d.f. $f_{\theta, u}(y) = \exp\{\theta^T y - \zeta_u(\theta)\}$ w.r.t. a σ -finite measure depending on ϕ . Furthermore, the conditional distribution of Y given $U = u$ has the p.d.f. $f_{\theta, u}(y) = \exp(\theta^T y - \zeta_u(\theta))$ w.r.t. a σ -finite measure depending on u . (ii) If η_0 is an interior point of the natural parameter space, then the m.g.f. of $P_{\eta_0} \circ T^{-1}$ is finite in a neighborhood of 0 and is given by $\psi_{\eta_0}(t) = \exp\{\zeta(\eta_0 + t) - \zeta(\eta_0)\}$.

Definition 4 (Location-scale families): Let P be a known probability measure on $(\mathbb{R}^k, \mathcal{B}^k)$, $\mathcal{V} \subset \mathbb{R}^k$, and \mathcal{M}_k be a collection of $k \times k$ symmetric positive definite matrices. The family $\{P_{(\mu, \Sigma)} : \mu \in \mathcal{V}, \Sigma \in \mathcal{M}_k\}$ is called a location-scale family (on \mathbb{R}^k), where $P_{(\mu, \Sigma)}(B) = P(\Sigma^{-1/2}(B - \mu)), B \in \mathcal{B}^k$. The parameters μ and $\Sigma^{1/2}$ are called the location and scale parameters, respectively.

Definition 5 (Statistics and their sampling distributions): Our data set is a realization of a sample

(random vector) X from an unknown population P . Statistic $T(X)$: A measurable function T of X ; $T(X)$ is a known value whenever X is known. A nontrivial statistic $T(X)$ is usually simpler than X . Finding the form of the distribution of T is one of the major problems in statistical inference and decision theory.

Example 1: Let X_1, \dots, X_n be i.i.d. random variables having a common distribution P . The sample mean and sample variance $\bar{X} = \frac{1}{n} \sum_{i=1}^n X_i$, $S^2 = \frac{1}{n-1} \sum_{i=1}^n (X_i - \bar{X})^2$ are two commonly used statistics.

Example 2 (Order statistics): Let $X = (X_1, \dots, X_n)$ with i.i.d. random components. Let $X_{(i)}$ be the i th smallest value of X_1, \dots, X_n . The statistics $X_{(1)}, \dots, X_{(n)}$ are called the order statistics.

2.2 Sufficiency and minimal sufficiency

Definition 1 (Sufficiency): Let X be a sample from an unknown population $P \in \mathcal{P}$, where \mathcal{P} is a family of populations. A statistic $T(X)$ is said to be sufficient for $P \in \mathcal{P}$ iff conditional distribution of X given T is known.

Theorem 1 (The factorization theorem): Suppose that X is a sample from $P \in \mathcal{P}$ and \mathcal{P} is a family of probability measures on $(\mathbb{R}^n, \mathcal{B}^n)$ dominated by a σ -finite measure ν . Then $T(X)$ is sufficient for $P \in \mathcal{P}$ iff there are nonnegative Borel functions h and g_p on the range of T such that $\frac{dP}{d\nu}(x) = g_p(T(x))h(x)$.

Theorem 2: If a family \mathcal{P} is dominated by a σ -finite measure, then \mathcal{P} is dominated by a probability measure $Q = \sum_{i=1}^{\infty} c_i P_i$, where c_i 's are nonnegative constants with $\sum_{i=1}^{\infty} c_i = 1$ and $P_i \in \mathcal{P}$.

Convention 1: If a statement holds except for outcomes in an event A satisfying $P(A) = 0$ for all $P \in \mathcal{P}$, then we say that the statement holds a.s. \mathcal{P} .

Definition 2 (Minimal sufficiency): Let T be a sufficient statistic for $P \in \mathcal{P}$. T is called a minimal sufficient statistic iff, for any other statistic S sufficient for $P \in \mathcal{P}$, there is a measurable function ψ such that $T = \psi(S)$ a.s. \mathcal{P} .

Theorem 3 (Existence and uniqueness): Minimal sufficient statistics exist when \mathcal{P} contains distributions on \mathbb{R}^k dominated by a σ -finite measure. If both T and S are minimal sufficient statistics, then by definition there is one-to-one measurable function ψ such that $T = \psi(S)$ a.s. \mathcal{P} .

Theorem 4: Let \mathcal{P} be a family of distributions on \mathbb{R}^k . (i) Suppose that $\mathcal{P}_0 \subset \mathcal{P}$ and a.s. \mathcal{P}_0 implies a.s. \mathcal{P} . If T is sufficient for $P \in \mathcal{P}$ and minimal sufficient for $P \in \mathcal{P}_0$, then T is minimal sufficient for $P \in \mathcal{P}$. (ii) Suppose that \mathcal{P} contains p.d.f.'s f_0, f_1, f_2, \dots w.r.t. a σ -finite ν . Let $f_{\infty}(x) = \sum_{i=0}^{\infty} c_i f_i(x)$, where $c_i > 0$ for all i and $\sum_{i=0}^{\infty} c_i = 1$, and let $T_i(x) = f_i(x)/f_{\infty}(x)$ when $f_{\infty}(x) > 0$, $i = 0, 1, 2, \dots$. Then $T(x) = (T_0, T_1, T_2, \dots)$ is minimal sufficient for $P \in \mathcal{P}$. Furthermore, if $\{x : f_i(x) > 0\} \subset \{x : f_0(x) > 0\}$ for all i , then we may replace $f_{\infty}(x)$ for $f_0(x)$, in which case $T(x) = (T_1, T_2, \dots)$ is minimal sufficient for $P \in \mathcal{P}$. (iii) Suppose that \mathcal{P} contains p.d.f.'s f_p w.r.t. a σ -finite measure and that there exists a sufficient statistic $T(x)$ such that, for any possible values x and y of X , $f_p(x) = f_p(y)\phi(x, y)$ for all P implies $T(x) = T(y)$, where ϕ is a measurable function. Then $T(x)$ is minimal sufficient for $P \in \mathcal{P}$.

2.3 Completeness

Definition 1 (Ancillary statistics): A statistic $V(x)$ is ancillary iff its distribution does not depend on any unknown quantity. A statistic $V(X)$ is first-order ancillary iff $\mathbb{E}[V(X)]$ does not depend on any unknown quantity.

Remark 1: If $V(x)$ is a non-trivial ancillary statistic, then $\sigma(V)$ does not contain any information about the unknown population P . If $T(x)$ is a statistic and $V(T(x))$ is a non-trivial ancillary statistic, it indicates that the reduced data set by T contains a non-trivial part that does not contain any information about θ and, hence, a further simplification of T may still be needed.

Definition 2 (Completeness): A statistic $T(x)$ is complete (or boundedly complete) for $P \in \mathcal{P}$ iff, for any Borel f (or bounded Borel f), $\mathbb{E}[f(T)] = 0$ for all $P \in \mathcal{P}$ implies $f = 0$ a.s. \mathcal{P} .

Remark 2: If T is complete (or boundedly complete) and $S = \psi(T)$ for a measurable ψ , then S is complete (or boundedly complete). A complete and sufficient statistic should be minimal sufficient. But a minimal sufficient statistic may be not complete.

Proposition 1: If P is in an exponential family of full rank with p.d.f.'s given by $f_\eta(x) = \exp\{\eta^T T(x) - \zeta(\eta)\}h(x)$, then $T(x)$ is complete and sufficient for $\eta \in \Xi$.

Example 1: Suppose that X_1, \dots, X_n are i.i.d. random variables having the $\mathcal{N}(\mu, \sigma^2)$ distribution, $\mu \in \mathbb{R}$, $\sigma > 0$. The joint p.d.f. of X_1, \dots, X_n is $(2\pi)^{-n/2} \exp\{\eta_1 T_1 + \eta_2 T_2 - n\zeta(\eta)\}$, where $T_1 = \sum_{i=1}^n X_i$, $T_2 = -\sum_{i=1}^n X_i^2$ and $\eta = (\eta_1, \eta_2) = (\frac{\mu}{\sigma^2}, \frac{1}{2\sigma^2})$. Hence, the family of distributions for $X = (X_1, \dots, X_n)$ is a natural exponential family of full rank ($\Xi = \mathbb{R} \times (0, \infty)$). Thus $T(X) = (T_1, T_2)$ is complete and sufficient for η .

Example 2: $T(x) = (X_{(1)}, \dots, X_{(n)})$ of i.i.d. random variables X_1, \dots, X_n is sufficient for $P \in \mathcal{P}$, where \mathcal{P} is the family of distributions on \mathbb{R} having Lebesgue p.d.f.'s. We can show that $T(x)$ is also complete for $P \in \mathcal{P}$.

Theorem 1 (Basu's theorem): Let V and T be two statistics of X from a population $P \in \mathcal{P}$. If V is ancillary and T is boundedly complete and sufficient for $P \in \mathcal{P}$, then V and T are independent w.r.t. any $P \in \mathcal{P}$.

Example 3: X_1, \dots, X_n is a random sample from uniform($\theta, \theta + 1$), $\theta \in \mathbb{R}$, and $T = (X_{(1)}, X_{(n)})$ is the minimal sufficient statistic for θ . We can show that T is not complete.

Theorem 2: Suppose that S is a minimal sufficient statistic and T is a complete and sufficient statistic. Then T must be minimal sufficient and S must be complete.

2.4 Statistical decision

Convention 1 (Basic elements): X : a sample from a population $P \in \mathcal{P}$. Decision: an action we take after observing X . \mathcal{A} : the set of allowable actions. $(\mathcal{A}, \mathcal{F}_{\mathcal{A}})$: the action space. \mathcal{X} : the range of X . Decision rule: a measurable function T from $(\mathcal{X}, \mathcal{F}_{\mathcal{X}})$ to $(\mathcal{A}, \mathcal{F}_{\mathcal{A}})$. If $X = x$ is observed, then we take the action $T(x) \in \mathcal{A}$.

Definition 1 (Loss function): $L(P, a)$: a function from $\mathcal{P} \times \mathcal{A}$ to $[0, \infty)$. $L(P, a)$ is Borel for each P . If $X = x$ is observed and our decision rule is T , then our loss is $L(P, T(x))$.

Definition 2 (Risk): The averaged loss $R_T(P) := \mathbb{E}[L(P, T(X))] = \int_{\mathcal{X}} L(P, T(X)) dP_X(x)$.

Definition 3 (Comparisons): For decision rules T_1 and T_2 , T_1 is as good as T_2 iff $R_{T_1}(P) \leq R_{T_2}(P)$ for any $P \in \mathcal{P}$ and is better than T_2 if, in addition, $R_{T_1}(P) < R_{T_2}(P)$ for some P . T_1 and T_2 are equivalent iff $R_{T_1}(P) = R_{T_2}(P)$ for all $P \in \mathcal{P}$. Optimal rule: If T^* is as good as any other rule in \mathcal{E} , a class of allowable decision rules, then T^* is \mathcal{E} -optimal.

Definition 4 (Randomized decision rules): A function δ on $\mathcal{X} \times \mathcal{F}_{\mathcal{A}}$; for every $A \in \mathcal{F}_{\mathcal{A}}$, $\delta(\cdot, A)$ is a Borel function and, for every $x \in \mathcal{X}$, $\delta(x, \cdot)$ is a probability measure on $(\mathcal{A}, \mathcal{F}_{\mathcal{A}})$. If $X = x$ is observed, we have a distribution of actions: $\delta(x, \cdot)$. A nonrandomized rule T is a special randomized decision rule with $\delta(x, \{a\}) = I_{\{a\}}(T(x))$, $a \in \mathcal{A}$, $x \in \mathcal{X}$. The loss function for a randomized rule δ is defined as $L(P, \delta, x) = \int_{\mathcal{A}} L(P, a) d\delta(x, a)$, which reduces to the same loss function when δ is nonrandomized. The risk of a randomized δ is then $R_{\delta}(P) = \mathbb{E}[L(P, \delta, X)] = \int_{\mathcal{X}} \int_{\mathcal{A}} L(P, a) d\delta(x, a) dP_X(x)$.

Example 1: $X = (X_1, \dots, X_n)$ is a vector of i.i.d. measurements for a parameter $\theta \in \mathbb{R}$. We want to estimate θ . Action space: $(\mathcal{A}, \mathcal{F}_{\mathcal{A}}) = (\mathbb{R}, \mathcal{B})$. A common loss function in this problem is the squared error loss $L(P, a) = (\theta - a)^2$, $a \in \mathcal{A}$. Let $T(X) = \bar{X}$, the sample mean. The loss for \bar{X} is $(\bar{X} - \theta)^2$. If the population has mean μ and variance $\sigma^2 < \infty$, then $R_{\bar{X}}(P) = (\mu - \theta)^2 + \frac{\sigma^2}{n}$. This problem is a special case of a general problem called estimation. In an estimation problem, a decision rule T is called an estimator.

Example 2: Let \mathcal{P} be a family of distributions, $\mathcal{P}_0 \subset \mathcal{P}$, $\mathcal{P}_1 = \{P \in \mathcal{P} : P \notin \mathcal{P}_0\}$. A hypothesis testing problem can be formulated as that of deciding which of the following two statements is true: $H_0 : P \in \mathcal{P}_0$ versus $H_1 : P \in \mathcal{P}_1$. H_0 is called the null hypothesis and H_1 is the alternative hypothesis. The action space for this problem contains only two elements, i.e., $\mathcal{A} = \{0, 1\}$, where 0 is accepting H_0 and 1 is rejecting H_0 . This problem is a special case of a general problem called hypothesis testing. A decision rule is called a test, which must have the form $I_C(X)$, where $C \in \mathcal{F}_{\mathcal{X}}$ is called the rejection or critical region.

Definition 5 (0-1 loss): $L(P, a) = 0$ if a correct decision is made and 1 if an incorrect decision is made, which leads to the risk $R_T(P) = \begin{cases} P(T(X) = 1) = P(X \in C) & P \in \mathcal{P}_0 \\ P(T(X) = 0) = P(X \notin C) & P \in \mathcal{P}_1 \end{cases}$.

Definition 6 (Admissibility): Let \mathcal{E} be a class of decision rules. A decision rule $T \in \mathcal{E}$ is called \mathcal{E} -admissible iff there does not exist any $S \in \mathcal{E}$ that is better than T (in terms of the risk).

Remark 1: An admissible decision rule is not necessarily good. For example, in an estimation problem a silly estimator $T(X) \equiv a$ constant may be admissible.

Proposition 1: Let $T(X)$ be a sufficient statistic for $P \in \mathcal{P}$ and let δ_0 be a decision rule. Then $\delta_1(t, A) = \mathbb{E}[\delta_0(X, A) | T = t]$, which is a randomized decision rule depending only on T , is equivalent to δ_0 if $R_{\delta_0}(P) < \infty$ for any $P \in \mathcal{P}$.

Theorem 1: Suppose that \mathcal{A} is a convex subset of \mathbb{R}^k and that for any $P \in \mathcal{P}$, $L(P, a)$ is a convex function of a . (i) Let δ be a randomized rule satisfying $\int_{\mathcal{A}} \|a\| d\delta(x, a) < \infty$ for any $x \in \mathcal{X}$ and let $T_1(x) = \int_{\mathcal{A}} a d\delta(x, a)$. Then $L(P, T_1(x)) \leq L(P, \delta, x)$ (or $L(P, T_1(x)) < L(P, \delta, x)$) if L is strictly convex in a for any $x \in \mathcal{X}$ and $P \in \mathcal{P}$. (ii) Rao-Blackwell theorem. Let T be a sufficient statistic for $P \in \mathcal{P}$, $T_0 \in \mathbb{R}^k$ be a nonrandomized rule satisfying $\mathbb{E}\|T_0\| < \infty$, and $T_1 = \mathbb{E}[T_0(X) | T]$. Then $R_{T_1}(P) \leq R_{T_0}(P)$ for any $P \in \mathcal{P}$. If L is strictly convex in a and T_0 is not a function of T ,

then T_0 is inadmissible.

Definition 7 (Unbiasedness): In an estimation problem, the bias of an estimator $T(X)$ of a parameter θ of the unknown population is defined to be $b_T(P) = \mathbb{E}[T(X)] - \theta$. An estimator $T(X)$ is unbiased for θ iff $b_T(P) = 0$ for any $P \in \mathcal{P}$.

Approach 1: Define a class \mathcal{E} of decision rules that have some desirable properties and then try to find the best rule in \mathcal{E} .

Approach 2: Consider some characteristic R_T of $R_T(P)$, for a given decision rule T , and then minimize R_T over $T \in \mathcal{E}$. Methods include the Bayes rule and the minimax rule.

2.5 Statistical inference

Definition 1 (Three components in statistical inference): Point estimators, hypothesis tests, confidence sets.

Definition 2 (Point estimators): Let $T(X)$ be an estimator of $\theta \in \mathbb{R}$. Bias: $b_T(P) = \mathbb{E}[T(X)] - \theta$. Mean squared error (mse): $\text{mse}_T(P) = \mathbb{E}[T(X) - \theta]^2 = [b_T(P)]^2 + \text{Var}(T(X))$. Bias and mse are two common criteria for the performance of point estimators, i.e., instead of considering risk functions, we use bias and mse to evaluate point estimators.

Definition 3 (Hypothesis tests): To test the hypotheses $H_0 : P \in \mathcal{P}_0$ versus $H_1 : P \in \mathcal{P}_1$, there are two types of errors we may commit: rejecting H_0 when H_0 is true (called the type I error) and accepting H_0 when H_0 is wrong (called the type II error). A test T : a statistic from \mathcal{X} to $\{0, 1\}$.

Theorem 1 (Probabilities of making two types of errors): Type I error rate: $\alpha_T(P) = P(T(X) = 1), P \in \mathcal{P}_0$. Type II error rate: $1 - \alpha_T(P) = P(T(X) = 0), P \in \mathcal{P}_1$. $\alpha_T(P)$ is also called the power function of T . Power function is $\alpha_T(\theta)$ if P is in a parametric family indexed by θ .

Example 1: Let X_1, \dots, X_n be i.i.d. from the $\mathcal{N}(\mu, \sigma^2)$ distribution with an unknown $\mu \in \mathbb{R}$ and a known σ^2 . Consider the hypotheses $H_0 : \mu \leq \mu_0$ versus $H_1 : \mu > \mu_0$, where μ_0 is a fixed constant. Since the sample mean \bar{X} is sufficient for $\mu \in \mathbb{R}$, it is reasonable to consider the following class of tests: $T_c(X) = I_{(c, \infty)}(\bar{X})$. By the property of the normal distributions, $\alpha_{T_c}(\mu) = P(T_c(X) = 1) = 1 - \phi(\frac{\sqrt{n}(c-\mu)}{\sigma})$. Since $\phi(t)$ is an increasing function of t , $\sup_{P \in \mathcal{P}_0} \alpha_{T_c}(\mu) = 1 - \phi(\frac{\sqrt{n}(c-\mu_0)}{\sigma})$. In fact, it is also true for $\sup_{P \in \mathcal{P}_1} [1 - \alpha_{T_c}(\mu)] = \phi(\frac{\sqrt{n}(c-\mu_0)}{\sigma})$. If we would like to use an α as the level of significance, then the most effective way is to choose a c_α such that $\alpha = \sup_{P \in \mathcal{P}_0} \alpha_{T_{c_\alpha}}(\mu)$, in which case c_α must satisfy $1 - \phi(\frac{\sqrt{n}(c_\alpha-\mu_0)}{\sigma}) = \alpha$, i.e., $c_\alpha = \sigma z_{1-\alpha} / \sqrt{n} + \mu_0$, where $z_a = \Phi^{-1}(a)$. It can be shown that for any test $T(X)$ satisfying $\sup_{P \in \mathcal{P}_0} \alpha_T(P) \leq \alpha$, $1 - \alpha_T(\mu) \geq 1 - \alpha_{T_{c_\alpha}}(\mu), \mu > \mu_0$.

Definition 4 (Significance tests): A common approach of finding an “optimal” test is to assign a small bound α to the type I error rate $\alpha_T(P), P \in \mathcal{P}_0$, and then to attempt to minimize the type II error rate $1 - \alpha_T(P), P \in \mathcal{P}_1$, subject to $\sup_{P \in \mathcal{P}_0} \alpha_T(P) \leq \alpha$. The bound α is called the level of significance. The left-hand side is called the size of the test T . The level of significance should be positive, otherwise no test satisfies.

Definition 5 (p-value): It is good practice to determine not only whether H_0 is rejected for a given α and a chosen test T_α , but also the smallest possible level of significance at which H_0 would be rejected for the computed $T_\alpha(x)$, i.e., $\hat{\alpha} = \inf\{\alpha \in (0, 1) : T_\alpha(x) = 1\}$. Such an $\hat{\alpha}$, which depends on x and the chosen test and is a statistic, is called the p -value for the test T_α .

Example 2: Let us calculate the p -value for T_{c_α} in Example 1. Note that $\alpha = 1 - \phi(\frac{\sqrt{n}(c_\alpha - \mu_0)}{\sigma}) > 1 - \Phi(\frac{\sqrt{n}(\bar{X} - \mu_0)}{\sigma})$ if and only if $\bar{X} > c_\alpha$ (or $T_{c_\alpha}(x) = 1$). Hence, $1 - \phi(\frac{\sqrt{n}(\bar{X} - \mu_0)}{\sigma}) = \inf\{\alpha \in (0, 1) : T_{c_\alpha}(x) = 1\} = \hat{\alpha}(X)$ is the p -value for T_{c_α} . It turns out that $T_{c_\alpha}(x) = I_{(0, \alpha)}(\hat{\alpha}(X))$.

Definition 6 (Confidence sets) θ : a k -vector of unknown parameters related to the unknown $P \in \mathcal{P}$. If a Borel set $C(X)$ (in the range of θ) depending only on the sample X such that $\inf_{P \in \mathcal{P}} P(\theta \in C(X)) \geq 1 - \alpha$, where α is a fixed constant in $(0, 1)$, then $C(X)$ is called a confidence set for θ with level of significance $1 - \alpha$. The left-hand side is called the confidence coefficient of $C(X)$, which is the highest possible level of significance for $C(X)$. A confidence set is a random element that covers the unknown θ with certain probability.

Example 3: Let X_1, \dots, X_n be i.i.d. from the $\mathcal{N}(\mu, \sigma^2)$ distribution with both $\mu \in \mathbb{R}$ and $\sigma^2 > 0$ unknown. Let $\theta = (\mu, \sigma^2)$ and $\alpha \in (0, 1)$ be given. Let \bar{X} be the sample mean and S^2 be the sample variance. Since (\bar{X}, S^2) is sufficient, we focus on $C(X)$ that is a function of (\bar{X}, S^2) . Since $\sqrt{n}(\bar{X} - \mu)/\sigma$ has the $\mathcal{N}(0, 1)$ distribution, $P(-\tilde{c}_\alpha \leq \frac{\bar{X} - \mu}{\sigma/\sqrt{n}} \leq \tilde{c}_\alpha) = \sqrt{1 - \alpha}$, where $\tilde{c}_\alpha = \Phi^{-1}(\frac{1 + \sqrt{1 - \alpha}}{2})$. Since the χ^2 distribution χ_{n-1}^2 is a known distribution, we can always find two constants $c_{1\alpha}$ and $c_{2\alpha}$ such that $P(c_{1\alpha} \leq \frac{(n-1)S^2}{\sigma^2} \leq c_{2\alpha}) = \sqrt{1 - \alpha}$. Then $P(-\tilde{c}_\alpha \leq \frac{\bar{X} - \mu}{\sigma/\sqrt{n}} \leq \tilde{c}_\alpha, c_{1\alpha} \leq \frac{(n-1)S^2}{\sigma^2} \leq c_{2\alpha}) = 1 - \alpha$. The LHS defines a set in the range of $\theta = (\mu, \sigma^2)$ bounded by two straight lines, $\sigma^2 = (n-1)S^2/c_{i\alpha}, i = 1, 2$, and a curve $\sigma^2 = n(\bar{X} - \mu)^2/\tilde{c}_\alpha^2$. This set is a confidence set for θ with confidence coefficient $1 - \alpha$.

Definition 7 (Randomized tests): Since the action space contains only two points, 0 and 1, for a hypothesis testing problem, any randomized test $\delta(X, A)$ is equivalent to a statistic $T(X) \in [0, 1]$ with $T(x) = \delta(x, \{1\})$ and $1 - T(X) = \delta(x, \{0\})$. A nonrandomized test is obviously a special case where $T(x)$ does not take any value in $(0, 1)$. For any randomized test $T(X)$, we define the type I error probability to be $\alpha_T(P) = \mathbb{E}[T(X)], P \in \mathcal{P}_0$, and the type II error probability to be $1 - \alpha_T(P) = \mathbb{E}[1 - T(X)], P \in \mathcal{P}_1$. For a class of randomized tests, we would like to minimize $1 - \alpha_T(P)$ subject to $\sup_{P \in \mathcal{P}_0} \alpha_T(P) = \alpha$.

Definition 8 (Consistency of point estimators): Let $X = (X_1, \dots, X_n)$ be a sample from $P \in \mathcal{P}$, $T_n(X)$ be an estimator of θ for every n , and $\{a_n\}$ be a sequence of positive constants, $a_n \rightarrow \infty$. (i) $T_n(x)$ is consistent for θ iff $T_n(x) \rightarrow_p \theta$ w.r.t. any P . (ii) $T_n(x)$ is a_n -consistent for θ iff $a_n[T_n(X) - \theta] = O_p(1)$ w.r.t. any P . (iii) $T_n(x)$ is strongly consistent for θ iff $T_n(x) \rightarrow_{a.s.} \theta$ w.r.t. any P . (iv) $T_n(X)$ is L_r -consistent for θ iff $T_n(x) \rightarrow_{L_r} \theta$ w.r.t. for any P for some fixed $r > 0$; if $r = 2$, L_2 -consistency is called consistency in mse.

Remark 1 (Consistency is an essential requirement): Like the admissibility, consistency is an essential requirement: any inconsistent estimators should not be used, but there are many consistent estimators and some may not be good. Thus, consistency should be used together with other criteria.

Remark 2 (Approximate and asymptotic bias): Unbiasedness is a criterion for point estimator. In some cases, however, there is no unbiased estimator. Furthermore, having a “slight” bias in some cases may not be a bad idea.

Definition 9: (i) Let ξ, ξ_1, ξ_2, \dots be random variables and $\{a_n\}$ be a sequence of positive numbers satisfying $a_n \rightarrow \infty$ or $a_n \rightarrow a > 0$. If $a_n \xi_n \rightarrow_d \xi$ and $\mathbb{E}|\xi| < \infty$, then $\mathbb{E}\xi/a_n$ is called an asymptotic expectation of ξ_n . (ii) For a point estimator T_n of θ , an asymptotic expectation of $T_n - \theta$, if it exists,

is called an asymptotic bias of T_n and denoted by $\tilde{b}_{T_n}(P)$. If $\lim_{n \rightarrow \infty} \tilde{b}_{T_n}(P) = 0$ for any P , then T_n is asymptotically unbiased.

Proposition 1 (Asymptotic expectation is essentially unique): For a sequence of random variables $\{\xi_n\}$, suppose both $\mathbb{E}\xi/a_n$ and $\mathbb{E}\eta/b_n$ are asymptotic expectations of ξ_n . Then, one of the following three must hold: (a) $\mathbb{E}\xi = \mathbb{E}\eta = 0$; (b) $\mathbb{E}\xi \neq 0, \mathbb{E}\eta = 0$, and $b_n/a_n \rightarrow 0$; (c) $\mathbb{E}\xi \neq 0, \mathbb{E}\eta \neq 0$, and $(\mathbb{E}\xi/a_n)/(\mathbb{E}\eta/b_n) \rightarrow 1$.

Example 4 (Functions of sample means): We consider the case where X_1, \dots, X_n are i.i.d. random k -vectors with finite $\Sigma = \text{Var}(X_1)$, $T_n = g(\bar{X})$, where g is a function on \mathbb{R}^k that is second-order differentiable at $\mu = \mathbb{E}X_1$. Consider T_n as an estimator of $\theta = g(\mu)$. By Taylor's expansion, $T_n - \theta = [\nabla g(\mu)]^T(\bar{X} - \mu) + 2^{-1}(\bar{X} - \mu)^T \nabla^2 g(\mu)(\bar{X} - \mu) + o_p(n^{-1})$. By the CLT, $2^{-1}n(\bar{X} - \mu) \nabla^2 g(\mu)(\bar{X} - \mu) \rightarrow_d 2^{-1}Z_\Sigma^T \nabla^2 g(\mu) Z_\Sigma$, where $Z_\Sigma = \mathcal{N}_k(0, \Sigma)$. Thus, $\frac{\mathbb{E}[Z_\Sigma^T \nabla^2 g(\mu) Z_\Sigma]}{2n} = \frac{\text{tr}(\nabla^2 g(\mu) \Sigma)}{2n}$ is the n^{-1} order asymptotic bias of $T_n = g(\bar{X})$.

Definition 10 (Asymptotic variance and amse): Let T_n be an estimator of θ for every n and $\{a_n\}$ be a sequence of positive numbers satisfying $a_n \rightarrow \infty$ or $a_n \rightarrow a > 0$. Assume that $a_n(T_n - \theta) \rightarrow_d Y$ with $0 < \mathbb{E}Y^2 < \infty$. (i) The asymptotic mean squared error of T_n , denoted by $\text{amse}_{T_n}(P)$, is defined as the asymptotic expectation of $(T_n - \theta)^2$, $\text{amse}_{T_n}(P) = \mathbb{E}Y^2/a_n^2$. The asymptotic variance of T_n is defined as $\sigma_{T_n}^2(P) = \text{Var}(Y)/a_n^2$. (ii) Let T'_n be another estimator of θ . The asymptotic relative efficiency of T'_n w.r.t. T_n is defined as $e_{T'_n, T_n} = \text{amse}_{T_n}(P)/\text{amse}_{T'_n}(P)$. (iii) T_n is said to be asymptotically more efficient than T'_n iff $\limsup_n e_{T'_n, T_n}(P) \leq 1$ for any P and < 1 for some P .

Proposition 2: Let T_n be an estimator of θ for every n and $\{a_n\}$ be a sequence of positive numbers satisfying $a_n \rightarrow \infty$ or $a_n \rightarrow a > 0$. If $a_n(T_n - \theta) \rightarrow_d Y$ with $0 < \mathbb{E}Y^2 < \infty$, then (i) $\mathbb{E}Y^2 \leq \liminf_n \mathbb{E}[a_n^2(T_n - \theta)^2]$ and (ii) $\mathbb{E}Y^2 = \lim_{n \rightarrow \infty} \mathbb{E}[a_n^2(T_n - \theta)^2]$ if and only if $\{a_n^2(T_n - \theta)^2\}$ is uniformly integrable.

Example 5: Let X_1, \dots, X_n be i.i.d. from the Poisson distribution $P(\theta)$ with an unknown $\theta > 0$. Consider the estimation of $\theta = P(X_i = 0) = e^{-\theta}$. Let $T_{1n} = F_n(0)$, where F_n is the empirical c.d.f. Then T_{1n} is unbiased and has $\text{mse}_{T_{1n}}(\theta) = e^{-\theta}(1 - e^{-\theta})/n$. Also, $\sqrt{n}(T_{1n} - \theta) \rightarrow_d \mathcal{N}(0, e^{-\theta}(1 - e^{-\theta}))$ by the CLT. Thus, in the case $\text{amse}_{T_{1n}}(\theta) = \text{mse}_{T_{1n}}(\theta)$. Consider $T_{2n} = e^{-\bar{X}}$. Note that $\mathbb{E}T_{2n} = e^{n\theta(e^{-1/n} - 1)}$, hence $nb_{T_{2n}}(\theta) \rightarrow \theta e^{-\theta}/2$. Using the CLT, we can show that $\sqrt{n}(T_{2n} - \theta) \rightarrow_d \mathcal{N}(0, e^{-2\theta}\theta)$. Then $\text{amse}_{T_{2n}}(\theta) = e^{-2\theta}\theta/n$. Thus, the asymptotic relative efficiency of T_{1n} w.r.t. T_{2n} is $e_{T_{1n}, T_{2n}} = \theta/(e^\theta - 1) < 1$. This shows that T_{2n} is asymptotically more efficient than T_{1n} .

3 Unbiased Estimation

3.1 UMVUE: functions of sufficient and complete statistics

Definition 1 (Estimable): If there exists an unbiased estimator of ϑ , then ϑ is called an estimable parameter.

Definition 2 (UMVUE): An unbiased estimator $T(X)$ of θ is called uniformly minimum variance unbiased estimator (UMVUE) iff $\text{Var}(T(X)) \leq \text{Var}(U(X))$ for any $P \in \mathcal{P}$ and any other unbiased estimator $U(X)$ of θ .

Theorem 1 (Lehmann-Scheffé theorem): Suppose that there exists a sufficient and complete

statistic $T(X)$ for $P \in \mathcal{P}$. If θ is estimable, i.e., there is a unique unbiased estimator of θ , then there is a unique UMVUE of θ that is of the form $h(T)$ with a Borel function h .

The first method (Directly solving for h): Need the distribution of T . Try some function h to see if $\mathbb{E}[h(T)]$ is related to θ . If $\mathbb{E}[h(T)] = \theta$ for all P , what should h be?

Example 1: Let X_1, \dots, X_n be i.i.d. from the uniform distribution on $(0, \theta)$, $\theta > 0$. Consider $\vartheta = \theta$. Since the sufficient and complete statistic $X_{(n)}$ has the Lebesgue p.d.f. $n\theta^{-n}x^{n-1}1_{(0,\theta)}(x)$, $\mathbb{E}X_{(n)} = n\theta^{-n} \int_0^\theta x^n dx = \frac{n}{n+1}\theta$. An unbiased estimator of θ is $(n+1)X_{(n)}/n$, which is the UMVUE. Consider now $\vartheta = g(\theta)$, where g is a differentiable function on $(0, \theta)$. An unbiased estimator $h(X_{(n)})$ of ϑ must satisfy $\theta^n g(\theta) = n \int_0^\theta h(x)x^{n-1}dx$ for all $\theta > 0$. Hence, the UMVUE of ϑ is $h(X_{(n)}) = g(X_{(n)}) + n^{-1}X_{(n)}g'(X_{(n)})$.

The second method (When a sufficient and complete statistic is available): Find an unbiased estimator of θ , say $U(X)$. Conditioning on a sufficient and complete statistic $T(X)$: $\mathbb{E}[U(X)|T]$ is the UMVUE of θ . We need to derive an explicit form of $\mathbb{E}[U(X)|T]$.

Example 2: Let X_1, \dots, X_n be i.i.d. from the exponential distribution $\text{Exp}(0, \theta)$. $F_\theta(x) = (1 - e^{-x/\theta})1_{(0,\theta)}(x)$. Consider the estimation of $\vartheta = 1 - F_\theta(t)$. \bar{X} is sufficient and complete for $\theta > 0$. $1_{(t,\infty)}(X_1)$ is unbiased for ϑ , $\mathbb{E}[1_{(t,\infty)}(X_1)] = P(X_1 > t) = \vartheta$. Hence $T(X) = \mathbb{E}[1_{(t,\infty)}(X_1)|\bar{X}] = P(X_1 > t|\bar{X})$ is the UMVUE of ϑ . By Basu's theorem, X_1/\bar{X} and \bar{X} are independent. Thus, $P(X_1 > t|\bar{X} = \bar{x}) = P(X_1/\bar{X} > t/\bar{x}|\bar{X} = \bar{x}) = P(X_1/\bar{X} > t/\bar{x})$. To compute this unconditional probability, we need the distribution of $X_1/\sum_{i=1}^n X_i = X_1/(X_1 + \sum_{i=2}^n X_i)$. Using the transformation technique and the fact that $\sum_{i=2}^n X_i$ is independent of X_1 and has a gamma distribution, we obtain that $X_1/\sum_{i=1}^n X_i$ has the Lebesgue p.d.f. $(n-1)(1-x)^{n-2}1_{(0,1)}(x)$. Hence $P(X_1 > t|\bar{X} = \bar{x}) = (n-1) \int_{t/(n\bar{x})}^1 (1-x)^{n-2}dx = (1 - \frac{t}{n\bar{x}})^{n-1}$ and the UMVUE of ϑ is $T(X) = (1 - \frac{t}{n\bar{X}})^{n-1}$.

Example 3: Let X_1, \dots, X_n be i.i.d. from an unknown population P in a nonparametric family \mathcal{P} . In many cases the vector of order statistics, $T = (X_{(1)}, \dots, X_{(n)})$, is sufficient and complete for $P \in \mathcal{P}$. Note that an estimator $\phi(X_1, \dots, X_n)$ is a function of T iff the function ϕ is symmetric in its n arguments. Hence, if T is sufficient and complete, then a symmetric unbiased estimator of any estimable ϑ is the UMVUE. Specific examples: \bar{X} is the UMVUE of $\vartheta = \mathbb{E}X_1$, S^2 is the UMVUE of $\text{Var}(X_1)$, $n^{-1} \sum_{i=1}^n X_i^2 - S^2$ is the UMVUE of $(\mathbb{E}X_1)^2$, $F_n(t)$ is the UMVUE of $P(X_1 \leq t)$ for any fixed t . The previous conclusions are not true if T is not sufficient and complete for $P \in \mathcal{P}$.

Remark 1 (Nonexistence of any UMVUE): If $n > 2$ and \mathcal{P} contains all symmetric distributions having Lebesgue p.d.f.'s and finite means, then there is no UMVUE for $\mu = \mathbb{E}X_1$.

Example 4 (Survey samples from a finite population): Let $\mathcal{P} = \{1, \dots, N\}$ be a finite population of interest. For each $i \in \mathcal{P}$, let y_i be a value of interest associated with unit i . Let $s = \{i_1, \dots, i_n\}$ be a subset of distinct elements of \mathcal{P} , which is a sample selected with selection probability $p(s)$, where p is known. The value y_i is observed if and only if $i \in s$. If $p(s)$ is constant, the sampling plan is called the simple random sampling without replacement. Consider the estimation of $Y = \sum_{i=1}^N y_i$, the population total as the parameter of interest. Let $X = (X_i, i \in s)$ be the vector such that $P(X_1 = y_{i_1}, \dots, X_n = y_{i_n}) = p(s)/n!$. Let \mathcal{Y} be the range of y_i , $\theta = (y_1, \dots, y_N)$ and $\Theta = \prod_{i=1}^N \mathcal{Y}$. Under simple random sampling without replacement, the population under consideration is a parametric family indexed by $\theta \in \Theta$.

Theorem 2 (Watson-Royall theorem): (i) If $p(s) > 0$ for all s , then the vector of order statistics $X_{(1)} \leq \dots \leq X_{(n)}$ is complete for $\theta \in \Theta$. (ii) Under simple random sampling without replacement, the vector of order statistics is sufficient for $\theta \in \Theta$. (iii) Under simple random sampling without replacement, for any estimable function of θ , its unique UMVUE is the unbiased estimator $g(X_1, \dots, X_n)$, where g is symmetric in its n arguments.

3.2 Characteristic of UMVUE and Fisher information bound

Remark 1: When a complete and sufficient statistic is not available, it is usually very difficult to derive a UMVUE. In some cases, the following result can be applied, if we have enough knowledge about unbiased estimators of 0.

Theorem 1: Let \mathcal{U} be the set of all unbiased estimators of 0 with finite variances and T be an unbiased estimator of θ with $\mathbb{E}(T^2) < \infty$. (i) A necessary and sufficient condition for $T(X)$ to be a UMVUE of θ is that $\mathbb{E}[T(X)U(X)] = 0$ for any $U \in \mathcal{U}$ and any $P \in \mathcal{P}$. (ii) Suppose that $T = h(\tilde{T})$, where \tilde{T} is a sufficient statistic for $P \in \mathcal{P}$ and h is a Borel function. Let $\mathcal{U}_{\tilde{T}}$ be the subset of \mathcal{U} consisting of Borel functions of \tilde{T} . Then a necessary and sufficient condition for T to be a UMVUE of θ is that $\mathbb{E}[T(X)U(X)] = 0$ for any $U \in \mathcal{U}_{\tilde{T}}$ and any $P \in \mathcal{P}$. The theorem can be used to find a UMVUE, check whether a particular estimator is a UMVUE and show the nonexistence of any UMVUE.

Theorem 2: (i) If T_j is a UMVUE of $\theta_j, j = 1, \dots, k$, then $\sum_{j=1}^k c_j T_j$ is a UMVUE of $\theta = \sum_{j=1}^k c_j \theta_j$ for any constants c_1, \dots, c_k . (ii) If T_1 and T_2 are two UMVUE's of θ , then $T_1 = T_2$ a.s. P for any $P \in \mathcal{P}$.

Example 1: Let X_1, \dots, X_n be i.i.d. from the uniform distribution on the interval $(0, \theta)$. We have shown that $(1+n^{-1})X_{(n)}$ is the UMVUE for θ when the parameter space is $\Theta = (0, \infty)$. Suppose now that $\Theta = [1, \infty)$. Then $X_{(n)}$ is not complete, although it is still sufficient for θ . We now illustrate how to use Theorem 1 to find a UMVUE of θ . Let $U(X_{(n)})$ be an unbiased estimator of 0. Since $X_{(n)}$ has the Lebesgue p.d.f $n\theta^{-n}x^{n-1}1_{(0,\theta)}(x)$, $0 = \int_0^1 U(x)x^{n-1}dx + \int_1^\theta U(x)x^{n-1}dx$ for all $\theta \geq 1$. This implies that $U(x) = 0$ a.e. Lebesgue measure on $[1, \infty)$ and $\int_0^1 U(x)x^{n-1}dx = 0$. Consider $T = h(X_{(n)})$. To have $\mathbb{E}(TU) = 0$, we must have $\int_0^1 h(x)U(x)x^{n-1}dx = 0$. Thus, we may consider the

following function: $h(x) = \begin{cases} c & 0 \leq x \leq 1 \\ bx & x > 1 \end{cases}$, where c and b are some constants. Since $\mathbb{E}[h(X_{(n)})] = \theta$,

we obtain that $\theta = cP(X_{(n)} \leq 1) + b\mathbb{E}[X_{(n)}1_{(1,\infty)}(X_{(n)})] = c\theta^{-n} + \frac{bn}{n+1}(\theta - \theta^{-n})$. Thus, $c = 1$ and

$b = (n+1)/n$. The UMVUE of θ is then $h(X_{(n)}) = \begin{cases} 1 & 0 \leq X_{(n)} \leq 1 \\ (1+n^{-1})X_{(n)} & X_{(n)} > 1 \end{cases}$.

Theorem 3 (Cramér-Rao lower bound): Let $X = (X_1, \dots, X_n)$ be a sample from $P \in \mathcal{P} = \{P_\theta : \theta \in \Theta\}$, where Θ is an open set in \mathbb{R}^k . Suppose that $T(X)$ is an estimator with $\mathbb{E}[T(X)] = g(\theta)$ being a differentiable function of θ ; P_θ has a p.d.f. f_θ w.r.t. a measure ν for all $\theta \in \Theta$; and f_θ is differentiable as a function of θ and satisfies $\frac{\partial}{\partial \theta} \int h(x)f_\theta(x)d\nu = \int h(x)\frac{\partial}{\partial \theta} f_\theta(x)d\nu, \theta \in \Theta$ for $h(x) \equiv 1$ and $h(x) = T(x)$. Then $\text{Var}(T(X)) \geq [\frac{\partial}{\partial \theta} g(\theta)]^T [I(\theta)]^{-1} \frac{\partial}{\partial \theta} g(\theta)$, where $I(\theta) = \mathbb{E}\{\frac{\partial}{\partial \theta} \log f_\theta(X) [\frac{\partial}{\partial \theta} \log f_\theta(X)]^T\}$ is assumed to be positive definite for any $\theta \in \Theta$ and is called the Fisher information matrix.

Proposition 1: (i) If X and Y are independent with the Fisher information matrices $I_X(\theta)$ and $I_Y(\theta)$, respectively, then the Fisher information about θ contained in (X, Y) is $I_X(\theta) + I_Y(\theta)$. (ii) Suppose that X has the p.d.f. f_θ that is twice differentiable in θ and $\frac{\partial}{\partial \theta} \int h(x) f_\theta(x) d\nu = \int h(x) \frac{\partial}{\partial \theta} f_\theta(x) d\nu$ holds with $h(x) \equiv 1$ and f_θ replaced by $\partial f_\theta / \partial \theta$. Then $I(\theta) = -\mathbb{E}[\frac{\partial^2}{\partial \theta \partial \theta^T} \log f_\theta(X)]$.

Remark 2: If $\theta = \psi(\eta)$ and ψ is differentiable, then the Fisher information that X contains about η is $\frac{\partial}{\partial \eta} \psi(\eta) I(\psi(\eta)) [\frac{\partial}{\partial \eta} \psi(\eta)]^T$. However, the Cramér-Rao lower bound is not affected by any one-to-one reparameterization.

Proposition 2: Suppose that the distribution of X is from an exponential family $\{f_\theta : \theta \in \Theta\}$, i.e., the p.d.f. of X w.r.t. a σ -finite measure is $f_\theta(x) = \exp\{\eta(\theta)^T T(X) - \xi(\theta)\} c(x)$, where Θ is an open subset of \mathbb{R}^k . (i) The regularity condition $\frac{\partial}{\partial \theta} \int h(x) f_\theta(x) d\nu = \int h(x) \frac{\partial}{\partial \theta} f_\theta(x) d\nu$ is satisfied for any h with $\mathbb{E}|h(X)| < \infty$ and $I(\theta) = -\mathbb{E}[\frac{\partial^2}{\partial \theta \partial \theta^T} \log f_\theta(X)]$. (ii) If $I(\eta)$ is the Fisher information matrix for the natural parameter η , then the variance-covariance matrix $\text{Var}(T) = I(\eta)$. (iii) If $I(\theta)$ is the Fisher information matrix for the parameter $\vartheta = \mathbb{E}[T(X)]$, then $\text{Var}(T) = [I(\vartheta)]^{-1}$.

3.3 U- and V-statistics

Definition 1 (U-statistics): Let X_1, \dots, X_n be i.i.d. from an unknown population P in a non-parametric family \mathcal{P} . If the vector of order statistic is sufficient and complete for $P \in \mathcal{P}$, then a symmetric unbiased estimator of an estimable θ is the UMVUE of θ . In many problems, parameters to be estimated are of the form $\theta = \mathbb{E}[h(X_1, \dots, X_m)]$ with a positive integer m and a Borel function h that is symmetric and satisfies $\mathbb{E}|h(X_1, \dots, X_m)| < \infty$ for any $P \in \mathcal{P}$. An effective way of obtaining an unbiased estimator of θ is to use $U_n = (C_n^m)^{-1} \sum_c h(X_{i_1}, \dots, X_{i_m})$, where \sum_c denotes the summation over the C_n^m combinations of m distinct elements $\{i_1, \dots, i_m\}$ from $\{1, \dots, n\}$. The statistic is called a U-statistic with kernel h of order m .

Example 1: Consider the estimation of μ^m , where $\mu = \mathbb{E}X_1$ and m is an integer > 0 . Using $h(x_1, \dots, x_m) = x_1 \cdots x_m$, we obtain the following U-statistic for μ^m : $U_n = (C_n^m)^{-1} \sum_c X_{i_1} \cdots X_{i_m}$. Consider next the estimation of $\sigma^2 = \mathbb{E}[(X_1 - X_2)^2/2]$, we obtain the following U-statistic with kernel $h(x_1, x_2) = (x_1 - x_2)^2/2$: $U_n = \frac{2}{n(n-1)} \sum_{1 \leq i < j \leq n} \frac{(X_i - X_j)^2}{2} = \frac{1}{n-1} (\sum_{i=1}^n X_i^2 - n\bar{X}^2) = S^2$, which is the sample variance.

Theorem 1 (Hoeffding's theorem): For a U-statistic U_n with $\mathbb{E}[h(X_1, \dots, X_m)]^2 < \infty$, $\text{Var}(U_n) = (C_n^m)^{-1} \sum_{k=1}^m C_m^k C_{n-m}^{m-k} \zeta_k$, where $\zeta_k = \text{Var}(h_k(X_1, \dots, X_k))$, $h_k(x_1, \dots, x_k) = \mathbb{E}[h(X_1, \dots, X_m) | X_1 = x_1, \dots, X_k = x_k] = \mathbb{E}[h(x_1, \dots, x_k, X_{k+1}, \dots, X_m)]$, $\tilde{h}_k = h_k - \mathbb{E}[h(X_1, \dots, X_m)]$.

Proposition 1: (i) $\frac{m^2}{n} \zeta_1 \leq \text{Var}(U_n) \leq \frac{m}{n} \zeta_m$; (ii) $(n+1)\text{Var}(U_{n+1}) \leq n\text{Var}(U_n)$ for any $n > m$; (iii) For any fixed m and $k = 1, \dots, m$, if $\zeta_j = 0$ for $j < k$ and $\zeta_k > 0$, then $\text{Var}(U_n) = \frac{k!(C_m^k)^2 \zeta_k}{n^k} + O(\frac{1}{n^{k+1}})$.

Example 2: Consider $h(x_1, x_2) = x_1 x_2$, the U-statistic unbiased for μ^2 , $\mu = \mathbb{E}X_1$. Note that $h_1(x_1) = \mu x_1$, $\tilde{h}_1(x_1) = \mu(x_1 - \mu)$. $\zeta_1 = \mathbb{E}[\tilde{h}_1(X_1)]^2 = \mu^2 \text{Var}(X_1) = \mu^2 \sigma^2$, $\tilde{h}(x_1, x_2) = x_1 x_2 - \mu^2$, and $\zeta_2 = \text{Var}(X_1 X_2) = (\mu^2 + \sigma^2)^2 - \mu^4$. Thus for $U_n = (C_n^2)^{-1} \sum_{1 \leq i < j \leq n} X_i X_j$, $\text{Var}(U_n) = (C_n^2)^{-1} (C_2^1 C_{n-2}^1 \zeta_1 + C_2^2 C_{n-2}^0 \zeta_2) = \frac{2}{n(n-1)} [2(n-2)\mu^2 \sigma^2 + (\mu^2 + \sigma^2)^2 - \mu^4] = \frac{4\mu^2 \sigma^2}{n} + \frac{2\sigma^4}{n(n-1)}$.

Remark 1 (Asymptotic distributions of U-statistics): For nonparametric \mathcal{P} , the exact distribution of U_n is hard to derive. We study the method of projection, which is particularly effective for studying asymptotic distributions of U-statistics.

Definition 2: Let T_n be a given statistic based on X_1, \dots, X_n . The projection of T_n on k_n random elements Y_1, \dots, Y_{k_n} is defined to be $\tilde{T}_n = \mathbb{E}(T_n) + \sum_{i=1}^{k_n} [\mathbb{E}(T_n|Y_i) - \mathbb{E}(T_n)]$.

Theorem 2: Let T_n be a symmetric statistics with $\text{Var}(T_n) < \infty$ for every n and \tilde{T}_n be the projection of T_n on X_1, \dots, X_n . Then $\mathbb{E}(T_n) = \mathbb{E}(\tilde{T}_n)$ and $\mathbb{E}(T_n - \tilde{T}_n)^2 = \text{Var}(T_n) - \text{Var}(\tilde{T}_n)$.

Example 3: For a U-statistic U_n , one can show that $\tilde{U}_n = \mathbb{E}(U_n) + \frac{m}{n} \sum_{i=1}^n \tilde{h}_1(X_i)$, where \tilde{U}_n is the projection of U_n on X_1, \dots, X_n and $\tilde{h}_1(x) = h_1(x) - \mathbb{E}[h(X_1, \dots, X_m)]$, $h_1(x) = \mathbb{E}[h(x, X_2, \dots, X_m)]$. Hence, if $\zeta_1 = \text{Var}(\tilde{h}_1(X_i)) > 0$, $\text{Var}(\tilde{U}_n) = m^2 \zeta_1 / n$ and $\mathbb{E}(U_n - \tilde{U}_n)^2 = O(n^{-2})$. If $\zeta_1 = 0$ but $\zeta_2 > 0$, then we can show that $\mathbb{E}(U_n - \tilde{U}_n)^2 = O(n^{-3})$. One may derive results for the cases where $\zeta_2 = 0$, but the case of either $\zeta_1 > 0$ or $\zeta_2 > 0$ is the most interesting case in applications.

Theorem 3: Let U_n be a U-statistic with $\mathbb{E}[h(X_1, \dots, X_m)]^2 < \infty$. (i) If $\zeta_1 > 0$, then $\sqrt{n}[U_n - \mathbb{E}(U_n)] \rightarrow_d \mathcal{N}(0, m^2 \zeta_1)$. (ii) If $\zeta_1 = 0$ but $\zeta_2 > 0$, then $n[U_n - \mathbb{E}(U_n)] \rightarrow_d \frac{m(m-1)}{2} \sum_{j=1}^{\infty} \lambda_j (\chi_{1j}^2 - 1)$, where χ_{1j}^2 's are i.i.d. random variables having the chi-square distribution χ_1^2 and λ_j 's are some constants (which may depend on P) satisfying $\sum_{j=1}^{\infty} \lambda_j^2 = \zeta_2$.

Proposition 2: $\mathbb{E}[\frac{m(m-1)}{2} \sum_{j=1}^{\infty} \lambda_j (\chi_{1j}^2 - 1)]^2 = \frac{m^2(m-1)^2}{2} \zeta_2$.

Definition 3 (V-statistics): Let X_1, \dots, X_n be i.i.d. from P . For every U-statistic U_n as an estimator $\theta = \mathbb{E}[h(X_1, \dots, X_m)]$, there is a closely related V-statistic defined by $V_n = \frac{1}{n^m} \sum_{i_1=1}^n \dots \sum_{i_m=1}^n h(X_{i_1}, \dots, X_{i_m})$. As an estimator of θ , V_n is biased; but the bias is small asymptotically. For a fixed n , V_n may be better than U_n in terms of the mse.

Proposition 3: (i) Assume that $\mathbb{E}|h(X_{i_1}, \dots, X_{i_m})| < \infty$ for all $1 \leq i_1 \leq \dots \leq i_m \leq m$. Then the bias of V_n satisfies $b_{V_n}(P) = O(n^{-1})$. (ii) Assume that $\mathbb{E}[h(X_{i_1}, \dots, X_{i_m})]^2 < \infty$ for all $1 \leq i_1 \leq \dots \leq i_m \leq m$. Then the variance of V_n satisfies $\text{Var}(V_n) = \text{Var}(U_n) + O(n^{-2})$.

Theorem 4: Let V_n be a V-statistic with $\mathbb{E}[h(X_{i_1}, \dots, X_{i_m})]^2 < \infty$ for all $1 \leq i_1 \leq \dots \leq i_m \leq m$. (i) If $\zeta_1 = \text{Var}(h_1(X_1)) > 0$, then $\sqrt{n}(V_n - \theta) \rightarrow_d \mathcal{N}(0, m^2 \zeta_1)$. (ii) If $\zeta_1 = 0$ but $\zeta_2 = \text{Var}(h_2(X_1, X_2)) > 0$, then $n(V_n - \theta) \rightarrow_d \frac{m(m-1)}{2} \sum_{j=1}^{\infty} \lambda_j \chi_{1j}^2$.

3.4 Construction of unbiased or approximately unbiased estimators and method of moments

Definition 1 (Survey samples from a finite population): Let $\mathcal{P} = \{1, \dots, N\}$ be a finite population of interest. For each $i \in \mathcal{P}$, let y_i be a value of interest associated with unit i . Let $s = \{i_1, \dots, i_n\}$ be a subset of distinct elements of \mathcal{P} , which is a sample selected with selection probability $p(s)$, where p is known. The value y_i is observed iff $i \in s$. $Y = \sum_{j=1}^N y_j$ is the unknown population total of interest. Define π_i = probability that $i \in s, i = 1, \dots, N$.

Theorem 1: (i) (Horvitz-Thompson). If $\pi_i > 0$ for $i = 1, \dots, N$ and π_i is known when $i \in s$, then $\hat{Y}_{ht} = \sum_{i \in s} y_i / \pi_i$ is an unbiased estimator of the population total Y . (ii) Define π_{ij} = probability that $i \in s$ and $j \in s, i = 1, \dots, N, j = 1, \dots, N$. Then $\text{Var}(\hat{Y}_{ht}) = \sum_{i=1}^N \sum_{j=i+1}^N (\pi_i \pi_j - \pi_{ij}) (\frac{y_i}{\pi_i} - \frac{y_j}{\pi_j})^2$.

Remark 1 (Deriving asymptotically unbiased estimators): An exactly unbiased estimator may not exist, or is hard to obtain. We often derive asymptotically unbiased estimators. Functions of sample means are popular estimators.

Remark 2 (Functions of unbiased estimators): If the parameter to be estimated is $\vartheta = g(\theta)$ with a vector-valued parameter θ and U_n is a vector of unbiased estimators of components of θ ,

then $T_n = g(U_n)$ is often asymptotically unbiased for ϑ . Note that $\mathbb{E}(T_n) = \mathbb{E}g(U_n)$ may not exist. Assume that g is differentiable and $c_n(U_n - \theta) \rightarrow_d Y$. Then $\text{amse}_{T_n}(P) = \mathbb{E}\{[\nabla g(\theta)]^T Y\}^2 / c_n^2$. Hence, T_n has a good performance in terms of amse if U_n is optimal in terms of mse.

Definition 2 (Method of moments): Consider a parametric problem where X_1, \dots, X_n are i.i.d. random variables from $P_\theta, \theta \in \Theta \subset \mathbb{R}^k$, and $\mathbb{E}|X_1|^k < \infty$. Let $\mu_j = \mathbb{E}X_1^j$ be the j th moment of P and let $\hat{\mu}_j = \frac{1}{n} \sum_{i=1}^n X_i^j$ be the j th sample moment, which is an unbiased estimator of $\mu_j, j = 1, \dots, k$. Typically, $\mu_j = h_j(\theta), j = 1, \dots, k$, for some functions h_j on \mathbb{R}^k . By substituting μ_j 's on the left-hand side by the sample moments $\hat{\mu}_j$, we obtain a moment estimator $\hat{\theta}$, i.e. $\hat{\theta}$ satisfies $\hat{\mu}_j = h_j(\hat{\theta}), j = 1, \dots, k$. This method of deriving estimators is called the method of moments.

Example 1: Let X_1, \dots, X_n be i.i.d. from a population P_θ indexed by the parameter $\theta = (\mu, \sigma^2)$, where $\mu = \mathbb{E}X_1 \in \mathbb{R}$ and $\sigma^2 = \text{Var}(X_1) \in (0, \infty)$. Since $\mathbb{E}X_1 = \mu$ and $\mathbb{E}X_1^2 = \sigma^2 + \mu^2$, setting $\hat{\mu}_1 = \mu$ and $\hat{\mu}_2 = \sigma^2 + \mu^2$ we obtain the moment estimator $\hat{\theta} = (\bar{X}, \frac{1}{n} \sum_{i=1}^n (X_i - \bar{X})^2)$.

4 Estimation in Parametric Models

4.1 Bayesian approach

Definition 1 (Bayesian approach): X is from a population in a parametric family $\mathcal{P} = P_\theta : \theta \in \Theta$, where $\theta \in \mathbb{R}^k$ for a fixed integer $k \geq 1$. θ is viewed as a realization of a random vector $\theta \in \Theta$ whose prior distribution is Π . Prior distribution: past experience, past data, or a statistician's belief (subjective). Sample $X \in \mathcal{X}$: from $P_\theta = P_{x|\theta}$, the conditional distribution of X given θ . Posterior distribution: updated prior distribution using observed $X = x$.

Theorem 1 (Bayes formula): Assume $\mathcal{P} = \{P_{x|\theta} : \theta \in \Theta\}$ is dominated by a σ -finite measure ν and $f_\theta(x) = dP_{x|\theta}/d\nu$ is a Borel function on $(\mathcal{X} \times \Theta, \sigma(\mathcal{B}_\mathcal{X} \times \mathcal{B}_\Theta))$. Let Π be a prior distribution on Θ . Suppose that $m(x) = \int_\Theta f_\theta(x) d\Pi > 0$. (i) The posterior distribution $P_{\theta|x} \ll \Pi$ and $dP_{\theta|x}/d\Pi = f_\theta(x)/m(x)$. (ii) If $\Pi \ll \lambda$ and $d\Pi/d\lambda = \pi(\theta)$ for a σ -finite measure λ , then $dP_{\theta|x}/d\lambda = f_\theta(x)\pi(\theta)/m(x)$.

Definition 2 (Bayes action): Let \mathcal{A} be an action space in a decision problem and $L(\theta, a) \geq 0$ be a loss function. For any $x \in \mathcal{X}$, a Bayes action w.r.t. Π is any $\delta(x) \in \mathcal{A}$ such that $\mathbb{E}[L(\theta, \delta(x))|X = x] = \min_{a \in \mathcal{A}} \mathbb{E}[L(\theta, a)|X = x]$ where the expectation is w.r.t. the posterior distribution $P_{\theta|x}$.

Definition 3 (Conjugate prior): An interesting phenomenon is that the prior and the posterior are in the same parametric family of distributions. Such a prior is called a conjugate prior.

Definition 4 (Generalized Bayes action): The minimization in Definition 4.1 is the same as the minimizing $\int_\Theta L(\theta, \delta(x)) f_\theta(x) d\Pi = \min_{a \in \mathcal{A}} \int_\Theta L(\theta, a) f_\theta(x) d\Pi$. This is still defined even if Π is not a probability measure but a σ -finite measure on Θ , in which case $m(x)$ may not be finite. If $\Pi(\Theta) \neq 1$, Π is called an improper prior. $\delta(x)$ is called a generalized Bayes action.

Definition 5 (Hyperparameters and empirical Bayes): A Bayes action depends on the chosen prior with a vector ξ of parameters called hyperparameters. If the hyperparameters ξ is unknown, one way to solve the problem is to estimate ξ using some historical data; the resulting Bayes action is called an empirical Bayes action. If there is no historical data, we may estimate ξ using data x and the resulting Bayes action is also called an empirical Bayes action. The simplest empirical Bayes method is to

estimate ξ by viewing x as a “sample” from the marginal distribution $P_{x|\xi}(A) = \int_{\Theta} P_{x|\theta}(A) d\Pi_{\theta|\xi}$, $A \in \mathcal{B}_X$, where $\Pi_{\theta|\xi}$ is a prior depending on ξ or from the marginal p.d.f. $m(x) = \int_{\Theta} f_{\theta}(x) d\Pi$, if $P_{x|\theta}$ has a p.d.f. f_{θ} . The method of moments can be applied to estimate ξ .

Example 1: Let $X = (X_1, \dots, X_n)$ and X_i 's be i.i.d. with an unknown mean $\mu \in \mathbb{R}$ and a known variance σ^2 . Assume the prior $\Pi_{\mu|\xi}$ has mean μ_0 and variance σ_0^2 , $\xi = (\mu_0, \sigma_0^2)$. To obtain a moment estimate of ξ , we need to calculate $\int_{\mathbb{R}^n} x_1 m(x) dx$ and $\int_{\mathbb{R}^n} x_1^2 m(x) dx$, $x = (x_1, \dots, x_n)$. These two integrals can be obtained without knowing $m(x)$. Note that $\int_{\mathbb{R}^n} x_1 m(x) dx = \int_{\Theta} \int_{\mathbb{R}^n} x_1 f_{\mu}(x) dx d\Pi_{\mu|\xi} = \int_{\mathbb{R}} \mu d\Pi_{\mu|\xi} = \mu_0$ and $\int_{\mathbb{R}^n} x_1^2 m(x) dx = \int_{\Theta} \int_{\mathbb{R}^n} x_1^2 f_{\mu}(x) dx d\Pi_{\mu|\xi} = \sigma^2 + \int_{\mathbb{R}} \mu^2 d\Pi_{\mu|\xi} = \sigma^2 + \mu_0^2 + \sigma_0^2$. Thus, by viewing x_1, \dots, x_n as a sample from $m(x)$, we obtain the moment estimates $\hat{\mu}_0 = \bar{x}$ and $\hat{\sigma}_0^2 = \frac{1}{n} \sum_{i=1}^n (x_i - \bar{x})^2 - \sigma^2$, where \bar{x} is the sample mean of x_i 's.

Definition 6 (Hierarchical Bayes): Instead of estimating hyperparameters, in the hierarchical Bayes approach we put a prior on hyperparameters. Let $\Pi_{\theta|\xi}$ be a prior with a hyperparameter vector ξ and let Λ be a prior on Ξ , the range of ξ . Then the “marginal” prior for θ is defined by $\Pi(B) = \int_{\Xi} \Pi_{\theta|\xi}(B) d\Lambda(\xi)$, $B \in \mathcal{B}_{\Theta}$. If the second-stage prior Λ also depends on some unknown hyperparameters, then one can go on to consider a third-stage prior. In most applications, however, two-stage priors are sufficient, since misspecifying a second-stage prior is much less serious than misspecifying a first-stage prior.

Example 2: If $\bar{X} \sim \mathcal{N}(\mu, \sigma^2/n)$ with a known σ^2 , the prior $\pi(\mu|\xi)$ is the p.d.f. of $\mathcal{N}(\xi, \sigma_0^2)$ with a known σ_0^2 , and the prior of ξ is $\mathcal{N}(\mu_0, \tau^2)$ with a known μ_0 and τ^2 , then the marginal prior p.d.f. of μ is $\mathcal{N}(\mu_0, \sigma_0^2 + \tau^2)$.

4.2 Bayes rule and computation

Theorem 1 (Admissibility of Bayes rule) In a decision problem, let $\delta(x)$ be a Bayes rule w.r.t. a prior Π . (i) If $\delta(X)$ is a unique Bayes rule, then $\delta(X)$ is admissible. (ii) If Θ is countable set, the Bayes risk $r_{\delta}(\Pi) < \infty$, and Π gives positive probability to each $\theta \in \Theta$, then $\delta(X)$ is admissible. (iii) Let \mathcal{E} be the class of decision rules having continuous risk functions. If $\delta(X) \in \mathcal{E}$, $r_{\delta}(\Pi) < \infty$, and Π gives positive probability to any open subset of Θ , then $\delta(X)$ is \mathcal{E} -admissible.

Theorem 2: Suppose that Θ is an open set of \mathbb{R}^k . In a decision problem, let \mathcal{E} be the class of decision rules having continuous risk functions. A decision rule $T \in \mathcal{E}$ is \mathcal{E} -admissible if there exists a sequence $\{\Pi_j\}$ of priors such that (a) the generalized Bayes risks $r_T(\Pi_j)$ are finite for all j ; (2) for any $\theta_0 \in \Theta$ and $\eta > 0$, $\lim_{j \rightarrow \infty} \frac{r_T(\Pi_j) - r_j^*(\Pi_j)}{\Pi_j(O_{\theta_0, \eta})} = 0$, where $r_j^*(\Pi_j) = \inf_{T \in \mathcal{E}} r_T(\Pi_j)$ and $O_{\theta_0, \eta} = \{\theta \in \Theta : \|\theta - \theta_0\| < \eta\}$ with $\Pi_j(O_{\theta_0, \eta}) < \infty$ for all j .

Proposition 1 (Bayes estimators are biased): If $\delta(X)$ is a Bayes estimator of $\vartheta = g(\theta)$ under the squared error loss, then $\delta(X)$ is not unbiased except in the trivial case where $r_{\delta}(\Pi) = 0$.

Theorem 3: Suppose that X has a p.d.f. $f_{\theta}(x)$ w.r.t. a σ -finite measure ν . Suppose that $\theta = (\theta_1, \theta_2)$, $\theta_j \in \Theta_j$, and that the prior has a p.d.f. $\pi(\theta) = \pi_{\theta_1|\theta_2}(\theta_1) \pi_{\theta_2}(\theta_2)$ where $\pi_{\theta_2}(\theta_2)$ is a p.d.f. w.r.t. a σ -finite measure ν_2 on Θ_2 and for any given θ_2 , $\pi_{\theta_1|\theta_2}(\theta_1)$ is a p.d.f. w.r.t. a σ -finite measure ν_1 on Θ_1 . Suppose further that if θ_2 is given, the Bayes estimator of $h(\theta_1) = g(\theta_1, \theta_2)$ under the squared error loss is $\delta(X, \theta_2)$. Then the Bayse estimator of $g(\theta_1, \theta_2)$ under the squared error loss is $\delta(X)$ with $\delta(x) = \int_{\Theta_2} \delta(x, \theta_2) p_{\theta_2|x}(\theta_2) d\nu_2$ where $p_{\theta_2|x}(\theta_2)$ is the posterior p.d.f. of θ_2 given $X = x$.

Remark 1: Often, Bayes actions or estimators have to be computed numerically. Typically we need to compute $\mathbb{E}_p(g) = \int_{\Theta} g(\theta)p(\theta)d\nu$ with some function g , where $p(\theta)$ is a p.d.f. w.r.t. a σ -finite measure ν on $(\Theta, \mathcal{B}_{\Theta})$ and $\Theta \subset \mathbb{R}^k$. There are many numerical methods for computing integrals $\mathbb{E}_p(g)$.

Definition 1 (The simple Monte Carlo method): Generate i.i.d. $\theta^{(1)}, \dots, \theta^{(m)}$ from a p.d.f. $h(\theta) > 0$ w.r.t. ν . By the SLLN, as $m \rightarrow \infty$, $\hat{\mathbb{E}}_p(g) = \frac{1}{m} \sum_{j=1}^m \frac{g(\theta^{(j)})p(\theta^{(j)})}{h(\theta^{(j)})} \rightarrow_{\text{a.s.}} \int_{\Theta} \frac{g(\theta)p(\theta)}{h(\theta)} h(\theta)d\nu = \mathbb{E}_p(g)$.

Remark 2: The simple Monte Carlo method may not work well because (i) the convergence of $\hat{\mathbb{E}}_p(g)$ is very slow when k (the dimension of Θ) is large; (ii) generating a random vector from some k -dimensional distribution may be difficult, if not impossible.

Remark 3 (More sophisticated MCMC methods): Different from the simple Monte Carlo in two aspects: (i) generating random vectors can be done using distributions whose dimensions are much lower than k ; (ii) $\theta^{(1)}, \dots, \theta^{(m)}$ are not independent, but form a homogeneous Markov chain.

Definition 2 (Gibbs sampler): Let $y = (y_1, y_2, \dots, y_d)$. y_j 's may be vectors with different dimensions. At step $t = 1, 2, \dots$, given $y^{(t-1)}$, generate $y_1^{(t)}$ from $P(y_2^{(t-1)}, \dots, y_d^{(t-1)} | y_1^{(t-1)})$, \dots , $y_j^{(t)}$ from $P(y_1^{(t)}, \dots, y_{j-1}^{(t)}, y_{j+1}^{(t-1)}, \dots, y_k^{(t-1)} | y_j^{(t-1)})$, \dots , $y_k^{(t)}$ from $P(y_1^{(t)}, \dots, y_{k-1}^{(t)} | y_k^{(t-1)})$.

4.3 Minimality and admissibility

Definition 1 (Minimax estimator): An estimator δ is minimax if $\sup_{\theta} R_{\delta}(\theta) = \inf_T \sup_{\theta} R_T(\theta)$.

Remark 1: A minimax estimator can be very conservative and unsatisfactory. It tries to do as well as possible in the worst case. A unique minimax estimator is admissible, since any estimator better than a minimax estimator is also minimax.

Theorem 1 (Minimality of a Bayes estimator): Let Π be a proper prior on Θ and δ be a Bayes estimator of θ w.r.t. Π . Suppose δ has constant risk on Θ_{Π} . If $\Pi(\Theta_{\Pi}) = 1$, then δ is minimax. If, in addition, δ is the unique Bayes estimator w.r.t. Π , then it is the unique minimax estimator.

Theorem 2: Let $\Pi_j, j = 1, 2, \dots$ be a sequence of priors and r_j be the Bayes risk of a Bayes estimator of θ w.r.t. Π_j . Let T be a constant risk estimator of θ . If $\liminf_j r_j \geq R_T$, then T is minimax.

Example 1: Let X_1, \dots, X_n be i.i.d. components having the $\mathcal{N}(\mu, \sigma^2)$ distribution with an known $\mu = \theta \in \mathbb{R}$ and a known σ^2 . If the prior is $\mathcal{N}(\mu_0, \sigma_0^2)$, then the posterior of θ given $X = x$ is $\mathcal{N}(\mu_*(x), c^2)$ with $\mu_*(x) = \frac{\sigma^2}{n\sigma_0^2 + \sigma^2} \mu_0 + \frac{n\sigma_0^2}{n\sigma_0^2 + \sigma^2} \bar{X}$ and $c^2 = \frac{\sigma_0^2}{\sigma^2} n\sigma_0^2 + \sigma^2$. We now show that \bar{X} is minimax under the squared error loss. For any decision rule T , $\sup_{\theta \in \mathbb{R}} R_T(\theta) \geq \int_{\mathbb{R}} R_T(\theta) d\Pi(\theta) \geq \int_{\mathbb{R}} R_{\mu_*}(\theta) d\Pi(\theta) = \mathbb{E}\{[\theta - \mu_*(X)]^2\} = \mathbb{E}\{\mathbb{E}\{[\theta - \mu_*(X)]^2 | X\}\} = \mathbb{E}(c^2) = c^2$. Since this result is true for any $\sigma_0^2 > 0$ and $c^2 \rightarrow \sigma^2/n$ as $\sigma_0^2 \rightarrow \infty$, $\sup_{\theta \in \mathbb{R}} R_T(\theta) \geq \frac{\sigma^2}{n} = \sup_{\theta \in \mathbb{R}} R_{\bar{X}}(\theta)$ where the equality holds because the risk of \bar{X} under the squared error loss is σ^2/n and independent of $\theta = \mu$. Thus, \bar{X} is minimax.

Theorem 3: Let Θ_0 be a subset of Θ and T be a minimax estimator of θ when Θ_0 is the parameter space. Then T is minimax estimator if $\sup_{\theta \in \Theta} R_T(\theta) = \sup_{\theta \in \Theta_0} R_T(\theta)$.

Theorem 4 (Admissibility in one-parameter exponential families): Suppose that X has the p.d.f. $c(\theta)e^{\theta T(x)}$ w.r.t. a σ -finite measure ν , where $T(x)$ is real-valued and $\theta \in (\theta_-, \theta_+) \subset \mathbb{R}$. Consider the estimation of $\theta = \mathbb{E}[T(X)]$ under the squared error loss. Let $\lambda \geq 0$ and γ be known constants

and let $T_{\lambda,\gamma}(X) = (T + \gamma\lambda)/(1 + \lambda)$. Then a sufficient condition for the admissibility of $T_{\lambda,\gamma}$ is that $\int_{\theta_0}^{\theta_+} \frac{e^{-\gamma\lambda\theta}}{[c(\theta)]^\lambda} d\theta = \int_{\theta_-}^{\theta_0} \frac{e^{-\gamma\lambda\theta}}{[c(\theta)]^\lambda} d\theta = \infty$, where $\theta_0 \in (\theta_-, \theta_+)$.

Theorem 5: Assume that X has the p.d.f. as described in Theorem 4 with $\theta_- = -\infty$ and $\theta_+ = \infty$. (i) As an estimator of $\theta = \mathbb{E}(T)$, $T(X)$ is admissible under the squared error loss and the loss $(a - \theta)^2/\text{Var}(T)$. (ii) Y is the unique minimax estimator of θ under the loss $(a - \theta)^2/\text{Var}(T)$.

Example 2: Let X_1, \dots, X_n be i.i.d. from $\mathcal{N}(0, \sigma^2)$ with an unknown $\sigma^2 > 0$ and let $Y = \sum_{i=1}^n X_i^2$. Consider the estimation of σ^2 . The risk of $Y/(n+2)$ is a constant under the loss $(a - \sigma^2)^2/\sigma^4$. We now apply Theorem 4 to show that $Y/(n+2)$ is admissible. Note that the joint p.d.f. of X_i 's is of the form $c(\theta)e^{\theta T(x)}$ with $\theta = -n/(4\sigma^2)$, $c(\theta) = (-2\theta/n)^{n/2}$, $T(X) = 2Y/n$, $\theta_- = -\infty$ and $\theta_+ = 0$. By Theorem 4, $T_{\lambda,\gamma} = (T + \gamma\lambda)/(1 + \lambda)$ is admissible under the squared error loss if, for some $c > 0$, $\int_{-\infty}^{-c} e^{-\gamma\lambda\theta} \left(\frac{-2\theta}{n}\right)^{-n\lambda/2} d\theta = \int_0^c e^{\gamma\lambda\theta} \theta^{-n\lambda/2} d\theta = \infty$. This means $T_{\lambda,\gamma}$ is admissible if $\gamma = 0$ and $\lambda = 2/n$, or if $\gamma > 0$ and $\lambda \geq 2/n$. In particular, $2Y/(n+2)$ is admissible for estimating $\mathbb{E}(T) = 2\mathbb{E}(Y)/n = 2\sigma^2$, under the squared error loss. It is easy to see that $Y/(n+2)$ is then an admissible estimator of σ^2 under the squared error loss and the loss $(a - \sigma^2)^2/\sigma^4$. Hence $Y/(n+2)$ is minimax under the loss $(a - \sigma^2)^2/\sigma^4$.

4.4 Simultaneous estimation and shrinkage estimators

Definition 1 (Simultaneous estimation): Estimation of a p -vector ϑ of parameters (functions of θ) under the decision theory approach.

Remark 1 (Difference from estimating ϑ component-by-component): A single loss function $L(\vartheta, a)$, instead of p loss functions.

Definition 2 (Squared error loss): A natural generalization of the squared error loss is $L(\theta, a) = \|a - \theta\|^2 = \sum_{i=1}^p (a_i - \theta_i)^2$.

Definition 3 (James-Stein estimator): We start with the simple case where X is from $\mathcal{N}_p(\theta, I_p)$ with an unknown $\theta \in \mathbb{R}^p$. James and Stein proposed the following class of estimators of θ having smaller risks than X when the squared error loss is used and $p \geq 3$: $\delta_c = X - \frac{p-2}{\|X-c\|^2}(X-c)$, where $c \in \mathbb{R}^p$ is fixed and the choice of c is discussed later.

Definition 4 (Extended James-Stein estimators): For the purpose of generalizing the results to more complicated situations, we consider the following extension of the James-Stein estimator: $\delta_{c,r} = X - \frac{r(p-2)}{\|X-c\|^2}(X-c)$, where $c \in \mathbb{R}^p$ and $r \in \mathbb{R}$ are known.

Motivation 1 (Shrink the observation toward a given point c): Suppose it were thought a priori likely, though not certain, that $\theta = c$. Then we might first test a hypothesis $H_0 : \theta = c$ and estimate θ by c if H_0 is accepted and by X otherwise. The best rejection region has the form $\|X - c\|^2 > t$ for some constant $t > 0$ so that we might estimate θ by $I_{(t,\infty)}(\|X - c\|^2)X + [1 - I_{(t,\infty)}(\|X - c\|^2)]c$. $\delta_{c,r}$ is a smoothed version of this estimator, since, for some function ψ , $\delta_{c,r} = \psi(\|X - c\|^2)X + [1 - \psi(\|X - c\|^2)]c$. Any estimator having this form is called a shrinkage estimator.

Motivation 2 (Empirical Bayes estimator): A Bayes estimator of θ is of the form $\delta = (1 - B)X + Bc$, where c is the prior mean of θ and B involves prior variances. $1 - B$ is “estimated” by $\psi(\|X - c\|^2)$. $\delta_{c,r}$ can be viewed as an empirical Bayes estimator.

Theorem 1 (Risks of shrinkage estimators): Suppose that X is from $\mathcal{N}_p(\theta, I_p)$ with $p \geq 3$. Then,

under the squared error loss, the risks of the following shrinkage estimators of θ , $\delta_{c,r} = X - \frac{r(p-2)}{\|X-c\|^2}(X-c)$, where $c \in \mathbb{R}^p$ and $r \in \mathbb{R}$ are known, are given by $R_{\delta_{c,r}}(\theta) = p - (2r - r^2)(p-2)^2\mathbb{E}(\|X-c\|^{-2})$.

Remark 2: The risk of $\delta_{c,r}$ is smaller than p , the risk of X for every value of θ when $p \geq 3$ and $0 < r < 2$. $\delta = \delta_{c,1}$ is better than any $\delta_{c,r}$ with $r \neq 1$.

Remark 3 (The improvement): To see that δ_c may have a substantial improvement over X in terms of risks, consider the special case where $\theta = c$. Since $\|X-c\|^2$ has the chi-square distribution χ_p^2 when $\theta = c$, $\mathbb{E}\|X-c\|^{-2} = (p-2)^{-1}$ and $R_{\delta_{c,1}}(\theta) = p - (2r - r^2)(p-1)^2\mathbb{E}(\|X-c\|^{-2}) = 2$. The ratio $R_X(\theta)/R_{\delta_c}(\theta)$ equals $p/2$ when $\theta = c$ and can be substantially larger than 1 near $\theta = c$ when p is large.

Remark 4 (Minimaxity and admissibility of δ_c): Since X is minimax, $\delta_{c,r}$ is minimax provided that $p \geq 3$ and $0 < r < 2$. Unfortunately, the James-Stein estimator δ_c with any c is also inadmissible. It is dominated by $\delta_c^+ = X - \min\{1, \frac{p-2}{\|X-c\|^2}\}(X-c)$. This estimator, however, is still inadmissible. Although neither the James-Stein estimator δ_c nor δ_c^+ is admissible, it is found that no substantial improvements over δ_c^+ are possible.

Definition 5 (Extension of Theorem 1 to $\text{Var}(X) = \sigma^2 D$): Consider the case where $\text{Var}(X) = \sigma^2 D$ with an unknown $\sigma^2 > 0$ and a known positive definite matrix D . If σ^2 is known, then an extended James-Stein estimator is $\tilde{\delta}_{c,r} = X - \frac{(p-2)r\sigma^2}{\|D^{-1}(X-c)\|^2}D^{-1}(X-c)$. Under the squared error loss, the risk of $\tilde{\delta}_{c,r}$ is $\sigma^2[\text{tr}(D) - (2r - r^2)(p-2)^2\mathbb{E}(\|D^{-1}(X-c)\|^{-2})]$. When σ^2 is unknown, we assume that there exists a statistic S_0^2 such that S_0^2 is independent of X and S_0^2/σ^2 has the chi-square distribution χ_m^2 . Replacing $r\sigma^2$ in $\tilde{\delta}_{c,r}$ by $\hat{\sigma}^2 = tS_0^2$ with a constant $t > 0$ leads to the following extended James-Stein estimator: $\tilde{\delta}_c = X - \frac{(p-2)\hat{\sigma}^2}{\|D^{-1}(X-c)\|^2}D^{-1}(X-c)$. From the risk formula for $\tilde{\delta}_{c,r}$ and the independence of $\hat{\sigma}^2$ and X , the risk of $\tilde{\delta}_c$ is $R_{\tilde{\delta}_c}(\theta) = \sigma^2\{\text{tr}(D) - [2tm - t^2m(m+2)](p-2)^2\sigma^2\kappa(\theta)\}$, where $\theta = (\theta, \sigma^2)$ and $\kappa(\theta) = \mathbb{E}(\|D^{-1}(X-c)\|^{-2})$. Replacing t by $1/(m+2)$ leads to $R_{\tilde{\delta}_c}(\theta) = \sigma^2[\text{tr}(D) - m(m+2)^{-1}(p-2)^2\sigma^2\mathbb{E}(\|D^{-1}(X-c)\|^{-2})]$, which is smaller than $\sigma^2\text{tr}(D)$ (the risk of X) for any fixed $\theta, p \geq 3$.

Example 1: Consider the general linear model $X = Z\beta + \epsilon$ with $\epsilon \sim \mathcal{N}_p(0, \sigma^2)$, $p \geq 3$, and a full rank Z . Consider the estimation of $\theta = \beta$ under the squared error loss. The LSE $\hat{\beta}$ is from $\mathcal{N}(\beta, \sigma^2 D)$ with a known matrix $D = (Z^T Z)^{-1}$, $S_0^2 = \text{SSR}$ is independent of $\hat{\beta}$, S_0^2/σ^2 has the chi-square distribution χ_{n-p}^2 . Hence, from the previous discussion, the risk of the shrinkage estimator $\hat{\beta} - \frac{(p-2)\hat{\sigma}^2}{\|Z^T Z(\hat{\beta}-c)\|^2}Z^T Z(\hat{\beta}-c)$ is smaller than that of $\hat{\beta}$ for any β and σ^2 , where $c \in \mathbb{R}^p$ is fixed and $\hat{\sigma}^2 = \text{SSR}/(n-p+2)$.

Definition 6 (Other shrinkage estimators): From the previous discussion, the James-Stein estimators improve X substantially when we shrink the observations toward a vector c that is near $\theta = \mathbb{E}X$. One may consider shrinking the observations toward the mean of the observations rather than a given point; that is, one may obtain a shrinkage estimator by replacing c in $\delta_{c,r}$ by $\bar{X}J_p$, where $\bar{X} = p^{-1}\sum_{i=1}^p X_i$ and J_p is the p -vectors of ones. However, we have to replace the factor $p-2$ in $\delta_{c,r}$ by $p-3$. This leads to shrinkage estimators $X - \frac{p-3}{\|X-\bar{X}J_p\|^2}(X-\bar{X}J_p)$ and $X - \frac{(p-3)\hat{\sigma}^2}{\|D^{-1}(X-\bar{X}J_p)\|^2}D^{-1}(X-\bar{X}J_p)$. These estimators are better than X (and, hence, are minimax) when $p \geq 4$, under the squared error loss.

Remark 5: The idea of shrinkage has been used in problems with high dimensions, e.g. LASSO.

4.5 Likelihood and maximum likelihood estimator (MLE)

Definition 1: Let $X \in \mathcal{X}$ be a sample with a p.d.f. f_θ w.r.t. a σ -finite measure ν , where $\theta \in \Theta \subset \mathbb{R}^k$. (i) For each $x \in \mathcal{X}$, $f_\theta(x)$ considered as a function of θ is called the likelihood function and denoted by $l(\theta)$. (ii) Let $\bar{\Theta}$ be the closure of Θ . A $\hat{\theta} \in \bar{\Theta}$ satisfying $l(\hat{\theta}) = \max_{\theta \in \bar{\Theta}} l(\theta)$ is called a maximum likelihood estimate (MLE) of θ . If $\hat{\theta}$ is a Borel function of X a.e. ν , then $\hat{\theta}$ is called a maximum likelihood estimator *MLE* of θ . (iii) Let g be a Borel function from Θ to \mathbb{R}^p , $p \leq k$. If $\hat{\theta}$ is an MLE of θ , then $\hat{\vartheta} = g(\hat{\theta})$ is defined to be an MLE of $\vartheta = g(\theta)$.

Remark 1 (Finding an MLE): Since $\log x$ is a strictly increasing function, $\hat{\theta}$ is an MLE if and only if it maximizes the log-likelihood function $\log l(\theta)$. If $l(\theta)$ is differentiable on Θ° , then possible candidates for MLE's are the values of $\theta \in \Theta^\circ$ satisfying $\frac{\partial \log l(\theta)}{\partial \theta} = 0$, which is called the likelihood equation or log-likelihood equation.

Example 1: Let X_1, \dots, X_n be i.i.d. binary random variables with $P(X_1 = 1) = p \in \Theta = (0, 1)$. When $(X_1, \dots, X_n) = (x_1, \dots, x_n)$ is observed, the likelihood function is $l(p) = \prod_{i=1}^n p^{x_i} (1-p)^{1-x_i} = p^{n\bar{x}} (1-p)^{n(1-\bar{x})}$, where $\bar{x} = n^{-1} \sum_{i=1}^n x_i$. Note that $\bar{\Theta} = [0, 1]$ and $\Theta^\circ = \Theta$. The likelihood equation is $\frac{n\bar{x}}{p} - \frac{n(1-\bar{x})}{1-p} = 0$. If $0 < \bar{x} < 1$, then this equation has a unique solution \bar{x} . The second-order derivative of $\log l(p)$ is $-\frac{n\bar{x}}{p^2} - \frac{n(1-\bar{x})}{(1-p)^2}$, which is always negative. Also, when p tends to 0 or 1 (the boundary of Θ), $l(p) \rightarrow 0$. Thus, \bar{x} is the unique MLE of p .

Definition 2 (The Newton-Raphson method): In applications, MLE's typically do not have analytic forms and some numerical methods have to be used to compute MLE's. A commonly used numerical method is the Newton-Raphson iteration method, which repeatedly computes $\hat{\theta}^{(t+1)} = \hat{\theta}^{(t)} - [\frac{\partial^2 \log l(\theta)}{\partial \theta \partial \theta^T} |_{\theta=\hat{\theta}^{(t)}}]^{-1} \frac{\partial \log l(\theta)}{\partial \theta} |_{\theta=\hat{\theta}^{(t)}}$, $t = 0, 1, \dots$, where $\hat{\theta}^{(0)}$ is an initial value and $\partial^2 \log l(\theta) / \partial \theta \partial \theta^T$ is assumed of full rank for every $\theta \in \Theta$.

Definition 3 (The Fisher-scoring method): If, at each iteration, we replace $[\frac{\partial^2 \log l(\theta)}{\partial \theta \partial \theta^T} |_{\theta=\hat{\theta}^{(t)}}]^{-1}$ by $[\mathbb{E}(\frac{\partial^2 \log l(\theta)}{\partial \theta \partial \theta^T}) |_{\theta=\hat{\theta}^{(t)}}]^{-1}$, where the expectation is taken under P_θ , then the method is known as the Fisher-scoring method.

4.6 Asymptotically efficient estimation

Definition 1 (Asymptotic comparison): Let $\{\hat{\theta}_n\}$ be a sequence of estimators of θ based on a sequence of samples $\{X = (X_1, \dots, X_n), n = 1, 2, \dots\}$. Suppose that as $n \rightarrow \infty$, $\hat{\theta}_n$ is asymptotically normal (AN) in the sense that $[V_n(\theta)]^{-1/2}(\hat{\theta}_n - \theta) \rightarrow_d \mathcal{N}_k(0, I_k)$, where, for each n , $V_n(\theta)$ is a $k \times k$ positive definite matrix depending on θ . If θ is one-dimensional, then $V_n(\theta)$ is the asymptotic variance as well as the amse of $\hat{\theta}_n$. When $k > 1$, $V_n(\theta)$ is called the asymptotic covariance matrix of $\hat{\theta}_n$ and can be used as a measure of asymptotic performance of estimators. If $\hat{\theta}_{j_n}$ is AN with asymptotic covariance matrix $V_{j_n}(\theta)$, $j = 1, 2$, and $V_{1n}(\theta) \leq V_{2n}(\theta)$ for all $\theta \in \Theta$, then $\hat{\theta}_{1n}$ is said to be asymptotically more efficient than $\hat{\theta}_{2n}$.

Theorem 1: Let X_1, \dots, X_n be i.i.d. from a p.d.f. f_θ w.r.t. a σ -finite measure ν on $(\mathbb{R}, \mathcal{B})$, where $\theta \in \Theta$ and Θ is an open set in \mathbb{R}^k . Suppose that for every x in the range of X_1 , $f_\theta(x)$ is twice continuously differentiable in θ and satisfies $\frac{\partial}{\partial \theta} \int \psi_\theta(x) d\nu = \int \frac{\partial}{\partial \theta} \psi_\theta(x) d\nu$ for $\psi_\theta(x) = f_\theta(x)$ and $= \partial f_\theta(x) / \partial \theta$; the Fisher information matrix $I_1(\theta) = \mathbb{E}\{\frac{\partial}{\partial \theta} \log f_\theta(X_1) [\frac{\partial}{\partial \theta} \log f_\theta(X_1)]^T\}$ is positive

definite; and for any given $\theta \in \Theta$, there exists a positive number c_θ and a positive function h_θ such that $\mathbb{E}[h_\theta(X_1)] < \infty$ and $\sup_{\gamma: \|\gamma - \theta\| < c_\theta} \|\frac{\partial^2 \log f_\gamma(x)}{\partial \gamma \partial \gamma^T}\| \leq h_\theta(x)$ for all x in the range of X_1 , where $\|A\| = \sqrt{\text{tr}(A^T A)}$ for any matrix A . If $\hat{\theta}_n$ is an estimator of θ and is AN with $V_n(\theta) = V(\theta)/n$, then there is a $\Theta_0 \subset \Theta$ with Lebesgue measure 0 such that the information inequality $V_n(\theta) \geq [I_n(\theta)]^{-1}$ holds if $\theta \notin \Theta_0$.

Definition 2 (Asymptotic efficiency): Assume that the Fisher information matrix $I_n(\theta)$ is well defined and positive definite for every n . A sequence of estimators $\{\hat{\theta}_n\}$ that is AN is said to be asymptotically efficient or asymptotically optimal if and only if $V_n(\theta) = [I_n(\theta)]^{-1}$.

Remark 1 (Estimating a function of θ): Suppose that we are interested in estimating $\vartheta = g(\theta)$, where g is a differentiable function from Θ to \mathbb{R}^p , $1 \leq p \leq k$. If $\hat{\theta}_n$ is AN, then $\hat{\vartheta}_n = g(\hat{\theta}_n)$ is asymptotically distributed as $\mathcal{N}_p(\vartheta, [\nabla g(\theta)]^T V_n(\theta) \nabla g(\theta))$. Thus, the information inequality becomes $[\nabla g(\theta)]^T V_n(\theta) \nabla g(\theta) \geq [I_n(\vartheta)]^{-1}$, where $I_n(\vartheta)$ is the Fisher information matrix about ϑ contained in X . If $p = k$ and g is one-to-one, then $[I_n(\vartheta)]^{-1} = [\nabla g(\theta)]^T [I_n(\theta)]^{-1} \nabla g(\theta)$ and, therefore, $\hat{\theta}_n$ is asymptotically efficient if and only if $\hat{\theta}_n$ is asymptotically efficient.

Theorem 2: Assume the conditions of Theorem 1. (i) Asymptotic existence and consistency. There is a sequence of estimators $\{\hat{\theta}_n\}$ such that $P(s_n(\hat{\theta}_n) = 0) \rightarrow 1$ and $\hat{\theta}_n \rightarrow_p \theta$, where $s_n(\gamma) = \frac{\partial \log l(\gamma)}{\partial \gamma}$. (ii) Asymptotic efficiency. Any consistent sequence $\tilde{\theta}_n$ of RLE (root of the likelihood equation)'s is asymptotically normal and asymptotically efficient.

Theorem 3: Assume the conditions of Theorem 1. Let $\pi(\gamma)$ be a prior p.d.f w.r.t. the Lebesgue measure on Θ and $p_n(\gamma)$ be the posterior p.d.f., given X_1, \dots, X_n , $n = 1, 2, \dots$. Assume that there exists an n_0 such that $p_{n_0}(\gamma)$ is continuous and positive for all $\gamma \in \Theta$, $\int p_{n_0}(\gamma) d\gamma = 1$ and $\int \|\gamma\| p_{n_0}(\gamma) d\gamma < \infty$. Suppose further that, for any $\epsilon > 0$, there exists a $\delta > 0$ such that $\lim_{n \rightarrow \infty} P(\sup_{\|\gamma - \theta\| \geq \epsilon} \frac{\log l(\gamma) - \log l(\theta)}{n} > -\delta) = 0$, $\lim_{n \rightarrow \infty} P(\sup_{\|\gamma - \theta\| \leq \delta} \frac{\|\nabla s_n(\gamma) - \nabla s_n(\theta)\|}{n} \geq \epsilon) = 0$, where $l(\gamma)$ is the likelihood function and $s_n(\gamma)$ is the score function. (i) Let $p_n^*(\gamma)$ be the posterior p.d.f of $\sqrt{n}(\gamma - T_n)$, where $T_n = \theta + [I_n(\theta)]^{-1} s_n(\theta)$ and θ is the true parameter value, and let $\psi(\gamma)$ be the p.d.f. of $\mathcal{N}_k(0, [I_1(\theta)]^{-1})$. Then $\int (1 + \|\gamma\|) |p_n^*(\gamma) - \psi(\gamma)| d\gamma \rightarrow_p 0$. (ii) The Bayes estimator of θ under the squared error loss is asymptotically efficient.

Proposition 1: The posterior p.d.f. is approximately normal with mean $\theta + [I_n(\theta)]^{-1} s_n(\theta)$ and covariance matrix $[I_n(\theta)]^{-1}$.

Remark 2: The results hold regardless of the prior being used, indicating that the effect of the prior declines as $n \rightarrow \infty$.

4.7 MLE in generalized linear models (GLM) and quasi-MLE

Definition 1 (The structure of a GLM): The sample $X = (X_1, \dots, X_n)$ has independent X_i 's and X_i has the p.d.f. $\exp\{\frac{\eta_i x_i - \zeta(\eta_i)}{\phi_i}\} h(x_i, \phi_i)$, $i = 1, \dots, n$, w.r.t. a σ -finite measure ν , where η_i and ϕ_i are unknown, $\phi_i > 0$, $\eta_i \in \Xi = \{\eta : 0 < \int h(x, \phi) e^{\eta x / \phi} d\nu(x) < \infty\} \subset \mathbb{R}$ for all i , ζ and h are known functions, and $\zeta''(\eta) > 0$ is assumed for all $\eta \in \Xi$. Note that the p.d.f. belongs to an exponential family if ϕ_i is known. As a consequence, $\mathbb{E}(X_i) = \zeta'(\eta_i)$ and $\text{Var}(X_i) = \phi_i \zeta''(\eta_i)$, $i = 1, \dots, n$. Define $\mu(\eta) = \zeta'(\eta)$. It is assumed that η_i is related to Z_i , the i th value of a p -value of covariates, through $\mu(\eta_i) = \beta^T Z_i$, $i = 1, \dots, n$, where β is a p -vector of unknown parameters and g , called a link

function, is a known one-to-one, third-order continuously differentiable function on $\{\mu(\eta) : \eta \in \Xi^\circ\}$. If $\mu = g^{-1}$, then $\eta_i = \beta^T Z_i$ and g is called the canonical or natural link function. If g is not canonical, we assume that $\frac{d}{d\eta}(g \circ \mu)(\eta) \neq 0$ for all η . In a GLM, the parameter of interest is β . We assume the range of β is $B = \{\beta : (g \circ \mu)^{-1}(\beta^T z) \in \Xi^\circ \text{ for all } z \in \mathcal{Z}\}$, where \mathcal{Z} is the range of Z_i 's. ϕ_i 's are called dispersion parameters and are considered to be nuisance parameters.

Proposition 1 (MLE in GLM): An MLE of β in a GLM is considered under assumption $\phi_i = \phi/t_i, i = 1, \dots, n$, with an unknown $\phi > 0$ and known positive t_i 's. Let $\theta = (\beta, \phi)$ and $\psi = (g \circ \mu)^{-1}$. $\log l(\theta) = \sum_{i=1}^n [\log h(x_i, \frac{\phi}{t_i}) + \frac{\psi(\beta^T Z_i)x_i - \zeta(\psi(\beta^T Z_i))}{\phi/t_i}]$, $\frac{\partial \log l(\theta)}{\partial \beta} = \frac{1}{\phi} \sum_{i=1}^n \{[x_i - \mu(\psi(\beta^T Z_i))] \psi'(\beta^T Z_i) t_i Z_i\} = 0$, $\frac{\partial \log l(\theta)}{\partial \phi} = \sum_{i=1}^n \{ \frac{\partial \log h(x_i, \phi/t_i)}{\partial \phi} - \frac{t_i [\psi(\beta^T Z_i)x_i - \zeta(\psi(\beta^T Z_i))]}{\phi^2} \} = 0$. From the first likelihood equation, an MLE of β , if it exists, can be obtained without estimating ϕ . The second likelihood equation, however, is usually difficult to solve. Some other estimators of ϕ are suggested by various researchers. Suppose there is a solution $\hat{\beta}$ to the likelihood equation. $\text{Var}(\frac{\partial \log l(\theta)}{\partial \beta}) = \frac{M_n(\beta)}{\phi}$, $\frac{\partial^2 \log l(\theta)}{\partial \beta \partial \beta^T} = \frac{R_n(\beta) - M_n(\beta)}{\phi}$, where $M_n(\beta) = \sum_{i=1}^n [\psi'(\beta^T Z_i)]^2 \zeta''(\psi(\beta^T Z_i)) t_i Z_i Z_i^T$, $R_n(\beta) = \sum_{i=1}^n [x_i - \mu(\psi(\beta^T Z_i))] \psi''(\beta^T Z_i) t_i Z_i Z_i^T$. Consider first the simple case of canonical g , $\psi'' = 0$ and $R_n = 0$. If $M_n(\beta)$ is positive definite for all β , then $-\log l(\theta)$ is strictly convex in β for any fixed ϕ and, therefore, $\hat{\beta}$ is the unique MLE of β . For noncanonical g , $R_n(\beta) \neq 0$ and $\hat{\beta}$ is not necessarily an MLE. If $R_n(\beta)$ is dominated by $M_n(\beta)$, i.e., $[M_n(\beta)]^{-1/2} R_n(\beta) [M_n(\beta)]^{-1/2} \rightarrow 0$ in some sense, then $-\log l(\theta)$ is convex and $\hat{\beta}$ is an MLE for large n . In a GLM, an MLE $\hat{\beta}$ usually does not have an analytic form and a numerical method such as the Newton-Raphson has to be applied.

Example 1: Consider the GLM with $\zeta(\eta) = \eta^2/2, \eta \in \mathbb{R}$. If g is the canonical link, then the model is the same as a linear model with independent ϵ_i 's distributed as $\mathcal{N}(0, \phi_i)$. Suppose now that g is noncanonical but $\phi_i \equiv \phi$. Then the model reduces to the one with independent X_i 's and $X_i = \mathcal{N}(g^{-1}(\beta^T Z_i), \phi), i = 1, \dots, n$. This type of model is called a nonlinear regression model (with normal errors) and an MLE of β under this model is also called a nonlinear LSE, since maximizing the log-likelihood is equivalent to minimizing the sum of squares $\sum_{i=1}^n [X_i - g^{-1}(\beta^T Z_i)]^2$. Under certain conditions the matrix $R_n(\beta)$ is dominated by $M_n(\beta)$ and an MLE of β exists.

Example 2 (The Poisson model): Consider the GLM with $\zeta(\eta) = e^\eta, \eta \in \mathbb{R}, \phi_i = \phi/t_i$. If $\phi_i = 1$, then X_i has the Poisson distribution with mean e^{η_i} . Under the canonical link $g(t) = \log t$, $M_n(\beta) = \sum_{i=1}^n e^{\beta^T Z_i} t_i Z_i Z_i^T$, which is positive definite if $\inf_i e^{\beta^T Z_i} > 0$ and the matrix $(\sqrt{t_1} Z_1, \dots, \sqrt{t_n} Z_n)$ is of full rank. There is one noncanonical link that deserves attention. Suppose that we choose a link function so that $[\psi'(t)]^2 \zeta''(\psi(t)) \equiv 1$. Then $M_n(\beta) = \sum_{i=1}^n t_i Z_i Z_i^T$ does not depend on β . It is shown that the asymptotic variance of the MLE $\hat{\beta}$ is $\phi [M_n(\beta)]^{-1}$. The fact that $M_n(\beta)$ does not depend on β makes the estimation of the asymptotic variance (and, thus, statistical inference) easy. Under the Poisson model, $\zeta''(t) = e^t$ and, therefore, we need to solve the differentiable equation $[\psi'(t)]^2 e^{\psi(t)} = 1$. A solution is $\psi(t) = 2 \log(t/2)$ and the link $g(\mu) = 2\sqrt{\mu}$.

Theorem 1: Consider the GLM with $\phi_i = \phi/t_i$ and t_i 's in a fixed interval $(t_0, t_\infty), 0 < t_0 \leq t_\infty < \infty$. Assume that the range of unknown parameter β is an open subset of \mathbb{R}^p ; at the true value of $\beta, 0 < \inf_i \phi(\beta^T Z_i) \leq \sup_i \phi(\beta^T Z_i) < \infty$, where $\phi(t) = [\psi'(t)]^2 \zeta''(\psi(t))$; as $n \rightarrow \infty, \max_{i \leq n} Z_i^T (Z^T Z)^{-1} Z_i \rightarrow 0$ and $\lambda_- [Z^T Z] \rightarrow \infty$, where Z is the $n \times p$ matrix whose i th row is the vector Z_i and $\lambda_- [A]$ is the smallest eigenvalue of A . (i) There is a unique sequence of estimators $\{\hat{\beta}_n\}$ such that $P(s_n(\hat{\beta}_n) = 0) \rightarrow$

1 and $\hat{\beta}_n \rightarrow_p \beta$, where $s_n(\beta) = \partial \log l(\beta, \phi) / \partial \phi$ is the score function. (ii) Let $I_n(\beta) = \text{Var}(s_n(\beta))$. Then $[I_n(\beta)]^{1/2}(\hat{\beta}_n - \beta) \rightarrow_d \mathcal{N}_p(0, I_p)$. (iii) If ϕ is known or the p.d.f. indexed by $\theta = (\beta, \phi)$ satisfies the conditions for f_θ in Theorem 1 of section 4.6, then $\hat{\beta}_n$ is asymptotically efficient.

Definition 2 (Quasi-MLE): If assumption ϕ_i is arbitrary, or the distribution assumption on X_i does not hold, but $\mathbb{E}(X_i) = \zeta'(\eta_i)$, $\text{Var}(X_i) = \phi_i \zeta''(\eta_i)$, $i = 1, \dots, n$ and $g(\mu(\eta_i)) = \beta^T Z_i$, $i = 1, \dots, n$ still hold, we estimate β by solving equation $G_n(\beta) = \sum_{i=1}^n \{[x_i - \mu(\psi(\beta^T Z_i))] \psi'(\beta^T Z_i) t_i Z_i\} = 0$, then the resulting estimator is called a quasi-MLE. This method is also called the method of generalized estimating equations (GEE). They are efficient if the GEE is a likelihood equation, and is robust if it is not.

Remark 1: The asymptotic existence and consistency of quasi-MLE can be shown using a similar argument to the proof of Theorem 2 of section 4.6.

4.8 Other asymptotically efficient estimators and pseudo MLE

Definition 1 (One-Step MLE): Let $s_n(\gamma)$ be the score function. Let $\hat{\theta}_n^{(0)}$ be an estimator of θ that may not be asymptotically efficient. The one-step MLE is the first iteration in computing an RLE using the Newton-Raphson method with $\hat{\theta}_n^{(0)}$ as the initial value, $\hat{\theta}_n^{(1)} = \hat{\theta}_n^{(0)} - [\nabla s_n(\hat{\theta}_n^{(0)})]^{-1} s_n(\hat{\theta}_n^{(0)})$. Without any further iteration, $\hat{\theta}_n^{(1)}$ is asymptotically efficient under some conditions.

Theorem 1: Assume that the conditions in Theorem 1 of section 4.6 hold and that $\hat{\theta}_n^{(0)}$ is \sqrt{n} -consistent for θ . (i) The one-step MLE $\hat{\theta}_n^{(1)}$ is asymptotically efficient. (ii) The one-step MLE obtained by replacing $\nabla s_n(\gamma)$ with its expected value, $-I_n(\gamma)$ (the Fisher-scoring method), is asymptotically efficient.

Definition 2 (Pseudo MLE): Let X_1, \dots, X_n be a random sample from a pdf in a family indexed by two parameters θ and π with likelihood $l(\theta, \pi)$. The method of pseudo MLE may be viewed as follows. Based on the sample, an estimate $\hat{\pi}$ of π is obtained using some technique other than MLE. The pseudo MLE of θ is then obtained by maximizing the likelihood $l(\theta, \hat{\pi})$.

Remark 1: π is viewed as a nuisance parameter. Pseudo MLE consists of replacing π by an estimate and solving a reduced system of likelihood equations, which works when a higher dimensional MLE is intractable but a lower dimensional MLE is feasible. The consistency and asymptotic normality hold under fairly standard regularity conditions.

Theorem 2 (Asymptotic existence and consistency of pseudo MLE): Assume the conditions in Theorem 1 of section 4.6. Assume also $\hat{\pi}$ is a consistent estimator of π_0 . As $n \rightarrow \infty$, with probability tending to 1, there exists $\hat{\theta}$ such that $\frac{\partial \log l(\hat{\theta}, \hat{\pi})}{\partial \theta} = 0$ and $\hat{\theta} \rightarrow_p \theta_0$ where θ_0 is the true value of θ .

5 Estimation in Non-Parametric Models

5.1 Empirical c.d.f. and empirical likelihoods

Definition 1 (Estimation in nonparametric models): Data $X = (X_1, \dots, X_n)$, where X_i 's are random d -vectors i.i.d. from an unknown c.d.f. F in a nonparametric family. We study mainly two topics: estimation of the c.d.f. F and estimation of $\theta = T(F)$, where T is a functional.

Definition 2 (Empirical c.d.f.): $F_n(t) = \frac{1}{n} \sum_{i=1}^n I_{(-\infty, t]}(X_i)$, $t \in \mathbb{R}^d$, where $(-\infty, a]$ denotes the set $(-\infty, a_1] \times \cdots \times (-\infty, a_d]$ for any $a = (a_1, \dots, a_d) \in \mathbb{R}^d$. F_n is the distribution putting mass n^{-1} at each X_i , $i = 1, \dots, n$.

Proposition 1 (Properties of empirical c.d.f.): (i) For any $t \in \mathbb{R}^d$, $nF_n(t)$ has the binomial distribution $B(F(t), n)$; (ii) $F_n(t)$ is unbiased variance $F(t)[1 - F(t)]/n$; (iii) $F_n(t)$ is the UMVUE under some nonparametric models; (iv) $F_n(t)$ is \sqrt{n} -consistent for $F(t)$.

Theorem 1: Define sup-norm distance $\rho_\infty(G_1, G_2) = \|G_1 - G_2\|_\infty = \sup_{t \in \mathbb{R}^d} |G_1(t) - G_2(t)|$, $G_j \in \mathcal{F}$. (i) When $d = 1$, there exists a positive constant C (not depending on F) such that $P(\rho_\infty(F_n, F) > z) \leq Ce^{-2nz^2}$, $z > 0, n = 1, 2, \dots$. (ii) When $d \geq 2$, for any $\epsilon > 0$, there exists a positive constant $C_{\epsilon, d}$ (not depending on F) such that $P(\rho_\infty(F_n, F) > z) \leq C_{\epsilon, d}e^{-(2-\epsilon)nz^2}$, $z > 0, n = 1, 2, \dots$.

Theorem 2: Let F_n be the empirical c.d.f. of i.i.d. X_1, \dots, X_n from a c.d.f. F on \mathbb{R}^d . (i) $\rho_\infty(F_n, F) \rightarrow_{a.s.} 0$ as $n \rightarrow \infty$; (ii) $\mathbb{E}[\sqrt{n}\rho_\infty(F_n, F)]^s = O(1)$ for any $s > 0$.

Theorem 3: Let F_n be the empirical c.d.f. based on i.i.d. random variables X_1, \dots, X_n from a c.d.f $F \in \mathcal{F}_1$. (i) $\rho_{L_p}(F_n, F) \rightarrow_{a.s.} 0$; (ii) $\mathbb{E}[\sqrt{n}\rho_{L_p}(F_n, F)] = O(1)$ if $1 < p < 2$ and $\int \{F(t)[1 - F(t)]\}^{p/2} dt < \infty$ if $p \geq 2$.

Theorem 4: For X_1, \dots, X_n i.i.d. from $F \in \mathcal{F}$, the empirical c.d.f. F_n maximizes the nonparametric likelihood function $l(G)$ over $G \in \mathcal{F}$.

Definition 3 (Empirical likelihoods): The nonparametric MLE can be extended to various situations with some modifications of $l(G)$ and/or constraints on p_i 's. Modifications of the likelihood $l(G)$ are called empirical likelihoods. An estimator obtained by maximizing an empirical likelihood is then called a maximum empirical likelihood estimator (MELE).

Remark 1 (Estimation of F with auxiliary information about F): In some cases we have some information about F . For instance, suppose that there is a known Borel function u from \mathbb{R}^d to \mathbb{R}^s such that $\int u(x)dF = 0$. It is reasonable to expect that any estimate \hat{F} of F has property $\int u(x)d\hat{F} = 0$, which is not true for the empirical c.d.f F_n , since $\int u(x)dF_n = \frac{1}{n} \sum_{i=1}^n u(X_i) \neq 0$ even if $\mathbb{E}[u(X_1)] = 0$. Using the method of empirical likelihoods, a natural solution is to put another constraint in the process of maximizing the likelihood. That is, we maximize $l(G)$ subject to $p_i > 0, i = 1, \dots, n, \sum_{i=1}^n p_i = 1$, and $\sum_{i=1}^n p_i u(x_i) = 0$ where $p_i = P_G(\{x_i\})$. Using the Lagrange multiplier method, it can be shown that an MELE of F is $\hat{F}(t) = \sum_{i=1}^n \hat{p}_i I_{(-\infty, t]}(X_i)$, where $\hat{p}_i = n^{-1}[1 + \lambda_n^T u(X_i)]^{-1}, i = 1, \dots, n$ and $\lambda_n \in \mathbb{R}^s$ is the Lagrange multiplier satisfying $\sum_{i=1}^n \hat{p}_i u(X_i) = \frac{1}{n} \sum_{i=1}^n \frac{u(X_i)}{1 + \lambda_n^T u(X_i)} = 0$.

Theorem 5: Let u be a Borel function on \mathbb{R}^d satisfying $\int u(x)dF = 0$ and \hat{F} be the MELE of F . Suppose that $U = \text{Var}(u(X_1))$ is positive definite. Then, for any m fixed distinct $t_1, \dots, t_m \in \mathbb{R}^d$, $\sqrt{n}[(\hat{F}(t_1), \dots, \hat{F}(t_m)) - (F(t_1), \dots, F(t_m))] \rightarrow_d \mathcal{N}_m(0, \Sigma_u)$, where $\Sigma_u = \Sigma - W^T U^{-1} W$, Σ is the covariance matrix of $\sqrt{n}[(F_n(t_1), \dots, F_n(t_m)) - (F(t_1), \dots, F(t_m))]$, $W = (W(t_1), \dots, W(t_m))$, and $W(t_j) = \mathbb{E}[u(X_1)I_{(-\infty, t_j]}(X_1)]$.

5.2 Profile likelihoods, GEE, and GMM

Definition 1 (Profile likelihoods): Let $l(\theta, \xi)$ be a likelihood (or empirical likelihood), where θ and ξ are not necessarily vector-valued. It may be difficult to maximize the likelihood $l(\theta, \xi)$ simultaneously over θ and ξ . For each fixed θ , let $\xi(\theta)$ satisfy $l(\theta, \xi(\theta)) = \sup_\xi l(\theta, \xi)$. The function

$l_p(\theta) = l(\theta, \xi(\theta))$ is called a profile likelihood function for θ . Suppose that $\hat{\theta}_p$ maximizes $l_p(\theta)$. Then $\hat{\theta}_p$ is called a maximum profile likelihood estimator of θ . Although this idea can be applied to parametric models, it is more useful in semi-parametric models, especially when θ is a parametric component and ξ is a nonparametric component.

Example 1 (Missing data): Assume that X_1, \dots, X_n are i.i.d. random variables from an unknown c.d.f. F and some X_i 's are missing. Let $\delta_i = 1$ if X_i is observed and $\delta_i = 0$ if X_i is missing. Suppose that (X_i, δ_i) are i.i.d. and let $\pi(x) = P(\delta_i = 1 | X_i = x)$. If X_i and δ_i are independent, i.e. $\pi(x) \equiv \pi$ does not depend on x , then the empirical c.d.f based on observed data, i.e., the c.d.f. putting mass r^{-1} to each observed X_i , where r is the number of observed X_i 's, is an unbiased and consistent estimator of F , provided that $\pi > 0$. On the other hand, if $\pi(x)$ depends on x (called nonignorable missingness), then the empirical c.d.f. based on observed data is a biased and inconsistent estimator of F . In fact, the empirical c.d.f. based on observed data is an unbiased estimator of $P(X_i \leq x | \delta_i = 1)$, which is generally different from the unconditional probability $F(x) = P(X_i \leq x)$. If both π and F are in parametric models, then we can apply the method of maximum likelihood. For example, if $\pi(x) = \pi_\theta(x)$ and $F(x) = F_\vartheta(x)$ has a p.d.f. f_ϑ , where θ and ϑ are vectors of unknown parameters, then a parametric likelihood of (θ, ϑ) is $l(\theta, \vartheta) = \prod_{i=1}^n [\pi_\theta(x_i) f_\vartheta(x_i)]^{\delta_i} (1 - \pi)^{1 - \delta_i}$, where $\pi = \int \pi_\theta(x) f_\vartheta(x) dx$. computationally, it may be difficult to maximizing this likelihood, since π is an integral. Suppose now that $\pi(x) = \pi_\theta(x)$ is the parametric component and F is the nonparametric component. Then an empirical likelihood can be defined as $l(\theta, G) = \prod_{i=1}^n [\pi_\theta(x_i) p_i]^{\delta_i} (1 - \pi)^{1 - \delta_i}$, $p_i = P_G(\{x_i\})$ subject to $p_i \geq 0$, $\sum_{i=1}^n \delta_i p_i = 1$, $\sum_{i=1}^n \delta_i p_i [\pi_\theta(x_i) - \pi] = 0$, $i = 1, \dots, n$. It can be shown that the logarithm of the profile empirical likelihood for (θ, π) with a Lagrange multiplier is $\sum_{i=1}^n \{\delta_i \log(\pi_\theta(x_i)) + (1 - \delta_i) \log(1 - \pi) - \delta_i \log(1 + \lambda[\pi_\theta(x_i) - \pi])\}$. Under some conditions, it can be shown that the estimators $\hat{\theta}$, $\hat{\pi}$ and $\hat{\lambda}$ obtained by maximizing this likelihood are consistent and asymptotically normal and that the empirical c.d.f. putting mass $\hat{p}_i = r^{-1} \{1 + \hat{\lambda}[\pi_\theta(X_i) - \hat{\pi}]\}^{-1}$ to each observed X_i is consistent for F . The result can be extended when there is an observed covariate.

Definition 2 (Generalized estimating equation (GEE)): Assume that X_1, \dots, X_n are independent random vectors, where the dimension of X_i is d_i , $i = 1, \dots, n$ ($\sup_i d_i < \infty$), and that we are interested in estimating θ , a k -vector of unknown parameters related to the unknown population. Let $\Theta \subset \mathbb{R}^k$ be the range of θ , ψ_i be a Borel function form $\mathbb{R}^{d_i} \times \Theta$ to \mathbb{R}^k , $i = 1, \dots, n$, and $s_n(\gamma) = \sum_{i=1}^n \psi_i(X_i, \gamma)$, $\gamma \in \Theta$. If θ is estimated by $\hat{\theta} \in \Theta$ satisfying $s_n(\hat{\theta}) = 0$, then $\hat{\theta}$ is called a GEE estimator. The equation $s_n(\gamma) = 0$ is called a GEE.

Remark 1 (Motivation): Usually GEE's are chosen so that $\mathbb{E}[s_n(\theta)] = \sum_{i=1}^n \mathbb{E}[\psi_i(X_i, \theta)] = 0$, where the expectation \mathbb{E} may be replaced by an asymptotic expectation if the exact expectation does not exist.

Proposition 1 (Consistency of GEE estimators): Suppose that X_1, \dots, X_n are i.i.d. from F and $\psi \equiv \psi$, a bounded and continuous function form $\mathbb{R}^d \times \Theta$ to \mathbb{R}^k . Let $g(t) = \int \psi(x, t) dF(x)$. Suppose that $g(\theta) = 0$ and $\partial g(t) / \partial t$ exists and is of full rank at $t = \theta$. Then $\hat{\theta}_n \rightarrow_p \theta$.

Definition 3 (Generalized method of moments (GMM)): Suppose that we have a set of $m \geq k$ functions $\psi_j(x, \theta)$, $j = 1, \dots, m$ such that $\mathbb{E}_\theta[\psi_j(X, \theta)] = 0$ for all j and ψ_j 's are not linearly independent, i.e., the $m \times m$ matrix whose (j, j') th element is $\mathbb{E}_\theta[\psi_j(X_i, \theta) \psi_{j'}(X_i, \theta)]$ is positive

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definite, which can usually be achieved by eliminating some redundant functions where ψ_j 's are linearly dependent. Let $G_n(\theta) = (\frac{1}{n} \sum_{i=1}^n \psi_1(x_i, \theta), \dots, \frac{1}{n} \sum_{i=1}^n \psi_m(x_i, \theta))^T, \theta \in \Theta$. if $m = k$, a solution to $G_n(\theta) = 0$ is a GEE estimator. If $m > k$, a solution to $G_n(\theta) = 0$ may not exist. Then we can minimize $G_n^T(\theta)G_n(\theta)$, using a data driven procedure.

Definition 4 (GMM algorithm): A GMM estimate of θ can be obtained using the following two-step algorithm (the second step is to gain efficiency). (1) Obtain $\hat{\theta}^{(1)}$ by minimizing $G_n^T(\theta)G_n(\theta)/2$ over $\theta \in \Theta$. (2) Let \hat{W} be the inverse matrix of the $m \times m$ matrix whose (j, j') element is equal to $\frac{1}{n} \sum_{i=1}^n \psi_j(x_i, \hat{\theta}^{(1)})\psi_{j'}(x_i, \hat{\theta}^{(1)})$. The GMM estimate $\hat{\theta}$ is obtained by minimizing $G_n^T(\theta)\hat{W}G_n(\theta)/2$ over $\theta \in \Theta$.

6 Hypothesis Tests

6.1 Neyman-Pearson lemma and monotone likelihood ratio

Definition 1 (Theory of testing hypotheses): X : a sample from a population $P \in \mathcal{P}$, a family of populations. Based on the observed X , we test a given hypothesis $H_0 : P \in \mathcal{P}_0$ vs $H_1 : P \in \mathcal{P}_1$ where \mathcal{P}_0 and \mathcal{P}_1 are two disjoint subsets of \mathcal{P} and $\mathcal{P}_0 \cup \mathcal{P}_1 = \mathcal{P}$. A test for a hypothesis is a statistic $T(X)$ taking values in $[0, 1]$. When $X = x$ is observed, we reject H_0 with probability $T(x)$. If $T(X) = 1$ or 0 a.s. \mathcal{P} , then $T(X)$ is a nonrandomized test; otherwise $T(X)$ is randomized. For a given test $T(X)$, the power function of $T(X)$ is defined to be $\beta_T(P) = \mathbb{E}[T(X)], P \in \mathcal{P}$, which is the type I error probability of $T(X)$ when $P \in \mathcal{P}_0$ and one minus the type II error probability of $T(X)$ when $P \in \mathcal{P}_1$.

Definition 2 (Significance tests): With a sample of a fixed size, we are not able to minimize two error probabilities simultaneously. Our approach involves maximizing the power $\beta_T(P)$ over all $P \in \mathcal{P}_1$ (i.e., minimizing the type II error probability) and over all tests T satisfying $\sup_{P \in \mathcal{P}_0} \beta_T(P) \leq \alpha$, where $\alpha \in [0, 1]$ is a given level of significance. The left-hand side of the last expression is defined to be the size of T .

Definition 3: A test T_* of size α is a uniformly most powerful (UMP) test if and only if $\beta_{T_*}(P) \geq \beta_T(P)$ for all $P \in \mathcal{P}_1$ and T of level α .

Proposition 1 (Using sufficient statistics): If $U(X)$ is a sufficient statistic for $P \in \mathcal{P}$, then for any test $T(X)$, $\mathbb{E}(T|U)$ has the same power function as T and, therefore, to find a UMP test we may consider tests that re functions of U only.

Theorem 1 (Neyman-Pearson lemma): Suppose that $\mathcal{P}_0 = \{P_0\}$ and $\mathcal{P}_1 = \{P_1\}$. Let f_j be the p.d.f of P_j w.r.t. a σ -finite measure ν (e.g., $\nu = P_0 + P_1$), $j = 0, 1$. (i) Existence of a UMP test. For

every α , there exists a UMP test of size α , which is $T_*(X) = \begin{cases} 1 & f_1(X) > cf_0(X) \\ \gamma & f_1(X) = cf_0(X) \text{ where } \gamma \in [0, 1] \\ 0 & f_1(X) < cf_0(X) \end{cases}$

and $c \geq 0$ are some constants chosen so that $\mathbb{E}[T_*(X)] = \alpha$ when $P = P_0$ ($c = \infty$ is allowed). (ii)

Uniqueness. If T_{**} is a UMP test of size α , then $T_{**}(X) = \begin{cases} 1 & f_1(X) > cf_0(X) \\ 0 & f_1(X) < cf_0(X) \end{cases}$ a.s. \mathcal{P} .

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Example 1: Suppose that X is a sample of size 1, $\mathcal{P}_0 = \{P_0\}$, and $\mathcal{P}_1 = \{P_1\}$, where P_0 is $\mathcal{N}(0, 1)$ and P_1 is the double exponential distribution $\text{DE}(0, 2)$ with the p.d.f $4^{-1}e^{-|x|/2}$. Since $P(f_1(X) = cf_0(X)) = 0$, there is a unique nonrandomized UMP test. By theorem 1, the UMP test $T_*(x) = 1$ if and only if $\frac{\pi}{8}e^{x^2-|x|} > c^2$ for some $c > 0$, which is equivalent to $|x| > t$ or $|x| < 1 - t$ for some $t > \frac{1}{2}$. Suppose that $\alpha < \frac{1}{3}$. To determine t , we use $\alpha = \mathbb{E}_0[T_*(X)] = P_0(|X| > t) + P_0(|X| < 1 - t)$. If $t \leq 1$, then $P_0(|X| > t) \geq P_0(|X| > 1) = 0.3374 > \alpha$. Hence t should be larger than 1 and $\alpha = P_0(|X| > t) = \Phi(-t) + 1 - \Phi(t)$. Thus, $t = \Phi^{-1}(1 - \alpha/2)$ and $T_*(X) = I_{(t, \infty)}(|X|)$. Note that it is not necessary to find out what c is.

Theorem 2: Suppose that there is a test T_* of size α such that for every $P_1 \in \mathcal{P}_1$, T_* is UMP for testing H_0 versus the hypothesis $P = P_1$. Then T_* is the UMP for testing H_0 versus H_1 .

Definition 4: Suppose that the distribution of X is in $\mathcal{P} = \{P_\theta : \theta \in \Theta\}$, a parametric family indexed by a real-valued θ , and that \mathcal{P} is dominated by a σ -finite measure ν . Let $f_\theta = dP_\theta/d\nu$. The family \mathcal{P} is said to have monotone likelihood ration in $Y(X)$ (a real-valued statistic) if and only if, for any $\theta_1 < \theta_2$, $f_{\theta_2}(x)/f_{\theta_1}(x)$ is a nondecreasing function of $T(x)$ for values x at which at least one of $f_{\theta_1}(x)$ and $f_{\theta_2}(x)$ is positive.

Example 2: Let θ be real-valued and $\eta(\theta)$ be a nondecreasing function of θ . Then the one-parameter exponential family with $f_\theta(x) = \exp\{\eta(\theta)Y(x) - \xi(\theta)\}h(x)$ has monotone likelihood ration in $Y(X)$.

Theorem 3: Suppose that X has a distribution in $\mathcal{P} = \{P_\theta : \theta \in \Theta\}$ ($\Theta \subset \mathbb{R}$) that has monotone likelihood ratio in $Y(X)$. Consider the problem of testing $H_0 : \theta \leq \theta_0$ versus $H_1 : \theta > \theta_0$, where θ_0 is

a given constant. (i) There exists a UMP test of size α , which is given by $T_*(X) = \begin{cases} 1 & Y(X) > c \\ \gamma & Y(X) = c \\ 0 & Y(X) < c \end{cases}$

where c and γ are determined by $\beta_{T_*}(\theta_0) = \alpha$, and $\beta_T(\theta) = \mathbb{E}[T(X)]$ is the power function of a test T . (ii) $\beta_{T_*}(\theta)$ is strictly increasing for all θ 's for which $0 < \beta_{T_*}(\theta) < 1$. (iii) For any $\theta < \theta_0$, T_* minimizes $\beta_T(\theta)$ among all tests T satisfying $\beta_T(\theta_0) = \alpha$. (iv) Assume that $P_\theta(f_\theta(X) = cf_{\theta_0}(X)) = 0$ for any $\theta > \theta_0$ and $c \geq 0$, where f_θ is the p.d.f. of P_θ . If T is a test with $\beta_T(\theta_0) = \beta_{T_*}(\theta_0)$, then for any $\theta > \theta_0$, either $\beta_T(\theta) < \beta_{T_*}(\theta)$ or $T = T_*$ a.s. P_θ . (v) For any fixed θ_1 , T_* is UMP for testing $H_0 : \theta \leq \theta_1$ versus $H_1 : \theta > \theta_1$, with size $\beta_{T_*}(\theta_1)$.

Theorem 4 (one-parameter exponential families): Suppose that X has a p.d.f. in a one-parameter exponential family with η being a strictly monotone function of θ . If η is increasing, then T_* given by Theorem 3 is UMP for testing $H_0 : \theta \leq \theta_0$ versus $H_1 : \theta > \theta_0$, where γ and c are determined by $\beta_{T_*}(\theta_0) = \alpha$. If η is decreasing or $H_0 : \theta \geq \theta_0$ ($H_1 : \theta < \theta_0$), the result is still valid by reversing inequalities in the definition of T_* .

6.2 UMP tests and unbiased tests

Theorem 1: Suppose that the distribution of X is in a parametric family P indexed by a real-valued θ and that P has monotone likelihood ratio in $Y(X)$. If ψ is a nondecreasing function of Y , then $g(\theta) = \mathbb{E}[\psi(Y)]$ is a nondecreasing function of θ .

Proposition 1 (Generalized Neyman-Pearson lemma): Let f_1, \dots, f_{m+1} be Borel functions on

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\mathbb{R}^p integrable w.r.t. a σ -finite ν . For given constants t_1, \dots, t_m , let \mathcal{T} be the class of Borel functions ϕ (from $\mathbb{R}^p \rightarrow [0, 1]$) satisfying $\int \phi f_i d\nu \leq t_i, i = 1, \dots, m$, and \mathcal{T}_0 be the set of ϕ 's in \mathcal{T} satisfying $\int \phi f_i d\nu = t_i, i = 1, \dots, m$. If there are constants c_1, \dots, c_m such that $\phi_*(x) = \begin{cases} 1 & f_{m+1}(x) > c_1 f_1(x) + \dots + c_m f_m(x) \\ 0 & f_{m+1}(x) < c_1 f_1(x) + \dots + c_m f_m(x) \end{cases}$ is a member of \mathcal{T}_0 , then ϕ_* maximizes $\int \phi f_{m+1} d\nu$ over $\phi \in \mathcal{T}_0$. If $c_i \geq 0$ for all i , then ϕ_* maximizes $\int \phi f_{m+1} d\nu$ over $\phi \in \mathcal{T}$.

Theorem 2: Let f_1, \dots, f_m and ν be given by Proposition 1. Then the set $M = \{(\int \phi f_1 d\nu, \dots, \int \phi f_m d\nu) : \phi \text{ is from } \mathbb{R}^p \rightarrow [0, 1]\}$ is convex and closed. If (t_1, \dots, t_m) is an interior point of M , then there exist constant c_1, \dots, c_m such that the function ϕ_* defined in Proposition 1 is in \mathcal{T}_0 .

Definition 1 (Two-sided hypotheses): The following hypotheses are called two-sided hypotheses: $H_0 : \theta \leq \theta_1$ or $\theta \geq \theta_2$ versus $H_1 : \theta_1 < \theta < \theta_2$, $H_0 : \theta_1 \leq \theta \leq \theta_2$ versus $H_1 : \theta < \theta_1$ or $\theta > \theta_2$, $H_0 : \theta = \theta_0$ versus $H_1 : \theta \neq \theta_0$, where $\theta_0, \theta_1, \theta_2$ are given constants and $\theta_1 < \theta_2$.

Theorem 3 (UMP tests for two-sided hypotheses): Suppose that X has a p.d.f in a one-parameter exponential family, i.e., the p.d.f is $f_\theta(x) = \exp\{\eta(\theta)Y(x) - \xi(\theta)\}h(x)$ w.r.t. a σ -finite measure, where η is a strictly increasing function of θ . (i) For testing hypotheses $H_0 : \theta \leq \theta_1$ or $\theta \geq \theta_2$

versus $H_1 : \theta_1 < \theta < \theta_2$, a UMP test of size α is $T_*(x) = \begin{cases} 1 & c_1 < Y(X) < c_2 \\ \gamma_i & Y(X) = c_i, i = 1, 2, \\ 0 & Y(X) < c_1 \text{ or } Y(X) > c_2 \end{cases}$, where

c_i 's and γ_i 's are determined by $\beta_{T_*}(\theta_1) = \beta_{T_*}(\theta_2) = \alpha$. (ii) T_* minimizes $\beta_T(\theta)$ over all $\theta < \theta_1, \theta > \theta_2$, and T satisfying $\beta_T(\theta_1) = \beta_T(\theta_2) = \alpha$. (iii) If T_* and T_{**} are two tests satisfying $T(x) =$

$\begin{cases} 1 & c_1 < Y(X) < c_2 \\ \gamma_i & Y(X) = c_i, i = 1, 2, \\ 0 & Y(X) < c_1 \text{ or } Y(X) > c_2 \end{cases}$ and $\beta_{T_*}(\theta_1) = \beta_{T_{**}}(\theta_1)$ and if the region $\{T_{**} = 1\}$ is to the right

of $\{T_* = 1\}$, then $\beta_{T_*}(\theta_1) < \beta_{T_{**}}(\theta)$ for $\theta > \theta_1$ and $\beta_{T_*}(\theta) > \beta_{T_{**}}(\theta)$ for $\theta < \theta_1$. If both T_* and T_{**}

satisfy $T(x) = \begin{cases} 1 & c_1 < Y(X) < c_2 \\ \gamma_i & Y(X) = c_i, i = 1, 2, \\ 0 & Y(X) < c_1 \text{ or } Y(X) > c_2 \end{cases}$ and $\beta_{T_*}(\theta_1) = \beta_{T_*}(\theta_2) = \alpha$, then $T_* = T_{**}$ a.s. \mathcal{P} .

Example 1: Let X_1, \dots, X_n be i.i.d. from $\mathcal{N}(\theta, 1)$. By Theorem 3, a UMP test for testing $H_0 : \theta \leq \theta_1$ or $\theta \geq \theta_2$ versus $H_1 : \theta_1 < \theta < \theta_2$ is $T_*(X) = I_{(c_1, c_2)}(\bar{X})$, where c_i 's are determined by $\Phi(\sqrt{n}(c_2 - \theta_1)) - \Phi(\sqrt{n}(c_1 - \theta_1)) = \alpha$ and $\Phi(\sqrt{n}(c_2 - \theta_2)) - \Phi(\sqrt{n}(c_2 - \theta_2)) = \alpha$.

Remark 1 (Nonexistence of UMP tests): Unfortunately, a UMP test does not exist in general for testing hypotheses $H_0 : \theta_1 \leq \theta \leq \theta_2$ versus $H_1 : \theta < \theta_1$ or $\theta > \theta_2$, $H_0 : \theta = \theta_0$ versus $H_1 : \theta \neq \theta_0$.

Definition 2: Let α be a given level of significance. A test T for $H_0 : P \in \mathcal{P}_0$ versus $P \in \mathcal{P}_1$ is said to be unbiased of level α if and only if $\beta_T(P) \leq \alpha, P \in \mathcal{P}_0$ and $\beta_T(P) \geq \alpha, P \in \mathcal{P}_1$. A test of size α is called a uniformly most powerful unbiased (UMPU) test if and only if it is UMP within the class of unbiased tests of level α .

Definition 3 (Similarity): Consider the hypotheses $H_0 : \theta \in \Theta_0$ vs $H_1 : \theta \in \Theta_1$. Let α be a given level of significance and let $\bar{\Theta}_{01}$ be the common boundary of Θ_0 and Θ_1 , i.e., the set of points θ that are points or limit points of both Θ_0 and Θ_1 . A test T is similar on $\bar{\Theta}_{01}$ if and only if $\beta_T(P) = \alpha$ for

all $\theta \in \bar{\Theta}_{01}$.

Remark 2: It is more convenient to work with similarity than to work with unbiasedness for testing $H_0 : \theta \in \Theta_0$ vs $H_1 : \theta \in \Theta_1$.

Theorem 4: Consider hypotheses $H_0 : \theta \in \Theta_0$ vs $H_1 : \theta \in \Theta_1$. Suppose that, for every T , $\beta_T(P)$ is continuous in θ . If T_* is uniformly most powerful among all similar tests and has size α , then T_* is a UMPU test.

6.3 Likelihood ratio and asymptotic tests

Definition 1: Let $l(\theta) = f_\theta(X)$ be the likelihood function. For testing $H_0 : \theta \in \Theta_0$ versus $H_1 : \theta \in \Theta_1$, a likelihood ratio (LR) test is any test that rejects H_0 if and only if $\lambda(X) < c$, where $c \in [0, 1]$ and $\lambda(X)$ is the likelihood ratio defined by $\lambda(X) = \sup_{\theta \in \Theta_0} l(\theta) / \sup_{\theta \in \Theta} l(\theta)$.

Remark 1 (Optimality): When a UMP or UMPU test exists, an LR test is often the same as this optimal test.

Proposition 1: Suppose that X has a p.d.f. in one-parameter exponential family: $f_\theta(x) = \exp\{\eta(\theta)Y(x) - \xi(\theta)\}h(x)$ w.r.t. a σ -finite measure ν , where η is a strictly increasing and differentiable function of θ . (i) For testing $H_0 : \theta \leq \theta_0$ versus $H_1 : \theta > \theta_0$, there is an LR test whose rejection region is the same as that of the UMP test T_* . (ii) For testing $H_0 : \theta \leq \theta_1$ or $\theta \geq \theta_2$ versus $H_1 : \theta_1 < \theta < \theta_2$, there is an LR test whose rejection is the same as that of the UMP test T_* . (iii) For testing the other two-sided hypotheses, there is an LR test whose rejection region is equivalent to $Y(X) < c_1$ or $Y(X) > c_2$ for some constants c_1 and c_2 .

Example 1: Consider the testing problem $H_0 : \theta = \theta_0$ versus $H_1 : \theta \neq \theta_0$ based on i.i.d. X_1, \dots, X_n from the uniform distribution $U(0, \theta)$. We now show that the UMP test with rejection region $X_{(n)} > \theta_0$ or $X_{(n)} \leq \theta_0 \alpha^{1/n}$ is an LR test. Note that $l(\theta) = \theta^{-n} I_{(X_{(n)}, \infty)}(\theta)$. Hence $\lambda(X) = \begin{cases} (X_{(n)}/\theta_0)^n & X_{(n)} \leq \theta_0 \\ 0 & X_{(n)} > \theta_0 \end{cases}$ and $\lambda(X) < c$ is equivalent to $X_{(n)} > \theta_0$ or $X_{(n)}/\theta_0 < c^{1/n}$. Taking $c = \alpha$ ensures that the LR test has size α .

Definition 2 (Asymptotic tests): Let $X = (X_1, \dots, X_n)$ be a sample from $P \in \mathcal{P}$ and $T_n(X)$ be a test for $H_0 : P \in \mathcal{P}_0$ versus $H_1 : P \in \mathcal{P}_1$. (i) If $\limsup_n \alpha_{T_n}(P) \leq \alpha$ for any $P \in \mathcal{P}_0$, then α is an asymptotic significance level of T_n . (ii) If $\lim_{n \rightarrow \infty} \sup_{P \in \mathcal{P}_0} \alpha_{T_n}(P)$ exists, it is called the limiting size of T_n . (iii) T_n is consistent iff the type II error probability converges to 0.

Remark 2: If \mathcal{P}_0 is not a parametric family, the limiting size of T_n may be 1. This is the reason why we consider the weaker requirement in (i).

Definition 3: If $\alpha \in (0, 1)$ is pre-assigned level of significance for the problem, then a consistent test T_n having an asymptotic significance level α is called asymptotically correct, and a consistent test having limiting size α is called strongly asymptotically correct.

Theorem 1 (Asymptotic distribution of likelihood ratio): Assume the conditions of Theorem 1 in section 4.6. Suppose that $H_0 : \theta = g(\vartheta)$, where ϑ is $(k - r)$ -vector of unknown parameters and g is a continuously differentiable function from \mathbb{R}^{k-r} to \mathbb{R}^k with a full rank $\partial g(\vartheta)/\partial \vartheta$. Under H_0 , $-2 \log \lambda_n \rightarrow_d \chi_r^2$, where $\lambda_n = \lambda(X)$ and χ_r^2 is a random variable having the chi-square distribution

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χ_r^2 . Consequently, the LR test with rejection region $\lambda_n < e^{-\chi_{r,\alpha}^2/2}$ has asymptotic significance level α , where $\chi_{r,\alpha}^2$ is the $(1 - \alpha)$ th quantile of the chi-square distribution χ_r^2 .

Theorem 2: Assume the conditions of Theorem 1 in section 4.6. (i) Under $H_0 : R(\theta) = 0$, where $R(\theta)$ is a continuously differentiable function from \mathbb{R}^k to \mathbb{R}^r , $\hat{\theta}$ is an MLE or RLE of θ , $\tilde{\theta}$ is an MLE or RLE of θ under $H_0 : R(\theta) = 0$, $W_n = [R(\hat{\theta})]^T \{[C(\hat{\theta})]^T [I_n(\hat{\theta})]^{-1} C(\hat{\theta})\}^{-1} R(\hat{\theta}) \rightarrow_d \chi_r^2$ and, therefore, the test rejects H_0 if and only if $W_n > \chi_{r,\alpha}^2$ has asymptotic significance level α , where $\chi_{r,\alpha}^2$ is the $(1 - \alpha)$ th quantile of the chi-square distribution χ_r^2 . (ii) The result in (i) still holds if W_n is replaced by $R_n = [s_n(\tilde{\theta})]^T [I_n(\tilde{\theta})]^{-1} s_n(\tilde{\theta})$.

6.4 Asymptotic chi-square tests

Definition 1 (Testing in multinomial distributions): Consider n independent trials with k possible outcomes for each trial. Let $p_j > 0$ be the probability that the j th outcome occurs in a given trial and X_j be the number of occurrences of the j th outcome in n trials. Then $X = (X_1, \dots, X_k)$ has the multinomial distribution with the parameter $p = (p_1, \dots, p_k)$. Let $\xi_i = (0, \dots, 0, 1, 0, \dots, 0)$, where the single nonzero component 1 is located in the j th position if the i th trial yields the j th outcome. Then ξ_1, \dots, ξ_n are i.i.d. and $X/n = \bar{\xi} = \sum_{i=1}^n \xi_i/n$. X/n is an unbiased estimator of p and by the CLT, $Z_n(p) = \sqrt{n}(\frac{X}{n} - p) = \sqrt{n}(\bar{\xi} - p) \rightarrow_d \mathcal{N}_k(0, \Sigma)$, where $\Sigma = \text{Var}(X/\sqrt{n})$ is a symmetric $k \times k$ matrix whose i th diagonal element is $p_i(1 - p_i)$ and (i, j) th off-diagonal element is $-p_i p_j$. We first consider the problem of testing $H_0 : p = p_0$ versus $H_1 : p \neq p_0$ where $p_0 = (p_{01}, \dots, p_{0k})$ is a known vector of cell probabilities.

Definition 2 (χ^2 tests): For testing $H_0 : p = p_0$ vs $H_1 : p \neq p_0$, a class of tests related to the asymptotic tests, a popular test is based on the following χ^2 -statistic: $\chi^2 = \sum_{j=1}^k \frac{(X_j - np_{0j})^2}{np_{0j}} = \|D(p_0)Z_n(p_0)\|^2$, where $D(c)$ with $c = (c_1, \dots, c_k)$ is the $k \times k$ diagonal matrix whose j th diagonal element is $c_j^{-1/2}$. Another popular test is based on the following modified χ^2 -statistic: $\tilde{\chi}^2 = \sum_{j=1}^k \frac{(X_j - np_{0j})^2}{X_j} = \|D(X/n)Z_n(p_0)\|^2$.

Theorem 1: Let $\phi = (\sqrt{p_1}, \dots, \sqrt{p_k})$ and Λ be a $k \times k$ projection matrix. (i) If $\Lambda\phi = a\phi$, then $[Z_n(p)]^T D(p)\Lambda D(p)Z_n(p) \rightarrow_d \chi_r^2$, where χ_r^2 has the chi-square distribution χ_r^2 with $r = \text{tr}(\Lambda) - a$. (ii) The same result holds if $D(p)$ in (i) is replaced by $D(X/n)$.

Example 1 (Goodness of fit tests): Let Y_1, \dots, Y_n be i.i.d. from F . Consider the problem of testing $H_0 : F = F_0$ versus $H_1 : F \neq F_0$, where F_0 is a known c.d.f. One way to test $H_0 : F = F_0$ is to partition the range of Y_1 into k disjoint events A_1, \dots, A_k and test $H_0 : p = p_0$ with $p_j = P_F(A_j)$ and $p_{0j} = P_{F_0}(A_j)$, $j = 1, \dots, k$. Let X_j be the number of Y_i 's in A_j , $j = 1, \dots, k$. Based on X_j 's, the χ^2 -tests discussed previously can be applied.

Definition 3 (Generalized χ^2 -statistics): The generalized χ^2 -statistics χ^2 and $\tilde{\chi}^2$ are defined to be the previously defined χ^2 -statistics with p_{0j} 's replaced by $p_j(\hat{\theta})$'s, where $\theta \in \Theta \subset \mathbb{R}^s$ ($s < k$) and $\hat{\theta}$ is an MLE of θ under H_0 .

Theorem 2: Under $H_0 : p = p(\theta)$, the generalized χ^2 -statistics converge in distribution to χ_{k-s-1}^2 .

Definition 4 (Testing independence): Testing independence of $\{A_j : j = 1, \dots, c\}$ and $\{B_i, i = 1, \dots, r\}$ is equivalent to testing hypotheses $H_0 : p_{ij} = p_i p_j$ for all i, j versus $H_1 : p_{ij} \neq p_i p_j$ for some i, j , where $p_{ij} = P(A_j \cap B_i) = \mathbb{E}[X_{ij}]/n$, $p_i = P(B_i)$ and $p_j = P(A_j)$, $i = 1, \dots, r$, $j = 1, \dots, c$. Under

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H_0 , MLE's of $p_{i\cdot}$ and $p_{\cdot j}$ are $\bar{X}_{i\cdot} = n_{i\cdot}/n$ and $\bar{X}_{\cdot j}m_j/n$, respectively, $i = 1, \dots, r, j = 1, \dots, c$. The number of free parameters is $rc-1$. Under H_0 , the number of free parameters is $r-1+c-1 = r+c-2$. The difference of the two is $(r-1)(c-1)$. Then the χ^2 -test rejects H_0 when $\chi^2 > \chi_{(r-1)(c-1), \alpha}^2$ where $\chi^2 = \sum_{i=1}^r \sum_{j=1}^c \frac{(X_{ij} - n\bar{X}_{i\cdot}\bar{X}_{\cdot j})^2}{n\bar{X}_{i\cdot}\bar{X}_{\cdot j}}$ and $\chi_{(r-1)(c-1), \alpha}^2$ is the $(1-\alpha)$ th quantile of the chi-square distribution of $\chi_{(r-1)(c-1)}^2$.

Remark 1 (Construction of asymptotic tests): A simple method of constructing asymptotic tests (for almost all problems, parametric or nonparametric) for testing $H_0 : \theta = \theta_0$ versus $H_1 : \theta \neq \theta_0$, where θ is a vector of parameters, is to use an asymptotically normally distributed estimator θ . Let $\hat{\theta}_n$ be an estimator of θ based on a sample size n from P . Suppose that under H_0 , $V_n^{-1/2}(\hat{\theta}_n - \theta) \rightarrow_d \mathcal{N}_k(0, I_k)$, where V_n is the asymptotic covariance matrix of $\hat{\theta}_n$. If V_n is known when $\theta = \theta_0$, then we define a test with rejection region $(\hat{\theta}_n - \theta_0)^T V_n^{-1}(\hat{\theta}_n - \theta_0) > \chi_{k, \alpha}^2$, where $\chi_{k, \alpha}^2$ is the $(1-\alpha)$ th quantile of the chi-squared distribution χ_k^2 . If V_n depends on the unknown population P even if H_0 is true ($\theta = \theta_0$), then we have to replace V_n by an estimator \hat{V}_n . If \hat{V}_n is consistent, then the resulting test still has asymptotic significance level α .

Theorem 3: Assume that $V_n^{-1/2}(\hat{\theta}_n - \theta) \rightarrow_d \mathcal{N}_k(0, I_k)$, holds for any P . Assume also that $\lambda_+[V_n] \rightarrow 0$, where $\lambda_+[V_n]$ is the largest eigenvalue of V_n . (i) The test having rejection region $(\hat{\theta}_n - \theta_0)^T V_n^{-1}(\hat{\theta}_n - \theta_0) > \chi_{k, \alpha}^2$ with a known V_n (or with V_n replaced by a consistent estimator \hat{V}_n) is consistent. (ii) If we choose $\alpha = \alpha_n \rightarrow 0$ as $n \rightarrow \infty$ and $\chi_{k, 1-\alpha_n}^2 \lambda_+[V_n] = o(1)$, then the test in (i) is Chernoff-consistent.

7 Confidence Sets

7.1 Pivotal quantities and confidence sets

Definition 1 (Confidence sets): X : a sample from a population $P \in \mathcal{P}$. $\theta = \theta(P)$: a functional from \mathcal{P} to $\Theta \subset \mathbb{R}^k$ for a fixed integer k . $C(X)$: a confidence set for θ , a set in \mathcal{B}_Θ depending only on X . The confidence coefficient of $C(X) : \inf_{P \in \mathcal{P}} P(\theta \in C(X))$. If the confidence coefficient of $C(X)$ is $\geq 1 - \alpha$ for fixed $\alpha \in (0, 1)$, then we say that $C(X)$ has confidence level $1 - \alpha$ or $C(X)$ is a level $1 - \alpha$ confidence set.

Definition 2: A known Borel function h of (X, θ) is called a pivotal quantity if and only if the distribution of $h(X, \theta)$ does not depend on P .

Construction: First, find two constants c_1 and c_2 such that $P(c_1 \leq h(x, \theta) \leq c_2) \geq 1 - \alpha$. Next, define $C(X) = \{\theta \in \Theta : c_1 \leq h(X, \theta) \leq c_2\}$. Then $C(X)$ is a level $1 - \alpha$ confidence set. The confidence coefficient of $C(X)$ may not be $1 - \alpha$. If $h(X, \theta)$ has a continuous c.d.f., then we can choose c_i 's such that the equality in the last expression holds and the confidence set $C(X)$ has confidence coefficient $1 - \alpha$.

Computation: When $h(X, \theta)$ and c_i 's are chosen, we need to compute the confidence set $C(X) = \{c_1 \leq h(X, \theta) \leq c_2\}$. This can be done by inverting $c_1 \leq h(X, \theta) \leq c_2$. For example, if θ is real-valued and $h(X, \theta)$ is monotone in θ when X is fixed, then $C(X) = \{\theta : \underline{\theta}(X) \leq \theta \leq \bar{\theta}(X)\}$ for some $\underline{\theta}(X) < \bar{\theta}(X)$, i.e., $C(X)$ is an interval (finite or infinite). If $h(X, \theta)$ is not monotone, then $C(X)$ may be a union of several intervals.

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Example 1 (Location-scale families): Suppose that X_1, \dots, X_n are i.i.d. with a Lebesgue p.d.f. $\frac{1}{\sigma}f(\frac{x-\mu}{\sigma})$, where $\mu \in \mathbb{R}, \sigma > 0$, and f is a known Lebesgue p.d.f. Consider first the case where σ is known and $\theta = \mu$. $\bar{X} - \mu$ is a pivotal quantity. Let c_1 and c_2 be constants such that $P(c_1 \leq \bar{X} - \mu \leq c_2) = 1 - \alpha$, then $C(X) = \{\mu : \bar{X} - c_2 \leq \mu \leq \bar{X} - c_1\}$. Consider next the case where μ is known and $\theta = \sigma$. S/σ is a pivotal quantity, where S^2 is the sample variance. Let c_1 and c_2 be chosen such that $P(c_1 \leq S/\sigma \leq c_2) = 1 - \alpha$. If both c_i 's are positive, then $C(X) = \{\sigma : S/c_2 \leq \sigma S/c_1\} = [S/c_2, S/c_1]$ is a finite interval. Similarly, if $c_1 = 0$ or $c_2 = \infty$, then $C(X) = [S/c_2, \infty)$ or $(0, S/c_1]$. When $\theta = \sigma$ and μ is also known, S/σ is still a pivotal quantity and, hence, confidence intervals of σ based on S are still valid. Finally, we consider the case where both μ and σ are unknown and $\theta = \mu$. There are still many different pivotal quantities, but the most commonly used pivotal quantity is $t(X) = \sqrt{n}(\bar{X} - \mu)/S$. The distribution of $t(X)$ does not depend on (μ, σ) . When f is normal, $t(X)$ has the t -distribution t_{n-1} . A confidence interval for μ based on $t(X)$ is of the form $\{\mu : c_1 \leq \sqrt{n}(\bar{X} - \mu)/S \leq c_2\} = [\bar{X} - c_2 S/\sqrt{n}, \bar{X} - c_1 S/\sqrt{n}]$, where c_i 's are chosen so that $P(c_1 \leq t(X) \leq c_2) = 1 - \alpha$.

Example 2: Let X_1, \dots, X_n be i.i.d. random variables from the uniform distribution $U(0, \theta)$. Consider the problem of finding a confidence set for θ . Note that $X_{(n)}/\theta$ has the Lebesgue p.d.f. $nx^{n-1}I_{(0,1)}(x)$. Hence c_i 's should satisfy $c_2^n - c_1^n = 1 - \alpha$. The resulting confidence interval for θ is $[c_2^{-1}X_{(n)}, c_1^{-1}X_{(n)}]$.

Proposition 1 (Existence of pivotal quantities in parametric problems): Let $T(X) = (T_1(X), \dots, T_s(X))$ and T_1, \dots, T_s be independent statistics. Suppose that each T_i has a continuous c.d.f. $F_{T_i, \theta}$ indexed by θ . Then $h(X, \theta) = \prod_{i=1}^s F_{T_i, \theta}(T_i(X))$ is a pivotal quantity.

Theorem 1: Suppose that P is in a parametric family indexed by a real-valued θ . Let $T(X)$ be a real-valued statistic with c.d.f. $F_{T, \theta}(t)$ and let α_1 and α_2 be fixed positive constants such that $\alpha_1 + \alpha_2 = \alpha < \frac{1}{2}$. (i) Suppose that $F_{T, \theta}(t)$ and $F_{T, \theta}(t-)$ are nonincreasing in θ for each fixed t . Define $\bar{\theta} = \sup\{\theta : F_{T, \theta}(T) \geq \alpha_1\}$ and $\underline{\theta} = \inf\{\theta : F_{T, \theta}(T-) \leq 1 - \alpha_2\}$. Then $[\underline{\theta}(T), \bar{\theta}(T)]$ is a level $1 - \alpha$ confidence interval for θ . (ii) If $F_{T, \theta}(t)$ and $F_{T, \theta}(t-)$ are nondecreasing in θ for each t , then the same result holds with $\underline{\theta} = \inf\{\theta : F_{T, \theta}(T) \geq \alpha_1\}$ and $\bar{\theta} = \sup\{\theta : F_{T, \theta}(T-) \leq 1 - \alpha_2\}$. (iii) If $F_{T, \theta}$ is a continuous c.d.f. for any θ , then $F_{T, \theta}(T)$ is a pivotal quantity and the confidence interval in (i) or (ii) has confidence coefficient $1 - \alpha$.

Definition 3 (Asymptotic criterion): In some problems, especially in nonparametric problems, it is difficult to find a reasonable confidence set with a given confidence coefficient or confidence level $1 - \alpha$. A common approach is to find a confidence set whose confidence coefficient or confidence level is nearly $1 - \alpha$ when the sample size n is large. A confidence set $C(X)$ for θ has asymptotic confidence level $1 - \alpha$ if $\liminf_n P(\theta \in C(X)) \geq 1 - \alpha$ for any $P \in \mathcal{P}$. If $\lim_{n \rightarrow \infty} P(\theta \in C(X)) = 1 - \alpha$ for any $P \in \mathcal{P}$, then $C(X)$ is a $1 - \alpha$ asymptotically correct confidence set.

Definition 4 (Asymptotically pivotal quantities): A known Borel function of (X, θ) , $h_n(X, \theta)$, is said to be asymptotically pivotal iff the limiting distribution of $h_n(X, \theta)$ does not depend on P . Like a pivotal quantity in constructing confidence sets with a given confidence coefficient or confidence level, an asymptotically pivotal quantity can be used in constructing asymptotically correct confidence sets.

Example 3 (Functions of means): Suppose that X_1, \dots, X_n are i.i.d. random vectors having

a c.d.f. F on \mathbb{R}^d and that the unknown parameter of interest $\theta = g(\mu)$, where $\mu = \mathbb{E}(X_1)$ and g is a known differentiable function from \mathbb{R}^d to \mathbb{R}^k , $k \leq d$. From the CLT, $\hat{\theta}_n = g(\bar{X})$ satisfies $V_n^{-1/2}(\hat{\theta}_n - \theta) \rightarrow_d \mathcal{N}_k(0, I_k)$, $V_n = [\nabla g(\mu)]^T \text{Var}(X_1) \nabla g(\mu)/n$. A consistent estimator of the asymptotic covariance matrix V_n is $\hat{V}_n = [\nabla g(\bar{X})]^T S^2 \nabla g(\bar{X})/n$. Thus, $C(X) = \{\theta : \|\hat{V}_n^{-1/2}(\hat{\theta}_n - \theta)\|^2 \leq \chi_{k,\alpha}^2\}$, is a $1 - \alpha$ asymptotically correct confidence set for θ .

7.2 Inverting acceptance regions of tests, UMA and UMAU confidence sets

Remark 1 (Confidence sets and hypothesis tests): Another popular method of constructing confidence sets is to use a close relationship between confidence sets and hypothesis tests. For any test T , the set $\{x : T(x) \neq 1\}$ is called the acceptance region. This terminology is not precise when T is a randomized test.

Theorem 1: For each $\theta_0 \in \Theta$, let T_{θ_0} be a test for $H_0 : \theta = \theta_0$ with significance level α and acceptance region $A(\theta_0)$. For each x in the range of X , define $C(x) = \{\theta : x \in A(\theta)\}$. Then $C(X)$ is a level $1 - \alpha$ confidence set for θ . If T_{θ_0} is nonrandomized and has size α for every θ_0 , then $C(X)$ has confidence coefficient $1 - \alpha$.

Proposition 1: Let $C(X)$ be a confidence set for θ with confidence level $1 - \alpha$. For any $\theta_0 \in \Theta$, define a region $A(\theta_0) = \{x : \theta_0 \in C(x)\}$. Then the test $T(X) = 1 - I_{A(\theta_0)}(X)$ has significance level α for testing $H_0 : \theta = \theta_0$ versus some H_1 .

Example 1: Suppose that X has the following p.d.f. in a one-parameter exponential family: $f_\theta(x) = \exp\{\eta(\theta)Y(x) - \xi(\theta)\}h(x)$, where θ is real-valued and $\eta(\theta)$ is nondecreasing in θ . Consider $H_0 : \theta = \theta_0$ and $H_1 : \theta > \theta_0$. The acceptance region of the UMP test of size α is $A(\theta_0) = \{x : Y(x) \leq c(\theta_0)\}$. It can be shown that $c(\theta)$ is nondecreasing in θ . Inverting $A(\theta)$, we obtain $C(X) = [\underline{\theta}(X), \infty)$ or $(\underline{\theta}(X), \infty)$, a one-sided confidence interval for θ with confidence level $1 - \alpha$. $\underline{\theta}(X)$ is called a lower confidence bound for θ . If $H_0 : \theta = \theta_0$ and $H_1 : \theta < \theta_0$ are considered, then $C(X) = \{\theta : Y(X) \geq c(\theta)\}$ and is of the form $(-\infty, \bar{\theta}(X)]$ or $(-\infty, \bar{\theta}(X))$. $\bar{\theta}(X)$ is called an upper confidence bound for θ .

Remark 2 (Confidence sets related to optimal tests): For a confidence set obtained by inverting the acceptance regions of some UMP or UMPU tests, it is expected that the confidence set inherits some optimality property.

Definition 1: Let $\theta \in \Theta$ be an unknown parameter and Θ' be a subset of Θ that does not contain the true parameter value θ . A confidence set $C(X)$ for θ with confidence coefficient $1 - \alpha$ is said to be Θ' -uniformly most accurate iff for any other confidence set $C_1(X)$ with confidence level $1 - \alpha$, $P(\theta' \in C(X)) \leq P(\theta' \in C_1(X))$ for all $\theta' \in \Theta'$. $C(X)$ is UMA iff it is Θ' -UMA with $\Theta' = \{\theta\}^c$.

Theorem 2: Let $C(X)$ be a confidence set for θ obtained by inverting the acceptance regions of nonrandomized tests T_{θ_0} for testing $H_0 : \theta = \theta_0$ versus $H_1 : \theta \in \Theta_{\theta_0}$. Suppose that for each θ_0 , T_{θ_0} is UMP of size α . Then $C(X)$ is Θ' -UMA with confidence coefficient $1 - \alpha$, where $\Theta' = \{\theta' : \theta \in \Theta_{\theta'}\}$.

Definition 2: Let $\theta \in \Theta$ be an unknown parameter, Θ' be a subset of Θ that does not contain the true parameter value θ , and $1 - \alpha$ be a given confidence level. (i) A level $1 - \alpha$ confidence set $C(X)$ is said to be Θ' -unbiased (unbiased when $\Theta' = \{\theta\}^c$) iff $P(\theta' \in C(X)) \leq 1 - \alpha$ for all $\theta' \in \Theta'$. (ii) Let $C(X)$ be a Θ' -unbiased confidence set with confidence coefficient $1 - \alpha$. If $P(\theta' \in C(X)) \leq P(\theta' \in C_1(X))$ for all $\theta' \in \Theta'$ holds for any other Θ' -unbiased confidence set $C_1(X)$

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with confidence level $1 - \alpha$, then $C(X)$ is Θ' -uniformly most accurate unbiased (UMAU). $C(X)$ is UMAU if and only if it is Θ' -UMAU with $\Theta' = \{\theta\}^c$.

Theorem 3: Let $C(X)$ be a confidence set for θ obtained by inverting the acceptance regions of nonrandomized tests T_{θ_0} for testing $H_0 : \theta = \theta_0$ versus $H_1 : \theta \in \Theta_{\theta_0}$. If T_{θ_0} is unbiased of size α for each θ_0 , then $C(X)$ is Θ' -unbiased with confidence coefficient $1 - \alpha$, where $\Theta' = \{\theta' : \theta \in \Theta_{\theta'}\}$. If T_{θ_0} is also UMPU for each θ_0 , then $C(X)$ is Θ' -UMAU.