Analysis

[Matrix] $c^Tc = ||c||^2 = c_1^2 + \cdots + c_k^2$, cc^T is $k \times k$ matrix with (i,j)th element as c_ic_j,

[Max, Min] $\max(a,b) = \frac{1}{2}(a+b+|a-b|)$, $\min(a,b) = \frac{1}{2}(a+b-|a-b|)$

Probability

[Moments] $\mu^k = E(X^k) = \int x^k f(x)dx$

[Deduce $X = 0$] If $X \geq 0$ a.s. and $EX = 0$ then $X = 0$ a.s.

[Variance, Covariance] $Var(X) = E[(X-EX)(X-EX)^T]$, $Cov(X,Y) = E[(X-EX)(Y-EY)^T]$, $Corr(X,Y) = Cov(X,Y)/(\sigma_X\sigma_Y)$, $E(a^TX) = a^TEX$, $Var(a^TX) = a^TVar(X)a$

[CHF] $\phi_X(t) = E\left[exp(\sqrt{-1}t^TX\right] = E\left[\cos(t^TX) + \sqrt{-1}\sin(t^TX)\right]\ \forall\ t \in \mathcal{R}^d$, well defined with $|\phi_X| \leq 1$

[MGF] $\psi_X(t) = E\left[exp(t^TX)\right]\ \forall\ t \in \mathcal{R}^d$,

[MGF properties] $\psi_{-X}(t) = \psi_X(-t)$, if $\psi(t) < \infty\ \forall\ ||t|| < \delta \Rightarrow E|X|^a < \infty\ \forall\ a > 1$ and $\phi_X(t) = \psi_X(\sqrt{-1}t)$

[Conditional Exp] $f_{X|Y}(x|Y=y) = \frac{f_{X,Y}(x,y)}{f_Y(y)} = \frac{f_{Y|X}(y|X=x)f_X(x)}{f_Y(y)}$

Integration

[MCT] $0 \leq f_1 \leq f_2 \leq \cdots \leq f_n$ and $\lim_n f_n = f$ a.e. $\Rightarrow \int \lim_n f_n d\nu = \lim_n \int f_n d\nu$

[Fatou] $f_n \geq 0 \Rightarrow \int \liminf_n f_n d\nu \leq \liminf_n \int f_n d\nu$

[DCT] $\lim_{n\to\infty} f_n = f$ and $|f_n| \leq g$ a.e. $\Rightarrow \int \lim_n f_n d\nu = \lim_n \int f_n d\nu$. g is an integrable function.

[Interchange Diff and Int] ① $\partial f(\omega,\theta)/\partial\theta$ exists in (a,b) ② $|\partial f(\omega,\theta)/\partial\theta| \leq g(\omega)$ a.e. $\Rightarrow$
① $\partial f(\omega,\theta)/\partial\theta$ integrable in (a,b) ② $\frac{d}{d\theta}\int f(\omega,\theta)d\nu(\omega) = \int \frac{\partial f(\omega,\theta)}{\partial\theta}d\nu(\omega)$

[Change of Var] $Y = g(X), X = g^{-1}(Y) = h(Y)$ and A_i disjoint, $f_Y(y) = \sum_{j:1\leq j\leq m, y\in g(A_j)}\left|\det\left(\frac{\partial h_j(y)}{\partial y}\right)\right| f_X(h_j(y))$. Simple version: $f_Y(y) = |det(\partial h(y)/\partial y)|f_X(h(y))$

Inequalities

[Cauchy-Schewarz] $Cov(X,Y)^2 \leq Var(X)Var(Y)$, and $E^2[XY] \leq EX^2EY^2$

[Jensen] φ is convex $\Rightarrow \varphi(EX) \leq E\varphi(X)$ e.g. $(EX)^{-1} < E(X^{-1})$ and $E(logX) < log(EX)$

[Chebyshev] If $\varphi(-x) = \varphi(x)$, and φ non-decreasing on $[0,\infty) \Rightarrow \varphi(t)P(|X| \geq t) \leq \int_{\{|X|\geq t\}}\varphi(X)dP \leq E\varphi(X) \forall\ t \geq 0$. e.g. $P(|X-\mu| \geq t) \leq \frac{\sigma_X^2}{t^2}$ and $P(|X| \geq t) \leq \frac{E|X|}{t}$

[Hölder] $p,q > 0$ and $1/p + 1/q = 1$ or $q = p/(p-1) \Rightarrow E|XY| \leq (E|X|^p)^{1/p}(E|Y|^q)^{1/q}$. Equality $\Leftrightarrow |X|^p$ and $|Y|^q$ linearly dependent

[Young] $ab \leq \frac{a^p}{p} + \frac{b^q}{q}$, equality $\Leftrightarrow a^p = b^q$

[Minkowski] $p \geq 1$, $(E|X+Y|^p)^{1/p} \leq (E|X|^p)^{1/p} + (E|Y|^p)^{1/p}$

[Lyapunov] for $0 < s < t$, $(E|X|^s)^{1/2} \leq (E|X|^t)^{1/t}$

[KL] $K(f_0,f_1) = E_0 \log\frac{f_0(X)}{f_1(X)} = \int \log\left(\frac{f_0(x)}{f_1(x)}\right)f_0(x)d\nu(x) \geq 0$ equality $\Leftrightarrow f_1(\omega) = f_0(\omega)$

Convergence

[a.s] $X_n \xrightarrow{\text{a.s.}} X$ if $P(\lim_{n\to\infty} X_n = X) = 1$. Can show $\forall\ \epsilon > 0, \sum_{i=1}^\infty P(|X_n - X| > \epsilon) < \infty$ via BC lemma

[Infinity often] $\{A_n\ i.o.\} = \cap_{n\geq 1}\cup_{j\geq n} A_j := \limsup_{n\to\infty} A_n$

[Borel-Cantelli lemmas]

(First BC) If $\sum_{n=1}^\infty P(A_n) < \infty$, then $P(A_n\ i.o.) = 0$

(Second BC) Given pairwisely independent events $\{A_n\}_{n=1}^\infty$, if $\sum_{n=1}^\infty P(A_n) = \infty$, then $P(A_n\ i.o.) = 1$

[L^p] $X_n \xrightarrow{L_p} X$ if $\lim_{n\to\infty} E|X_n - X|^p = 0$, given $p > 0$, $E|X|^p < \infty$ and $E|X_n|^p < \infty$

[Probability] $X_n \xrightarrow{P} X$ if $\forall\ \epsilon > 0 \lim_{n\to\infty} P(|X_n - X| > \epsilon) = 0$. Can show $E(X_n) = X$, $\lim_{n\to\infty} Var(X_n) = 0$

[Distribution] $X_n \xrightarrow{D} X$ if $\lim_{n\to} F_n(x) = F(x)$ for every $x \in \mathcal{R}$ at which F is continuous

[Relationships between convergence]

① $L^p \Rightarrow L^q \Rightarrow P$ ② $a.s. \Rightarrow P$, $P \Rightarrow D$ ③ $X_n \to_D C \Rightarrow X_n \to_P C$ ④ If $X_n \to_P X \Rightarrow \exists$ sub-seq s.t. $X_{n_j} \to_{\text{a.s.}} X$.

[Continuous mapping] If $g: \mathcal{R}^k \to \mathcal{R}$ is continuous and $X_n \xrightarrow{*} X$, then $g(X_n) \xrightarrow{*} g(X)$, where $*$ is either ⓐ a.s. ⓑ P ⓒ D.

[Convengence properties]

① Unique in limit: $X = Y$ if $X_n \to X$ and $X_n \to Y$ for ⓐ a.s., ⓑ P, ⓒ L^p. ⓓ If $F_n \to F$ and $F_n \to G$, then $F(t) = G(t)\ \forall\ t$

② Concatenation: $(X_n, Y_n) \to (X, Y)$ when ⓐ P ⓑ a.s. ⓒ $(X_n, Y_n) \xrightarrow{D} (X, c)$ only when c is constant.

③ Linearity: $(aX_n + bY_n) \to aX + bY$ when ⓐ a.s. ⓑ P ⓒ L^p ⓓ NOT for distribution.

④ Cramér-Wold device: for k-random vectors, $X_n \xrightarrow{D} X \Leftrightarrow c^TX_n \xrightarrow{D} c^TX$ for every $c \in \mathcal{R}^k$

[Lévy continuity] $X_n \xrightarrow{D} X \Leftrightarrow \phi_{X_n} \to \phi_X$ pointwise

[Scheffés theorem] If $\lim_{n\to\infty} f_n(x) = f(x) \Rightarrow \lim_{n\to\infty}\int|f_n(x) - f(x)|d\nu = 0$ and $P_{f_n} \to P_f$. Useful to check pdf converge in distribution.

[Slutsky's theorem] If $X_n \xrightarrow{D} X$ and $Y_n \xrightarrow{D} c$ for constant c. Then $X_n + Y_n \xrightarrow{D} X + c$, $X_nY_n \xrightarrow{D} cX$, $X_n/Y_n \xrightarrow{D} X/c$ if $c \neq 0$

[Skorohod's theorem] If $X_n \xrightarrow{D} X$, then $\exists\ Y, Y_1, Y_2, \cdots$ s.t. $P_{Y_n} = P_{X_n}$, $P_Y = P_X$ and $Y_n \xrightarrow{\text{a.s.}} Y$

[δ-method - first order] If $\{a_n\} > 0$ and $\lim_{n\to\infty} a_n = \infty$ and $a_n(X_n - c) \xrightarrow{D} Y$ and $c \in \mathcal{R}$ and $g'(c)$ exists at c, then $a_n[g(X_n) - g(c)] \xrightarrow{D} g'(c)Y$

[δ-method - higher order] If $g^{(j)}(c) = 0$ for all $1 \leq j \leq m-1$ and $g^{(m)}(c) \neq 0$. Then $a_n^m[g(X_n) - g(c)] \xrightarrow{D} \frac{1}{m!}g^{(m)}(c)Y^m$

[δ-method - multivariate] If X_i, Y are k-vectors rvs and $c \in \mathcal{R}^k$ and $a_n[g(X_n) - g(c)] \xrightarrow{D} \nabla g(c)^TY$

[Stochastic order - Real] for a constant $c > 0$ and all n, ① $a_n = O(b_n) \Leftrightarrow |a_n| \leq c|b_n|$ ② $a_n = o(b_n) \Leftrightarrow \lim_{n\to\infty} a_n/b_n = 0$

[Stochastic order - RV] ① $X_n = O_{\text{a.s.}}(Y_n) \Leftrightarrow P\{|X_n| = O(|Y_n|)\} = 1$ ② $X_n = o_{\text{a.s.}}(Y_n) \Leftrightarrow X_n/Y_n \xrightarrow{\text{a.s.}} 0$, ③ $\forall\ \epsilon > 0, \exists C_\epsilon > 0, n_\epsilon \in \mathcal{N} s.t.$ $X_n = O_P(Y_n) \Leftrightarrow \sup_{n\geq n_\epsilon} P(\{\omega \in \Omega : |X_n(\omega)| \geq C_\epsilon|Y_n(\omega)|\}) < \epsilon$ ④ If $X_n = O_P(1)$, $\{X_n\}$ is bounded in probability. ⑤ $X_n = o_P(Y_n) \Leftrightarrow X_n/Y_n \xrightarrow{P} 0$

[Stochastic Order Properties] ① If $X_n \xrightarrow{\text{a.s.}} X$, then $\{\sup_{n\geq k}|X_n|\}_k$ is $O_p(1)$. ② If $X_n \xrightarrow{D} X$ for a rvs, then $X_n = O_P(1)$ (tightness). ③ If $E|X_n| = O(a_n)$, then $X_n = O_P(a_n)$ ④ If $E|X_n| = o(a_n)$, then $X_n = o_P(a_n)$

[SLLN, iid] $E|X_1| < \infty \Leftrightarrow \frac{1}{n}\sum_{i=1}^n X_1 \xrightarrow{\text{a.s.}} EX_1$

[SLLN, non-identical but independent] If $\exists\, p \in [1,2]$ s.t. $\sum_{i=1}^{\infty} \frac{E|X_i|^p}{i^p} < \infty$, then $\frac{1}{n}\sum_{i=1}^{n}(X_i - EX_i) \xrightarrow{\text{a.s.}} 0$
[USLLN, idd] Suppose ① $U(x,\theta)$ is continuous in θ for any fixed x ② for each θ, $\mu(\theta) = EU(X,\theta)$ is finite ③ Θ is compact ④ There exists function $M(x)$ s.t. $EM(X) < \infty$ and $|U(x,\theta) \leq M(x)|$ for all x, θ. Then $P\left\{\lim_{n\to\infty}\sup_{\theta\in\Theta}\left|\frac{1}{n}\sum_{i=1}^{n}U(X_j,\theta) - \mu(\theta)\right| = 0\right\} = 1$
[WLLN, iid] $a_n = E(X_1 I_{\{|X_1|\leq n\}}) \in [-n, n]$ $nP(|X_1| > n) \to 0 \Leftrightarrow \frac{1}{n}\sum_{i=1}^{n} X_i - a_n \xrightarrow{P} 0$
[WLLN, non-identical but independent] If $\exists\, p \in [1,2]$ s.t. $\lim_{n\to\infty}\frac{1}{n^p}\sum_{i=1}^{n} E|X_i|^p = 0$, then $\frac{1}{n}\sum_{i=1}^{n}(X_i - EX_i) \xrightarrow{P} 0$
[Weak Convergency] $\int f d\nu_n \to \int f d\nu$ for every bounded and continous real function f. $X_n \xrightarrow{D} X \Leftrightarrow E[h(X_n)] \to E[h(X)]$
[CLT, iid] Suppose $\Sigma = VarX_1 < \infty$, then $\frac{1}{\sqrt{n}}\sum_{i=1}^{n}(X_i - EX_i) \xrightarrow{D} N(0,\Sigma)$
[CLT, non-identical but independent] Suppose ① $k_n \to \infty$ as $n \to \infty$ ② (Lindeberg's condition) $0 < \sigma_n^2 = Var\left(\sum_{j=1}^{k_n} X_{nj}\right) < \infty$. ③ If for any $\epsilon > 0$, $\frac{1}{\sigma_n^2}\sum_{j=1}^{k_n} E\left\{(X_{nj} - EX_{nj})^2 I_{\{|X_{nj}-EX_{nj}|>\epsilon\sigma_n\}}\right\} \to 0$. Then $\frac{1}{\sigma_n}\sum_{j=1}^{k_n}(X_{nj} - EX_{nj}) \xrightarrow{D} N(0,1)$
[Check Lindeberg condition] Option ① (Lyapunov condition) $\frac{1}{\sigma_n^{2+\delta}}\sum_{j=1}^{k_n} E|X_{nj} - EX_{nj}|^{2+\delta} \to 0$ for some $\delta > 0$
Option ② (Uniform boundedness) If $|X_{nj}| \leq M$ for all n and j and $\sigma_n^2 = \sum_{j=1}^{k_n} Var(X_{nj}) \to \infty$
[Feller's condition] Ensures Lindeberg's condition is sufficient and necessary (else only sufficient). $\lim_{n\to\infty}\max_{j\leq k_n}\frac{Var(X_{nj})}{\sigma_n^2} = 0$

Exponential Families

[NEF] $f_\eta(X) = \exp\left\{\eta^T T(X) - \mathcal{C}(\eta)\right\} h(x)$, where $\eta = \eta(\theta)$ and $\mathcal{C}(\eta) = \log\left\{\int_\Omega \exp\left\{\eta^T T(X)\right\} h(X) dX\right\}$. NEF is full rank if Ξ contains open set in $\mathcal{R}^p$, $\Xi = \{\eta(\theta) : \theta \in \Theta\} \subset \mathcal{R}^p$. Suppose $X_i \sim f_i$ independently with f_i Exp Fam, then joint distribution X is also Exp Fam.
[Showing non Exp Fam] For an exp fam P_θ, there is nonzero measure λ s.t. $\frac{dP_\theta}{d\lambda}(\omega) > 0$ λ-a.e. and for all θ. Consider $f = \frac{dP_\theta}{d\lambda} I_{(t,\infty)}(x)$, $\int f d\lambda = 0, f \geq 0 \Rightarrow f = 0$. Since $\frac{dP_\theta}{d\lambda} > 0$ by assumption, then $I_{(t,\infty)}(x) = 0 \Rightarrow v([t,\infty)) = 0$. Since t is arbitary, consider $v(\mathcal{R}) = 0$ (contradiction)
[NEF MGF] Suppose η_0 is interior point on Ξ, then $\psi_{\eta_0}(t) = \exp\left\{\mathcal{C}(\eta_0 + t) - \mathcal{C}(\eta_0)\right\}$ and is finite in neighborhood of $t = 0$.
[NEF Moments] Let $A(\theta) = \mathcal{C}(\eta_0(\theta))$, $\frac{dA(\theta)}{d\theta} = \frac{d\mathcal{C}(\eta_0(\theta))}{d\eta_0(\theta)} \cdot \frac{d\eta_0(\theta)}{d\theta}$, $T(x) = \frac{\partial \mathcal{C}(\eta)}{\partial \eta}$ ⓐ $E_{\eta_0} T = \frac{d\psi_{\eta_0}}{dt}|_{t=0} = \frac{d\mathcal{C}}{d\eta_0} = \frac{A'(\theta)}{\eta_0'(\theta)}$, ⓑ $E_{\eta_0} T^2 = \mathcal{C}''(\eta_0) + \mathcal{C}'(\eta_0)^2$, ⓒ $Var(T) = \mathcal{C}''(\eta_0) = \frac{A''(\theta)}{[\eta_0(\theta)]^2} - \frac{\eta_0(\theta)'' A'(\theta)}{[\eta_0(\theta)']^3} = \frac{\partial^2 \mathcal{C}(\eta)}{\partial\eta\partial\eta^T}$
[NEF Differential] $G(\eta) := E_\eta(g) = \int g(\omega)\exp\left\{\eta^T T(\omega) - \mathcal{C}(\eta)\right\} h(\omega) d\nu(\omega)$ for η in interior of Ξ_g ① G is continuous and has continuous derivatives of all orders. ② Derivatives can be computed by differentiation under the integral sign. $\frac{dG(\eta)}{d\eta} = E_\eta\left[g(\omega)\left(T(\omega) - \frac{\partial}{\partial\eta}\xi(\eta)\right)\right]$ where Ξ_g is set η such that $\int |g(\omega)|\exp\left\{\eta^T T(\omega) - \mathcal{C}(\eta)\right\} h(\omega) d\nu(\omega) < \infty$
[NEF Min Suff] ① If there exists $\Theta_0 = \{\theta_0, \theta_1, \cdots, \theta_p\} \subset \Theta$ s.t. vectors $\eta_i = \eta(\theta_i) - \eta(\theta_0), i \in [1,p]$ are linearly independent in $\mathcal{R}^p$, then T is also minimal sufficient. Check $det([\eta_1, \cdots, \eta_p])$ is non-zero ② $\Xi = \{\eta(\theta) : \theta \in \Theta\}$ contains $(p+1)$ points that do not lie on the same hyperplane ③ Ξ is full rank.
[NEF complete and sufficient] If $\mathcal{P}$ is NEF of full rank then $T(X)$ is complete and sufficient for $\eta \in \Xi$
[NEF MLE] $\hat{\theta} = \eta^{-1}(\hat{\eta})$ or solution of $\frac{\partial\eta(\theta)}{\partial\theta} T(x) = \frac{\partial\xi(\theta)}{\partial\theta}$
[NEF Fisher Info] If $\underline{I}(\eta)$ is fisher info natural parameter η, then $Var(T) = \underline{I}(\eta)$. Let $\psi = E[T(X)]$. Suppose $\bar{I}(\psi)$ is fisher info matrix for parameter ψ, then $Var(T) = [\bar{I}(\psi)]^{-1}$
[NEF RLEs] RLE regularity condition (1, 2, 3, 4) holds due to proposition 3.2 and theorem 2.1, and result for (3). Only need to check condition on Fisher Info, then when n is large, there exists $\hat{\eta}_n$ s.t. $g(\hat{\eta}_n) = \hat{\mu}_n$ and $\hat{\eta}_n \to_{\text{a.s.}} \eta$ $\sqrt{n}(\hat{\eta}_n - \eta) \to_D N\left(0, \left[\frac{\partial^2}{\partial\eta\partial\eta^T}\mathcal{C}(\eta)\right]^{-1}\right)$
Where $g(\eta) = \frac{\partial\mathcal{C}(\eta)}{\partial\eta}$ and $\hat{\mu}_n = \frac{1}{n}\sum_{i=1}^{n} T(X_i)$
[UMP NEF] ⓐ UMP $T(Y) = I(Y > c)$ ⓘ $\eta(\theta)$ increasing and $H_1 : \theta \geq \theta_0$ ⓘⓘ $\eta(\theta)$ decreasing and $H_1 : \theta \leq \theta_0$ ⓑ Reverse inequalities $T(Y) = I(Y < c)$ ⓘ $\eta(\theta)$ increasing and $H_1 : \theta \leq \theta_0$ ⓘⓘ $\eta(\theta)$ decreasing and $H_1 : \theta \geq \theta_0$
[UMP Normal results] Given $X_i \sim N(\mu, \sigma^2)$ and $H_0 : \sigma^2 = \sigma_0^2$ ⓐ $S^2 = \frac{1}{n-1}\sum_{i=1}^{n}(X_i - \bar{X})^2$ independent to $\bar{X}$ ⓑ $V = \frac{1}{\sigma_0^2}\sum_{i=1}^{n}(X_i - \bar{X})^2 = \frac{(n-1)S^2}{\sigma^2} \sim \chi^2_{n-1}$ ⓒ $t = \frac{\sqrt{n}(\bar{X}-\mu)/\sigma}{\sqrt{V/(n-1)}} = \frac{Z}{\sqrt{V/(n-1)}} \sim t_{(n-1)}$ (only if $X_i \sim N$)
[UMPU NEF $\eta(\theta) = \theta$] Require: ① suff stat Y for θ ② suff and complete U for φ such that φ is full-rank
[UMPU NEF $H_0 : \theta \leq \theta_1$ or $\theta \geq \theta_2$ $H_1 : \theta_1 < \theta < \theta_2$] $T(Y,U) = I(c_1(U) < Y < c_2(U))$ s.t. $E_{\theta_1}[T(Y,U)|U = u] = E_{\theta_2}[T(Y,U)|U = u] = \alpha$
[UMPU NEF $H_0 : \theta_1 \leq \theta \leq \theta_2$ $H_1 : \theta < \theta_1$ or $\theta > \theta_2$] $T(Y,U) = I(Y < c_1(U)$ or $Y > c_2(U))$ s.t. $E_{\theta_1}[T(Y,U)|U = u] = E_{\theta_2}[T(Y,U)|U = u] = \alpha$
[UMPU NEF $H_0 : \theta = \theta_0$ $H_1 : \theta \neq \theta_0$] $T(Y,U) = I(Y < c_1(U)$ or $Y > c_2(U))$ s.t. $E_{\theta_0}[T_*(Y,U)|U = u] = \alpha$ and $E_{\theta_0}[T_*(Y,U)Y|U = u] = \alpha E_{\theta_0}(Y|U = u)$
[UMPU NEF $H_0 : \theta \leq \theta_0$ $H_1 : \theta > \theta_0$] $T(Y,U) = I(Y > c(U))$ s.t. $E_{\theta_0}[T(Y,U)|U = u] = \alpha$
[UMPU Normal] Require UMPU NEF ①, ② and ③ $V(Y,U)$ independent of U under H_0
[UMPU Normal $H_0 : \theta \leq \theta_1$ or $\geq \theta_2$ $H_1 : \theta_1 < \theta < \theta_2$] ④ V to be increasing in Y $T(V) = I(c_1 < V < c_2)$ s.t. $E_{\theta_1}[T(V)] = E_{\theta_2}[T(V)] = \alpha$
[UMPU Normal $H_0 : \theta_1 \leq \theta \leq \theta_2$ $H_1 : \theta < \theta_1$ or $\theta > \theta_2$] ④ V to be increasing in $Y \Rightarrow T(V) = I(V < c_1$ or $V > c_2)$ s.t. $E_{\theta_1}[T(V)] = E_{\theta_2}[T(V)] = \alpha$
[UMPU Normal $H_0 : \theta = \theta_0$ $H_1 : \theta \neq \theta_0$] ④ $V(Y,U) = a(u)Y + bU \Rightarrow T(V) = I(V < c_1$ or $V > c_2)$ s.t. $E_{\theta_0}[T(V)] = \alpha$ and $E_{\theta_0}[T(V)V] = \alpha E_{\theta_0}(V)$
[UMPU Normal $H_0 : \theta \leq \theta_0$ $H_1 : \theta > \theta_0$] ④ V to be increasing in $Y \Rightarrow T(V) = I(V > c)$ s.t. $E_{\theta_0}[T(V)] = \alpha$
[MLR for one-param exp fam] $\eta(\theta)$ nondecreasing in $\theta \Rightarrow \eta'(\theta) > 0$.

Statistics

[Sufficiency] $T(X)$ is sufficient for $P \in \mathcal{P} \Leftrightarrow P_X(x|Y = y)$ is known and does not depend on P. T sufficient for $\mathcal{P}_0$ but not ncessarily $\mathcal{P}_1$, $\mathcal{P}_0 \subset \mathcal{P} \subset \mathcal{P}_1$.
[Factorization theorem] $T(X)$ is sufficient for $P \in \mathcal{P} \Leftrightarrow$ there are non-negative Borel functions h with ① $h(x)$ does not depend on P ② $g_P(t)$ which depends on P s.t. $\frac{dP}{d\nu}(x) = g_P(T(x))h(x)$
[Minimal sufficiency] T is minimal sufficient $\Leftrightarrow T = \psi(S)$ for any other sufficient statistics S. Min suff is unique and usually exist.
[Min Suff-Method 1] (Theorem A) Suppose $\mathcal{P}_0 \subset \mathcal{P}$ and $\mathcal{P}_0$-a.s. implies $\mathcal{P}$-a.s. If T is sufficient for $P \in \mathcal{P}$ and minimal sufficient

for $P \in \mathcal{P}_0$, then T is minimal sufficient for $P \in \mathcal{P}$ (Theorem B) Suppose $\mathcal{P}$ contains PDFs $f_0, f_1, \cdots$ w.r.t a σ-finite measure. ⓐ Define $f_\infty(x) = \sum_{i=0}^{\infty} c_i f_i(x)$ and $T_i(x) = f_i(x)/f_\infty(x)$, then $T(X) = (T_0(X), T_1(X), \cdots)$ is minimal sufficient for $\mathcal{P}$. Where $c_i > 0, \sum_{i=0}^{\infty} c_i = 1, f_\infty(x) > 0$. ⓑ If $\{x : f_i(x) > 0\} \subset \{x : f_0(x) > 0\}$ for all i, then $T(X) = (f_1(x)/f_0(x), f_2(x)/f_0(x), \cdots$ is minimal sufficient for $\mathcal{P}$

[Min Suff-Method 2] (Theorem C) If ⓐ $T(X)$ is sufficient, and ⓑ $\exists\ \phi$ s.t. for $\forall\ x, y.\ f_P(x) = f_P(y)\phi(x, y)\ \forall\ P \in \mathcal{P} \Rightarrow T(x) = T(y)$. Then $T(X)$ is minimal sufficient for $\mathcal{P}$

[Ancillary statistics] A statistics $V(X)$ is ancillary for $\mathcal{P}$ if its distribution does not depend on population $P \in \mathcal{P}$ (First-order ancillary) if $E_P[V(X)]$ does not depend on $P \in \mathcal{P}$

[Completeness] $T(X)$ is complete for $P \in \mathcal{P} \Leftrightarrow$ for any Borel function g, $E_P g(T) = 0$ implies $g(T) = 0$, boundedly complete $\Leftrightarrow g$ is bounded. Completeness + Sufficiency $\Rightarrow$ Minimal Sufficiency

[Basu's theorem] If V is ancillary and T is boundedly complete and sufficient, then V and T are independent w.r.t any $P \in \mathcal{P}$

[Completeness for Varying Support] $\int_0^\theta g(x)x^{n-1}dx = 0 \implies g(\theta)\theta^{n-1} = 0, \implies g(\theta) = g(X_{(n)}) = 0$ and thus $X_{(n)}$ is complete

Fisher information $I(\theta) = E\left(\frac{\partial}{\partial\theta}\log f_\theta(X)\right)^2 = \int\left(\frac{\partial}{\partial\theta}\log f_\theta(X)\right)^2 f_\theta(X)d\nu(x) = E\left\{\frac{\partial}{\partial\theta}\log f_\theta(X)\left[\frac{\partial}{\partial\theta}\log f_\theta(X)\right]^T\right\}$

[Parameterization] If $\theta = \psi(\eta)$ and ψ' exists, $\tilde{I}(\eta) = \psi'(\eta)^2 I(\psi(\eta))$

[Twice differentiable] Suppose f_θ is twice differentiable in θ and $\int \frac{\partial^2}{\partial\theta^2} f_\theta(x) I_{f_\theta(x)>0} d\nu = 0$, then $I(\theta) = -E\left[\frac{\partial^2}{\partial\theta\partial\theta^T}\log f_\theta(X)\right]$

[Independent samples] If $\int \frac{\partial}{\partial\theta} f_\theta(x)d\nu = 0$ holds, then $I_{(X,Y)}(\theta) = I_X(\theta) + I_Y(\theta)$, and $I_{(X_1,\cdots,X_n)}(\theta) = nI_{X_1}(\theta)$

Comparing decision rules

[Compare decision rules] ⓐ as good as if $R_{T_1}(P) \le P_{T_2}(P)$. $\forall\ P \in \mathcal{P}$ ⓑ better if $R_{T_1}(P) < R_{T_2}(P)$ for some $P \in \mathcal{P}$ (and T_2 is dominated by T_1). ⓒ equivalent if $R_{T_2}(P) = R_{T_2}(P)$ for all $P \in \mathcal{P}$

[Optimal] T_* is $\mathcal{J}$-optimal if T_* is as good as any other rule in $\mathcal{J}$,

[Admissibility] $T \in \mathcal{J}$ is $\mathcal{J}$-admissible if no $S \in \mathcal{J}$ is better than T in terms of the risk.

[Minimaxity] $T_* \in \mathcal{J}$ is $\mathcal{J}$-minimax if $\sup_{P\subset\mathcal{P}} R_{T_*}(P) \le \sup_{P\subset\mathcal{P}} R_T(P)$ for any $T \in \mathcal{J}$

[Bayes Risk] A form of averaging $R_T(P)$ over $P \in \mathcal{P}$. Bayes risk $r_T(\Pi) = \int_{\mathcal{P}} R_T(P)d\Pi(P)$, $R_T(\Pi)$ is Bayes risk of T wrt a known probability measure Π.

[Bayes rule] T_* is $\mathcal{J}$-Bayes rule wrt Π if $r_{T_*}(\Pi) \le r_T(\Pi)$ for any $T \in \mathcal{J}$.

[Finding Bayes rule] Let $\tilde{\theta} \sim \pi$, $X|\tilde{\theta} \sim P_{\tilde{\theta}}$, then $r_\pi(T) = E\left[L(\tilde{\theta}, T(X)\right] = E\left[E\left[L(\tilde{\theta}, T(X)\right]|X\right]$ where E is taken jointly over $(\tilde{\theta}, X)$. Then find $T_*(x)$ that minimises the conditional risk.

[Rao-Blackwell] ⓐ Suppose $L(P, a)$ is convex and T is sufficient and S_0 is decision rule satisfying $E_P|||S_0|| < \infty$ for all $P \in \mathcal{P}$. Let $S_1 = E[S_0(X)|T]$, then $R_{S_1}(P) \le R_{S_0}(P)$. ⓑ If $L(P, a)$ is strictly convex in a, and S_0 is not a funciton of T, then S_0 is inadmissible and dominated by S_1.

MOM

[MoM] $\mu_j = E_\theta X^j = h_j(\theta), \implies \hat{\theta} = h_j^{-1}(\hat{\mu}_j)$. Provided h_j^{-1} exists and $\hat{\mu}_j = \frac{1}{n}\sum_{i=1}^n X_i^j$.

[MOM asymptotic] θ_n is unique if $h^{-1}(X)$ exists. Strongly consistent if h^{-1} is continuous via SLLN and continuous mapping. If h^{-1} is differentiable and $E|X_1|^{2k} < \infty$ then use CLT and δ-method. V_μ is $k \times k$ with $(i, j) = \mu_{i+j} - \mu_i\mu_j$ $\sqrt{n}(\hat{\theta}_n - \theta) \xrightarrow{D} N(0, [\nabla g]^T V_\mu \nabla g)$ MOM is $\sqrt{n}$-consistent, and if $k = 1$ $amse_{\hat{\theta}_n}(\theta) = g'(\mu_1)^2\sigma^2/n$, $\sigma^2 = \mu_2 - \mu_1^2$

MLE

[MLE] $\hat{\theta} = \arg\max_\theta L(\theta)$. Consider (a) boundary opint (b) $\partial L(\theta)/\partial\theta = 0$ and $\partial^2 L(\theta)/\partial\theta^2 < 0$ (Concave), note MLE may not exist

[MLE Consistency] Suppose ① Θ is compact ② $f(x|\theta)$ is continuous in θ for all x ③ There exists a function $M(x)$ s.t. $E_{\theta_0}[M(X)] < \infty$ and $|\log f(x|\theta) - \log f(x|\theta_0)| \le M(x)$ for all x, θ ④ identifiability holds $f(x|\theta) = f(x|\theta_0)$ ν-a.e. $\Rightarrow \theta = \theta_0$. Then MLE estimate $\hat{\theta}_n \xrightarrow{\text{a.s}} \theta_0$

[RLE] [Roots of the Likelihood Equation] θ that solves $\frac{\partial}{\partial\theta}\log L_n(\theta) = 0$

[RLE regularity conditions] Suppose ① Θ is open subset of $\mathcal{R}^k$ ② $f(x|\theta)$ is twice continuously differentiable in θ for all x, and $\frac{\partial}{\partial\theta}\int f(x|\theta)d\nu = \int \frac{\partial}{\partial\theta} f(x|\theta)d\nu$, $\frac{\partial}{\partial\theta}\int \frac{\partial}{\partial\theta^T} f(x|\theta)d\nu = \int \frac{\partial^2}{\partial\theta\partial\theta^T} f(x|\theta)d\nu$. ③ $\Psi(x, \theta) = \frac{\partial^2}{\partial\theta\partial\theta^T}\log f(x|\theta)$, there exists a constant c and non-negative function H s.t. $EH(X) < \infty$ and $\sup_{||\theta-\theta_*||<c} ||\Psi(x, \theta)|| \le H(x)$. ④ Identifiable

[RLE consistency] Under RLE regularity conditions, there exists a sequence of $\hat{\theta}_n$ s.t. $\frac{\partial}{\partial\theta}\log L_n(\hat{\theta}_n) = 0$ and $\hat{\theta}_n \to_{\text{a.s.}} \theta_*$.

[RLE asymptotic normality] Assume RLE regularity conditions, and $I(\theta) = \int \frac{\partial}{\partial\theta}\log f(x|\theta)\left[\frac{\partial}{\partial\theta}\log f(x|\theta)\right]^T d\nu(x)$ is positive definite and $\theta = \theta_*$. Then any consistent sequence $\{\tilde{\theta_n}\}$ of RLE it holds $\sqrt{n}(\tilde{\theta_n} - \theta_*) \xrightarrow{D} N\left(0, \frac{1}{I(\theta_*)}\right)$

[One-step MLE] Often asym efficient, useful to adjust an non asym efficient estimators provided $\hat{\theta}_n^{(0)}$ is $\sqrt{n}$-consistent. $\hat{\theta}_n^{(1)} = \hat{\theta}_n^{(0)} - \left[\nabla s_n(\hat{\theta}_n^{(0)})\right]^{-1} s_n(\hat{\theta}_n^{(0)})$

Unbiased Estimators

[UMVUE] $T(X)$ is UMVUE for $\theta \Leftrightarrow Var(T(X) \le Var(U(X))$ for any $P \in \mathcal{P}$ and any other unbiased estimator $U(X)$ of θ

[Lehmann-Scheffé] If $T(X)$ is sufficient and complete for θ. If θ is estimable, then there is a unique unbiased estimator of θ that is of the form $h(T)$.

[UMVUE method1] Using Lehmann-Scheffé, suppose T is sufficient and complete manipulate $E(h(T)) = \theta$ to get $\hat{\theta}$.

[UMVUE method2] Using Rao-Blackwellization. Find ① unbiased estimator of $\theta = U(X)$ ② sufficient and complete statistics $T(X)$ ③ then $E(U|T)$ is the UMVUE of θ by Lehmann-Scheffé.

[UMVUE method3] Useful when no complete and sufficient statistics. Can use to find UMVUE, check if estimator is UMVUE, show nonexistence of UMVUE. $T(X)$ is UMVUE $\Leftrightarrow E[T(X)U(X)] = 0$

ⓐ T is unbiased estimator of η with finite variance, $\mathcal{U}$ is set of all unbiased estimators of 0 with finite variances. ⓑ $T = h(S)$, where S is sufficient and h is Borel function, $\mathcal{U}_S$ is subset of $\mathcal{U}$ consisting of Borel functions of S.

[Using method3] ① Find $U(x)$ via $E[U(x)] = 0$ ② Construct $T = h(S)$ s.t. T is unbiased ③ Find T via $E[TU] = 0$

[Corollary] If T_j is UMVUE of η_j with finite variances, then $T = \sum_{j=1}^k c_j T_j$ is UMVUE of $\eta = \sum_{j=1}^k c_j\eta_j$. If T_1, T_2 are UMVUE of η with finite variances, then $T_1 = T_2$ a.s. $P, P \in \mathcal{P}$

[Cramér-Rao Lower Bound] Suppose ① Θ is an open set and P_θ has pdf f_θ ② f_θ is differentiable and $\frac{\partial}{\partial\theta}\int f_\theta(x)d\nu = \int \frac{\partial}{\partial\theta} f_\theta(x)d\nu = 0$. ③ $g(\theta)$ is differentiable and $T(X)$ is unbiased estimator of $g(\theta)$ s.t. $g'(\theta) = \frac{\partial}{\partial\theta}\int T(x)f_\theta(x)d\nu = \int T(x)\frac{\partial}{\partial\theta} f_\theta(x)d\nu, \theta \in \Theta$. Then

$Var(T(X)) \geq \frac{g'(\theta)^2}{I(\theta)} = \left[\frac{\partial}{\partial\theta}g(\theta)\right]^T [I(\theta)]^{-1}\frac{\partial}{\partial\theta}g(\theta)$

[CR LB for biasd estimator] $Var(T) \geq \frac{[g'(\theta)+b'(\theta)]^2}{I(\theta)}$

[CR LB iff] CR achieve equality ⓐ $\Leftrightarrow T = \left[\frac{g'(\theta)}{I(\theta)}\right]\frac{\partial}{\partial\theta}\log f_\theta(X) + g(\theta)$ ⓑ $\Leftrightarrow f_\theta(X) = \exp(\eta(\theta)T(x) - \xi(\theta))h(x)$, s.t. $\xi'(\theta) = g(\theta)\eta'(\theta)$ and $I(\theta) = \eta'(\theta)g'(\theta)$

[UMVUE asymptotic] Typically consistent, exactly unbiased, ratio of mse over Cramér-Rao LB converges to 1 (asym they are the same).

Other estimators

[Upper semi-continuous (usc)] $\lim_{\rho\to 0}\left\{\sup_{||\theta'-\theta||<\rho} f(x|\theta')\right\} = f(x|\theta)$

[USC in θ] Suppose (1) Θ is compact with metric $d(\cdot,\cdot)$ (2) $f(x|\theta)$ is usc in θ and for all x (3) there exists a function $M(x)$ s.t. $E_{\theta_0}|M(X)| < \infty$ and $\log f(x|\theta) - \log f(x|\theta_0) \leq M(x)$ for all x and θ (4) for all $\theta \in \Theta$ and sufficiency small $\rho > 0$, $\sup_{d(\theta',\theta)<\rho} f(x|\theta')$ is measurable in x (5) identifiable $f(x|\theta) = f(x|\theta_0)$ ν-a.e. $\Rightarrow \theta = \theta_0$. Then $d(\hat\theta_n, \theta_0) \to_{\text{a.s.}} 0$

[Asym Covariance Matrix] $V_n(\theta)$ is $k\times k$ positive definite matrix called asym covariance matrix. $V_n(\theta)$ is usually in form of $n^{-\delta}V(\theta)$, higher δ means faster convergence. $[V_n(\theta)]^{-1/2}(\hat\theta_n - \theta) \to_D N_k(0, I_k)$

[Information Inequalities] $A \preccurlyeq B$ means $B - A$ is positive semi-definite. Suppose two estimators $\hat\theta_{1n}, \hat\theta_{2n}$ satisfy asym covariance matrix with $V_{1n}(\theta), V_{2n}(\theta)$. $\hat\theta_{1n}$ is asym more efficient thant $\hat\theta_{2n}$ if (1) $V_{1n}(\theta) \preccurlyeq V_{2n}(\theta)$ for all $\theta \in \Theta$ and all large n (2) $V_{1n}(\theta) \prec V_{2n}(\theta)$ for at least one $\theta \in \Theta$ But note $\hat\theta_n$ is asym unbiased but CR LB might not hold even if regularity condition is satisfied.

[M-estimators] General method to find $\hat\theta_n$ maximises criterion function $S_\theta(x)$, for MLE $s_\theta(x) = \log f(x|\theta)$. $E_{\theta_0}s_\theta(X) < E_{\theta_0}s_{\theta_0}(X)$ $\forall \theta \neq \theta_0$. $\theta \mapsto S_n(\theta) = \frac{1}{n}\sum_{i=1}^n s_\theta(X_i)$

[Consistency of M-estimators] $S_n(\theta)$ is random function while $S(\theta)$ is fixed s.t. $\sup_{\theta\in\Theta}|S_n(\theta) - S(\theta)| \to_P 0$ and for every $\rho > 0$ $\sup_{\theta:d(\theta,\theta_0)\geq\rho} S(\theta) < S(\theta_0)$. Then any sequence of estimators $\hat\theta_n$ with $S_n(\hat\theta_n) \geq S_n(\theta_0) - o_P(1)$ converges in probability to θ_0

[Hodges' estimator] $X_i \sim N(\theta, 1)$, $\hat\theta_n = \bar X_n$ if $\bar X_n \geq n^{-1/4}$ and $t\bar X_n$ otherwise. $V_n(\theta) = 1/n$ if $\theta \neq 0$ and t^2/n otherwise. if $\theta \neq 0$: $\sqrt{n}(\hat\theta - \theta) = \sqrt{n}(\bar X_n - \theta) - (1-t)\sqrt{n}\bar X_n I_{|\bar\theta_n|<n^{-1/4}}$ if $\theta = 0$: $= t\sqrt{n}(\bar X_n - \theta) + (1-t)\sqrt{n}\bar X_n I_{|\bar X_n|\geq n^{-1/4}}$

[Super-efficiency] Point where UMVUE failed Hodeges' estiamtor in information inequality (2). But under the basic regularity condition and if Fisher Information is positive definite at $\theta = \theta_*$, if $\hat\theta_n$ satisfies Asym covariance matrix, then there is a $\Theta_0 \subset \Theta$ with Lebesgue measure 0 s.t. information inequality (2) holds for any $\theta \notin \Theta_0$

[Asym efficiency] Assume Fisher Info $I_n(\theta)$ is well-defined and positive definite for every n, seq of estimators $\{\hat\theta_n\}$ satisfies asym cov matrix is asym efficient or asym optimal if and only if $V_n(\theta) = [I_n(\theta)]^{-1}$.

Asymptotics

[Consistency of point estimators] ⓐ consistent $T_n(X) \xrightarrow{P} \theta$ ⓑ strongly consistent $T_n(X) \xrightarrow{\text{a.s.}} \theta$ ⓒ a_n-consistent $a_n(T_n(X) - \theta) = O_P(1)$, $\{a_n\} > 0$ and diverge to ∞ ⓓ L_r-consistent $T_n(X) \xrightarrow{L^P} \theta$ for some fixed $r > 0$.

[Remark on consistency] A combination of LLN, CLT, Slustky's, continuous mapping, δ-method are used. If T_n is (strongly) consistent for θ and g is continuous at θ then $g(T_n)$ is (strongly) consistent for $g(\theta)$

[Affine estimator] Consider $T_n = \sum_{i=1}^n c_{ni}X_i$ ① If $c_{ni} = c_i/n$ s.t $\frac{1}{n}\sum_{i=1}^n c_i \to 1$ and $\sup_i |c_i| < \infty$ then T_n is strongly consistent. ② If population variance is finite, then T_n is consistent in mse $\Leftrightarrow \sum_{i=1}^n c_{ni} \to 1$ and $\sum_{i=1}^n c_{ni}^2 \to 0$

[Asymptotic distribution] $\{a_n\} > 0$ and either ⓐ $a_n \to \infty$ ⓑ $a_n \to a > 0$, s.t. $a_n(T_n - \theta) \xrightarrow{D} Y$. When estimator's expectations or second moment are not well defined, we need asymptotic behaviours.

[Asymptotic bias] $\tilde b_{T_n} = EY/a_n$, asymptotically unbiased if $\lim_{n\to\infty}\tilde b_{T_n}(P) = 0$, $b_{T_n}(P) := ET_n(X) - \theta$

[Asymptotic expectation] If $a_n\xi_n \to^D \xi$, $E|\xi| < \infty$, then asymptotic expectation of ξ_n is $E\xi/a_n$

[Asymptotic MSE] asymptotic expectation of $(T_n - \theta)^2$ or $\text{amse}_{T_n}(P) = EY^2/a_n^2$ (Remark) $EY^2 \leq \liminf_{n\to\infty} E[a_n^2(T_n - v)^2]$ (amse is no greater than exact mse)

[Asymptotic variance] $\sigma^2_{T_n}(P) = Var(Y)/a_n^2$

[Asym Relative Efficiency] $e_{T_{1n},T_{2n}} = amse_{T_{2n}(P)}/amse_{T_{1n}(P)}$. Note efficiency of estimator T refers to $1/[I(\theta)MSE_T(\theta)]$

[δ-method corollary] If $a_n \to \infty$, g is differentiable at θ, $U_n = g(T_n)$. Then ⓐ amse of U_n is $[g'(\theta)^2EY^2]/a_n^2$ ⓑ asym var of U_n is $[g'(\theta)^2Var(Y)]/a_n^2$

[Quantiles asymptotic] $F(\theta) = \gamma \in (0,1)$ and $\hat\theta_n := \lfloor\gamma n\rfloor$-th order statistics, $F'(\theta) > 0$ and exists. $\sqrt{n}(\hat\theta_n - \theta) \xrightarrow{D} N\left(0, \frac{\gamma(1-\gamma)}{[F'(\theta)]^2}\right)$

Hypothesis testing

[Hypothesis tests] Let $\mathcal{P}$ be a family of distributions, $\mathcal{P}_0 \subset \mathcal{P}, \mathcal{P}_1 = \mathcal{P}\backslash\mathcal{P}_0$. Hypothesis testing decides between $H_0 : P \in \mathcal{P}_0, H_1 : P \in \mathcal{P}_1$. Action space $\mathcal{A} = \{0,1\}$, decision rule is called a test $T : \mathcal{X} \to \{0,1\} \Rightarrow T(X) = I_C(X)$ for some $C \subset \mathcal{X}$. C is called the region/critical region.

[$0-1$ loss] Common loss function for hypo test, $L(P,j) = 0$ for $P \in \mathcal{P}_j$ and $= 1$ for $P \in \mathcal{P}_{1-j}, j \in \{0,1\}$ Risk $R_T(P) = P(T(X) = 1) = P(X \in C)$ if $P \in \mathcal{P}_0$ or $P(T(X) = 0) = P(X \notin C)$ if $P \in \mathcal{P}_1$

[Type I and II errors] Type I: H_0 is rejected when H_0 is true. $\beta_T(\theta_0) = E_{H_0}(T) \leq \alpha$ (within controlled with size α) Error rate: $\alpha_T(P) = P(T(X) = 1), P \in \mathcal{P}_0$ Type II: H_0 is accepted when H_0 is false. $1 - \beta_T(\theta)$ for $\theta \in \Theta_1$ Error rate: $1 - \alpha_T(P) = P(T(X) = 1), P \in \mathcal{P}_1$

[Power function of T] $\alpha_T(P)$, Type I and Type II error rates cannot be minimized simultaneously.

[Significance level] Under Neyman-Pearson framework, assign pre-specified bound α (significance level of test): $\sup_{P\subset\mathcal{P}_0} P(T(X) = 1) \leq \alpha$

[size of test] α' is the size of the test $\sup_{P\subset\mathcal{P}_0} P(T(X) = 1) = \alpha'$

[NP Test] Steps ① Find joint distribution $f(X)$ and determine MLR and/or NEF ② Formulate hypothesis H_0, H_1 - simple/composite about θ and not $f(\theta)$ ③ Form N-P test structure T_* ④ Find test distribution and rejection region.

[Generalised NP] ϕ is the T (Test framework) $\max_\phi \int \phi f_{m+1} d\nu$ s.t. $\int \phi f_i d\nu \leq t_i$ $\forall$ $i \in (1,m)$, (Required condition) If $\exists$ $c_1, \cdots, c_m$ s.t. $\phi_*(x) = I[f_{m+1}(x) > \sum_{i=1}^m c_i f_i(x)]$, then ϕ_* maximises objective function with equality constraint. If $c_i \geq 0$ then ϕ_* maximises with inequality constraint.

[UMP] ① $H_0 : P = p_0$ $H_1 : P = p_1 \Rightarrow T(X) = I(p_1(X) > cp_0(X))$, $\beta_T(p_0) = \alpha$ ② $H_0 : \theta \leq \theta_0$ $H_1 : \theta > \theta_0 \Rightarrow T(Y) = I(Y > c)$, $\beta_T(\theta_0) = \alpha$ ③ $H_0 : \theta \leq \theta_1$ or $\theta \geq \theta_2$ $H_1 : \theta_1 < \theta < \theta_2$, $\Rightarrow T(Y) = I(c_1 < Y < c_2)$, $\beta_T(\theta_1) = \beta_T(\theta_2) = \alpha$

UMP Satisfy (1) pre-set size $\alpha = E_{H_0}(T)$ (2) max power $\beta_T(P) = E_{H_1}(T)$

[No UMP] $H_0 : \theta = \theta_1, H_1 : \theta \neq \theta_1$ and $H_0 : \theta \in (\theta_1, \theta_2)$ $H_1 : \theta \notin (\theta_1, \theta_2)$

[N-P lemma] NP test has non-trival power $\alpha < \beta_{H_1}(T)$ unless $P_0 = P_1$, and is unique up to γ (randomised test)
[Show T_* is UMP in simple hypothesis] UMP when $E_1[T_*] - E_1[T] \geq 0$, *key equation*: $(T_* - T)(f_1 - cf_0) \geq 0$. $\Rightarrow \int (T_* - T)(f_1 - cf_0) = \beta_{H_1}(T_*) - \beta_{H_1}(T) \geq 0$.
[UMP unique up to randomised test in simple hypothesis] $(T_* - T)(f_1 - cf_0) \geq 0, \int (T_* - T)(f_1 - cf_0) = 0 \Rightarrow (T_* - T)(f_1 - cf_0) = 0$ and $T_* = T$
[Composite hypothesis] Simple $\Rightarrow$ Composite when $\beta_T(\theta_0) \geq \beta_T(\theta \in H_0)$ and/or $\beta_T(\theta_0) \leq \beta_T(\theta \in H_1)$ (or does not depend on θ). For MLR this is satisfied, others need to check.
[Monotone Likelihood Ratio] $\theta_2 > \theta_2$, increasing likelihood ratio in Y if $g(Y) = \frac{f_{\theta_2}(Y)}{f_{\theta_1}(Y)} > 1$ or $g'(Y) > 0$.
[Simultaneous] (Bonferroni) adjust each paramter level to $\alpha_t = \alpha/k$ (Bootstrap) Monte Carlo percentile estimate

Asymptotic test

[LR test] $\lambda(X) = \frac{\sup_{\theta \in \theta_0} \ell(\theta)}{\sup_{\theta \in \Theta} \ell(\theta)}$ Rejects $H_0 \Leftrightarrow \lambda(X) < c \in [0,1]$. *1-param Exp Fam LR test is also UMP.*
Assume MLE regularity condition, under H_0, $-2\log\lambda(X) \to \chi^2_r$, where $r := dim(\theta)$ $T(X) = I\left[\lambda(X) < \exp(-\chi^2_{r,1-\alpha}/2)\right]$ where $\chi^2_{r,1-\alpha}$ is the $(1-\alpha)$th quantile of χ^2_r.
[Wald's test] $W_n = R(\hat\theta)^T \{C(\hat\theta)^T I_n^{-1}(\hat\theta) C(\hat\theta)\}^{-1} R(\hat\theta)$, where $C(\theta) = \partial R(\theta)/\partial\theta$, $I_n(\theta)$ is fisher info for $X_1, \cdots, X_n$, $\hat\theta$ is unrestricted MLE/RLE of θ. [Wald's test - easy case] if $H_0 : \theta = \theta_0 \Rightarrow R(\theta) = \theta - \theta_0$, and $W_n = (\hat\theta - \theta_0)^T I_n(\hat\theta)(\hat\theta - \theta_0)$
[Rao's score test] $Q_n = s_n(\tilde\theta)^T I_n^{-1}(\tilde\theta) s_n(\tilde\theta)$. where score function $s_n(\theta) = \partial \log \ell(\theta)/\partial\theta$, $\tilde\theta$ is MLE/RLE of θ under $H_0 : R(\theta) = 0$. [Asymptotic Tests] Same test structure for LR, Wald', Rao's score test. $H_0 : R(\theta) = 0$, $\lim_{n\to\infty} W_n, Q_n \sim \chi^2_r$, $T(X) = I(W_n > \chi^2_{r,1-\alpha})$ or $I(Q_n > \chi^2_{r,1-\alpha})$

Non-param tests

[Sign test] $X_i \sim^{iid} F$, u is fixed constant, $p = F(u)$, $\triangle_i = I(X_i - u \leq 0)$, $P(\triangle_i = 1) = p$, $p_0 \in (0,1)$ $H_0 : p \leq p_0$ $H_1 : p > p_0$ $\Rightarrow T(Y) = I(Y > m)$, $Y = \sum_{i=1}^n \triangle_i \sim Bin(n,p)$, m, γ s.t. $\alpha = E_{p_0}[T(Y)]$ $H_0 : p = p_0$ $H_1 : p \neq p_0 \Rightarrow T(Y) = I(Y < c_1$ or $Y > c_2)$, $E_{p_0}[T] = \alpha$ and $E_{p_0}[TY] = \alpha n p_0$
[Permutation test] $X_{i1}, \cdots, X_{in_i} \sim^{iid} F_i$, $i = 1,2$ $H_0 : F_1 = F_2$ $H_1 : F_1 \neq F_2$, $\Rightarrow T(X)$ with $\frac{1}{n!}\sum_{z \in \pi(x)} T(z) = \alpha$ $\pi(x)$ is set of $n!$ points obtained from x by permuting components of x E.g. $T(X) = I(h(X) > h_m)$, $h_m := (m+1)^{th}$ largest $\{h(z : z \in \pi(x)\}$ e.g $h(X) = |\bar X_1 - \bar X_2|$ or $|S_1 - S_2|$
[Rank test] $X_i \sim^{iid} F$, $Rank(X_i) = \#\{X_j : X_j \leq X_i\}$, $H_0 : F$ symm ard 0, $H_1 : H_0$ false, R^o_+ vector of ordered R_+. (Wilcoxon) $T(X) = I[W(R^o_+) < c_1$ or $W(R^o_+ > c_2)]$, $W(R^o_+) = J(R^o_{+1}/n) + \cdots + J(R^o_{+n_*}/n)$ c_1, c_2 are $(m+1)^{th}$ smallest/largest of $\{W(y) : y \in \mathcal{Y}\}$, $\gamma = \alpha 2^n/2 - m$
[KS test] $X_i \sim^{iid} F$ $H_0 : F = F_0$, $H_1 : F \neq F_0$, $\Rightarrow T(X) = I(D_n(F_0) > c)$, $D_n(F) = \sup_{x \in \mathcal{R}} |F_n(x) - F(x)|$ With F_n Emp CDF, and for any $d, n > 0$, $P(D_n(F) > d) \leq 2\exp(-2nd^2)$,
[Cramer-von test] Modified KS with $T(X) = I(C_n(F_0) > c)$, $C_n(F) = \int \{F_n(x) - F(x)\}^2 dF(x)$ $nC_n(F_0) \xrightarrow{D} \sum_{j=1}^\infty \lambda_j \chi^2_{1j}$, with $\chi^2_{1j} \sim \chi^2_1$ and $\lambda_j = j^{-2}\pi^{-2}$
[Empirical LR] $X_i \sim^{iid} F$, $H_0 : \Lambda(F) = t_0$ $H_1 : \Lambda(F) \neq t_0$, $\Rightarrow T(X) = I(ELR_n(X) < c)$ $ELR_n(X) = \frac{\ell(\hat F_0)}{\ell(\hat F)}$, $\ell(G) = \prod_{i=1}^n P_G(\{x_i\})$, $G \in \mathcal{F}$. ($\mathcal{F} :=$ collection of CDFs, $P_G :=$ measure induced by CDF G)

Confidence set $C(X) : X \to \mathcal{B}(\Theta)$, Require $\inf_{P \in \mathcal{P}} P(\theta \in C(X)) \geq 1 - \alpha$, that is confidencen coeff should be more than level
[CI via pivotal qty] $C(X) = \{\theta : c_1 \leq \mathcal{R}(X,\theta) \leq c_2\}$, *not dependent on* P common pivotal qty: $(X_i - \mu)/\sigma$
[invert accept region] $C(X) = \{\theta : x \in A(\theta)\}$, Acceptance region $A(\theta) = \{x : T_{\theta_0}(x) \neq 1\}$. $H_0 : \theta = \theta_0$, any H_1 satisfy
[Shortest CI] require unimodal: $f'(x_0) = 0$ $f'(x) < 0, x < x_0$ and $f'(X) > 0, x > x_0$
[Pivotal $(T-\theta)/U$, f unimodal at x_0] Interval $[T - b_* U, T - a_* U]$, shortest when $f(a_*) = f(b_*) > 0$ $a_* \leq x_0 \leq b_*$
[Pivotal T/θ, $x^2 f(x)$ unimodal at x_0] Interval $[b_*^{-1}T, a_*^{-1}T_*]$ shortest when $a_*^2 f(a_*) = b_*^2 f(b_*) > 0$ $a_* \leq x_0 \leq b_*$
[General CI] Require $f > 0$, integrable, unimodal at x_0, (Objective) $\min b - a$ s.t. $\int_a^b f(x)dx$ and $a \leq b$ (Solution) a_*, b_* satisfy ① $a_* \leq x_0 \leq b_*$ ② $f(a_*) = f(b_*) > 0$ ③ $\int_{a_*}^{b_*} f(x)dx = 1 - \alpha$ forms the shortest CI, note it has to exactly the formulation above.
[Asymptotic CI] Require $\lim_{n\to} P(\theta \in C(X)) \geq 1 - \alpha$,
[Asymptotic pivotal] $\mathcal{R}_n(X,\theta) = \hat V_n^{-1/2}(\hat\theta_n - \theta)$ does not depend on P in limit
[Asymptotic LR CI] $C(X) = \left\{\theta : \ell(\theta, \hat\varphi) \geq exp(-\chi^2_{r,1-\alpha} - \alpha/2)\ell(\hat\theta)\right\}$
[Asymptotic Wald CI] $C(X) = \left\{\theta : (\hat\theta - \theta)^T \left[C^T \left(I_n(\hat\theta)\right)^{-1} C\right]^{-1} (\hat\theta - \theta) \leq \chi^2_{r,1-\alpha}\right\}$
[Asymptotic Rao CI] $C(X) = \left\{\theta : [s_n(\theta, \hat\varphi)]^T [I_n(\theta, \hat\varphi)]^{-1} [s_n(\theta, \hat\varphi)] \leq \chi^2_{r,1-\alpha}\right\}$

Bayesian

[Bayes formula] $\frac{dP_{\theta|X}}{d\Pi} = \frac{f_\theta(X)}{m(X)}$.
[Bayes action $\delta(x)$] $\arg\min_a E[L(\theta,a)|X = x]$, when $L(\theta,a) = (\theta - a)^2$, $\delta(x) = E(\theta|X = x)$, and bayes risk $r_\delta(\theta) = Var(\theta|X)$
[Generalised Bayes action] $\arg\min_a \int_\Theta L(\theta,a) f_\theta(x) d\Pi$, works for improper prior where $\Pi(\Theta) \neq 1$
[Interval estimation - Credible sets] $P_{\theta|x}(\theta \in C) = \int_C p_x(\theta) d\lambda \geq 1 - \alpha$
[HPD highest posterior dentsity] $C(x) = \{\theta : p_x(\theta) \geq c_\alpha\}$, often shortest length credible set. Is a horizontal line in the posterior density plot. Might not have exact confidence level $1 - \alpha$.
[Hierachical Bayes] With hyper-priors as hyper-parameters on the priors.
[Empirical Bayes] Estimate hyper-paramter via data using MoM (no MLE as not independent). $X_i \sim N(\mu, \sigma^2)$, $\mu|\xi \sim N(\mu_0, \sigma_0^2)$, σ^2 known, $\xi = (\mu_0, \sigma_0^2)$, Using MoM $E_\xi(X|\xi) = E_\xi(E[X|\mu,\xi]) = E_\xi(\mu|\xi) = \mu_0 \approx \bar X$, $E_\xi(X^2|\xi) = E_\xi(\mu^2 + \sigma^2|\xi) = \sigma^2 + \mu_0^2 + \sigma_0^2 \approx \frac{1}{n}\sum X_i^2$ $\Rightarrow \sigma_0^2 = \frac{1}{n}\sum_{i=1}^n (X_i - \bar X)^2 - \sigma^2$
[Normal posterior] Normal posterior $N(\mu_*(x), c^2)$ with prior unknown μ and known σ^2: $\mu_*(x) = \frac{\sigma^2}{\sigma^2 + n\sigma_0^2}\mu_0 + \frac{n\sigma_0^2}{\sigma^2 + n\sigma_0^2}\bar x$, $c^2 = \frac{\sigma_0^2\sigma^2}{n\sigma_0^2 + \sigma^2}$ $C(x) = [\mu_*(x) - cz_{1-\alpha/2},\ \mu_*(x) + cz_{1-\alpha/2}]$.
[Decision theory] (Admissibility) (1) $\delta(X)$ unique $\Rightarrow$ admissible, (2, 3) $r_\delta(\Pi) < \infty$, $\Pi(\theta) > 0$ for all θ and δ is Bayes action with respect to $\Pi \Rightarrow$ admissible. *Not true for improper priors*, Improper priors require excessive risk ignorable, take limit and observe if risk is admissible. (Bias) Under squared error loss, $\delta(X)$ is biased unless $r_\delta(\Pi) = 0$. *Not applicable to improper priors.* (Minimax) If T is

(unique) Bayes estimator under Π and $R_T(\theta) = \sup_{\theta'} R_T(\theta')$ π-a.e., , then T is (unique) minimax. *Limit of Bayes estimators* If T has constant risk and $\liminf_j r_j \geq R_T$, then T is minimax.
[Admissibility] $\delta(X)$ is a Bayes rule with prior Π, δ is admissible if (1) if δ is unique (2) If Θ is countable, $\Pi(\theta) > 0$ $\forall\Theta$. Note, not true for generalised Bayes rules unless limit is Bayes rule.
[Simul est] Simultaneous estimate vector-valued $\mathcal{V}$ with e.g. squared loss $L(\theta, a) = \|a - \theta\|^2 = \sum_{i=1}^{p}(a_i - \theta_i)^2$
[Bayes Asymptotic Property] (Posterior Consistency) $X \sim P_{\theta_0}$ and $\Pi(U|X_n) \xrightarrow{P_{\theta_0}} 1$ for all open U containing θ_0. (Wald type consistency) Assume $p_\theta(x)$ is continuous, measurable, θ_* is unique maximizer then MLE converge to true parameter θ^* P_* a.s. Furthermore, if θ^* is in the support of the prior, then posterior converges to θ^* in probability. (Posterior Robustness) all priors that lead to consistent posteriors are equivalent.
[Bernstein-von Mises] Assume MLE regularity conditions, posterior $T_n = \sqrt{n}(\tilde{\theta_n} - \hat{\theta_n}) \sim \mathcal{N}(\hat{\theta}_n, V^*/n)$ asymptotically. (Well-specified) $V^* = I(\theta^*)^{-1} = E_*\left[-\nabla_\theta^2 \log p_{\theta^*}(Y)\right]^{-1}$ (same as MLE, with θ^* as true parameter, CI = CR) $\sqrt{n}\left(\hat{\theta}_n - E_\theta[\theta|X_1, \cdots, X_n]\right) \xrightarrow{P} 0$ (If MLE has asym normality, so is posterior mean) (Mis-specified) $V^* = \mathbb{E}_*\left[-\nabla_\theta^2 \log p_{\theta_*}(Y)\right]^{-1}$, θ_* is projection of θ^* onto parameter space, or unique maximizer of $\ell^*(\theta) = E_*[\log p_\theta(Y)]$
[MLE asymptotic variance under model misspecification] $\mathbb{E}_*\left[-\nabla_\theta^2 \log p_{\theta^*}(Y)\right]^{-1} \mathrm{Var}_*\left(\nabla \log p_{\theta^*}(Y)\right) \mathbb{E}_*\left[-\nabla_\theta^2 \log p_{\theta^*}(Y)\right]^{-1}$ (differ from MLE, with θ_* the projection of P_* to parameter space)

Linear Model

[Linear Model] $X = Z\beta + \epsilon$ (or $X_i = Z_i^T\beta + \epsilon_i$) Estimate with $b = \min_b \|X - Zb\|^2 = \|X - Z\hat{\beta}\|^2$, [Generalised inverse] Moore-Penrose inverse $A^+AA^+ = A^+$, $A = (Z^TZ)$ [Projection matrix] $P_Z = Z(Z^TZ)^-Z^T$, $P_Z^2 = P_Z$, $P_Z Z = Z$, $rank(P_Z) = tr(P_Z) = r$
[LM Solution] (solution = normal equation) $Z^Z b = Z^T X$ (when Z is full rank): $\hat{\beta} = (Z^TZ)^{-1}Z^TX$ (when Z is not full rank): $\hat{\beta} = (Z^TZ)^-Z^TX$
[LM assumptions] (A1 Gaussian noise) $\epsilon \sim N_n(0, \sigma^2 I_n)$ (A2 homoscedastic noise) $E(\epsilon) = 0$, $Var(\epsilon) = \sigma^2 I_n$ (A3 general noise) $E(\epsilon) = 0$, $Var(\epsilon) = \Sigma$
[Estimable $\ell\beta$] Estimate linear combination of coefficient (General) necessary and Sufficient condition: $\ell \in R(Z) = R(Z^TZ)$ (under A3) LSE $\ell^T\hat{\beta}$ is unique and unbiased (under A1) if $\ell \notin R(Z)$, $\ell^T\beta$ not estimable
[LM property under A1] 1 LSE $\ell^T\hat{\beta}$ is UMVUE of $\ell^T\beta$, 2 UMVUE of $\hat{\sigma}^2 = (n-r)^{-1}\|X - Z\hat{\beta}\|^2$, r is rank of Z 3 $\ell\hat{\beta}$ and $\hat{\sigma}^2$ are independent, $\ell^T\hat{\beta} \sim N(\ell^T\beta, \sigma^2\ell^T(Z^TZ) - \ell)$, $(n-r)\hat{\sigma}/\sigma^2 \sim \chi^2_{n-r}$
[LM property under A2] LSE $\ell^T\hat{\beta}$ is BLUE (Best Linear Unbiased Estimator, best as in min var)
[LM property under A3] Following are equivalent: a $\ell^T\hat{\beta}$ is BLUE for $\ell^T\beta$ (also UMVUE), b $E[\ell^T\hat{\eta}^TX) = 0]$, any η is s.t. $E[\eta^TX] = 0$ c $Z^T var(\epsilon)U = 0$, for U s.t. $Z^TU = 0$, $R(U^T) + R(Z^T) = R^n$ d $Var(\epsilon) = Z\Lambda_1 Z^T + U\Lambda_2 U^T$, for some Λ_1, Λ_2, U s.t. $Z^TU = 0$, $R(U^T) + R(Z^T) = R^n$ e $Z(Z^TZ)^-Z^TVar(\epsilon)$ is symmetric
[LM consistency] $\lambda_+[A]$ is the largest eigenvalue of $A_n = (Z^TZ)^-$. Suppose $\sup_n \lambda_+[Var\epsilon)] < \infty$ and $\lim_{n\to\infty}\lambda_+[A_n] = 0$, $\ell^T\hat{\beta}$ is consistent in MSE.
[LM asymptotic normality] $\ell^T(\hat{\beta} - \beta)/\sqrt{Var(\ell^T\hat{\beta})} \xrightarrow{D} N(0,1)$. sufficient condition: $\lambda_+[A_n] \to 0$ and $Z_n^T A_n Z_n \to 0$ as $n \to \infty$ and there exist $\{a_n\}$ s.t. $a_n \to \infty$, $a_n/a_{n+1} \to 1$ and Z^TZ/a_n converge to positive definite matrix.
[LM Hypothesis testing] Under A1, $\ell \in R(Z)$, θ_0 fixed constant
[LM hypothesis testing - simple] $\ell \in R(Z)$, a $H_0 : \ell^T\beta \leq \theta_0$, $H_1 : \ell^T\beta > \theta_0$, b $H_0 : \ell^T\beta = \theta_0$, $H_1 : \ell^T\beta \neq \theta_0$, Under H_0 : $t(X) = \frac{\ell^T\hat{\beta} - \theta_0}{\sqrt{\ell^T(Z^TZ)^-\ell}\sqrt{SSR/(n-r)}} \sim t_{n-r}$, UMPU reject $t(X) > t_{n-r,\alpha}$ or $|t(X)| > t_{n-r,\alpha/2}$
[LM hypothesis testing - multiple] $L_{s\times p}$, $s \leq r$ and all rows $= \ell_j \in R(Z)$ a $H_0 : L\beta = 0$, $H_1 : L\beta \neq 0$ Under H_0: $W = \frac{(\|X - Z\hat{\beta}_0\|^2 - \|X - Z\hat{\beta}\|^2)/s}{\|X - Z\hat{\beta}\|^2/(n-r)} \sim F_{s,n-r}$ with non-central param $\sigma^{-2}\|Z\beta - \Pi_0 Z\beta\|^2$, reject $W > F_{s,n-r,1-\alpha}$
[LM confidence set] Pivotal qty: $\mathcal{R}(X, \beta) = \frac{(\hat{\beta} - \beta)^T Z^T Z(\hat{\beta} - \beta)/p}{\|X - Z\hat{\beta}\|^2/(n-p)} \sim F_{p,n-p}$, where $\hat{\beta}$ is LSE of β, $C(X) = \{\beta : \mathcal{R}(X, \beta) \leq F_{p,n-p,1-\alpha}\}$