

# Contents

<b>I</b>	<b>Calculus</b>	<b>1</b>
<b>1</b>	<b>Limit Theory</b>	<b>2</b>
1.1	Function . . . . .	2
<b>2</b>	<b>Differential Calculus</b>	<b>3</b>
<b>3</b>	<b>Integral Calculus</b>	<b>4</b>
<b>II</b>	<b>Real Analysis</b>	<b>5</b>
<b>4</b>	<b>Measure Theory</b>	<b>6</b>
4.1	Semi-algebras, Algebras and Sigma-algebras . . . . .	6
4.2	Measure . . . . .	8
<b>5</b>	<b>Lebesgue Integration</b>	<b>9</b>
5.1	Properties of the Integral . . . . .	9
5.2	Product Measures . . . . .	11
<b>III</b>	<b>Functional Analysis</b>	<b>12</b>
<b>IV</b>	<b>Matrix Theory</b>	<b>13</b>
<b>6</b>	<b>Matrix Norms</b>	<b>14</b>
6.1	Matrix Norms Induced by Vector Norms . . . . .	14
<b>7</b>	<b>Matrix Decompositions</b>	<b>15</b>
7.1	Spectral Decomposition . . . . .	15
7.2	Singular Value Decomposition . . . . .	16
7.2.1	Relationship to Matrix Norm . . . . .	17

<b>V</b>	<b>Convex Optimization</b>	<b>19</b>
<b>8</b>	<b>Convex Sets</b>	<b>20</b>
8.1	Affine and Convex Sets . . . . .	20
8.1.1	Affine Sets . . . . .	20
8.1.2	Convex Sets . . . . .	20
8.1.3	Cones . . . . .	21
8.2	Some Important Examples . . . . .	21
8.3	Generalized Inequalities . . . . .	22
8.3.1	Definition of Generalized Inequalities . . . . .	22
8.3.2	Properties of Generalized Inequalities . . . . .	23
<b>9</b>	<b>Convex Optimization Problems</b>	<b>24</b>
9.1	Generalized Inequality Constraints . . . . .	24
9.1.1	Conic Form Problems . . . . .	24
9.1.2	Semidefinite Programming . . . . .	24
9.2	Vector Optimization . . . . .	24
<b>10</b>	<b>Unconstrained Minimization</b>	<b>25</b>
10.1	Definition of Unconstrained Minimization . . . . .	25
10.2	General Descent Method . . . . .	27
10.3	Gradient Descent Method . . . . .	27
10.4	Steepest Descent Method . . . . .	27
10.5	Newton's Method . . . . .	27
<b>11</b>	<b>Exercises for Convex Optimization</b>	<b>30</b>
11.1	Convex Sets . . . . .	30
<b>VI</b>	<b>Probability Theory</b>	<b>31</b>
<b>12</b>	<b>Random Variables</b>	<b>32</b>
12.1	Probability Space . . . . .	32
12.2	Random Variables . . . . .	32
12.3	Distributions . . . . .	33
12.3.1	Definition of Distributions . . . . .	33
12.3.2	Properties of Distributions . . . . .	33
12.3.3	Families of Distributions . . . . .	34
12.4	Expected Value . . . . .	36
12.5	Independence . . . . .	36
12.5.1	Definition of Independence . . . . .	36
12.5.2	Sufficient Conditions for Independence . . . . .	37

12.5.3	Independence, Distribution, and Expectation . . . . .	37
12.5.4	Sums of Independent Random Variables . . . . .	38
12.6	Moments . . . . .	38
12.7	Characteristic Functions . . . . .	39
12.7.1	Definition of Characteristic Functions . . . . .	39
12.7.2	Properties of Characteristic Functions . . . . .	39
12.7.3	The Inversion Formula . . . . .	39
12.7.4	Moments and Derivatives . . . . .	40
<b>13</b>	<b>Convergence of Random Variables</b>	<b>41</b>
13.1	Modes of Convergence . . . . .	41
13.1.1	Convergence in Mean . . . . .	41
13.1.2	Convergence in Probability . . . . .	41
13.1.3	Convergence in Distribution . . . . .	42
13.1.4	Almost Sure Convergence . . . . .	45
13.1.5	Convergence in Uninform . . . . .	45
13.1.6	Asymptotic Notation . . . . .	46
13.2	Relationships of Modes . . . . .	47
<b>14</b>	<b>Law of Large Numbers</b>	<b>50</b>
14.1	Weak Law of Large Numbers . . . . .	50
14.2	Strong Law of Large Numbers . . . . .	52
14.2.1	Borel-Cantelli Lemmas . . . . .	52
14.2.2	Strong Law of Large Numbers . . . . .	52
14.3	Uniform Law of Large Numbers . . . . .	53
<b>15</b>	<b>Central Limit Theorems</b>	<b>55</b>
15.1	Central Limit Theorem . . . . .	55
15.1.1	The De Moivre-Laplace Theorem . . . . .	55
15.1.2	Central Limit Theorem . . . . .	57
15.1.3	Berry-Esseen Theorem . . . . .	57
15.2	CLT for independent non-identical Random Variables . . . . .	58
15.3	CLT for Dependent Random Variables . . . . .	58
<b>16</b>	<b>Multivariate Extensions</b>	<b>59</b>
16.1	Multivariate Distributions . . . . .	59
16.1.1	Multivariate Normal Distribution . . . . .	59
16.1.2	Wishart Distribution . . . . .	60
16.1.3	Hotelling's T-squared Distribution . . . . .	60
16.2	Convergence of Random Vectors . . . . .	60

<b>17 Exercises for Probability Theory and Examples</b>	<b>64</b>
17.1 Measure Theory . . . . .	64
17.2 Laws of Large Numbers . . . . .	65
17.3 Central Limit Theorems . . . . .	65
 <b>VII Stochastic Process</b>	 <b>67</b>
<b>18 Martingales</b>	<b>68</b>
18.1 Conditional Expectation . . . . .	68
18.2 Martingales . . . . .	68
18.3 Doob's Inequality . . . . .	70
18.4 Uniform Integrability . . . . .	70
18.5 Optional Stopping Theorems . . . . .	70
 <b>19 Markov Chains</b>	 <b>71</b>
19.1 Markov Chain . . . . .	71
19.2 Markov Properties . . . . .	73
19.3 Recurrence and Transience . . . . .	73
19.4 Stationary Measures . . . . .	75
19.5 Asymptotic Behavior . . . . .	75
19.6 Ergodic Theorems . . . . .	75
 <b>20 Brownian Motion</b>	 <b>76</b>
20.1 Markov Properties . . . . .	77
20.2 Martingales . . . . .	77
20.3 Sample Paths . . . . .	78
20.4 Itô Stochastic Calculus . . . . .	80
 <b>21 Exercises for Probability Theory and Examples</b>	 <b>83</b>
21.1 Martingales . . . . .	83
21.2 Markov Chains . . . . .	83
21.3 Ergodic Theorems . . . . .	83
21.4 Brownian Motion . . . . .	83
21.5 Applications to Random Walk . . . . .	83
21.6 Multidimensional Brownian Motion . . . . .	83
 <b>VIII Empirical Process</b>	 <b>84</b>
<b>22</b>	<b>85</b>
22.1 Concentration by Entropic Techniques . . . . .	85

22.2	Some Matrix Calculus and Covariance Estimation . . . . .	90
------	--	----

## **IX Random Matrix Theory 92**

### **23 Preliminary 93**

23.1	Empirical Spectral Measure . . . . .	93
23.2	Stieltjes Transform . . . . .	94
23.3	Matrix Equivalents . . . . .	96
23.4	Resolvent and Perturbation Identities . . . . .	97

### **24 Wigner Matrix 99**

### **25 Sample Covariance Matrix 100**

25.1	Eigenvalues and Singular Values . . . . .	101
25.2	Laguerre Orthogonal Ensemble . . . . .	101
25.3	Marčenko-Pastur Theorem . . . . .	106
25.4	Limits of Extreme Eigenvalues . . . . .	109

## **X Statistics Inference 111**

### **26 Statistical Theory 112**

26.1	Populations and Samples . . . . .	112
26.2	Statistics . . . . .	112
26.2.1	Sufficient Statistics . . . . .	112
26.2.2	Complete Statistics . . . . .	112
26.3	Estimators . . . . .	113

### **27 Point Estimation 115**

27.1	Maximum Likelihood Estimator . . . . .	115
27.1.1	Consistency . . . . .	115
27.1.2	Fisher Information . . . . .	117
27.1.3	Asymptotic Normality . . . . .	118
27.1.4	Efficiency . . . . .	119
27.2	Modified Likelihood Estimator . . . . .	119
27.2.1	Marginal Likelihood . . . . .	119
27.2.2	Conditional Likelihood . . . . .	119
27.2.3	Profile Likelihood . . . . .	121
27.2.4	Quasi Likelihood . . . . .	121
27.3	Minimum-Variance Unbiased Estimator . . . . .	121
27.4	Accuracy of Estimators . . . . .	123

<b>28 Interval Estimation</b>	<b>126</b>
28.1 Confidence Interval . . . . .	126
28.2 Pivot . . . . .	126
28.3 Likelihood Interval . . . . .	126
28.4 Prediction Interval . . . . .	126
28.5 Tolerance Interval . . . . .	126
<b>29 Testing Hypotheses</b>	<b>127</b>
29.1 Testing Hypotheses . . . . .	127
29.2 Parametric Tests . . . . .	127
29.3 Specific Tests . . . . .	127
29.3.1 Goodness of Fit . . . . .	127
29.3.2 Rank statistics . . . . .	127
<b>30 Bayesian Inference</b>	<b>128</b>
30.1 Bayes Estimator . . . . .	128
30.1.1 Single-Prior Bayes . . . . .	130
30.1.2 Hierarchical Bayes . . . . .	131
30.1.3 Empirical Bayes . . . . .	132
30.1.4 Bayes Prediction . . . . .	132
<b>31 Nonparametric Statistics</b>	<b>133</b>
31.1 Probability Distribution . . . . .	133
31.1.1 Cumulative Distribution Function . . . . .	133
31.1.2 Probability Density Function . . . . .	134
31.2 Kernel Methods . . . . .	137
31.2.1 Positive Definite Kernels . . . . .	137
<b>32 Minimax Theory</b>	<b>141</b>
32.1 Fano's Inequality . . . . .	141
32.2 Minimax Rate . . . . .	143
32.3 Applications . . . . .	145
<b>33 Multivariate Extensions</b>	<b>146</b>
33.1 Applications . . . . .	146
33.1.1 Mean vector . . . . .	146
33.1.2 Difference of two mean vectors . . . . .	147
33.1.3 Simple Linear Regression . . . . .	148
33.1.4 Multinomial One-Sample Test . . . . .	149
33.1.5 Contingency Table . . . . .	149

## **XI Computational Statistics 150**

### **34 Random Generator 151**

34.1 Uniform Random Number Generation . . . . . 151

34.2 Non-uniform Random Number Generation . . . . . 151

34.2.1 Inversion Method . . . . . 151

34.2.2 Rejection Sampling Method . . . . . 152

34.3 Markov Chain Monte Carlo . . . . . 153

34.3.1 Metropolis-Hastings Sampling . . . . . 153

34.3.2 Gibbs Sampling . . . . . 154

### **35 Monte Carlo Integration 155**

35.1 Monte Carlo Integration . . . . . 155

35.2 Importance Sampling . . . . . 155

### **36 Bootstrap 157**

36.1 Bootstrap Principle . . . . . 157

36.2 Standard Error Estimation . . . . . 158

36.3 Bias Estimation . . . . . 158

36.4 Confidence Interval Estimation . . . . . 158

36.5 Hypothesis Testing . . . . . 159

36.6 Jackknife . . . . . 159

## **XII Regression Analysis 161**

### **37 Linear Regression 162**

### **38 Generalized Linear Model 163**

38.1 Introduction . . . . . 163

38.2 Binary Data . . . . . 165

38.3 Polytomous Data . . . . . 165

38.4 Count Data . . . . . 167

### **39 Quantile Regression 168**

### **40 Survival Analysis 169**

40.1 General Formulation . . . . . 169

40.2 Estimation of Survival Function . . . . . 171

40.3 Proportional Hazards Model . . . . . 172

<b>41 Nonparametric Regression</b>	<b>173</b>
41.1 Uniform Stability of Regularized Kernel Model . . . . .	173
41.1.1 Introduction . . . . .	173
41.1.2 Some Notations and Concepts . . . . .	174
41.1.3 Uniform Stability of Regularized Kernel Model . . . . .	175
<b>42 High Dimensional Regression Analysis</b>	<b>181</b>
42.1 Lasso for Linear Regression . . . . .	181
42.1.1 Numerical Algorithms . . . . .	181
42.1.2 Selection of the Tuning Parameter . . . . .	182
42.2 Theory for the Lasso . . . . .	182
42.3 Other Lasso-Type Estimators . . . . .	184
42.3.1 Adaptive Lasso . . . . .	184
42.3.2 Elastic Net . . . . .	184
42.3.3 Group Lasso . . . . .	184
42.3.4 Fused Lasso . . . . .	184
42.4 Nonconvex Penalties . . . . .	184
42.4.1 SCAD . . . . .	185
42.4.2 MCP . . . . .	185
<b>XIII Statistics Applications</b>	<b>186</b>
<b>43 Missingness Data</b>	<b>187</b>
43.1 The Problem of Missing Data . . . . .	187
43.1.1 Missingness Mechanisms . . . . .	188
43.1.2 Commonly Used Methods for Missing Data . . . . .	189
43.2 Likelihood-Based Inference with Missing Data . . . . .	189
43.2.1 Ignorable Missingness Mechanism . . . . .	189
43.2.2 Expectation-Maximization Algorithm . . . . .	192
43.3 Missing Not At Random Models . . . . .	198
43.3.1 Normal Models for MNAR Missing Data . . . . .	199
<b>44 Treatment-effects Analysis</b>	<b>200</b>
44.1 Evaluations . . . . .	200
44.1.1 Average Treatment Effect . . . . .	200
44.1.2 Mann-Whitney Statistic . . . . .	200
44.1.3 Distribution-type Index . . . . .	200
<b>45 Graphical Lasso</b>	<b>201</b>



<b>46 Semi-supervised Learning</b>	<b>203</b>
46.1 Assumptions . . . . .	203
 <b>XIV Machine Learning</b>	 <b>204</b>
<b>47 Support Vector Machine</b>	<b>205</b>
<b>48 Linear Discriminant Analysis</b>	<b>207</b>
<b>49 K-Nearest Neighbor</b>	<b>208</b>
<b>50 Decision Tree</b>	<b>209</b>
<b>51 Kalman Filter</b>	<b>210</b>
 <b>XV Deep Learning</b>	 <b>211</b>
<b>52 Mixture of Experts</b>	<b>212</b>
 <b>XVI Generative Models</b>	 <b>213</b>
<b>53 Diffusion Model</b>	<b>214</b>
53.1 Introduction . . . . .	214
53.1.1 Denoising Diffusion Probabilistic Model . . . . .	214
53.2 Effort in Inference . . . . .	214

# Part I

## Calculus

# Chapter 1

## Limit Theory

### 1.1 Function

#### Definition 1.1.1 (Mapping)

Let  $X : \Omega_1 \rightarrow \Omega_2$  be a mapping.

1. For every subset  $B \in \Omega_2$ , the inverse image of B is

$$X^{-1}(B) = \{\omega : \omega \in \Omega_1, X(\omega) \in B\} := \{X \in B\}.$$

2. For every class

#### Definition 1.1.2 (Closed Function)

A function  $f : \mathbb{R}^n \rightarrow \mathbb{R}$  is said to be closed, if for each  $\alpha \in \mathbb{R}$ , the sublevel set

$$\{\mathbf{x} \in \text{dom } f \mid f(\mathbf{x}) \leq \alpha\} \tag{1.1}$$

is closed.

## Chapter 2

# Differential Calculus

## Chapter 3

# Integral Calculus

# Part II

## Real Analysis

# Chapter 4

## Measure Theory

### 4.1 Semi-algebras, Algebras and Sigma-algebras

#### Definition 4.1.1 (Semi-algebra)

A nonempty class  $\mathcal{S}$  of subsets of  $\Omega$  is an **semi-algebra** on  $\Omega$  that satisfy

1. if  $A, B \in \mathcal{S}$ , then  $A \cap B \in \mathcal{S}$ .
2. if  $A \in \mathcal{S}$ , then  $A^C$  is a finite disjoint union of sets in  $\mathcal{S}$ , i.e.,

$$A^C = \sum_{i=1}^n A_i, \text{ where } A_i \in \mathcal{S}, A_i \cap A_j = \emptyset, i \neq j.$$

#### Definition 4.1.2 (Algebra)

A nonempty class  $\mathcal{A}$  of subsets of  $\Omega$  is an **algebra** on  $\Omega$  that satisfy

1. if  $A \in \mathcal{A}$ , then  $A^C \in \mathcal{A}$ .
2. if  $A_1, A_2 \in \mathcal{A}$ , then  $A_1 \cup A_2 \in \mathcal{A}$ .

#### Definition 4.1.3 ( $\sigma$ -algebra)

A nonempty class  $\mathcal{F}$  of subsets of  $\Omega$  is a  **$\sigma$ -algebra** on  $\Omega$  that satisfy

1. if  $A \in \mathcal{F}$ , then  $A^C \in \mathcal{F}$ .
2. if  $A_i \in \mathcal{F}$  is a countable sequence of sets, then  $\cup_i A_i \in \mathcal{F}$ .

**Example** (Special  $\sigma$ -algebra). 1. **Trivial  $\sigma$ -algebra**  $:= \{\emptyset, \Omega\}$ . This is smallest  $\sigma$ -algebra.

2. **Power Set**  $:=$  all subsets of  $\Omega$ , denoted by  $\mathcal{P}(\Omega)$ . This is the largest  $\sigma$ -algebra.

3. **The smallest  $\sigma$ -algebra containing  $A \in \Omega$**   $:= \{\emptyset, A, A^C, \Omega\}$ .

It is easy to define (Lebesgue) measure on the semi-algebra  $\mathcal{S}$ , and then easily

to extend it to the algebra  $\overline{\mathcal{S}}$ , finally, we can extend it further to some  $\sigma$ -algebra (mostly consider the smallest one containing  $\mathcal{S}$ ).

**Lemma 4.1.1**

If  $\mathcal{S}$  is a semi-algebra, then

$$\overline{\mathcal{S}} = \{\text{finite disjoint unions of sets in } \mathcal{S}\}$$

is an algebra, denoted by  $\mathcal{A}(\mathcal{S})$ , called **the algebra generated by  $\mathcal{S}$** .

*Proof.* Let  $A, B \in \overline{\mathcal{S}}$ , then  $A = \sum_{i=1}^n A_i, B = \sum_{j=1}^m B_j$  with  $A_i, B_j \in \mathcal{S}$ .

**Intersection:** For  $A_i \cap B_j \in \mathcal{S}$  by the definition of semi-algebra  $\mathcal{S}$ , thus

$$A \cap B = \sum_{i=1}^n \sum_{j=1}^m A_i \cap B_j \in \overline{\mathcal{S}}.$$

So  $\overline{\mathcal{S}}$  is closed under (finite) intersection.

**Complement:** For DeMorgan's Law,  $A_i^C \in \mathcal{S}$  by the definition of semi-algebra  $\mathcal{S}$  and  $\overline{\mathcal{S}}$  closed under (finite) intersection that we just shown, thus

$$A^C = \left(\sum_{i=1}^n A_i\right)^C = \cap_{i=1}^n A_i^C \in \overline{\mathcal{S}}.$$

So  $\overline{\mathcal{S}}$  is closed under complement.

**Union:** For DeMorgan's Law and  $\overline{\mathcal{S}}$  closed under (finite) intersection and complement that we just shown, thus

$$A \cup B = (A^C \cap B^C)^C \in \overline{\mathcal{S}}.$$

So  $\overline{\mathcal{S}}$  is closed under (finite) union.

Hence,  $\overline{\mathcal{S}}$  is an algebra. □

**Theorem 4.1.1**

For any class  $\mathcal{A}$ , there exists a unique minimal  $\sigma$ -algebra containing  $\mathcal{A}$ , denoted by  $\sigma(\mathcal{A})$ , called **the  $\sigma$ -algebra generated by  $\mathcal{A}$** . In other words,

1.  $\mathcal{A} \subset \sigma(\mathcal{A})$ .
2. For any  $\sigma$ -algebra  $\mathcal{B}$  with  $\mathcal{A} \subset \mathcal{B}$ ,  $\sigma(\mathcal{A}) \subset \mathcal{B}$ .

and  $\sigma(\mathcal{A})$  is unique.

*Proof.* **Existence:**

**Uniqueness:** □

**Example** (Borel  $\sigma$ -algebras generated from semi-algebras). 1.



## 4.2 Measure

### Definition 4.2.1 (Measure)

**Measure** is a nonnegative countably additive set function, that is, a function  $\mu : \mathcal{A} \rightarrow \mathbb{R}$  with

1.  $\mu(A) \geq \mu(\emptyset) = 0$  for all  $A \in \mathcal{A}$ .
2. if  $A_i \in \mathcal{A}$  is a countable sequence of disjoint sets, then

$$\mu(\cup_i A_i) = \sum_i \mu(A_i).$$

### Definition 4.2.2 (Measure Space)

If  $\mu$  is a measure on a  $\sigma$ -algebra  $\mathcal{A}$  of subsets of  $\Omega$ , the triplet  $(\Omega, \mathcal{A}, \mu)$  is a **measure space**.

*Remark.* A measure space  $(\Omega, \mathcal{A}, \mu)$  is a **probability space**, if  $P(\Omega) = 1$ .

**Property.** Let  $\mu$  be a measure on a  $\sigma$ -algebra  $\mathcal{A}$

1. **monotonicity** if  $A \subset B$ , then  $\mu(A) \leq \mu(B)$ .
2. **subadditivity** if  $A \subset \cup_{m=1}^{\infty} A_m$ , then  $\mu(A) \leq \sum_{m=1}^{\infty} \mu(A_m)$ .
3. **continuity from below** if  $A_i \uparrow A$  (i.e.  $A_1 \subset A_2 \subset \dots$  and  $\cup_i A_i = A$ ), then  $\mu(A_i) \uparrow \mu(A)$ .
4. **continuity from above** if  $A_i \downarrow A$  (i.e.  $A_1 \supset A_2 \supset \dots$  and  $\cap_i A_i = A$ ), then  $\mu(A_i) \downarrow \mu(A)$ .

*Proof.*

□

# Chapter 5

## Lebesgue Integration

### 5.1 Properties of the Integral

#### Theorem 5.1.1 (Jensen's Inequality)

Let  $(\Omega, \mathcal{A}, \mu)$  be a probability space. If  $f$  is a real-valued function that is  $\mu$ -integrable, and if  $\varphi$  is a convex function on the real line, then:

$$\varphi\left(\int_{\Omega} f \, d\mu\right) \leq \int_{\Omega} \varphi(f) \, d\mu. \quad (5.1)$$

*Proof.* Let  $x_0 = \int_{\Omega} f \, d\mu$ . Since the existence of subderivatives for convex functions,  $\exists a, b \in \mathbb{R}$ , such that,

$$\forall x \in \mathbb{R}, \varphi(x) \geq ax + b \text{ and } ax_0 + b = \varphi(x_0).$$

Then, we got

$$\int_{\Omega} \varphi(f) \, d\mu \geq \int_{\Omega} af + b \, d\mu = a \int_{\Omega} f \, d\mu + b = ax_0 + b = \varphi\left(\int_{\Omega} f \, d\mu\right).$$

□

#### Theorem 5.1.2 (Hölder's Inequality)

Let  $(\Omega, \mathcal{F}, \mu)$  be a measure space and let  $p, q \in [1, \infty]$  with  $1/p + 1/q = 1$ . Then, for all measurable functions  $f$  and  $g$  on  $\Omega$ ,

$$\int_{\Omega} |f \cdot g| \, d\mu \leq \|f\|_p \|g\|_q. \quad (5.2)$$

*Proof.*

□

**Theorem 5.1.3 (Minkowski's Inequality)**

Let  $(\Omega, \mathcal{F}, \mu)$  be a measure space and let  $p \in [1, \infty]$ . Then, for all measurable functions  $f$  and  $g$  on  $\Omega$ ,

$$\|f + g\|_p \leq \|f\|_p + \|g\|_p. \quad (5.3)$$

*Proof.* Since  $\varphi(x) = x^p$  is a convex function for  $p \in [1, \infty)$ . By it's definition,

$$|f + g|^p = \left| 2 \cdot \frac{f}{2} + 2 \cdot \frac{g}{2} \right|^p \leq \frac{1}{2}|2f|^p + \frac{1}{2}|2g|^p = 2^{p-1}(|f|^p + |g|^p).$$

Therefore,

$$|f + g|^p < 2^{p-1}(|f|^p + |g|^p) < \infty.$$

By Hölder's Inequality (5.1.2),

$$\begin{aligned} \|f + g\|_p^p &= \int |f + g|^p d\mu \\ &= \int |f + g| \cdot |f + g|^{p-1} d\mu \\ &\leq \int (|f| + |g|) |f + g|^{p-1} d\mu \\ &= \int |f| |f + g|^{p-1} d\mu + \int |g| |f + g|^{p-1} d\mu \\ &\leq \left( \left( \int |f|^p d\mu \right)^{\frac{1}{p}} + \left( \int |g|^p d\mu \right)^{\frac{1}{p}} \right) \left( \int |f + g|^{(p-1)(\frac{p}{p-1})} d\mu \right)^{1-\frac{1}{p}} \\ &= (\|f\|_p + \|g\|_p) \frac{\|f + g\|_p^p}{\|f + g\|_p} \end{aligned}$$

which means, as  $p \in [1, \infty)$ ,

$$\|f + g\|_p \leq \|f\|_p + \|g\|_p.$$

When  $p = \infty$ ,

$a$

□

**Theorem 5.1.4 (Bounded Convergence Theorem)****Theorem 5.1.5 (Fatou's Lemma)****Theorem 5.1.6 (Monotone Convergence Theorem)**

## 5.2 Product Measures

Theorem 5.2.1 (Fubini's Theorem)

**Part III**

**Functional Analysis**

# **Part IV**

## **Matrix Theory**

# Chapter 6

## Matrix Norms

### 6.1 Matrix Norms Induced by Vector Norms

# Chapter 7

## Matrix Decompositions

### 7.1 Spectral Decomposition

#### Definition 7.1.1 (Eigenvectors and Eigenvalues)

A (non-zero) vector  $\mathbf{v}$  of dimension  $n$  is an **eigenvector** of a square  $n \times n$  matrix  $\mathbf{A}$ , if

$$\mathbf{A}\mathbf{v} = \lambda\mathbf{v} \quad (7.1)$$

where  $\lambda$  is a scalar, termed the **eigenvalue** corresponding to  $\mathbf{v}$ .

#### Definition 7.1.2 (Spectral Decomposition)

For any  $n \times n$  matrix with  $n$  linearly independent eigenvectors  $\mathbf{q}_i, i = 1, \dots, n$ . Then  $\mathbf{A}$  can be factorized as

$$\mathbf{A} = \mathbf{Q}\mathbf{\Lambda}\mathbf{Q}^{-1}$$

where  $\mathbf{Q}$  is the square  $n \times n$  matrix whose  $i$ -th column is the eigenvector  $\mathbf{q}_i$  of  $\mathbf{A}$ , and  $\mathbf{\Lambda}$  is the diagonal matrix whose diagonal elements are the corresponding eigenvalues,  $\mathbf{\Lambda} = \lambda_i$ . This factorization is called eigendecomposition or sometimes spectral decomposition.

**Example** (Real Symmetric Matrices). As a special case, for every  $n \times n$  real symmetric matrix, the eigenvalues are real and the eigenvectors can be chosen as real and orthonormal. Thus a real symmetric matrix  $\mathbf{A}$  can be decomposed as

$$\mathbf{A} = \mathbf{Q}\mathbf{\Lambda}\mathbf{Q}' \quad (7.2)$$

where  $\mathbf{Q}$  is an orthogonal matrix whose columns are eigenvectors of  $\mathbf{A}$ , and  $\mathbf{\Lambda}$  is a diagonal matrix whose entries are the eigenvalues of  $\mathbf{A}$ .



## 7.2 Singular Value Decomposition

### Definition 7.2.1 (Singular Value Decomposition)

For any matrix  $\mathbf{A} \in \mathbb{R}^{m \times n}$ , we have

$$\mathbf{A} = \mathbf{U}\mathbf{\Sigma}\mathbf{V}' \quad (7.3)$$

where

- $\mathbf{U} \in \mathbb{R}^{m \times m}$  is an orthogonal matrix whose columns are the eigenvectors of  $\mathbf{A}\mathbf{A}'$
- $\mathbf{V} \in \mathbb{R}^{n \times n}$  is an orthogonal matrix whose columns are the eigenvectors of  $\mathbf{A}'\mathbf{A}$
- $\mathbf{\Sigma} \in \mathbb{R}^{m \times n}$  is an all-zero matrix except for the first  $r$  diagonal elements

$$\sigma_i = \Sigma_{ii}, \quad i = 1, 2, \dots, r$$

which are called singular values, which are the square roots of the eigenvalues of  $\mathbf{A}'\mathbf{A}$  and of  $\mathbf{A}\mathbf{A}'$  (these two matrices have the same eigenvalues)

*Remark.* We assume above that the singular values are sorted in descending order and the eigenvectors are sorted according to descending order of their eigenvalues.

*Proof.* Without loss of generality, we assume  $m \geq n$ . Since for the case  $n > m$ , can then be established by transposing the SVD of  $\mathbf{A}'$ ,

$$\mathbf{A} = (\mathbf{A}')' = (\mathbf{U}'\mathbf{\Sigma}\mathbf{V})' = \mathbf{V}'(\mathbf{U}'\mathbf{\Sigma})' = \mathbf{V}'\mathbf{\Sigma}\mathbf{U}$$

For  $m \geq n$ , suppose  $\text{rank}(\mathbf{A}) = r$ , and then  $\text{rank}(\mathbf{A}'\mathbf{A}) = r$  and the spectral decomposition of  $\mathbf{A}'\mathbf{A}$  be

$$\mathbf{A}'\mathbf{A}\mathbf{V} = \mathbf{V} \text{diag}(\sigma_1^2, \dots, \sigma_r^2, 0, \dots, 0)$$

where  $\sigma_i^2$  are the eigenvalues of  $\mathbf{A}'\mathbf{A}$  and the columns of  $\mathbf{V}$ , denoted  $\mathbf{v}^{(i)}$ , are the corresponding orthonormal eigenvectors.

Let

$$\mathbf{u}^{(i)} = \frac{\mathbf{A}\mathbf{v}^{(i)}}{\sigma_i}$$

then

$$\begin{aligned} \mathbf{A}'\mathbf{u}^{(i)} &= \frac{\mathbf{A}'\mathbf{A}\mathbf{v}^{(i)}}{\sigma_i} = \sigma_i \mathbf{v}^{(i)} \Rightarrow \\ \mathbf{A}\mathbf{A}'\mathbf{u}^{(i)} &= \sigma_i \mathbf{A}\mathbf{v}^{(i)} = \sigma_i^2 \mathbf{u}^{(i)} \end{aligned}$$

implying that  $\mathbf{u}^{(i)}$  are eigenvectors of  $\mathbf{A}\mathbf{A}'$  corresponding to eigenvalues  $\sigma_i^2$ .

Since the eigenvectors  $\mathbf{v}^{(i)}$  are orthonormal, then so are the eigenvectors  $\mathbf{u}^{(i)}$

$$(\mathbf{u}^{(i)})' \mathbf{u}^{(j)} = \frac{(\mathbf{v}^{(i)})' \mathbf{A}' \mathbf{A} \mathbf{v}^{(j)}}{\sigma_i^2} = (\mathbf{v}^{(i)})' \mathbf{v}^{(j)} = \begin{cases} 1 & i = j \\ 0 & i \neq j \end{cases}$$

We have thus far a matrix  $\mathbf{V}$  whose columns are eigenvectors of  $\mathbf{A}'\mathbf{A}$  with eigenvalues  $\sigma_i^2$ , and a matrix  $\mathbf{U}$  whose columns are  $r$  eigenvectors of  $\mathbf{A}\mathbf{A}'$  corresponding to eigenvalues  $\sigma_i^2$ .

We augment the eigenvectors  $\mathbf{u}^{(i)}, i = 1, \dots, r$  with orthonormal vectors  $\mathbf{u}^{(i)}, i = r+1, \dots, m$  that span  $\text{null}(\mathbf{A}\mathbf{A}')$ , and together  $\mathbf{u}^{(i)}, i = 1, \dots, n$  are a full orthonormal set of eigenvectors of  $\mathbf{A}\mathbf{A}'$  with eigenvalues  $\sigma_i^2$  (with  $\sigma_i = 0$  for  $i > r$ ).

Since

$$[\mathbf{U}'\mathbf{A}\mathbf{V}]_{ij} = (\mathbf{u}^{(i)})' \mathbf{A} \mathbf{v}^{(j)} = \begin{cases} \sigma_j (\mathbf{u}^{(i)})' \mathbf{u}^{(j)} & i \leq r \\ 0 & i > r \end{cases}$$

we get

$$\mathbf{U}'\mathbf{A}\mathbf{V} = \mathbf{\Sigma}$$

where

$$\mathbf{\Sigma} = \begin{pmatrix} \text{diag}(\sigma_1, \dots, \sigma_n) \\ \mathbf{0} \end{pmatrix}, \quad \sigma_i = 0 \forall r < i \leq n$$

Consequently, we get the desired decompositions

$$\mathbf{A} = \mathbf{U}\mathbf{\Sigma}\mathbf{V}'$$

□

### 7.2.1 Relationship to Matrix Norm

#### Theorem 7.2.1

For any matrix  $\mathbf{A} \in \mathbb{R}^{m \times n}$ ,

$$\|\mathbf{A}\|_2 = \sigma_{\max}(\mathbf{A}) \quad (7.4)$$

*Proof.* For any matrix  $\mathbf{A} \in \mathbb{R}^{m \times n}$ , the SVD implies that,

$$\|\mathbf{A}\|_2 = \sup_{\mathbf{x} \neq 0} \frac{\|\mathbf{A}\mathbf{x}\|_2}{\|\mathbf{x}\|_2} = \sup_{\mathbf{x} \neq 0} \frac{\|\mathbf{U}\mathbf{\Sigma}\mathbf{V}'\mathbf{x}\|_2}{\|\mathbf{x}\|_2}$$

Since  $\mathbf{U}$  is unitary, that is,

$$\|\mathbf{U}\mathbf{x}\|_2^2 = \mathbf{x}^\top \mathbf{U}'\mathbf{U}\mathbf{x} = \|\mathbf{x}\|_2^2, \quad \forall \mathbf{x} \in \mathbb{R}^m$$

thus,

$$= \sup_{\mathbf{x} \neq 0} \frac{\|\Sigma \mathbf{V}' \mathbf{x}\|_2}{\|\mathbf{x}\|_2}$$

Let  $\mathbf{y} = \mathbf{V}' \mathbf{x}$ , and since  $\mathbf{V}$  is unitary, we have

$$\|\mathbf{y}\|_2 = \|\mathbf{V}' \mathbf{x}\|_2 = \|\mathbf{x}\|_2 = 1$$

thus,

$$= \sup_{\mathbf{y} \neq 0} \frac{\|\Sigma \mathbf{y}\|_2}{\|\mathbf{V} \mathbf{y}\|_2} = \sup_{\mathbf{y} \neq 0} \frac{\left( \sum_{i=1}^r \sigma_i^2 |y_i|^2 \right)^{\frac{1}{2}}}{\left( \sum_{i=1}^r |y_i|^2 \right)^{\frac{1}{2}}} \leq \sigma_{\max}(\mathbf{A})$$

which takes "=", if  $\mathbf{y} = (1, 0, \dots, 0)'$ . □

### Theorem 7.2.2

For any matrix  $\mathbf{A} \in \mathbb{R}^{m \times n}$ , suppose  $\text{rank}(\mathbf{A}) = n$ , then

$$\min_{\|\mathbf{x}\|_2=1} \|\mathbf{A} \mathbf{x}\|_2 = \sigma_n(\mathbf{A}) \tag{7.5}$$

*Proof.* The proof process is analogous to the above theorem. □

*Remark.* If  $\text{rank}(\mathbf{A}) < n$ , then there is an  $\mathbf{x}$  such that the minimum is zero.

# Part V

## Convex Optimization

# Chapter 8

## Convex Sets

### 8.1 Affine and Convex Sets

#### 8.1.1 Affine Sets

##### Definition 8.1.1 (Affine Set)

A nonempty set  $C$  is said to be **affine set**, if

$$\forall x_1, x_2 \in C, \theta \in \mathbb{R}, \theta x_1 + (1 - \theta)x_2 \in C.$$

#### 8.1.2 Convex Sets

##### Definition 8.1.2 (Convex Set)

A nonempty set  $C$  is said to be **convex set**, if

$$\forall x_1, x_2 \in C, \theta \in [0, 1], \theta x_1 + (1 - \theta)x_2 \in C.$$

##### Definition 8.1.3 (Convex Hull)

The **convex hull** of said to be set  $C$ , denoted by  $\text{conv } C$  is a set of all convex combinations of points in  $C$ ,

$$\text{conv } C = \{\theta_1 x_1 + \cdots + \theta_k x_k | x_i \in C; \theta_i \geq 0, i = 1, \dots, k; \theta_1 + \cdots + \theta_k = 1\}.$$

*Remark.* The convex hull  $\text{conv } C$  is always convex, which is the minimal convex set that contains  $C$ .

### 8.1.3 Cones

#### Definition 8.1.4 (Cone)

A nonempty set  $C$  is said to be **cone**, if

$$\forall x \in C, \theta \geq 0, \theta x \in C.$$

#### Definition 8.1.5 (Convex Cone)

A nonempty set  $C$  is said to be **convex cone**, if

$$\forall x_1, x_2 \in C, \theta_1, \theta_2 \geq 0, \theta_1 x_1 + \theta_2 x_2 \in C.$$

## 8.2 Some Important Examples

#### Definition 8.2.1 (Hyperplane)

A hyperplane is defined to be

$$\{x | a^\top x = b\},$$

where  $a \in \mathbb{R}^n, a \neq 0, b \in \mathbb{R}$ .

#### Definition 8.2.2 (Halfspace)

A hyperplane is defined to be

$$\{x | a^\top x \leq b\},$$

where  $a \in \mathbb{R}^n, a \neq 0, b \in \mathbb{R}$ .

#### Definition 8.2.3 ((Euclidean) Ball)

A (Euclidean) ball in  $\mathbb{R}^n$  with center  $x_c$  and radius  $r$  is defined to be

$$B(x_c, r) = \{x | \|x - x_c\|_2 \leq r\} = \{x_c + ru | \|u\|_2 \leq 1\},$$

where  $r > 0$ .

**Definition 8.2.4 (Ellipsoid)**

A Ellipsoid in  $\mathbb{R}^n$  with center  $x_c$  is defined to be

$$\mathcal{E} = \{x | (x - x_c)^\top P^{-1}(x - x_c) \leq 1\} = \{x_c + Au | \|u\| \leq 1\},$$

where  $P \in \mathbb{S}_{++}^n$  (symmetric positive definite).

## 8.3 Generalized Inequalities

### 8.3.1 Definition of Generalized Inequalities

**Definition 8.3.1 (Proper Cone)**

A cone  $K \subseteq \mathbb{R}^n$  is said to be a proper cone, if

- $K$  is convex.
- $K$  is closed.
- $K$  is solid (nonempty interior).
- $K$  is pointed (contains no line).

**Definition 8.3.2 (Generalized Inequalities)**

The partial ordering on  $\mathbb{R}^n$  defined by proper cone  $K$ , if

$$y - x \in K, \tag{8.1}$$

which can be denoted by

$$x \preceq_K y \text{ or } y \succeq_K x. \tag{8.2}$$

The strict partial ordering on  $\mathbb{R}^n$  defined by proper cone  $K$ , if

$$y - x \in \text{int } K, \tag{8.3}$$

which can be denoted by

$$x \prec_K y \text{ or } y \succ_K x. \tag{8.4}$$

*Remark.* When  $K = \mathbb{R}_+$ , the partial ordering  $\preceq_K$  is the usual ordering  $\leq$  on  $\mathbb{R}$ , and the strict partial ordering  $\prec_K$  is the usual strict ordering  $<$  on  $\mathbb{R}$ .

### 8.3.2 Properties of Generalized Inequalities

**Theorem 8.3.1 (Properties of Generalized Inequalities)**

A generalized inequality  $\preceq_K$  has the following properties:

- Preserved under addition:
- Transitive:
- Preserved under nonnegative scaling:
- Reflexive:
- Antisymmetric:
- Preserved under limits:

A strict generalized inequality  $\prec_K$  has the following properties:



# Chapter 9

## Convex Optimization Problems

### 9.1 Generalized Inequality Constraints

#### Definition 9.1.1 (With Generalized Inequality Constraints)

A convex optimization problem with generalized inequality constraints is defined to be

$$\begin{aligned} \min_x \quad & f_0(x) \\ \text{s.t.} \quad & f_i(x) \preceq_{K_i} 0, \quad i = 1, \dots, m \\ & Ax = b \end{aligned} \tag{9.1}$$

where  $f_0 : \mathbb{R}^n \rightarrow \mathbb{R}$ ,  $K_i \in \mathbb{R}^{k_i}$  are proper convexes, and  $f_i : \mathbb{R}^n \rightarrow \mathbb{R}^{k_i}$  are  $K_i$ -convex.

#### 9.1.1 Conic Form Problems

##### Definition 9.1.2 (Conic Form Problem)

A conic form problem is defined to be

$$\begin{aligned} \min \quad & c^T x \\ \text{s.t.} \quad & Fx + g \preceq_K 0 \\ & Ax = b \end{aligned} \tag{9.2}$$

#### 9.1.2 Semidefinite Programming

### 9.2 Vector Optimization

# Chapter 10

## Unconstrained Minimization

### 10.1 Definition of Unconstrained Minimization

#### Definition 10.1.1 (Unconstrained Minimization Problem)

The unconstrained minimization problem is defined to be

$$\min_{\mathbf{x}} f(\mathbf{x}) \quad (10.1)$$

where  $f : \mathbb{R}^n \rightarrow \mathbb{R}$  is convex and twice continuously differentiable.

Assume the problem is solvable, i.e., there exists an optimal point  $\mathbf{x}^*$ , such that,

$$f(\mathbf{x}^*) = \inf_{\mathbf{x}} f(\mathbf{x})$$

and denote it by  $p^*$ . Since  $f$  is differentiable and convex, the point  $\mathbf{x}^*$  is optimal, if and only if

$$\nabla f(\mathbf{x}^*) = 0 \quad (10.2)$$

Solving (10.1) is equal to finding the solution of (10.2), thus (10.1) can be solved by analytic solution of (10.2) in a few cases, but usually can be solved by an iterative algorithm, i.e.,

$$\exists \mathbf{x}^{(0)}, \mathbf{x}^{(1)}, \dots \in \text{dom } f, \text{ s.t. } f(\mathbf{x}^{(k)}) \rightarrow p^*, \text{ as } k \rightarrow \infty$$

This algorithm is terminated when  $f(\mathbf{x}^{(k)}) - p^* \leq \epsilon$ , where  $\epsilon > 0$  is some specified tolerance.

*Remark.* The initial point  $\mathbf{x}^{(0)}$  must lie in  $\text{dom } f$ , and the sublevel set

$$S = \{\mathbf{x} \in \text{dom } f \mid f(\mathbf{x}) \leq f(\mathbf{x}^{(0)})\}$$

must be closed. Any closed function (Definition 1.1.2)

**Example** (Quadratic Minimization). The general convex quadratic minimization problem has the form

$$\min_{\mathbf{x}} \frac{1}{2} \mathbf{x}^\top \mathbf{P} \mathbf{x} + \mathbf{q}' \mathbf{x} + r \quad (10.3)$$

where  $\mathbf{P} \in \mathbb{S}_+^n$ ,  $\mathbf{q} \in \mathbb{R}^n$ , and  $r \in \mathbb{R}$ . The optimality condition is

$$\mathbf{P} \mathbf{x}^* + \mathbf{q} = \mathbf{0} \quad (10.4)$$

which is a set of linear equations.

1. If  $\mathbf{P} \succ 0$ , exists a unique solution  $\mathbf{x}^* = -\mathbf{P}^{-1} \mathbf{q}$ .
2. If  $\mathbf{P}$  is not positive definite, any solution of (10.4) is optimal for (10.3).
3. If (10.4) does not have a solution, then (10.3) is unbounded.

*Proof.*

1. Obviously.
2. Since  $\mathbf{P} \not\succ 0$ , i.e.,

$$\exists \mathbf{v}, \text{ s.t. } \mathbf{v}' \mathbf{P} \mathbf{v} < 0$$

Let  $\mathbf{x} = t \mathbf{v}$ , we have

$$f(\mathbf{x}) = t^2 (\mathbf{v}' \mathbf{P} \mathbf{v} / 2) + t (\mathbf{q}' \mathbf{v}) + r$$

which converges to  $-\infty$  as  $t \rightarrow \infty$ .

3. Since (10.4) does not have a solution, i.e.,

$$\mathbf{q} \notin \mathcal{R}(\mathbf{P})$$

Let

$$\mathbf{q} = \tilde{\mathbf{q}} + \mathbf{v}$$

where  $\tilde{\mathbf{q}}$  is the Euclidean projection of  $\mathbf{q}$  onto  $\mathcal{R}(\mathbf{P})$ , and  $\mathbf{v} = \mathbf{q} - \tilde{\mathbf{q}}$ . And  $\mathbf{v}$  is nonzero and orthogonal to  $\mathcal{R}(\mathbf{P})$ , i.e.,  $\mathbf{v}' \mathbf{P} \mathbf{v} = 0$ . If we take  $\mathbf{x} = t \mathbf{v}$ , we have

$$f(\mathbf{x}) = t \mathbf{q}' \mathbf{v} + r = t (\tilde{\mathbf{q}} + \mathbf{v})' \mathbf{v} + r = t (\mathbf{v}' \mathbf{v}) + r$$

which is unbounded below. □

*Remark.* The least-squares problem is a special case of quadratic minimization, that,

$$\min_{\mathbf{x}} \|\mathbf{A} \mathbf{x} - \mathbf{b}\|_2^2 = \mathbf{x}^\top (\mathbf{A}' \mathbf{A}) \mathbf{x} - 2 (\mathbf{A}' \mathbf{b})' \mathbf{x} + \mathbf{b}' \mathbf{b}$$

The optimality condition is

$$\mathbf{A}' \mathbf{A} \mathbf{x}^* = \mathbf{A}' \mathbf{b}$$

are called the normal equations of the least-squares problem.

**Example** (Unconstrained Geometric Programming). The unconstrained geometric program in convex form

$$\min_{\mathbf{x}} f(\mathbf{x}) = \log \left( \sum_{i=1}^m \exp(\mathbf{a}'_i \mathbf{x} + b_i) \right)$$

The optimality condition is

$$\nabla f(x^*) = \frac{\sum_{i=1}^m \exp(\mathbf{a}'_i \mathbf{x}^* + b_i) \mathbf{a}_i}{\sum_{j=1}^m \exp(\mathbf{a}'_j \mathbf{x}^* + b_j)} = \mathbf{0}$$

which has no analytical solution, so we must resort to an iterative algorithm. For this problem,  $\text{dom } f = \mathbb{R}^n$ , so any point can be chosen as the initial point  $\mathbf{x}^{(0)}$ .

**Example** (Analytic Center of Linear Inequalities). Consider the optimization problem

$$\min_{\mathbf{x}} f(x) = - \sum_{i=1}^m \log(\mathbf{b}_i - \mathbf{a}_i^T \mathbf{x})$$

where the domain of  $f$  is the open set

$$\text{dom } f = \{\mathbf{x} \mid \mathbf{a}'_i \mathbf{x} < \mathbf{b}_i, i = 1, \dots, m\}$$

#### Definition 10.1.2 (Strong Convexity)

## 10.2 General Descent Method

## 10.3 Gradient Descent Method

## 10.4 Steepest Descent Method

## 10.5 Newton's Method

(Smoothness Hessian) Suppose the Hessians of  $f$  are Lipschitz continuous, i.e.,

$$\|\nabla^2 f(\mathbf{x}) - \nabla^2 f(\mathbf{y})\|_2 \leq L \|\mathbf{x} - \mathbf{y}\|_2 \quad (10.5)$$

**Algorithm 1:** Damped Newton Method**Input:** Initial point  $\mathbf{x}_0 \in \text{dom} f$ , tolerance  $\epsilon > 0$ **Output:****1 repeat****2**     Compute the Newton step and decrement

$$\Delta \mathbf{x}_{\text{nt}} := -\nabla^2$$

**3 until;****Theorem 10.5.1**

Under the condition , there exist  $0 < \eta < m^2/L$  and  $\gamma > 0$ , for the damped Newton method, we have

- If  $\|\nabla^2 f(\mathbf{x}^{(k)})\| \geq \eta$ , then

$$f(\mathbf{x}^{(k+1)}) - f(\mathbf{x}^{(k)}) \leq -\gamma.$$

- If  $\|\nabla^2 f(\mathbf{x}^{(k)})\| < \eta$ , then the backtracking line search select  $t^{(k)} = 1$ , and

$$\frac{L}{2m^2} \|\nabla f(\mathbf{x}^{(k+1)})\|_2 \leq \left( \frac{L}{2m^2} \|\nabla f(\mathbf{x}^{(k)})\|_2 \right)^2.$$

Method	Descent Direction	Step Length	Features
Steepest			
Steepest (MG)			
Steepest (CD)			
Steepest (BB)			
Newton			
Newton (LM)			
Newton (Mixed)			
Quasi-Newton (SR1)			
Quasi-Newton (DFP)			
Quasi-Newton (BFGS)			
Quasi-Newton (LBFGS)			

**Example** (Extended Rosenbrock Function).

$$\min_{\mathbf{x}} f(\mathbf{x}) = \sum_{i=1}^n r_i^2(\mathbf{x}) \quad (10.6)$$

where  $n$  is even, and

$$r_i(\mathbf{x}) = \begin{cases} 10(x_{2k} - x_{2k-1}^2), & i = 2k - 1 \\ 1 - x_{2k-1}, & i = 2k \end{cases} \quad (10.7)$$

The minimum point is  $\mathbf{x}^* = (1, 1, \dots, 1)'$ , the initial point is  $\mathbf{x}_0 = (-1.2, 1, \dots, -1.2, 1)'$ .

# Chapter 11

## Exercises for Convex Optimization

### 11.1 Convex Sets

**Exercise.** Solution set of a quadratic inequality Let  $C \subseteq \mathbb{R}^n$  be the solution set of a quadratic inequality,

$$C = \{x \in \mathbb{R}^n | x^T A x + b^T x + c \leq 0\}$$

with  $A \in \mathbb{S}^n$ ,  $b \in \mathbb{R}^n$ , and  $c \in \mathbb{R}$ .

1. Show that  $C$  is convex if  $A \succeq 0$ .

*Proof.* 1. We have to show that  $\theta x + (1 - \theta)y \in C$  for all  $\theta \in [0, 1]$  and  $x, y \in C$ .

$$\begin{aligned} & (\theta x + (1 - \theta)y)^T A (\theta x + (1 - \theta)y) + b^T (\theta x + (1 - \theta)y) + c \\ &= \theta^2 x^T A x + \theta(1 - \theta)(y^T A x + x^T A y) + (1 - \theta)^2 y^T A y + \theta b^T x + (1 - \theta)b^T y + c \\ &= \theta^2(x^T A x + b^T x + c) + (1 - \theta)^2(y^T A y + b^T y + c) - \theta^2(b^T x + c) \\ & \quad - (1 - \theta)^2(b^T y + c) + \theta(1 - \theta)(y^T A x + x^T A y) + \theta b^T x + (1 - \theta)b^T y + c \\ &\leq -\theta^2(b^T x + c) - (1 - \theta)^2(b^T y + c) + \theta(1 - \theta)(y^T A x + x^T A y) \\ & \quad + \theta b^T x + (1 - \theta)b^T y + c \\ &= \theta(1 - \theta)[(b^T x + c) + (b^T y + c) + x^T A x + y^T A y] \\ &\leq \theta(1 - \theta)(-x^T A x - y^T A y + x^T A x + y^T A y) \leq 0 \end{aligned}$$

Therefore,  $\theta x + (1 - \theta)y \in C$ , which shows that  $C$  is convex if  $A \succeq 0$ .

□

# **Part VI**

## **Probability Theory**



# Chapter 12

## Random Variables

### 12.1 Probability Space

#### Definition 12.1.1 (Probability Space)

A probability space is a triple  $(\Omega, \mathcal{F}, P)$  consisting of:

1. the sample space  $\Omega$ : an arbitrary non-empty set.
2. the  $\sigma$ -algebra  $\mathcal{F} \subseteq 2^\Omega$ : a set of subsets of  $\Omega$ , called events.
3. the probability measure  $P : \mathcal{F} \rightarrow [0, 1]$ : a function on  $\mathcal{F}$  which is a measure function.

### 12.2 Random Variables

#### Definition 12.2.1 (Random Variable)

A random variable is a measurable function  $X : \Omega \rightarrow S$  from a set of possible outcomes  $(\Omega, \mathcal{F})$  to a measurable space  $(S, \mathcal{S})$ , that is,

$$X^{-1}(B) \equiv \{\omega : X(\omega) \in B\} \in \mathcal{F} \quad \forall B \in \mathcal{S}. \quad (12.1)$$

Typically,  $(S, \mathcal{S}) = (R^d, \mathcal{R}^d)$  ( $d > 1$ ).

How to prove that functions are measurable?

#### Theorem 12.2.1

If  $\{\omega : X(\omega) \in A\} \in \mathcal{F}$  for all  $A \in \mathcal{A}$  and  $\mathcal{A}$  generates  $\mathcal{S}$ , then  $X$  is measurable.

- 1.

## 12.3 Distributions

### 12.3.1 Definition of Distributions

#### Definition 12.3.1 (Distribution)

A distribution of random variable  $X$  is a probability function  $P : \mathcal{R} \rightarrow \mathbb{R}$  by setting

$$\mu(A) = P(X \in A) = P\left(X^{-1}(A)\right), \quad \text{for } A \in \mathcal{R}. \quad (12.2)$$

#### Definition 12.3.2 (Distribution Function)

The distribution of a random variable  $X$  is usually described by giving its **distribution function**,

$$F(x) = P(X \leq x). \quad (12.3)$$

#### Definition 12.3.3 (Density Function)

If the distribution function  $F(x) = P(X \leq x)$  has the form

$$F(x) = \int_{-\infty}^x f(y) dy,$$

that  $X$  has density function  $f$ .

### 12.3.2 Properties of Distributions

#### Theorem 12.3.1 (Properties of Distribution Function)

Any distribution function  $F$  has the following properties,

1.  $F$  is nondecreasing.
2.  $\lim_{x \rightarrow \infty} F(x) = 1, \lim_{x \rightarrow -\infty} F(x) = 0$ .
3.  $F$  is right continuous, i.e.,  $\lim_{y \downarrow x} F(y) = F(x)$ .
4. If  $F(x-) = \lim_{y \uparrow x} F(y)$ , then  $F(x-) = P(X < x)$ .
5.  $P(X = x) = F(x) - F(x-)$ .

*Proof.*

□

#### Theorem 12.3.2

If  $F$  satisfies (1), (2), and (3) in Theorem 12.3.1, then it is the distribution function of some random variable.

*Proof.*

□

**Theorem 12.3.3**

A distribution function has at most countably many discontinuities

*Proof.*

□

**12.3.3 Families of Distributions****Exponential Family****Definition 12.3.4 (Exponential Family)**

An exponential family of probability distributions is those distributions whose density is defined to be

$$f(y | \theta, \phi) = \exp \left[ \frac{y\theta - b(\theta)}{a(\phi)} + c(y, \phi) \right] \quad (12.4)$$

**Property.** The exponential family has the following properties,

$$E(Y) = b'(\theta) \quad \text{Var}(Y) = b''(\theta)a(\phi).$$

*Proof.*

□

**Table 12.1:** Common Distributions of Exponential Family

Distribution	Parameter(s)	$\theta$	$\phi$	$b(\theta)$	$a(\phi)$	$c(y, \phi)$	$E(Y)$	$\text{Var}(Y)$
Normal	$N(\mu, \sigma^2)$	$\mu$	$\sigma^2$	$\frac{\theta^2}{2}$	$\phi$	$-\frac{1}{2} \left[ \frac{y^2}{\phi} + \log(2\pi\phi) \right]$	$\theta$	$\phi$
Bernoulli	$\text{Bern}(p)$	$\log\left(\frac{p}{1-p}\right)$	1	$\log(1 + e^\theta)$	1	0	$\frac{e^\theta}{1+e^\theta}$	$\frac{e^\theta}{(1+e^\theta)^2}$
Poisson	$P(\mu)$	$\log(\mu)$	1	$e^\theta$	1	$-\log(y!)$	$e^\theta$	$e^\theta$
Gamma	$\Gamma(\alpha, \beta)$	$\log\left(\frac{\alpha}{\beta}\right)$	1	$-\log(-\theta)$	1	$-\log(\Gamma(\alpha)) + (\alpha - 1)\log(y) - y$	$\frac{\alpha}{\beta}$	$\frac{\alpha}{\beta^2}$

## 12.4 Expected Value

### Definition 12.4.1 (Expectation)

### Theorem 12.4.1 (Bounded Convergence theorem)

### Theorem 12.4.2 (Fatou's Lemma)

If  $X_n \geq 0$ , then

$$\liminf_{n \rightarrow \infty} EX_n \geq E \left( \liminf_{n \rightarrow \infty} X_n \right). \quad (12.5)$$

### Theorem 12.4.3 (Monotone Convergence theorem)

If  $0 \leq X_n \uparrow X$ , then

$$EX_n \uparrow EX. \quad (12.6)$$

### Theorem 12.4.4 (Dominated Convergence theorem)

If  $X_n \rightarrow X$  a.s.,  $|X_n| \leq Y$  for all  $n$ , and  $EY < \infty$ , then

$$EX_n \rightarrow EX. \quad (12.7)$$

## 12.5 Independence

### 12.5.1 Definition of Independence

#### Definition 12.5.1 (Independence)

1. Two events  $A$  and  $B$  are independent if  $P(A \cap B) = P(A)P(B)$ .
2. Two random variables  $X$  and  $Y$  are independent if for all  $C, D \in \mathcal{R}$

$$P(X \in C, Y \in D) = P(X \in C)P(Y \in D). \quad (12.8)$$

3. Two  $\sigma$ -fields  $\mathcal{F}$  and  $\mathcal{G}$  are independent if for all  $A \in \mathcal{F}$  and  $B \in \mathcal{G}$  the events  $A$  and  $B$  are independent.

The second definition is a special case of the third.

**Theorem 12.5.1**

1. If  $X$  and  $Y$  are independent then  $\sigma(X)$  and  $\sigma(Y)$  are independent.
2. Conversely, if  $\mathcal{F}$  and  $\mathcal{G}$  are independent,  $X \in \mathcal{F}$  and  $Y \in \mathcal{G}$ , then  $X$  and  $Y$  are independent.

The first definition is, in turn, a special case of the second.

**Theorem 12.5.2**

1. If  $A$  and  $B$  are independent, then so are  $A^c$  and  $B$ ,  $A$  and  $B^c$ , and  $A^c$  and  $B^c$ .
2. Conversely, events  $A$  and  $B$  are independent if and only if their indicator random variables  $1_A$  and  $1_B$  are independent.

The definition of independence can be extended to the infinite collection.

**Definition 12.5.2**

An infinite collection of objects ( $\sigma$ -fields, random variables, or sets) is said to be independent if every finite subcollection is,

1.  $\sigma$ -fields  $\mathcal{F}_1, \mathcal{F}_2, \dots, \mathcal{F}_n$  are independent if whenever  $A_i \in \mathcal{F}_i$  for  $i = 1, \dots, n$ , we have

$$P(\cap_{i=1}^n A_i) = \prod_{i=1}^n P(A_i). \quad (12.9)$$

2. Random variables  $X_1, \dots, X_n$  are independent if whenever  $B_i \in \mathcal{R}$  for  $i = 1, \dots, n$  we have

$$P(\cap_{i=1}^n \{X_i \in B_i\}) = \prod_{i=1}^n P(X_i \in B_i). \quad (12.10)$$

3. Sets  $A_1, \dots, A_n$  are independent if whenever  $I \subset \{1, \dots, n\}$  we have

$$P(\cap_{i \in I} A_i) = \prod_{i \in I} P(A_i). \quad (12.11)$$

## 12.5.2 Sufficient Conditions for Independence

## 12.5.3 Independence, Distribution, and Expectation

**Theorem 12.5.3**

Suppose  $X_1, \dots, X_n$  are independent random variables and  $X_i$  has distribution  $\mu_i$ , then  $(X_1, \dots, X_n)$  has distribution  $\mu_1 \times \dots \times \mu_n$ .

**Theorem 12.5.4**

If  $X_1, \dots, X_n$  are independent and have

1.  $X_i \geq 0$  for all  $i$ , or
2.  $E|X_i| < \infty$  for all  $i$ .

then

$$E\left(\prod_{i=1}^n X_i\right) = \prod_{i=1}^n EX_i \quad (12.12)$$

**12.5.4 Sums of Independent Random Variables****Theorem 12.5.5 (Convolution for Random Variables)**

1. If  $X$  and  $Y$  are independent,  $F(x) = P(X \leq x)$ , and  $G(y) = P(Y \leq y)$ , then

$$P(X + Y \leq z) = \int F(z - y) dG(y). \quad (12.13)$$

2. If  $X$  and  $Y$  are independent,  $X$  with density  $f$  and  $Y$  with distribution function  $G$ , then  $X + Y$  has density

$$h(x) = \int f(x - y) dG(y). \quad (12.14)$$

Suppose  $Y$  has density  $g$ , the last formula can be written as

$$h(x) = \int f(x - y)g(y) dy. \quad (12.15)$$

3. If  $X$  and  $Y$  are independent, integral-valued random variables, then

$$P(X + Y = n) = \sum_m P(X = m)P(Y = n - m). \quad (12.16)$$

**12.6 Moments****Lemma 12.6.1**

If  $Y > 0$  and  $p > 0$ , then

$$E(Y^p) = \int_0^\infty py^{p-1}P(Y > y) dy. \quad (12.17)$$

## 12.7 Characteristic Functions

### 12.7.1 Definition of Characteristic Functions

#### Definition 12.7.1 (Characteristic Function)

If  $X$  is a random variable, we define its characteristic function (ch.f) by

$$\varphi(t) = E(e^{itX}) = E(\cos tX) + iE(\sin tX). \quad (12.18)$$

*Remark.* Euler Equation.

### 12.7.2 Properties of Characteristic Functions

#### Theorem 12.7.1 (Properties of Characteristic Function)

Any characteristic function has the following properties:

1.  $\varphi(0) = 1$ ,
2.  $\varphi(-t) = \overline{\varphi(t)}$ ,
3.  $|\varphi(t)| = |Ee^{itX}| \leq E|e^{itX}| = 1$ ,
4.  $\varphi(t)$  is uniformly continuous on  $(-\infty, \infty)$ ,
5.  $Ee^{it(aX+b)} = e^{itb}\varphi(at)$ ,
6. If  $X_1$  and  $X_2$  are independent and have ch.f.'s  $\varphi_1$  and  $\varphi_2$ , then  $X_1 + X_2$  has ch.f.  $\varphi_1(t)\varphi_2(t)$ .

*Proof.*

□

### 12.7.3 The Inversion Formula

The characteristic function uniquely determines the distribution. This and more is provided by:

#### Theorem 12.7.2 (The Inversion Formula)

Let  $\varphi(t) = \int e^{itx} \mu(dx)$  where  $\mu$  is a probability measure. If  $a < b$ , then

$$\lim_{T \rightarrow \infty} (2\pi)^{-1} \int_{-T}^T \frac{e^{-ita} - e^{-itb}}{it} \varphi(t) dt = \mu(a, b) + \frac{1}{2} \mu(\{a, b\}) \quad (12.19)$$

*Proof.*

□



**Theorem 12.7.3**

If  $\int |\varphi(t)| dt < \infty$ , then  $\mu$  has bounded continuous density

$$f(y) = \frac{1}{2\pi} \int e^{-ity} \varphi(t) dt. \quad (12.20)$$

*Proof.*

□

**12.7.4 Moments and Derivatives****Theorem 12.7.4**

If  $\int |x|^n \mu(dx) < \infty$ , then its characteristic function  $\varphi$  has a continuous derivative of order  $n$  given by

$$\varphi^{(n)}(t) = \int (ix)^n e^{itx} \mu(dx). \quad (12.21)$$

**Theorem 12.7.5**

If  $E|X|^2 < \infty$  then

$$\varphi(t) = 1 + itEX - t^2 E(X^2)/2 + o(t^2). \quad (12.22)$$

**Theorem 12.7.6**

If  $\limsup_{h \downarrow 0} \{\varphi(h) - 2\varphi(0) + \varphi(-h)\}/h^2 > -\infty$ , then

$$E|X|^2 < \infty. \quad (12.23)$$

# Chapter 13

## Convergence of Random Variables

### 13.1 Modes of Convergence

#### 13.1.1 Convergence in Mean

##### Definition 13.1.1 (Convergence in Mean)

A sequence  $\{X_n\}$  of real-valued random variables **converges in the r-th mean** ( $r \geq 1$ ) towards the random variable  $X$ , if

1. The r-th absolute moments  $E(|X_n|^r)$  and  $E(|X|^r)$  of  $\{X_n\}$  and  $X$  exist,
2.  $\lim_{n \rightarrow \infty} E(|X_n - X|^r) = 0$ .

Convergence in the r-th mean is denoted by

$$X_n \xrightarrow{L^r} X. \quad (13.1)$$

#### 13.1.2 Convergence in Probability

##### Definition 13.1.2 (Convergence in Probability)

A sequence  $\{X_n\}$  of real-valued random variables **converges in probability** towards the random variable  $X$ , if

$$\forall \varepsilon > 0, \quad \lim_{n \rightarrow \infty} P(|X_n - X| > \varepsilon) = 0. \quad (13.2)$$

Convergence in probability is denoted by

$$X_n \xrightarrow{p} X. \quad (13.3)$$

### 13.1.3 Convergence in Distribution

#### Definition 13.1.3 (Convergence in Distribution)

A sequence  $\{X_n\}$  of real-valued random variables is said to **converge in distribution**, or **converge weakly**, or **converge in law** to a random variable  $X$ , if

$$\lim_{n \rightarrow \infty} F_n(x) = F(x), \quad (13.4)$$

for every number at  $x \in \mathbb{R}$  which  $F$  is continuous. Here  $F_n$  and  $F$  are the cumulative distribution functions of random variables  $X_n$  and  $X$ , respectively. Convergence in distribution is denoted as

$$X_n \xrightarrow{d} X, \text{ or } X_n \Rightarrow X. \quad (13.5)$$

- Convergence in Distribution is the weakest form of convergence typically discussed since it is implied by all other types of convergence mentioned in this chapter.
- Convergence in Distribution does not imply that the sequence of corresponding probability density functions will also converge. However, according to Scheff's theorem, convergence of the probability density functions implies convergence in distribution.

#### Theorem 13.1.1 (Portmanteau Lemma)

$\{X_n\}$  converges in distribution to  $X$ , if and only if any of the following statements are true,

- $P(X_n \leq x) \rightarrow P(X \leq x)$ , for all continuity points of the distribution of  $X$ .
- $Ef(X_n) \rightarrow Ef(X)$ , for all bounded, continuous (Lipschitz) functions  $f$ .
- $\liminf_{n \rightarrow \infty} P(X_n \in G) \geq P(X_\infty \in G)$ , for all open sets  $G$ .
- $\limsup_{n \rightarrow \infty} P(X_n \in K) \leq P(X_\infty \in K)$ , for all closed sets  $K$ .
- $\lim_{n \rightarrow \infty} P(X_n \in A) = P(X_\infty \in A)$ , for all Borel sets  $A$  with  $P(X_\infty \in \partial A) = 0$ .

*Proof.*

□

**Continuous Mapping Theorem****Theorem 13.1.2 (Continuous Mapping Theorem)**

Let  $g$  be a measurable function and  $D_g = \{x : g \text{ is discontinuous at } x\}$  with  $P(X \in D_g) = 0$ , then,

$$\begin{aligned} X_n \xrightarrow{d} X &\Rightarrow g(X_n) \xrightarrow{d} g(X), \\ X_n &\xrightarrow{p} X \Rightarrow g(X_n) \xrightarrow{p} g(X), \\ X_n &\xrightarrow{a.s.} X \Rightarrow g(X_n) \xrightarrow{a.s.} g(X). \end{aligned} \tag{13.6}$$

If in addition  $g$  is bounded, then

$$Eg(X_n) \rightarrow Eg(X). \tag{13.7}$$

*Proof.*

□

**Slutsky's Theorem****Theorem 13.1.3 (Slutsky's Theorem)**

Let  $X_n, Y_n$  be sequences of random variables. If  $X_n \xrightarrow{d} X$  and  $Y_n \xrightarrow{p} c$ , then

1.  $X_n + Y_n \xrightarrow{d} X + c$ .
2.  $X_n Y_n \xrightarrow{d} cX$ .
3.  $X_n/Y_n \xrightarrow{d} X/c$ , provided that  $c$  is invertible.

*Proof.*

□

*Remark.* However that convergence in distribution of  $X_n \xrightarrow{d} X$  and  $Y_n \xrightarrow{d} Y$  does in general not imply convergence in distribution of  $X_n + Y_n \xrightarrow{d} X + Y$  or of  $X_n Y_n \xrightarrow{d} XY$ .

### The Delta Methods

#### Theorem 13.1.4 (Delta Method)

Let  $\{X_n\}$  be a sequence of random variables with

$$\sqrt{n}(X_n - \theta) \xrightarrow{d} N(0, \sigma^2) \quad (13.8)$$

where  $\theta$  and  $\sigma$  are finite, then for any function  $g$  with the property that  $g'(\theta)$  exists and is non-zero valued,

$$\sqrt{n}[g(X_n) - g(\theta)] \xrightarrow{d} N(0, \sigma^2 \cdot [g'(\theta)]^2) \quad (13.9)$$

*Proof.* Under the assumption that  $g'(\theta)$  is continuous.

Since,  $g'(\theta)$  exists, with the first-order Taylor Approximation, that

$$g(X_n) = g(\theta) + g'(\tilde{\theta})(X_n - \theta)$$

where  $\tilde{\theta}$  lies between  $X_n$  and  $\theta$ . Since  $X_n \xrightarrow{p} \theta$ , and  $|\tilde{\theta} - \theta| < |X_n - \theta|$ , then

$$\tilde{\theta} \xrightarrow{p} \theta$$

Since  $g'(\theta)$  is continuous, by Continuous Mapping Theorem (13.1.2),

$$g'(\tilde{\theta}) \xrightarrow{p} g'(\theta)$$

and,

$$\begin{aligned} \sqrt{n}(g(X_n) - g(\theta)) &= \sqrt{n}g'(\tilde{\theta})(X_n - \theta) \\ \sqrt{n}(X_n - \theta) &\xrightarrow{d} N(0, \sigma^2) \end{aligned}$$

by Slutsky's Theorem (13.1.3),

$$\sqrt{n}[g(X_n) - g(\theta)] \xrightarrow{d} N(0, \sigma^2 \cdot [g'(\theta)]^2)$$

□

#### Theorem 13.1.5 (Second-order Delta Method)

*Remark.* We can approximate the moments of a function  $f(\cdot)$  of a random variable  $X$  using Taylor expansions, provided that  $f(\cdot)$  is sufficiently differentiable and that the moments of  $X$  are finite. Suppose  $\mu = \mathbb{E}(X)$ , and  $\sigma^2 = \text{Var}(X)$ , with the Taylor expansions for the functions of random variables,

$$f(X) = f[\mu + (X - \mu)] \approx f(\mu) + f'(\mu)(X - \mu) \quad (13.10)$$

Thus,

$$\mathbb{E}[f(X)] \approx \mathbb{E}[f(\mu)], \quad \text{Var}[f(X)] \approx [f'(\mu)]^2 \cdot \sigma^2 \quad (13.11)$$

**Lèvy' s Continuity Theorem****Theorem 13.1.6 (Lèvy' s Continuity Theorem)**

Let  $\mu_n, 1 \leq n \leq \infty$  be probability measures with ch.f.  $\varphi_n$ .

1. If  $\mu_n \xrightarrow{d} \mu_\infty$ , then  $\varphi_n(t) \rightarrow \varphi_\infty(t)$  for all  $t$ .
2. If  $\varphi_n(t)$  converges pointwise to a limit  $\varphi(t)$  that is continuous at 0, then the associated sequence of distributions  $\mu_n$  is tight and converges weakly to the measure  $\mu$  with characteristic function  $\varphi$ .

*Proof.*

□

**Cramér-Wold Theorem****Theorem 13.1.7 (Cramér-Wold Theorem)****13.1.4 Almost Sure Convergence****Definition 13.1.4 (Almost Sure Convergence)**

A sequence  $\{X_n\}$  of real-valued random variables converges **almost sure** or **almost everywhere** or **with probability 1** or **strongly** towards the random variable  $X$ , if

$$P\left(\lim_{n \rightarrow \infty} X_n = X\right) = 1. \quad (13.12)$$

Almost sure convergence is denoted by

$$X_n \xrightarrow{a.s.} X. \quad (13.13)$$

*Remark.*

**13.1.5 Convergence in Uninform****Definition 13.1.5 (Convergence in Uninform)**

### 13.1.6 Asymptotic Notation

#### Definition 13.1.6

A sequence  $\{A_n\}$  of real-valued random variables is of smaller order in probability than a sequence  $\{B_n\}$ , if

$$\frac{A_n}{B_n} \xrightarrow{p} 0. \quad (13.14)$$

Smaller order in probability is denoted by

$$A_n = o_p(B_n). \quad (13.15)$$

Particularly,

$$A_n = o_p(1) \iff A_n \xrightarrow{p} 0. \quad (13.16)$$

#### Definition 13.1.7

A sequence  $\{A_n\}$  of real-valued random variables is of smaller order than or equal to a sequence  $\{B_n\}$  in probability, if

$$\forall \varepsilon > 0 \exists M_\varepsilon, \quad \lim_{n \rightarrow \infty} P(|A_n| \leq M_\varepsilon |B_n|) \geq 1 - \varepsilon. \quad (13.17)$$

Smaller order than or equal to in probability is denoted by

$$A_n = O_p(B_n). \quad (13.18)$$

#### Definition 13.1.8

A sequence  $\{A_n\}$  of real-valued random variables is of the same order as a sequence  $\{B_n\}$  in probability, if

$$\forall \varepsilon > 0 \exists m_\varepsilon < M_\varepsilon, \quad \lim_{n \rightarrow \infty} P\left(m_\varepsilon < \frac{|A_n|}{|B_n|} < M_\varepsilon\right) \geq 1 - \varepsilon. \quad (13.19)$$

The same order in probability is denoted by

$$A_n \asymp_p B_n. \quad (13.20)$$

## 13.2 Relationships of Modes

### Lemma 13.2.1

If  $p > 0$  and  $E|Z_n|^p \rightarrow 0$ , then

$$Z_n \xrightarrow{p} 0. \quad (13.21)$$

*Proof.* □

### Theorem 13.2.1

If  $X_n \xrightarrow{p} X$ , then

$$X_n \xrightarrow{d} X, \quad (13.22)$$

and that, conversely, if  $X_n \xrightarrow{d} c$ , where  $c$  is a constant, then

$$X_n \xrightarrow{p} c. \quad (13.23)$$

*Proof.* 1.  $\forall \varepsilon > 0$ , at fixed point  $x$ , since if  $X_n \leq x$  and  $|X_n - X| \leq \varepsilon$ , then  $X \leq x + \varepsilon$ , then

$$\{X \leq x + \varepsilon\} \subset \{X_n \leq x\} \cup \{|X_n - X| > \varepsilon\},$$

similarly, if  $X \leq x - \varepsilon$  and  $|X_n - X| \leq \varepsilon$ , then  $X_n \leq x$ , then

$$\{X_n \leq x\} \subset \{X \leq x - \varepsilon\} \cup \{|X_n - X| > \varepsilon\},$$

then, by the union bound,

$$\begin{aligned} P(X \leq x + \varepsilon) &\leq P(X_n \leq x) + P(|X_n - X| > \varepsilon), \\ P(X_n \leq x) &\leq P(X \leq x - \varepsilon) + P(|X_n - X| > \varepsilon). \end{aligned}$$

So, we got

$$\begin{aligned} P(X \leq x + \varepsilon) - P(|X_n - X| > \varepsilon) &\leq P(X_n \leq x) \\ &\leq P(X \leq x - \varepsilon) + P(|X_n - X| > \varepsilon) \end{aligned}$$

As  $n \rightarrow \infty$ ,  $P(|X_n - X| > \varepsilon) \rightarrow 0$ , then

$$\begin{aligned} P(X \leq x - \varepsilon) &\leq \lim_{n \rightarrow \infty} P(X_n \leq x) \leq P(X \leq x + \varepsilon) \\ &\Rightarrow F(x - \varepsilon) \leq \lim_{n \rightarrow \infty} F_n(x) \leq F(x + \varepsilon) \end{aligned}$$

By the property of distribution (Theorem 12.3.1), as  $\varepsilon \rightarrow 0$ , then

$$\lim_{n \rightarrow \infty} F_n(x) = F(x),$$

which means,

$$X_n \xrightarrow{d} X.$$



2. Since  $X_n \xrightarrow{d} c$ , where  $c$  is a constant, then  $\forall \varepsilon > 0$ ,

$$\begin{aligned} \lim_{n \rightarrow \infty} P(X_n \leq c + \varepsilon) = 1 &\Rightarrow \lim_{n \rightarrow \infty} P(X_n > c + \varepsilon) = 0 \\ \lim_{n \rightarrow \infty} P(X_n \leq c - \varepsilon) &= 0. \end{aligned}$$

Therefore,

$$P(|X_n - c| < \varepsilon) = 0,$$

which means

$$X_n \xrightarrow{p} c.$$

□

**Theorem 13.2.2**

If  $X_n \xrightarrow{a.s.} X$ , then

$$X_n \xrightarrow{p} X. \quad (13.24)$$

*Proof.*

□

**Theorem 13.2.3**

$X_n \xrightarrow{p} X$  if and only if for all subsequence  $X_{n(m)}$  exists a further subsequence  $X_{n(m_k)}$ , such that

$$X_{n(m_k)} \xrightarrow{a.s.} X. \quad (13.25)$$

**Lemma 13.2.2**

If  $F_n \xrightarrow{d} F_\infty$ , then there are random variables  $Y_n, 1 \leq n \leq \infty$ , with distribution  $F_n$  so that

$$Y_n \xrightarrow{a.s.} Y_\infty. \quad (13.26)$$

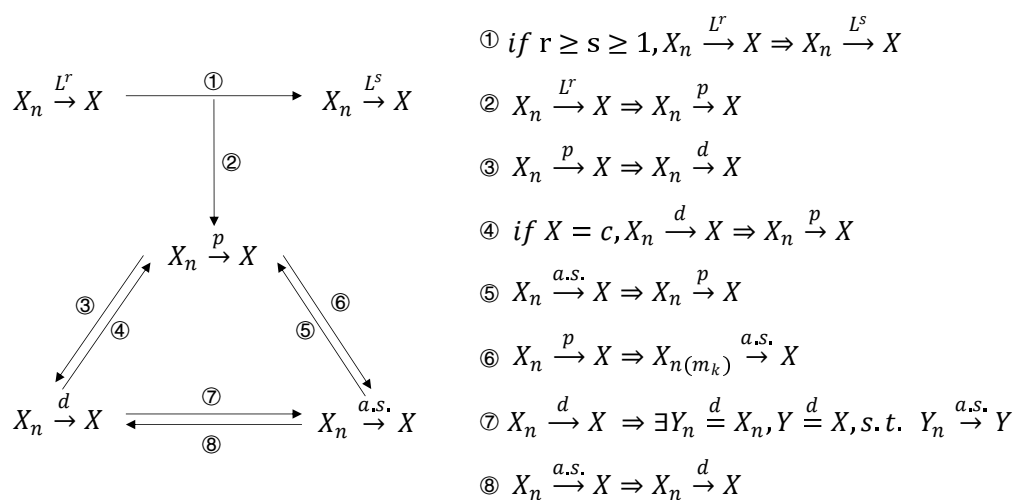


Figure 13.1: Relations of Convergence of Random Variables

# Chapter 14

## Law of Large Numbers

### 14.1 Weak Law of Large Numbers

#### Theorem 14.1.1 (Weak Law of Large Numbers with Finite Variances)

Let  $X_1, X_2, \dots$  be iid random variables with  $EX_i = \mu$  and  $\text{Var}(X_i) \leq C < \infty$ . Suppose  $S_n = X_1 + X_2 + \dots + X_n$ , then

$$S_n/n \xrightarrow{L^2} \mu, \quad S_n/n \xrightarrow{p} \mu. \quad (14.1)$$

*Proof.*

□

#### Theorem 14.1.2 (Weak Law of Large Numbers without iid)

Let  $X_1, X_2, \dots$  be random variables, Suppose  $S_n = X_1 + X_2 + \dots + X_n$ ,  $\mu_n = ES_n$ ,  $\sigma_n^2 = \text{Var}(S_n)$ , if  $\sigma_n^2/b_n^2 \rightarrow 0$ , then

$$\frac{S_n - \mu_n}{b_n} \xrightarrow{p} 0. \quad (14.2)$$

*Proof.*

□

**Theorem 14.1.3 (Weak Law of Large Numbers for Triangular Arrays)**

For each  $n$ , let  $X_{n,m}$ ,  $1 \leq m \leq n$ , be independent random variables. Suppose  $b_n > 0$  with  $b_n \rightarrow \infty$ ,  $\bar{X}_{n,m} = X_{n,m}I_{(X_{n,m} \leq b_n)}$ , if

1.  $\sum_{m=1}^n P(|X_{n,m}| > b_n) \rightarrow 0$ , and
2.  $b_n^{-2} \sum_{m=1}^n E\bar{X}_{n,m}^2 \rightarrow 0$ .

Suppose  $S_n = X_{n,1} + \cdots + X_{n,n}$  and  $a_n = \sum_{m=1}^n E\bar{X}_{n,m}$ , then

$$\frac{S_n - a_n}{b_n} \xrightarrow{p} 0. \quad (14.3)$$

*Proof.* □

**Theorem 14.1.4 (Weak Law of Large Numbers by Feller)**

Let  $X_1, X_2, \dots$  be iid random variables with

$$\lim_{x \rightarrow 0} xP(|X_i| > x) = 0. \quad (14.4)$$

Suppose  $S_n = X_1 + X_2 + \cdots + X_n$ ,  $\mu_n = E(X_1 I_{(|X_1| < n)})$ , then

$$S_n/n - \mu_n \xrightarrow{p} 0. \quad (14.5)$$

*Proof.* □

**Theorem 14.1.5 (Weak Law of Large Numbers)**

Let  $X_1, X_2, \dots$  be iid random variables with  $E|X_i| < \infty$ . Suppose  $S_n = X_1 + X_2 + \cdots + X_n$ ,  $\mu = EX_i$ , then

$$S_n/n \xrightarrow{p} \mu. \quad (14.6)$$

*Proof.* □

*Remark.*  $E|X_i| = \infty$

## 14.2 Strong Law of Large Numbers

### 14.2.1 Borel-Cantelli Lemmas

#### Lemma 14.2.1 (Borel-Cantelli Lemma)

If  $\sum_{n=1}^{\infty} P(A_n) < \infty$ , then

$$P(A_n \text{ i.o.}) = 0. \quad (14.7)$$

#### Lemma 14.2.2 (The Second Borel-Cantelli Lemma)

If  $\{A_n\}$  are independent with  $\sum_{n=1}^{\infty} P(A_n) = \infty$ , then,

$$P(A_n \text{ i.o.}) = 1. \quad (14.8)$$

#### Corollary 14.2.1

Suppose  $\{A_n\}$  are independent with  $P(A_n) < 1, \forall n$ . If  $P(\cup_{n=1}^{\infty} A_n) = 1$  then

$$\sum_{n=1}^{\infty} P(A_n) = \infty, \quad (14.9)$$

and hence  $P(A_n \text{ i.o.}) = 1$

*Proof.*

□

### 14.2.2 Strong Law of Large Numbers

#### Theorem 14.2.1 (Strong Law of Large Numbers)

Let  $X_1, X_2, \dots$  be iid random variables with  $E|X_i| < \infty$ . Suppose  $S_n = X_1 + X_2 + \dots + X_n$ ,  $\mu = EX_i$ , then

$$S_n/n \xrightarrow{a.s.} \mu. \quad (14.10)$$

### 14.3 Uniform Law of Large Numbers

#### Theorem 14.3.1 (Uniform Law of Large Numbers)

Suppose

1.  $\Theta$  is compact.
2.  $g(X_i, \theta)$  is continuous at each  $\theta \in \Theta$  almost sure.
3.  $g(X_i, \theta)$  is dominated by a function  $G(X_i)$ , i.e.  $|g(X_i, \theta)| \leq G(X_i)$ .
4.  $EG(X_i) < \infty$ .

Then

$$\sup_{\theta \in \Theta} \left| n^{-1} \sum_{i=1}^n g(X_i, \theta) - Eg(X_i, \theta) \right| \xrightarrow{p} 0. \quad (14.11)$$

*Proof.* Suppose

$$\Delta_\delta(X_i, \theta_0) = \sup_{\theta \in B(\theta_0, \delta)} g(X_i, \theta) - \inf_{\theta \in B(\theta_0, \delta)} g(X_i, \theta).$$

Since (i)  $\Delta_\delta(X_i, \theta_0) \xrightarrow{a.s.} 0$  by condition (2), (ii)  $\Delta_\delta(X_i, \theta_0) \leq 2 \sup_{\theta \in \Theta} |g(X_i, \theta)| \leq 2G(X_i)$  by condition (3) and (4). Then

$$E\Delta_\delta(X_i, \theta_0) \rightarrow 0, \delta \rightarrow 0.$$

So, for all  $\theta \in \Theta$  and  $\varepsilon > 0$ , there exists  $\delta_\varepsilon(\theta)$  such that

$$E[\Delta_{\delta_\varepsilon(\theta)}(X_i, \theta)] < \varepsilon.$$

Since  $\Theta$  is compact, we can find a finite subcover, such that  $\Theta$  is covered by

$$\cup_{k=1}^K B(\theta_k, \delta_\varepsilon(\theta_k)).$$

$$\begin{aligned} & \sup_{\theta \in \Theta} \left[ n^{-1} \sum_{i=1}^n g(X_i, \theta) - Eg(X_i, \theta) \right] \\ &= \max_k \sup_{\theta \in B(\theta_k, \delta_\varepsilon(\theta_k))} \left[ n^{-1} \sum_{i=1}^n g(X_i, \theta) - Eg(X_i, \theta) \right] \\ &\leq \max_k \left[ n^{-1} \sum_{i=1}^n \sup_{\theta \in B(\theta_k, \delta_\varepsilon(\theta_k))} g(X_i, \theta) - E \inf_{\theta \in B(\theta_k, \delta_\varepsilon(\theta_k))} g(X_i, \theta) \right] \end{aligned}$$

Since

$$E \left| \sup_{\theta \in B(\theta_k, \delta_\varepsilon(\theta_k))} g(X_i, \theta) \right| \leq EG(X_i) < \infty,$$

by the Weak Law of Large Numbers (Theorem 14.1.5),

$$\begin{aligned}
 &= o_p(1) + \max_k \left[ E \sup_{\theta \in B(\theta_k, \delta_\varepsilon(\theta_k))} g(X_i, \theta) - E \inf_{\theta \in B(\theta_k, \delta_\varepsilon(\theta_k))} g(X_i, \theta) \right] \\
 &= o_p(1) + \max_k E \Delta_{\delta_\varepsilon(\theta_k)}(X_i, \theta_k) \\
 &\leq o_p(1) + \varepsilon
 \end{aligned}$$

By analogous argument,

$$\inf_{\theta \in \Theta} \left[ n^{-1} \sum_{i=1}^n g(X_i, \theta) - E g(X_i, \theta) \right] \geq o_p(1) - \varepsilon.$$

The desired result follows from the above equation by the fact that  $\varepsilon$  is chosen arbitrarily.  $\square$

# Chapter 15

## Central Limit Theorems

### 15.1 Central Limit Theorem

#### 15.1.1 The De Moivre-Laplace Theorem

##### Lemma 15.1.1 (Stirling's Formula)

$$n! \sim \sqrt{2\pi n} n^{n+\frac{1}{2}} e^{-n} \quad n \rightarrow \infty. \quad (15.1)$$

*Proof.*

□

##### Lemma 15.1.2

If  $c_j \rightarrow 0$ ,  $a_j \rightarrow \infty$  and  $a_j c_j \rightarrow \lambda$ , then

$$(1 + c_j)^{a_j} \rightarrow e^\lambda. \quad (15.2)$$

*Proof.*

□

##### Theorem 15.1.1 (The De Moivre-Laplace Theorem)

Let  $X_1, X_2, \dots$  be iid with  $P(X_1 = 1) = P(X_1 = -1) = 1/2$  and let  $S_n = X_1 + \dots + X_n$ . If  $a < b$ , then as  $m \rightarrow \infty$

$$P(a \leq S_m/\sqrt{m} \leq b) \rightarrow \int_a^b (2\pi)^{-1/2} e^{-x^2/2} dx. \quad (15.3)$$

*Proof.* If  $n$  and  $k$  are integers

$$P(S_{2n} = 2k) = \binom{2n}{n+k} 2^{-2n}$$



By lemma 15.1.1, we have

$$\begin{aligned} \binom{2n}{n+k} &= \frac{(2n)!}{(n+k)!(n-k)!} \\ &\sim \frac{(2n)^{2n}}{(n+k)^{n+k}(n-k)^{n-k}} \cdot \frac{(2\pi(2n))^{1/2}}{(2\pi(n+k))^{1/2}(2\pi(n-k))^{1/2}} \end{aligned}$$

Hence,

$$\begin{aligned} P(S_{2n} = 2k) &= \binom{2n}{n+k} 2^{-2n} \\ &\sim \left(1 + \frac{k}{n}\right)^{-n-k} \cdot \left(1 - \frac{k}{n}\right)^{-n+k} \\ &\quad \cdot (\pi n)^{-1/2} \cdot \left(1 + \frac{k}{n}\right)^{-1/2} \cdot \left(1 - \frac{k}{n}\right)^{-1/2} \\ &= \left(1 - \frac{k^2}{n^2}\right)^{-n} \cdot \left(1 + \frac{k}{n}\right)^{-k} \cdot \left(1 - \frac{k}{n}\right)^k \\ &\quad \cdot (\pi n)^{-1/2} \cdot \left(1 + \frac{k}{n}\right)^{-1/2} \cdot \left(1 - \frac{k}{n}\right)^{-1/2} \end{aligned}$$

Let  $2k = x\sqrt{2n}$ , i.e.,  $k = x\sqrt{\frac{n}{2}}$ . By lemma 15.1.2, we have

$$\begin{aligned} \left(1 - \frac{k^2}{n^2}\right)^{-n} &= \left(1 - x^2/2n\right)^{-n} \rightarrow e^{x^2/2} \\ \left(1 + \frac{k}{n}\right)^{-k} &= (1 + x/\sqrt{2n})^{-x\sqrt{n/2}} \rightarrow e^{-x^2/2} \\ \left(1 - \frac{k}{n}\right)^k &= (1 - x/\sqrt{2n})^{x\sqrt{n/2}} \rightarrow e^{-x^2/2} \end{aligned}$$

For this choice of  $k$ ,  $k/n \rightarrow 0$ , so

$$\left(1 + \frac{k}{n}\right)^{-1/2} \cdot \left(1 - \frac{k}{n}\right)^{-1/2} \rightarrow 1.$$

Putting things together, we have

$$P(S_{2n} = 2k) \sim (\pi n)^{-1/2} e^{-x^2/2}, \frac{2k}{\sqrt{2n}} \rightarrow x.$$

Therefore,

$$P(a\sqrt{2n} \leq S_{2n} \leq b\sqrt{2n}) = \sum_{m \in [a\sqrt{2n}, b\sqrt{2n}] \cap 2\mathbb{Z}} P(S_{2n} = m)$$

Let  $m = x\sqrt{2n}$ , we have that this is

$$\approx \sum_{x \in [a, b] \cap (2\mathbb{Z}/\sqrt{2n})} (2\pi)^{-1/2} e^{-x^2/2} \cdot (2/n)^{1/2}$$

where  $2\mathbb{Z}/\sqrt{2n} = \{2z/\sqrt{2n} : z \in \mathbb{Z}\}$ . As  $n \rightarrow \infty$ , the sum just shown is

$$\approx \int_a^b (2\pi)^{-1/2} e^{-x^2/2} dx.$$

To remove the restriction to even integers, observe  $S_{2n+1} = S_{2n} \pm 1$ .

Let  $m = 2n$ , as  $m \rightarrow \infty$ ,

$$P(a \leq S_m/\sqrt{m} \leq b) \rightarrow \int_a^b (2\pi)^{-1/2} e^{-x^2/2} dx.$$

□

### 15.1.2 Central Limit Theorem

#### Theorem 15.1.2 (Classic Central Limit Theorem)

Let  $X_1, X_2, \dots$  be iid with  $\mathbb{E}X_i = \mu$ ,  $\text{Var}(X_i) = \sigma^2 < \infty$ . Let  $\bar{X} = \frac{1}{n} \sum_{i=1}^n X_i$ , then

$$n^{1/2} \frac{\bar{X} - \mu}{\sigma} \xrightarrow{d} \mathcal{N}(0, 1). \quad (15.4)$$

*Proof.*

□

#### Theorem 15.1.3 (The Linderberg-Feller Central Limit Theorem)

For each  $n$ , let  $X_{n,m}, 1 \leq m \leq n$ , be independent random variables with  $\mathbb{E}X_{n,m} = 0$ . If

1.  $\sum_{m=1}^n \mathbb{E}X_{n,m}^2 \rightarrow \sigma^2 > 0$ .
2.  $\forall \epsilon > 0, \lim_{n \rightarrow \infty} \sum_{m=1}^n \mathbb{E}(|X_{n,m}|^2; |X_{n,m}| > \epsilon) = 0$

Then  $S_n = X_{n,1} + \dots + X_{n,n} \xrightarrow{d} \sigma\chi$  as  $n \rightarrow \infty$ .

### 15.1.3 Berry-Esseen Theorem

#### Theorem 15.1.4 (Berry-Esseen Theorem)

Let  $X_1, X_2, \dots, X_n$  be iid with distribution  $F$ , which has a mean  $\mu$  and a finite third moment  $\sigma^3$ , then there exists a constant  $C$  (independent of  $F$ ),

$$|G_n(x) - \Phi(x)| \leq \frac{C}{\sqrt{n}} \frac{\mathbb{E}|X_1 - \mu|^3}{\sigma^3}, \quad \forall x. \quad (15.5)$$

**Corollary 15.1.1**

Under the assumptions of Theorem 51,

$$G_n(x) \rightarrow \Phi(x) \text{ as } n \rightarrow \infty$$

for any sequence  $F_n$  with mean  $\xi_n$  and variance  $\sigma_n^2$  for which

$$\frac{E_n |X_1 - \xi_n|^3}{\sigma_n^3} = o(\sqrt{n})$$

and thus in particular if (72) is bounded. Here  $E_n$  denotes the expectation under  $F_n$ .

## 15.2 CLT for independent non-identical Random Variables

**Theorem 15.2.1 (Liapounov CLT)****Theorem 15.2.2**

Let  $Y_1, Y_2, \dots$  be iid with  $\mathbb{E}(Y_i) = 0$ ,  $\text{Var}(Y_i) = \sigma^2 > 0$ , and  $\mathbb{E}|Y_i^3| = \gamma < \infty$ . If

$$\left( \sum_{i=1}^n |d_{ni}|^3 \right)^2 = o \left( \sum_{i=1}^n d_{ni}^2 \right)^3, \quad (15.6)$$

then

$$\frac{\sum_{i=1}^n d_{ni} Y_i}{\sigma \sqrt{\sum_{i=1}^n d_{ni}^2}} \xrightarrow{d} N(0, 1).$$

**Corollary 15.2.1**

The sufficient condition (15.6) is equivalent to

$$\max_{i=1, \dots, n} (d_{ni}^2) = o \left( \sum_{i=1}^n d_{ni}^2 \right). \quad (15.7)$$

## 15.3 CLT for Dependent Random Variables

# Chapter 16

## Multivariate Extensions

### 16.1 Multivariate Distributions

#### 16.1.1 Multivariate Normal Distribution

##### Definition 16.1.1 (Multivariate Normal Distribution)

The multivariate normal distribution of a  $p$ -dimensional random vector  $\mathbf{X}$  can be written as:

$$\mathbf{X} \sim \mathcal{N}(\boldsymbol{\mu}, \boldsymbol{\Sigma})$$

where  $\boldsymbol{\mu}$  is a  $p$ -dimensional mean vector and  $\boldsymbol{\Sigma}$  is a  $p \times p$  covariance matrix. Furthermore, the probability density function of  $\mathbf{X}$  is:

$$p(\mathbf{X}) = \frac{1}{(2\pi)^{p/2} |\boldsymbol{\Sigma}|^{1/2}} \exp \left( -\frac{1}{2} (\mathbf{X} - \boldsymbol{\mu})^\top \boldsymbol{\Sigma}^{-1} (\mathbf{X} - \boldsymbol{\mu}) \right).$$

##### Theorem 16.1.1

Suppose  $\mathbf{X} \sim \mathcal{N}(\boldsymbol{\mu}, \boldsymbol{\Sigma})$ , then:

1.  $\boldsymbol{\Sigma}^{-1/2}(\mathbf{X} - \boldsymbol{\mu}) \sim \mathcal{N}(\mathbf{0}, \mathbf{I})$ .
2.  $(\mathbf{X} - \boldsymbol{\mu})^\top \boldsymbol{\Sigma}^{-1} (\mathbf{X} - \boldsymbol{\mu}) \sim \chi_p^2$ .

### 16.1.2 Wishart Distribution

#### Definition 16.1.2 (Wishart Distribution)

The Wishart distribution is a generalization of the chi-squared distribution to multiple dimensions. If  $\mathbf{Z}$  is a  $p \times n$  matrix with each column drawn from a multivariate normal distribution  $\mathcal{N}(\mathbf{0}, \Sigma)$ , then the quadratic form  $\mathbf{X}$  has a Wishart distribution (with parameters  $\Sigma$ , and  $n$ ):

$$\mathbf{X} = \mathbf{Z}\mathbf{Z}^\top \sim W_p(\Sigma, n).$$

Furthermore, the probability density function of  $\mathbf{X}$  is:

$$p(\mathbf{X}) = \frac{|\mathbf{X}|^{(n-p-1)/2} \exp\left(-\frac{1}{2}\text{tr}(\Sigma^{-1}\mathbf{X})\right)}{2^{np/2} |\Sigma|^{n/2} \Gamma_p(n/2)}$$

### 16.1.3 Hotelling's T-squared Distribution

#### Definition 16.1.3 (Hotelling's $T^2$ Distribution)

If the vector  $\mathbf{d}$  is Gaussian multivariate-distributed with zero mean and unit covariance matrix  $\mathcal{N}(\mathbf{0}_p, \mathbf{I}_p)$  and  $\mathbf{M}$  is a  $p \times p$  matrix with unit scale matrix and  $m$  degrees of freedom with a Wishart distribution  $W(\mathbf{I}_p, m)$ , then the quadratic form  $X$  has a Hotelling distribution (with parameters  $p$  and  $m$ ):

$$X = m\mathbf{d}^\top \mathbf{M}^{-1} \mathbf{d} \sim T^2(p, m).$$

Furthermore, if a random variable  $X$  has Hotelling's  $T$ -squared distribution,  $X \sim T_{p,m}^2$ , then:

$$\frac{m-p+1}{pm} X \sim F_{p, m-p+1}$$

where  $F_{p, m-p+1}$  is the  $F$ -distribution with parameters  $p$  and  $m-p+1$ .

## 16.2 Convergence of Random Vectors

Let  $\mathbf{X}^{(n)}$  be a sequence of random vectors with cdf  $H_n$  converging in law to  $\mathbf{X}$  with cdf  $H$ . One then often needs to know whether for some set  $S$  in  $R^k$ ,

$$P(\mathbf{X}^{(n)} \in S) \rightarrow P(\mathbf{X} \in S). \quad (16.1)$$

That (16.1) need not be true for all  $S$  is seen from the case  $k = 1$ ,  $S = \{x : x \leq a\}$ . Then (16.1) can only be guaranteed when  $a$  is a continuity point of  $H$ .

**Theorem 16.2.1**

A sufficient condition for (16.1) to hold is that

$$P(\mathbf{X} \in \partial S) = 0.$$

**Example** (Multinomial).

**Example** (Difference of Means). Let  $X_1, \dots, X_m$  and  $Y_1, \dots, Y_n$  be independently distributed according to distributions  $F$  and  $G$ , with means  $\xi$  and  $\eta$  and finite variances  $\sigma^2$  and  $\tau^2$ , respectively. Then

$$\sqrt{m}(\bar{X} - \xi) \xrightarrow{d} \mathcal{N}(0, \sigma^2), \quad \sqrt{n}(\bar{Y} - \eta) \xrightarrow{d} \mathcal{N}(0, \tau^2).$$

If  $\frac{m}{m+n} \rightarrow \lambda$  ( $0 < \lambda < 1$ ), it follows that

$$\begin{aligned} \sqrt{m+n}(\bar{X} - \xi) &= \sqrt{\frac{m+n}{m}} \sqrt{m}(\bar{X} - \xi) \xrightarrow{d} N\left(0, \frac{\sigma^2}{\lambda}\right), \\ \sqrt{m+n}(\bar{Y} - \eta) &\xrightarrow{d} N\left(0, \frac{\tau^2}{1-\lambda}\right), \end{aligned}$$

and hence that

$$(\sqrt{m+n}(\bar{X} - \xi), \sqrt{m+n}(\bar{Y} - \eta)) \xrightarrow{d} (X, Y),$$

where  $X$  and  $Y$  are independent random variables with distributions  $N(0, \frac{\sigma^2}{\lambda})$  and  $N(0, \frac{\tau^2}{1-\lambda})$ , respectively. Since  $Y - X$  is a continuous function of  $(X, Y)$ , it follows that

$$\sqrt{m+n}[(\bar{Y} - \bar{X}) - (\eta - \xi)] \xrightarrow{d} N\left(0, \frac{\sigma^2}{\lambda} + \frac{\tau^2}{1-\lambda}\right),$$

or, equivalently, that

$$\frac{(\bar{Y} - \bar{X}) - (\eta - \xi)}{\sqrt{\frac{\sigma^2}{m} + \frac{\tau^2}{n}}} \xrightarrow{d} N(0, 1).$$

More specifically, consider the probability

$$P\{\sqrt{m+n}[(\bar{Y} - \bar{X}) - (\eta - \xi)] \leq z\}.$$

By Theorem 16.2.1, since  $P(Y - X = z) = 0$ , it follows that

$$P\{(Y - X) \leq z\} = \Phi\left(\frac{z}{\sqrt{\frac{\sigma^2}{\lambda} + \frac{\tau^2}{1-\lambda}}}\right).$$

**Theorem 16.2.2**

A necessary and sufficient condition for

$$(X_1^{(n)}, \dots, X_k^{(n)}) \xrightarrow{d} (X_1, \dots, X_k),$$

is that for any constants  $c_1, \dots, c_k$ , we have

$$\sum_{i=1}^k c_i X_i^{(n)} \xrightarrow{d} \sum_{i=1}^k c_i X_i.$$

**Example** (Orthogonal Linear Combinations). Let  $Y_1, Y_2, \dots$  be iid with mean  $\mathbb{E}(Y_i) = 0$  and variance  $\text{Var}(Y_i) = \sigma^2$ , and consider the joint distribution of the linear combinations

$$X_1^{(n)} = \sum_{j=1}^n a_{nj} Y_j, \quad X_2^{(n)} = \sum_{j=1}^n b_{nj} Y_j$$

satisfying the orthogonality conditions

$$\sum_{j=1}^n a_{nj}^2 = \sum_{j=1}^n b_{nj}^2 = 1, \quad \sum_{j=1}^n a_{nj} b_{nj} = 0. \quad (16.2)$$

Then if

$$\max_j a_{jn}^2 \rightarrow 0, \quad \max_j b_{jn}^2 \rightarrow 0, \quad n \rightarrow \infty.$$

it follows that

$$(X_1^{(n)}, X_2^{(n)}) \xrightarrow{d} (X_1, X_2),$$

with  $(X_1, X_2)$  independently distributed, each according to the normal distribution  $N(0, \sigma^2)$ .

*Proof.* To prove this result, it is by Theorem 16.2.2 enough to show that

$$c_1 X_1^{(n)} + c_2 X_2^{(n)} = \sum (c_1 a_{nj} + c_2 b_{nj}) Y_j \xrightarrow{d} c_1 X_1 + c_2 X_2,$$

where the distribution of  $c_1 X_1 + c_2 X_2$  is  $N(0, [c_1^2 + c_2^2] \sigma^2)$ . The sum on the left side of (5.1.24) is of the form  $\sum d_{nj} Y_j$  with

$$d_{nj} = c_1 a_{nj} + c_2 b_{nj}$$

It follows from (16.2) that

$$\sum_{j=1}^n d_{nj}^2 = c_1^2 + c_2^2$$

furthermore

$$\max d_{nj}^2 \leq 2 \max [c_1^2 a_{nj}^2 + c_2^2 b_{nj}^2] \leq 2 [c_1^2 \max a_{nj}^2 + c_2^2 \max b_{nj}^2].$$

Thus,

$$\max_j d_{nj}^2 / \sum_{j=1}^n d_{nj}^2 \rightarrow 0.$$

According to Theorems 15.2.2, it thus follows that

$$\sum d_{nj} Y_j \xrightarrow{d} N(0, (c_1^2 + c_2^2) \sigma^2).$$

Thus, by Theorem 16.2.2, we have

$$(X_1^{(n)}, X_2^{(n)}) \xrightarrow{d} (X_1, X_2),$$

with  $(X_1, X_2)$  independently distributed, each according to the normal distribution  $N(0, \sigma^2)$ .  $\square$



# Chapter 17

## Exercises for Probability Theory and Examples

### 17.1 Measure Theory

**Exercise.** 1. Show that if  $\mathcal{F}_1 \subset \mathcal{F}_2 \subset \dots$  are  $\sigma$ -algebras, then  $\cup_i \mathcal{F}_i$  is an algebra.

2. Give an example to show that  $\cup_i \mathcal{F}_i$  need not be a  $\sigma$ -algebra.

*Proof.* 1. **Complement:** Suppose  $A \in \cup_i \mathcal{F}_i$ , since  $\mathcal{F}_1 \subset \mathcal{F}_2 \subset \dots$ , assume  $A \in \mathcal{F}_i$ . And each  $\mathcal{F}_i$  is  $\sigma$ -algebra,

$$A^c \in \mathcal{F}_i \subset \cup_i \mathcal{F}_i.$$

**Finite Union:** Suppose  $A_1, A_2 \in \cup_i \mathcal{F}_i$ , since  $\mathcal{F}_1 \subset \mathcal{F}_2 \subset \dots$ , assume  $A_1 \in \mathcal{F}_i, A_2 \in \mathcal{F}_j$ , such that,

$$A_1, A_2 \in \mathcal{F}_{\max(i,j)}.$$

Since each  $\mathcal{F}_i$  is  $\sigma$ -algebra,

$$A_1 \cup A_2 \in \mathcal{F}_i \subset \cup_i \mathcal{F}_i.$$

2. Let  $\mathcal{F}_i$  be a Borel Set of  $[1, 2 - \frac{1}{i}]$ . Suppose  $A_i = [1, 2 - \frac{1}{i}] \in \mathcal{F}_i$ ,

$$\cup_i A_i = [1, 2) \notin \cup_i \mathcal{F}_i.$$

□

## 17.2 Laws of Large Numbers

## 17.3 Central Limit Theorems

**Exercise.** Let  $g \geq 0$  be continuous. If  $X_n \xrightarrow{d} X_\infty$ , then

$$\liminf_{n \rightarrow \infty} E g(X_n) \geq E g(X_\infty).$$

*Proof.* Let  $Y_n \stackrel{d}{=} X_n, 1 \leq n \leq \infty$  with  $Y_n \xrightarrow{a.s.} Y_\infty$  (Lemma 13.2.2). Since  $g \geq 0$  be continuous,  $g(Y_n) \xrightarrow{a.s.} g(Y_\infty)$  and  $g(Y_n) \geq 0$  (Theorem 13.1.2), and the Fatou's Lemma (12.4.2) implies,

$$\begin{aligned} \liminf_{n \rightarrow \infty} E g(X_n) &= \liminf_{n \rightarrow \infty} E g(Y_n) \geq E \left( \liminf_{n \rightarrow \infty} g(Y_n) \right) \\ &= E g(Y_\infty) = E g(X_\infty). \end{aligned}$$

□

**Exercise.** Suppose  $g, h$  are continuous with  $g(x) > 0$ , and  $|h(x)|/g(x) \rightarrow 0$  as  $|x| \rightarrow \infty$ . If  $F_n \xrightarrow{d} F$  and  $\int g(x) dF_n(x) \leq C < \infty$ , then

$$\int h(x) dF_n(x) \rightarrow \int h(x) dF(x).$$

*Proof.*

$$\begin{aligned} \left| \int h(x) dF_n(x) - \int h(x) dF(x) \right| &= \left| \int_{x \in [-M, M]} h(x) dF_n(x) + \int_{x \notin [-M, M]} h(x) dF_n(x) \right. \\ &\quad \left. - \int_{x \in [-M, M]} h(x) dF(x) - \int_{x \notin [-M, M]} h(x) dF(x) \right| \\ &\leq \left| \int_{x \in [-M, M]} h(x) dF_n(x) - \int_{x \in [-M, M]} h(x) dF(x) \right| \\ &\quad + \left| \int_{x \notin [-M, M]} h(x) dF_n(x) - \int_{x \notin [-M, M]} h(x) dF(x) \right|. \end{aligned}$$

Let  $X_n, 1 \leq n < \infty$ , with distribution  $F_n$ , so that  $X_n \xrightarrow{a.s.} X$  (Lemma 13.2.2).

$$\left| \int_{x \in [-M, M]} h(x) dF_n(x) - \int_{x \in [-M, M]} h(x) dF(x) \right| = \left| E(h(X_n) - h(X)) I_{x \in [-M, M]} \right|.$$

By Continuity Mapping Theorem (13.1.2),  $\lim_{n \rightarrow \infty} \left| E(h(X_n) - h(X)) I_{x \in [-M, M]} \right| = 0$ .

Since

$$h(x)I_{x \notin [-M, M]} \leq g(x) \sup_{x \notin [-M, M]} \frac{h(x)}{g(x)},$$

and by Exercise 17.3

$$Eg(X) \leq \liminf_{n \rightarrow \infty} Eg(X_n) = \liminf_{n \rightarrow \infty} \int g(x) dF_n(x) \leq C < \infty,$$

$$\begin{aligned} \left| \int_{x \notin [-M, M]} h(x) dF_n(x) - \int_{x \notin [-M, M]} h(x) dF(x) \right| &= \left| E(h(X_n) - h(X)) I_{x \notin [-M, M]} \right| \\ &\leq 2E \max(h(X_n), h(X)) I_{x \notin [-M, M]} \leq 2C \sup_{x \notin [-M, M]} \frac{h(x)}{g(x)}. \end{aligned}$$

Hence, let  $M \rightarrow \infty$ ,

$$\lim_{n \rightarrow \infty} \left| \int h(x) dF_n(x) - \int h(x) dF(x) \right| \leq 2C \sup_{x \notin [-M, M]} \frac{h(x)}{g(x)} \rightarrow 0,$$

which means,

$$\int h(x) dF_n(x) \rightarrow \int h(x) dF(x).$$

□

**Exercise.** Let  $X_1, X_2, \dots$  be iid with  $EX_i = 0$  and  $EX_i^2 = \sigma^2 \in (0, \infty)$ . Then

$$\sum_{m=1}^n X_m / \left( \sum_{m=1}^n X_m^2 \right)^{1/2} \xrightarrow{d} \chi.$$

**Exercise.** Show that if  $|X_i| \leq M$  and  $\sum_n \text{Var}(X_n) = \infty$ , then

$$(S_n - ES_n) / \sqrt{\text{Var}(S_n)} \xrightarrow{d} \chi.$$

**Exercise.** Suppose  $EX_i = 0$ ,  $EX_i^2 = 1$  and  $E|X_i|^{2+\delta} \leq C$  for some  $0 < \delta, C < \infty$ . Show that

$$S_n / \sqrt{n} \xrightarrow{d} \chi.$$

# Part VII

## Stochastic Process

# Chapter 18

## Martingales

### 18.1 Conditional Expectation

#### Definition 18.1.1 (Conditional Expectation)

**Example.** 1. If  $X \in \mathcal{F}$ , then

$$E(X | \mathcal{F}) = X.$$

2. If  $X$  is independent of  $\mathcal{F}$ , then

$$E(X | \mathcal{F}) = E(X).$$

3. If  $\Omega_1, \Omega_2, \dots$  is a finite or infinite partition of  $\Omega$  into disjoint sets, each of which has a positive probability, and let  $\mathcal{F} = \sigma(\Omega_1, \Omega_2, \dots)$ , then for each  $i$ ,

$$E(X | \mathcal{F}) = \frac{E(X; \Omega_i)}{P(\Omega_i)}.$$

**Property.**

### 18.2 Martingales

Let  $\mathcal{F}_n$  be a filtration, i.e., an increasing sequence of  $\sigma$ -fields.

**Definition 18.2.1 (Martingale)**

A sequence  $\{X_n\}$  of real-valued random variables is said to be a martingale for  $\mathcal{F}_n$ , if

1.  $X_n$  is integrable, i.e.,  $E|X_n| < \infty$
2.  $X_n$  is adapted to  $\mathcal{F}_n$ , i.e.,  $\forall n, X_n \in \mathcal{F}_n$
3.  $X_n$  satisfies the martingale condition, i.e.,

$$E(X_{n+1} | \mathcal{F}_n) = X_n, \quad \forall n \quad (18.1)$$

*Remark.* If in the last definition,  $=$  is replaced by  $\leq$  or  $\geq$ , then  $X$  is said to be a supermartingale or submartingale, respectively.

**Example** (Linear Martingale).

**Example** (Quadratic Martingale).

**Example** (Exponential Martingale).

**Example** (Random Walk). Suppose  $X_n = X_0 + \xi_1 + \cdots + \xi_n$ , where  $X_0$  is constant,  $\xi_m$  are independent and have  $E\xi_m = 0, \sigma_m^2 = E\xi_m^2 < \infty$ . Let  $\mathcal{F}_n = \sigma(\xi_1, \dots, \xi_n)$  for  $n \geq 1$  and take  $\mathcal{F}_0 = \{\emptyset, \Omega\}$ . Show  $X_n$  is a martingale, and  $X_n^2$  is a submartingale.

*Proof.* It is obvious that,

$$E|X_n| < \infty, \quad X_n \in \mathcal{F}_n$$

Since  $\xi_{n+1}$  is independent of  $\mathcal{F}_n$ , so using the linearity of conditional expectation, (4.1.1), and Example 4.1.4,

$$E(X_{n+1} | \mathcal{F}_n) = E(X_n | \mathcal{F}_n) + E(\xi_{n+1} | \mathcal{F}_n) = X_n + E\xi_{n+1} = X_n$$

So  $X_n$  is a martingale, and Theorem 4.2.6 implies  $X_n^2$  is a submartingale.  $\square$

*Remark.* If we let  $\lambda = x^2$  and apply Theorem 4.4.2 to  $X_n^2$ , we get Kolmogorov's maximal inequality, Theorem 2.5.5:

$$P\left(\max_{1 \leq m \leq n} |X_m| \geq x\right) \leq x^{-2} \text{var}(X_n) \quad (18.2)$$

**Theorem 18.2.1 (Orthogonality of Martingale Increments)****Theorem 18.2.2 (Conditional Variance Formula)**

Definition 18.2.2 (Predictable Sequence)

Definition 18.2.3 (Stopping Time)

Theorem 18.2.3 (Martingale Convergence Theorem)

## 18.3 Doob's Inequality

Theorem 18.3.1 (Doob's Decomposition)

Theorem 18.3.2 (Doob's Inequality)

Theorem 18.3.3 ( $L^p$  Maximum Inequality)

## 18.4 Uniform Integrability

## 18.5 Optional Stopping Theorems

# Chapter 19

## Markov Chains

### 19.1 Markov Chain

#### Definition 19.1.1 (Markov Chain, Simple)

A sequence  $\{X_n\}$  of real-valued random variables is said to be a Markov chain, if for any states  $i_0, \dots, i_{n-1}, i$ , and  $j$

$$P(X_{n+1} = j \mid X_n = i, X_{n-1} = i_{n-1}, \dots, X_0 = i_0) = P(X_{n+1} = j \mid X_n = i) \quad (19.1)$$

and the transition probability is

$$p(i, j) = P(X_{n+1} = j \mid X_n = i) \quad (19.2)$$

**Example** (Random Walk). Suppose  $X_n = X_0 + \xi_1 + \dots + \xi_n$ , where  $X_0$  is constant,  $\xi_m \in \mathbb{Z}^d$  are independent with distribution  $\mu$ . Show  $X_n$  is a Markov chain with transition probability,

$$p(i, j) = \mu(\{j - i\})$$

*Proof.* Since  $\xi_m$  are independent with distribution  $\mu$ ,

$$\begin{aligned} & P(X_{n+1} = j \mid X_n = i, X_{n-1} = i_{n-1}, \dots, X_0 = i_0) \\ &= P(X_n + \xi_{n+1} = j \mid X_n = i) = P(\xi_{n+1} = j - i) = \mu(\{j - i\}) \end{aligned}$$

□



**Definition 19.1.2 (Branching Processes)**

Let  $\xi_i^n, i, n \geq 1$ , be iid nonnegative integer-valued random variables. Define a sequence  $Z_n, n \geq 0$  by  $Z_0 = 1$  and

$$Z_{n+1} = \begin{cases} \xi_1^{n+1} + \cdots + \xi_{Z_n}^{n+1} & Z_n > 0 \\ 0 & Z_n = 0 \end{cases} \quad (19.3)$$

$Z_n$  is called a Branching process.

*Remark.* The idea behind the definitions is that  $Z_n$  is the number of individuals in the  $n$ -th generation, and each member of the  $n$ -th generation gives birth independently to an identically distributed number of children.

**Example** (Branching Processes). Show branching process is a Markov chain with transition probability,

$$p(i, j) = P\left(\sum_{k=1}^i \xi_k = j\right)$$

*Proof.* Since  $\xi_k^n$  are independent with identical distribution,

$$\begin{aligned} & P(Z_{n+1} = j \mid Z_n = i, Z_{n-1} = i_{n-1}, \dots, Z_0 = i_0) \\ &= P\left(\sum_{k=1}^{Z_n} \xi_k^{n+1} = j \mid Z_n = i\right) = P\left(\sum_{k=1}^i \xi_k = j\right) \end{aligned}$$

□

Suppose  $(S, \mathcal{S})$  is a measurable space, which will be the state space for our Markov chain.

**Definition 19.1.3 (Transition Probability)**

A function  $p : S \times \mathcal{S} \rightarrow \mathbb{R}$  is said to be a transition probability, if

1. For each  $x \in S$ ,  $A \rightarrow p(x, A)$  is a probability measure on  $(S, \mathcal{S})$
2. For each  $A \in \mathcal{S}$ ,  $x \rightarrow p(x, A)$  is a measurable function

**Definition 19.1.4 (Markov Chain)**

A sequence  $\{X_n\}$  of real-valued random variables with transition probability  $p$  is said to be a Markov chain with respect to  $\mathcal{F}_n$ , if

$$P(X_{n+1} \in B \mid \mathcal{F}_n) = p(X_n, B) \quad (19.4)$$

*Remark.* Given a transition probability  $p$  and an initial distribution  $\mu$  on  $(S, \mathcal{S})$ , the consistent set of finite dimensional distributions is

$$P(X_j \in B_j, 0 \leq j \leq n) = \int_{B_0} \mu(dx_0) \int_{B_1} p(x_0, dx_1) \cdots \int_{B_n} p(x_{n-1}, dx_n) F \quad (19.5)$$

## 19.2 Markov Properties

### Definition 19.2.1 (Shift Operator)

### Theorem 19.2.1 (Markov Property)

### Corollary 19.2.1 (Chapman-Kolmogorov Equation)

### Theorem 19.2.2 (Strong Markov Property)

## 19.3 Recurrence and Transience

Let  $T_y^0 = 0$ , and for  $k \geq 1$ , and

$$T_y^k = \inf \{n > T_y^{k-1} : X_n = y\} \quad (19.6)$$

then  $T_y^k$  is the time of the  $k$ -th return to  $y$ , where  $T_y^1 > 0$ , so any visit at time 0 does not count.

Let

$$\rho_{xy} = P_x(T_y < \infty) \quad (19.7)$$

and we have

$$P_x(T_y^k < \infty) = \rho_{xy} \rho_{yy}^{k-1} \quad (19.8)$$

*Proof.*

□

Let

$$N(y) = \sum_{n=1}^{\infty} 1_{(X_n=y)} \quad (19.9)$$

be the number of visits to  $y$  at positive times.

### Definition 19.3.1 (Recurrent)

A state  $y$  is said to be recurrent if  $\rho_{yy} = 1$ .

**Property.** The recurrent state  $y$  has the following properties

1.  $y$  is recurrent if and only if

$$E_y N(y) = \infty.$$

2. If  $x$  is recurrent and  $\rho_{xy} > 0$ , then  $y$  is recurrent and  $\rho_{yx} = 1$ .

**Definition 19.3.2**

A state  $y$  is said to be transient if  $\rho_{yy} < 1$ .

**Property.** The transient state  $y$  has the following properties

1. If  $y$  is transient, then

$$E_x N(y) < \infty, \quad \forall x.$$

*Proof.*

$$\begin{aligned} E_x N(y) &= \sum_{k=1}^{\infty} P_x(N(y) \geq k) = \sum_{k=1}^{\infty} P_x(T_y^k < \infty) \\ &= \sum_{k=1}^{\infty} \rho_{xy} \rho_{yy}^{k-1} = \frac{\rho_{xy}}{1 - \rho_{yy}} < \infty \end{aligned}$$

□

**Definition 19.3.3 (Closed State Set)**

A set  $C$  of states is said to be closed, if

$$x \in C, \rho_{xy} > 0 \Rightarrow y \in C. \quad (19.10)$$

**Definition 19.3.4 (Irreducible State Set)**

A set  $D$  of states is said to be irreducible, if

$$x, y \in D \Rightarrow \rho_{xy} > 0. \quad (19.11)$$

**Theorem 19.3.1**

Let  $C$  be a finite closed set, then

1.  $C$  contains a recurrent state.
2. If  $C$  is irreducible, then all states in  $C$  are recurrent.

**Theorem 19.3.2**

Suppose  $C_x = \{y : \rho_{xy} > 0\}$ , then  $C_x$  is an irreducible closed set.

*Proof.* If  $y, z \in C_x$ , then  $\rho_{yz} \geq \rho_{yx}\rho_{xz} > 0$ . If  $\rho_{yw} > 0$ , then  $\rho_{xw} \geq \rho_{xy}\rho_{yw} > 0$ , so  $w \in C_x$ . □

**Example** (A Seven-state Chain). Consider the transition probability,

	1	2	3	4	5	6	7
1	0.3	0	0	0	0.7	0	0
2	0.1	0.2	0.3	0.4	0	0	0
3	0	0	0.5	0.5	0	0	0
4	0	0	0	0.5	0	0.5	0
5	0.6	0	0	0	0.4	0	0
6	0	0	0	0.1	0	0.1	0.8
7	0	0	0	1	0	0	0

try to identify the recurrent states and those that are transient.

*Proof.*  $\{2, 3\}$  are transition states, and  $\{1, 4, 5, 6, 7\}$  are recurrent states.  $\square$

*Remark.* Suppose  $S$  is finite, for  $x \in S$ ,

1.  $x$  is transient, if

$$\exists y, \rho_{xy} > 0, \text{ s.t. } \rho_{yx} = 0$$

2.  $x$  is recurrent, if

$$\forall y, \rho_{xy} > 0, \text{ s.t. } \rho_{yx} > 0$$

## 19.4 Stationary Measures

## 19.5 Asymptotic Behavior

## 19.6 Ergodic Theorems

### Definition 19.6.1 (Stationary Sequence)

### Theorem 19.6.1 (Ergodic Theorem)

**Example.**

# Chapter 20

## Brownian Motion

### Definition 20.0.1 (Brownian Motion (1))

A real-valued stochastic process  $B(t), t \geq 0$  is said to be Brownian motion, if

1. for any  $0 = t_0 \leq t_1 \leq \dots \leq t_n$  the increments

$$B(t_1) - B(t_0), \dots, B(t_n) - B(t_{n-1})$$

are independent

2. for any  $s, t \geq 0$  and Borel sets  $A \in \mathbb{R}$ ,

$$P(B(s+t) - B(s) \in A) = \int_A (2\pi t)^{-1/2} \exp(-x^2/2t) dx \quad (20.1)$$

3. the sample paths  $t \rightarrow B(t)$  are a.s. continuous.

**Property.** For a one-dimensional Brownian motion, if  $B(0) = 0$ , then we have the following properties

1.  $EB_t = 0, \text{Var}(B_t) = t, \quad t \geq 0.$
2.  $\text{Cov}(B_s, B_t) = s, \text{Corr}(B_s, B_t) = \sqrt{s/t}, \quad \forall 0 \leq s \leq t.$

*Proof.* 1. Since  $B_t = B_t - B_0 \sim N(0, t)$ , then we have

$$EB_t = 0, \text{Var}(B_t) = t$$

2. Suppose  $0 \leq s \leq t$ ,

$$\text{Cov}(B_s, B_t) = E[(B_s - EB_s)(B_t - EB_t)] = EB_s B_t$$

Let  $B_t = (B_t - B_s) + B_s$ , we have

$$\begin{aligned} EB_s B_t &= E[B_s \cdot ((B_t - B_s) + B_s)] \\ &= E[B_s \cdot (B_t - B_s)] + EB_s^2 \end{aligned}$$

Since  $B_s = B_s - B_0$  and  $B_t - B_s$  are independent,

$$E[B_s \cdot (B_t - B_s)] = EB_s \cdot E[B_t - B_s] = 0$$

Thus

$$\text{Cov}(B_s, B_t) = EB_s^2 = s$$

And

$$\text{Corr}(B_s, B_t) = \frac{\text{Cov}(B_s, B_t)}{\sigma_{B_s}\sigma_{B_t}} = \frac{s}{\sqrt{st}} = \sqrt{\frac{s}{t}}$$

□

A second equivalent definition of Brownian motion is as follows,

### Definition 20.0.2 (Brownian Motion (2))

A real-valued stochastic process  $B(t), t \geq 0$ , **starting from 0**, is said to be Brownian motion, if

1.  $B(t)$  is a Gaussian process<sup>a</sup>
2.  $\forall s, t \geq 0, EB_s = 0$  and  $EB_s B_t = s \wedge t$
3. the sample paths  $t \rightarrow B(t)$  are a.s. continuous

---

<sup>a</sup>Gaussian process, i.e., all its finite-dimensional distributions are multivariate normal.

## 20.1 Markov Properties

## 20.2 Martingales

**Example** (Quadratic Martingale). Suppose  $B_t$  is a Brownian motion, then

$$B_t^2 - t$$

is a martingale.

*Proof.* Let  $B_t^2 = (B_s + B_t - B_s)^2$ , we have

$$\begin{aligned} E_x(B_t^2 | \mathcal{F}_s) &= E_x(B_s^2 + 2B_s(B_t - B_s) + (B_t - B_s)^2 | \mathcal{F}_s) \\ &= B_s^2 + 2B_s E_x(B_t - B_s | \mathcal{F}_s) + E_x((B_t - B_s)^2 | \mathcal{F}_s) \\ &= B_s^2 + 0 + (t - s) \end{aligned}$$

since  $B_t - B_s$  is independent of  $\mathcal{F}_s$  and has mean 0 and variance  $t - s$ . □

**Example** (Exponential Martingale). Suppose  $B_t$  is a Brownian motion, then

$$\exp\left(\theta B_t - \left(\theta^2 t/2\right)\right)$$

is a martingale.

*Proof.* Let  $B_t = B_t - B_s + B_s$ , then

$$\begin{aligned} E_x(\exp(\theta B_t) \mid \mathcal{F}_s) &= \exp(\theta B_s) E(\exp(\theta(B_t - B_s)) \mid \mathcal{F}_s) \\ &= \exp(\theta B_s) \exp(\theta^2(t-s)/2) \end{aligned}$$

since  $B_t - B_s$  is independent of  $\mathcal{F}_s$  and has mean 0 and variance  $t - s$ . Thus

$$\begin{aligned} E_x(\exp(\theta B_t - (\theta^2 t/2)) \mid \mathcal{F}_s) &= E_x(\exp(\theta B_t) \mid \mathcal{F}_s) \cdot \exp(-(\theta^2 t/2)) \\ &= \exp(\theta B_s - (\theta^2 s/2)) \end{aligned}$$

□

### Theorem 20.2.1 (Lévy's Martingale Characterization)

Let  $B(t), t \geq 0$ , be a real-valued stochastic process and let  $\mathcal{F}_t = \sigma(B_s, s \leq t)$  be the filtration generated by it. Then  $B(t)$  is a Brownian motion if and only if

1.  $B(0) = 0$  a.s.
2. the sample paths  $t \rightarrow B(t)$  are continuous a.s.
3.  $B(t)$  is a martingale with respect to  $\mathcal{F}_t$
4.  $|B(t)|^2 - t$  is a martingale with respect to  $\mathcal{F}_t$

## 20.3 Sample Paths

Let  $0 = t_0^n < t_1^n < \dots < t_n^n = T$ , where  $t_i^n = \frac{iT}{n}$  be a partition of the interval  $[0, T]$  into  $n$  equal parts, and

$$\Delta_i^n B = B(t_{i+1}^n) - B(t_i^n) \quad (20.2)$$

be the corresponding increments of the Brownian motion  $B(t)$ .

### Theorem 20.3.1

$$\lim_{n \rightarrow \infty} \sum_{i=0}^{n-1} (\Delta_i^n B)^2 = T \quad \text{in } L^2 \quad (20.3)$$

*Proof.* Since the increments  $\Delta_i^n B$  are independent and

$$E(\Delta_i^n B) = 0, \quad E((\Delta_i^n B)^2) = \frac{T}{n}, \quad E((\Delta_i^n B)^4) = \frac{3T^2}{n^2}$$

it follows that

$$\begin{aligned} E\left(\left[\sum_{i=0}^{n-1} (\Delta_i^n B)^2 - T\right]^2\right) &= E\left(\left[\sum_{i=0}^{n-1} \left((\Delta_i^n B)^2 - \frac{T}{n}\right)\right]^2\right) \\ &= \sum_{i=0}^{n-1} E\left[\left((\Delta_i^n B)^2 - \frac{T}{n}\right)^2\right] \\ &= \sum_{i=0}^{n-1} \left[E((\Delta_i^n B)^4) - \frac{2T}{n}E((\Delta_i^n B)^2) + \frac{T^2}{n^2}\right] \\ &= \sum_{i=0}^{n-1} \left[\frac{3T^2}{n^2} - \frac{2T^2}{n^2} + \frac{T^2}{n^2}\right] \\ &= \frac{2T^2}{n} \rightarrow 0, \quad n \rightarrow \infty \end{aligned}$$

□

### Definition 20.3.1 (Variation)

The variation of a function  $f : [0, T] \rightarrow \mathbb{R}$  is defined to be

$$\limsup_{\Delta t \rightarrow 0} \sum_{i=0}^{n-1} |f(t_{i+1}) - f(t_i)| \quad (20.4)$$

where  $t = (t_0, t_1, \dots, t_n)$  is a partition of  $[0, T]$ , i.e.  $0 = t_0 < t_1 < \dots < t_n = T$ , and where

$$\Delta t = \max_{i=0, \dots, n-1} |t_{i+1} - t_i| \quad (20.5)$$

### Theorem 20.3.2

The variation of the paths of  $B(t)$  is infinite a.s..

*Proof.* Consider the sequence of partitions  $t^n = (t_0^n, t_1^n, \dots, t_n^n)$  of  $[0, T]$  into  $n$  equal parts. Then

$$\sum_{i=0}^{n-1} |\Delta_i^n B|^2 \leq \left(\max_{i=0, \dots, n-1} |\Delta_i^n B|\right) \sum_{i=0}^{n-1} |\Delta_i^n B|$$

Since the paths of  $B(t)$  are a.s. continuous on  $[0, T]$ ,

$$\lim_{n \rightarrow \infty} \left(\max_{i=0, \dots, n-1} |\Delta_i^n B|\right) = 0 \quad \text{a.s.}$$



By Theorem 20.3.1, we have

$$\lim_{n \rightarrow \infty} \sum_{i=0}^{n-1} (\Delta_i^n B)^2 = T \text{ in } L^2$$

Since every sequence of random variables convergent in  $L^2$  has a subsequence convergent a.s. There is a subsequence  $t^{n_k} = (t_0^{n_k}, t_1^{n_k}, \dots, t_{n_k}^{n_k})$  of partitions such that

$$\lim_{k \rightarrow \infty} \sum_{i=0}^{n_k-1} |\Delta_i^{n_k} B|^2 = T \quad \text{a.s.}$$

Since

$$\sum_{i=0}^{n_k-1} |\Delta_i^{n_k} B| \geq \frac{\sum_{i=0}^{n_k-1} |\Delta_i^{n_k} B|^2}{\max_{i=0, \dots, n_k-1} |\Delta_i^{n_k} B|}$$

hence,

$$\lim_{k \rightarrow \infty} \sum_{i=0}^{n_k-1} |\Delta_i^{n_k} B| = \infty \quad \text{a.s.}$$

while

$$\lim_{k \rightarrow \infty} \Delta t^{n_k} = \lim_{k \rightarrow \infty} \frac{T}{n_k} = 0$$

□

## 20.4 Itô Stochastic Calculus

### Definition 20.4.1 (Itô Stochastic Integral)

For any  $T > 0$  we shall denote by  $M_T^2$  the space of all stochastic processes  $f(t), t \geq 0$  such that

$$1_{[0, T)} f \in M^2$$

The Itô stochastic integral (from 0 to  $T$ ) of  $f \in M_T^2$  is defined by

$$I_T(f) = I(1_{[0, T)} f) \quad (20.6)$$

which can be denoted by

$$\int_0^T f(t) dB(t) \quad (20.7)$$

**Property.** The Itô Stochastic Integral has the following properties:

1. Linearity: For  $\forall f, g \in M_t^2, \forall \alpha, \beta \in \mathbb{R}$ ,

$$\int_0^t (\alpha f(r) + \beta g(r)) dB(r) = \alpha \int_0^t f(r) dB(r) + \beta \int_0^t g(r) dB(r) \quad (20.8)$$

2. Isometry: For  $\forall f \in M_t^2$ ,

$$E \left( \left| \int_0^t f(r) dB(r) \right|^2 \right) = E \left( \int_0^t |f(r)|^2 dr \right) \quad (20.9)$$

3. Martingale Property: For  $\forall f \in M_t^2$  and  $\forall 0 \leq s < t$ ,

$$E \left( \int_0^t f(r) dB(r) \mid \mathcal{F}_s \right) = \int_0^s f(r) dB(r) \quad (20.10)$$

*Proof.*

□

#### Definition 20.4.2 (Itô Process)

A stochastic process  $\xi(t), t \geq 0$  is said to be an Itô process if it has a.s. continuous paths and can be represented as

$$\xi(T) = \xi(0) + \int_0^T a(t) dt + \int_0^T b(t) dB(t) \quad \text{a.s.} \quad (20.11)$$

where  $b(t)$  is a process belonging to  $M_T^2$  for all  $T > 0$  and  $a(t)$  is a process adapted to the filtration  $\mathcal{F}_t$  such that

$$\int_0^T |a(t)| dt < \infty \quad \text{a.s.} \quad (20.12)$$

for all  $T \geq 0$ . The Itô process is denoted by

$$d\xi(t) = a(t) dt + b(t) dB(t) \quad (20.13)$$

*Remark.* The class of all adapted processes  $a(t)$  satisfying 20.12 for some  $T > 0$  will be denoted by  $\mathcal{L}_T^1$ .

**Theorem 20.4.1 (Itô Formula)**

Suppose  $F(t, x)$  is a real-valued function with continuous partial derivatives  $F'_t(t, x)$ ,  $F'_x(t, x)$  and  $F''_{xx}(t, x)$  for all  $t \geq 0$  and  $x \in \mathbb{R}$ .

1. If  $\xi(t)$  be an Itô process

$$\xi(t) = \xi(0) + \int_0^t a(s) ds + \int_0^t b(s) dB(s)$$

and the process  $b(t)F'_x(t, \xi(t))$  belongs to  $M_T^2$  for all  $T \geq 0$ . Then  $F(t, \xi(t))$  is an Itô process such that

$$\begin{aligned} dF(t, \xi(t)) = & \left( F'_t(t, \xi(t)) + F'_x(t, \xi(t))a(t) + \frac{1}{2}F''_{xx}(t, \xi(t))b(t)^2 \right) dt \\ & + F'_x(t, \xi(t))b(t) dB(t) \end{aligned} \quad (20.14)$$

2. If  $\xi(t)$  be an Brownian Motion, such that  $\xi(t) = B(t)$ , and the process  $F'_x(t, B(t))$  belongs to  $M_T^2$  for all  $T \geq 0$ . Then  $F(t, B(t))$  is an Itô process such that

$$dF(t, B(t)) = \left( F'_t(t, B(t)) + \frac{1}{2}F''_{xx}(t, B(t)) \right) dt + F'_x(t, B(t)) dB(t) \quad (20.15)$$

**Example (Exponential Martingale).** Show that the exponential martingale

$$X(t) = e^{B(t)} e^{-\frac{t}{2}}$$

is an Itô process, and satisfies the equation

$$dX(t) = X(t) dB(t)$$

*Proof.* Let  $F(t, x) = e^x e^{-\frac{t}{2}}$ , then we have

$$F'_t(t, x) = -\frac{1}{2}F(t, x), \quad F'_x(t, x) = F(t, x), \quad F''_{xx}(t, x) = F(t, x)$$

thus, by Itô Formula, we have

$$\begin{aligned} dX(t) = dF(t, B(t)) &= \left( F'_t(t, B(t)) + \frac{1}{2}F''_{xx}(t, B(t)) \right) dt + F'_x(t, B(t)) dB(t) \\ &= \left( -\frac{1}{2}F(t, B(t)) + \frac{1}{2}F(t, B(t)) \right) dt + F(t, B(t)) dB(t) \\ &= X(t) dB(t) \end{aligned}$$

□

**Example.**

**Example.**

# Chapter 21

## Exercises for Probability Theory and Examples

21.1 Martingales

21.2 Markov Chains

21.3 Ergodic Theorems

21.4 Brownian Motion

21.5 Applications to Random Walk

21.6 Multidimensional Brownian Motion

# **Part VIII**

## **Empirical Process**

# Chapter 22

## 22.1 Concentration by Entropic Techniques

### Definition 22.1.1 (Entropy)

The entropy of a random variable  $X$  for the convex function  $\phi(\cdot)$  is defined as

$$H_\phi(X) = \mathbb{E}[\phi(X)] - \phi(\mathbb{E}[X]) \quad (22.1)$$

**Example.** 1. For  $\phi(u) = u^2$ ,  $H_\phi(X) = \text{Var}(X)$   
 2. For  $\phi(u) = -\log(u)$  ( $u > 0$ ), and for  $X$  real-valued random variable, we have that  $Z := \exp(\lambda X)$  is a non-negative random variable, and

$$H_\phi(Z) = -\lambda \mathbb{E}[X] + \log(\mathbb{E}[\exp(\lambda X)]) = \log \mathbb{E} e^{\lambda(X - \mathbb{E}[X])}. \quad (22.2)$$

### Definition 22.1.2

Let  $\Omega$  be a finite sample space and denote  $\mathcal{M}(\Omega)$  as the set of all probability measures (vectors) on  $\Omega$ .

1. The **relative entropy** with respect to  $q \in \mathcal{M}(\Omega)$  is defined as the mapping  $H(\cdot \mid q) : \mathcal{M}(\Omega) \rightarrow [0, \infty] : p \mapsto H(p \mid q)$ , where

$$H(p \mid q) = \begin{cases} \sum_{\omega \in \Omega} p(\omega) \log \left( \frac{p(\omega)}{q(\omega)} \right) & \text{if } p \ll q \\ +\infty & \text{otherwise} \end{cases} \quad (22.3)$$

2. The **Shannon entropy** of a  $\Omega$ -valued random variable  $X$  with distribution  $p \in \mathcal{M}(\Omega)$  is defined as

$$\mathcal{H}(p) = - \sum_{\omega \in \Omega} p(\omega) \log(p(\omega)). \quad (22.4)$$

### Proposition 22.1.1 (Duality formula of the Entropy)

**Lemma 22.1.1**

Let  $\Omega$  be a finite sample space and denote  $\mathcal{M}(\Omega)$  as the set of probability measures on  $\Omega$ . Let  $q \in \mathcal{M}(\Omega)$ , then the relative entropy  $H(\cdot \mid q)$  is strictly convex, continuous and

$$H(p \mid q) = 0 \iff p = q. \quad (22.5)$$

*Proof.* To prove this theorem, we analyze the properties of the relative entropy (also known as the Kullback-Leibler divergence). Let  $\Omega$  be a finite sample space,  $\mathcal{M}(\Omega)$  denote the set of probability measures on  $\Omega$ , and let  $q \in \mathcal{M}(\Omega)$  be a fixed probability measure. For any  $p \in \mathcal{M}(\Omega)$ , the relative entropy  $H(p \mid q)$  is defined as

$$H(p \mid q) = \sum_{\omega \in \Omega} p(\omega) \log \frac{p(\omega)}{q(\omega)},$$

where we assume  $q(\omega) > 0$  for all  $\omega \in \Omega$  (otherwise, the term is taken as 0).

**Non-negativity of Relative Entropy** First, observe that for any  $p \in \mathcal{M}(\Omega)$ , by Jensen's inequality, we have

$$\sum_{\omega \in \Omega} p(\omega) \log \frac{p(\omega)}{q(\omega)} \geq 0,$$

with equality if and only if  $p = q$ . Thus, we obtain  $H(p \mid q) \geq 0$ , and  $H(p \mid q) = 0 \iff p = q$ . This proves

$$H(p \mid q) = 0 \iff p = q.$$

**Strict Convexity of Relative Entropy** To prove that  $H(\cdot \mid q)$  is strictly convex, we consider the dependence of the relative entropy  $H(p \mid q)$  on  $p$ . Let  $p_1, p_2 \in \mathcal{M}(\Omega)$  with  $p_1 \neq p_2$ . For  $0 < \lambda < 1$ , define

$$p_\lambda = \lambda p_1 + (1 - \lambda) p_2.$$

Using the definition of relative entropy, we have

$$H(p_\lambda \mid q) = \sum_{\omega \in \Omega} p_\lambda(\omega) \log \frac{p_\lambda(\omega)}{q(\omega)}.$$

Applying Jensen's inequality for strictly convex functions, we obtain

$$H(p_\lambda \mid q) < \lambda H(p_1 \mid q) + (1 - \lambda) H(p_2 \mid q).$$

This shows that  $H(\cdot \mid q)$  is strictly convex.

**Continuity of Relative Entropy** Finally, we prove that the relative entropy  $H(p \mid q)$  is continuous with respect to  $p$ . Since  $\Omega$  is finite, in this finite-dimensional vector space,  $p \mapsto H(p \mid q)$  is a sum of a finite number of terms, each of which is a continuous function of  $p(\omega) \log(p(\omega)/q(\omega))$ . Thus,  $H(p \mid q)$  is continuous with respect to  $p$ . □

$$H(\exp(\lambda X)) = \lambda M'_X(\lambda) - M_X(\lambda) \log M_X(\lambda). \quad (22.6)$$

**Proposition 22.1.2 (Herbst argument)**

Let  $X$  be a random variable and suppose that for  $\sigma > 0$ ,

$$\mathcal{H}(e^{\lambda X}) \leq \frac{\lambda^2 \sigma^2}{2} M_X(\lambda), \quad (22.7)$$

for  $\lambda \in I$  with interval  $I$  being either  $\mathbb{R}$  or  $[0, \infty)$ . Then,

$$\log \mathbb{E} e^{\lambda(X - \mathbb{E}[X])} \leq \frac{\lambda^2 \sigma^2}{2}, \quad \forall \lambda \in I. \quad (22.8)$$

**Proposition 22.1.3 (Bernstein entropy bound)**

Suppose there exists  $B > 0$  and  $\sigma > 0$  such that

**Definition 22.1.3 (Separately convex)**

A function  $f : \mathbb{R}^n \rightarrow \mathbb{R}$  is said to be **separately convex** if for all  $k \in [n]$ , the function  $f(\cdot, x_{-k})$  is convex for all  $x_{-k} \in \mathbb{R}^{n-1}$ .

**Definition 22.1.4 (Lipschitz continuous)**

**Definition 22.1.5 (Globally Lipschitz continuous)**



**Lemma 22.1.2 (Entropy bound for univariate functions)**

Let  $X$  and  $Y$  two independent, identically distributed  $\mathbb{R}$ -valued random variables. Denote by  $\mathbb{E}_{X,Y}$  the expectation with respect to  $X$  and  $Y$ . For any function  $g : \mathbb{R} \rightarrow \mathbb{R}$  the following statements hold:

1.  $\forall \lambda > 0$ ,  $\mathcal{H}(e^{\lambda g(X)}) \leq \lambda^2 \mathbb{E}_{X,Y} [(g(X) - g(Y))^2 e^{\lambda g(X)} \mathbf{I}\{g(X) \geq g(Y)\}]$ .
2. If in addition the random variable  $X$  is supported on  $[a, b]$ ,  $a < b$ , and the function  $g$  is convex and Lipschitz continuous, then

$$\mathcal{H}(e^{\lambda g(X)}) \leq \lambda^2 (b - a)^2 \mathbb{E} [(g'(X))^2 e^{\lambda g(X)}], \quad \forall \lambda > 0.$$

*Proof.* Using the fact that  $X$  and  $Y$  are independent and identically distributed, we have

$$\mathcal{H}(e^{\lambda g(X)}) = \mathbb{E}_X [\lambda g(X) e^{\lambda g(X)}] - \mathbb{E}_X [\lambda g(X)] \log \mathbb{E}_Y e^{\lambda g(Y)}.$$

□

**Lemma 22.1.3 (Tensorisation of the entropy)**

Let  $X_1, \dots, X_n$  be independent real-valued random variables and  $f : \mathbb{R}^n \rightarrow \mathbb{R}$  be a given function. Then, for all  $\lambda > 0$ ,

$$\mathcal{H}(e^{\lambda f(X_1, \dots, X_n)}) \leq \sum_{i=1}^n \mathcal{H}(e^{\lambda f_i(X_i)} \mid \mathbf{X}_{-i}),$$

where  $f_i : \mathbb{R} \rightarrow \mathbb{R}$  is defined as  $f_i(x) = f(x_1, \dots, x_{i-1}, x, x_{i+1}, \dots, x_n)$ .

*Proof.* According to the variational representation of the entropy, we have

$$\mathcal{H}(e^{\lambda f(\mathbf{X})}) = \sup_{g \in \mathcal{G}} \left\{ \mathbb{E}[g(\mathbf{X}) e^{\lambda f(\mathbf{X})}] \right\},$$

where  $\mathcal{G} = \{g : \Omega \rightarrow \mathbb{R} : e^g \leq 1\}$ .

For each  $i \in [n]$ , define  $\mathbf{X}_i = (X_i, \dots, X_n)$  and for any  $g \in \mathcal{G}$  define  $g^i$ ,  $i \in [n]$ :

$$\begin{aligned} g^1(\mathbf{X}) &= g(\mathbf{X}) - \log \mathbb{E} [e^{g(\mathbf{X})} \mid \mathbf{X}_2], \\ g^i(\mathbf{X}_i) &= \log \frac{\mathbb{E} [e^{g(\mathbf{X})} \mid \mathbf{X}_i]}{\mathbb{E} [e^{g(\mathbf{X})} \mid \mathbf{X}_{i+1}]}, \quad i \in [n-1]. \end{aligned}$$

It is easy to see that by the above construction, we have

$$\sum_{i=1}^n g^i(\mathbf{X}) = g(\mathbf{X}) - \log \mathbb{E} [e^{g(\mathbf{X})}] \geq g(\mathbf{X}), \quad (22.9)$$

and

$$\mathbb{E} \left[ \exp \left( g^i(\mathbf{X}_i) \mid X_{i+1} \right) \right] = 1.$$

Within the variational representation of the entropy, we have

$$\begin{aligned} \mathbb{E} \left[ g(\mathbf{X}) e^{\lambda f(\mathbf{X})} \right] &\stackrel{(22.9)}{\leq} \mathbb{E} \left[ \sum_{i=1}^n g^i(\mathbf{X}) e^{\lambda f(\mathbf{X})} \right] \\ &= \sum_{i=1}^n \mathbb{E} \left[ g^i(\mathbf{X}) e^{\lambda f(\mathbf{X})} \right] \\ &\leq \end{aligned}$$

□

### Theorem 22.1.1 (Tail-bound for Lipschitz functions)

Let  $\mathbf{X} \in \mathbb{R}^n$  be a random vector with independent coordinates  $X_i$  supported on the interval  $[a, b]$ ,  $a < b$  and let  $f : \mathbb{R}^n \rightarrow \mathbb{R}$  be separately convex and  $L$ -Lipschitz continuous with respect to the Euclidean norm. Then, for all  $t > 0$ ,

$$\mathbb{P}(f(\mathbf{X}) - \mathbb{E}[f(\mathbf{X})] \geq t) \leq \exp \left( -\frac{t^2}{4L^2(b-a)^2} \right). \quad (22.10)$$

*Proof.* For  $i \in [n]$ , and every  $\mathbf{x}_{-i} \in \mathbb{R}^{n-1}$ , the function  $f_i(x)$  is convex, and thus by Lemma 22.1.2, for all  $\lambda > 0$  that for every fixed  $\mathbf{x}_{-i}$ , we have

$$\mathcal{H} \left( e^{\lambda f_i(X_i)} \mid \mathbf{X}_{-i} \right) \leq \lambda^2 (b-a)^2 \mathbb{E} \left[ (f'_i(X_i))^2 e^{\lambda f_i(X_i)} \mid \mathbf{X}_{-i} \right].$$

□

**Example** (Operator norm of a random matrix). Let  $M \in \mathbb{R}^{n \times d}$  be a random matrix with independent identically distributed mean-zero random entries  $M_{ij}$  supported on the interval  $[-1, 1]$ .

$$\|M\| = \max_{v \in \mathbb{S}^{d-1}} \|Mv\|_2 = \max_{u \in \mathbb{S}^{n-1}, v \in \mathbb{S}^{d-1}} \langle u, Mv \rangle. \quad (22.11)$$

The operator norm is maximin/supremum of functions that are linear in the entries of  $M$ , and thus any such function is convex and as such separately convex.

Moreover, for any  $M, M' \in \mathbb{R}^{n \times d}$ , we have

$$\begin{aligned}
 |||M|| - |||M'| || &= \left| \max_{u \in \mathbb{S}^{n-1}, v \in \mathbb{S}^{d-1}} \langle u, Mv \rangle - \max_{u \in \mathbb{S}^{n-1}, v \in \mathbb{S}^{d-1}} \langle u, M'v \rangle \right| \\
 &\leq \max_{u \in \mathbb{S}^{n-1}, v \in \mathbb{S}^{d-1}} |\langle u, Mv \rangle - \langle u, M'v \rangle| \\
 &\leq \max_{u \in \mathbb{S}^{n-1}, v \in \mathbb{S}^{d-1}} \|u\|_2 \|v\|_2 \|M - M'\|_F \\
 &\leq \|M - M'\|_F,
 \end{aligned}$$

where  $\|\cdot\|_F$  denotes the Frobenius norm, which is equivalent to the Euclidean norm for matrices. Thus, the operator norm is 1-Lipschitz continuous with respect to the Euclidean norm. Then by Theorem 22.1.1, we have

$$\mathbb{P}(|\|M|| - \mathbb{E}[||M||]| \geq t) \leq \exp\left(-\frac{t^2}{16}\right). \quad (22.12)$$

*Proof.* □

## 22.2 Some Matrix Calculus and Covariance Estimation

### Theorem 22.2.1 (Matrix Bernstein Inequality)

Let  $X_1, \dots, X_n$  be independent, mean-zero random systematic matrices in  $\mathbb{R}^{d \times d}$  such that  $\|X_i\| \leq K$  almost surely for all  $i \in [n]$ . Then for all  $t > 0$ , we have

$$\mathbb{P}\left(\left\|\sum_{i=1}^n X_i\right\| \geq t\right) \leq 2d \exp\left(-\frac{t^2}{\sigma^2 + Kt/3}\right),$$

where  $\sigma^2 = \|\sum_{i=1}^n \mathbb{E}[X_i^2]\|$  is the norm of the matrix variance of the sum.

*Proof.* Denote  $S_n = \sum_{i=1}^n X_i$  and  $\lambda_{\max}(S_n)$  as the largest eigenvalue of  $S_n$ . Then, we have

$$\|S_n\| = \max\{\lambda_{\max}(S_n), -\lambda_{\min}(S_n)\}.$$

Since

$$\mathbb{P}(\lambda_{\max}(S_n) \geq t) = \mathbb{P}(\exp(\lambda \lambda_{\max}(S_n)) \geq \exp(\lambda t)) \leq \frac{\mathbb{E}[\exp(\lambda \lambda_{\max}(S_n))]}{\exp(\lambda t)}.$$

□

**Lemma 22.2.1 (Bound on MGF)**

Let  $X$  be an  $d \times d$  symmetric mean-zero random matrix such that  $\|X\| \leq K$  almost surely. Then, for  $|\lambda| < 3/K$ , we have

$$\mathbb{E}[\exp(\lambda X)] \preceq \exp\left(g(\lambda)\mathbb{E}[X^2]\right),$$

where  $g(\lambda) = \frac{\lambda^2/2}{1-|\lambda|K/3}$ .

*Proof.*

□

**Proposition 22.2.1 (Expectation Bound via the Bernstein Inequality)**

Under the conditions of Theorem 22.2.1, we have the tail bound

$$\mathbb{P}\left(\left\|\sum_{i=1}^n X_i\right\| \geq t\right) \leq 2d \exp\left(-\frac{t^2}{\sigma^2 + Kt/3}\right).$$

Then,

$$\mathbb{E}\left[\left\|\sum_{i=1}^n X_i\right\|\right] \leq$$

# **Part IX**

## **Random Matrix Theory**

# Chapter 23

## Preliminary

### 23.1 Empirical Spectral Measure

#### Definition 23.1.1 (Empirical Spectral Measure)

For a symmetric matrix  $\mathbf{M} \in \mathbb{R}^{n \times n}$ , the spectral measure or empirical spectral measure or empirical spectral distribution (ESD)  $\mu_{\mathbf{M}}$  of  $\mathbf{M}$  is defined as the normalized counting measure of the eigenvalues  $\lambda_1(\mathbf{M}), \dots, \lambda_n(\mathbf{M})$  of  $\mathbf{M}$ , i.e.,

$$\mu_{\mathbf{M}} := \frac{1}{n} \sum_{i=1}^n \delta_{\lambda_i(\mathbf{M})} \quad (23.1)$$

where  $\delta_x$  is a Dirac measure for any (measurable) set, that

$$\delta_x(A) := \mathbf{1}_A(x) = \begin{cases} 0, & x \notin A \\ 1, & x \in A \end{cases}$$

Since  $\int \mu_{\mathbf{M}}(dx) = 1$ , the spectral measure  $\mu_{\mathbf{M}}$  of a matrix  $\mathbf{M} \in \mathbb{R}^{n \times n}$  (random or not) is a probability measure.

*Remark.* Many important statistics in multivariate analysis can be expressed as functionals of the ESD, such as, for  $\mathbf{M}$  be an  $n \times n$  positive definite matrix, then

$$\det(\mathbf{M}) = \prod_{i=1}^n \lambda_i = \exp \left( n \int_0^\infty \log x \mu_{\mathbf{M}}(dx) \right) \quad (23.2)$$

## 23.2 Stieltjes Transform

### Definition 23.2.1 (Resolvent)

For a symmetric matrix  $\mathbf{M} \in \mathbb{R}^{n \times n}$ , the resolvent  $\mathbf{Q}_{\mathbf{M}}(z)$  of  $\mathbf{M}$  is defined as

$$\mathbf{Q}_{\mathbf{M}}(z) := (\mathbf{M} - z\mathbf{I}_n)^{-1} \quad (23.3)$$

where  $z \in \mathbb{C}$  not eigenvalue of  $\mathbf{M}$ .

### Definition 23.2.2 (Stieltjes Transform)

For a real probability measure  $\mu$  with support  $\text{supp}(\mu)$ , the Stieltjes transform  $m_\mu(z)$  is defined as

$$m_\mu(z) := \int \frac{1}{t - z} \mu(dt) \quad (23.4)$$

where  $z \in \mathbb{C} \setminus \text{supp}(\mu)$ .

**Property.** The Stieltjes transform  $m_\mu$  has numerous interesting properties:

1. it is complex analytic on its domain of definition  $\mathbb{C} \setminus \text{supp}(\mu)$ .
2. it is bounded  $|m_\mu(z)| \leq 1/\text{dist}(z, \text{supp}(\mu))$ .
3. it satisfies  $\Im[z] > 0 \Rightarrow \Im[m_\mu(z)] > 0$ .
4. it is an increasing function on all connected components of its restriction to  $\mathbb{R} \setminus \text{supp}(\mu)$ .
5. if  $\text{supp}(\mu)$  is bounded,  $\lim_{x \rightarrow \pm\infty} m_\mu(x) = 0$ .

*Remark.* Most of the results involve Stieltjes transforms  $m_\mu(z)$  of a real probability measure with support  $\text{supp}(\mu) \subset \mathbb{R}$ . Since Stieltjes transforms are such that

$$m_\mu(z) > 0, \forall z < \inf \text{supp}(\mu), \quad m_\mu(z) < 0, \forall z > \sup \text{supp}(\mu), \quad \Im[z]\Im[m_\mu(z)] > 0, \text{ if } z \in \mathbb{C} \setminus \mathbb{R}$$

it will be convenient in the following to consider the set of scalar pairs

$$\mathcal{Z}(\mathcal{A}) = \{(z, m) \in \mathcal{A} \times \mathbb{C}, (\Im[z]\Im[m] > 0 \text{ if } \Im[z] \neq 0) \text{ or } (m > 0 \text{ if } z < \inf \mathcal{A}^c \cap \mathbb{R}) \\ \text{or } (m < 0 \text{ if } z > \sup \mathcal{A}^c \cap \mathbb{R})\}$$

As a transform,  $m_\mu$  has an inverse formula to recover  $\mu$ , as per the following result.

**Theorem 23.2.1 (Inverse Stieltjes Transform)**

For  $a, b$  continuity points of the probability measure  $\mu$ , we have

$$\mu([a, b]) = \frac{1}{\pi} \lim_{y \downarrow 0} \int_a^b \Im [m_\mu(x + iy)] dx \quad (23.5)$$

Specially, if  $\mu$  has a density  $f$  at  $x$ , then

$$f(x) = \frac{1}{\pi} \lim_{y \downarrow 0} \Im [m_\mu(x + iy)] \quad (23.6)$$

And, if  $\mu$  has an isolated mass at  $x$ , then

$$\mu(\{x\}) = \lim_{y \downarrow 0} -iy m_\mu(x + iy) \quad (23.7)$$

*Proof.*

$$\begin{aligned} \frac{1}{\pi} \int_a^b \Im [m_\mu(x + iy)] dx &= \frac{1}{\pi} \int_a^b \left\{ \int \Im \left[ \frac{1}{(t - x) - iy} \right] \mu(dt) \right\} dx \\ &= \frac{1}{\pi} \int_a^b \left[ \int \frac{y}{(t - x)^2 + y^2} \mu(dt) \right] dx \end{aligned}$$

By Fubini theorem,

$$\begin{aligned} &= \frac{1}{\pi} \int \left[ \int_a^b \frac{y}{(t - x)^2 + y^2} dx \right] \mu(dt) \\ &= \frac{1}{\pi} \int \left[ \arctan \left( \frac{b - t}{y} \right) - \arctan \left( \frac{a - t}{y} \right) \right] \mu(dt) \end{aligned}$$

Since

$$\left| \frac{y}{(t - x)^2 + y^2} \right| \leq \frac{1}{y}, \quad \forall y > 0$$

by the dominated convergence theorem,

$$\frac{1}{\pi} \lim_{y \downarrow 0} \int_a^b \Im [m_\mu(x + iy)] dx = \frac{1}{\pi} \int \lim_{y \downarrow 0} \left[ \arctan \left( \frac{b - t}{y} \right) - \arctan \left( \frac{a - t}{y} \right) \right] \mu(dt)$$

as  $y \downarrow 0$ , the difference in brackets converges either to  $\pm\pi$  or 0 depending on the relative position of  $a, b$  and  $t$ , thus

$$= \int 1_{[a, b]} \mu(dt) = \mu([a, b])$$



Thus, if  $\mu$  has a density  $f$  at  $x$ , then

$$f(x) = \frac{1}{\pi} \lim_{y \downarrow 0} \Im [m_\mu(x + iy)]$$

When  $\mu$  has an isolated mass at  $x$ , i.e.,  $\mu(dt) = a\delta_x(t)$ , similarly, since

$$|y(t - x)| \leq \frac{1}{2} (y^2 + (t - x)^2)$$

by dominated convergence theorem,

$$\lim_{y \downarrow 0} -iy m_\mu(x + iy) = - \lim_{y \downarrow 0} \int \frac{iy(t - x)\mu(dt)}{(t - x)^2 + y^2} + \lim_{y \downarrow 0} \int \frac{y^2\mu(dt)}{(t - x)^2 + y^2} = a$$

□

*Remark.* The important relation between the empirical spectral measure  $\mu_{\mathbf{M}}$  of  $\mathbf{M} \in \mathbb{R}^{n \times n}$ , the Stieltjes transform  $m_{\mu_{\mathbf{M}}}(z)$  and the resolvent  $\mathbf{Q}_{\mathbf{M}}(z)$  lies in the fact that

$$m_{\mu_{\mathbf{M}}}(z) = \frac{1}{n} \sum_{i=1}^n \int \frac{\delta_{\lambda_i(\mathbf{M})}(t)}{t - z} = \frac{1}{n} \sum_{i=1}^n \frac{1}{\lambda_i(\mathbf{M}) - z} = \frac{1}{n} \text{tr } \mathbf{Q}_{\mathbf{M}}(z) \quad (23.8)$$

The resolvent  $\mathbf{Q}_{\mathbf{M}}$  provides access to scalar observations of the eigenspectrum of  $\mathbf{M}$  through its linear functionals. Cauchy's integral formula provides a connection between the linear functionals of the eigenvalues of  $\mathbf{M}$  and the Stieltjes transform  $m_{\mu_{\mathbf{M}}}(z)$  through

$$\frac{1}{n} \sum_{i=1}^n f(\lambda_i(\mathbf{M})) = -\frac{1}{2\pi i n} \oint_{\Gamma} f(z) \text{tr}(\mathbf{Q}_{\mathbf{M}}(z)) dz = -\frac{1}{2\pi i} \oint_{\Gamma} f(z) m_{\mu_{\mathbf{M}}}(z) dz \quad (23.9)$$

for all  $f$  complex analytic in a compact neighborhood of  $\text{supp}(\mu_{\mathbf{M}})$ , by choosing the contour  $\Gamma$  to enclose  $\text{supp}(\mu_{\mathbf{M}})$  (i.e., all the eigenvalues  $\lambda_i(\mathbf{M})$ ).

## 23.3 Matrix Equivalents

### Definition 23.3.1 (Deterministic Equivalent)

$\bar{\mathbf{Q}} \in \mathbb{R}^{n \times n}$  is said to be deterministic equivalent for the symmetric random matrix  $\mathbf{Q} \in \mathbb{R}^{n \times n}$ , if for a (sequences of) deterministic matrix  $\mathbf{A} \in \mathbb{R}^{n \times n}$  and vectors  $\mathbf{a}, \mathbf{b} \in \mathbb{R}^n$  of unit norms (operator and Euclidean, respectively),

$$\frac{1}{n} \text{tr } \mathbf{A}(\mathbf{Q} - \bar{\mathbf{Q}}) \rightarrow 0, \quad \mathbf{a}'(\mathbf{Q} - \bar{\mathbf{Q}})\mathbf{b} \rightarrow 0, \quad \text{as } n \rightarrow \infty \quad (23.10)$$

where the convergence is either in probability or almost sure.

*Remark.* A practical use of deterministic equivalents is to establish that, for a random matrix  $\mathbf{M}$  of interest, suppose

$$\frac{1}{n} \operatorname{tr} (\mathbf{Q}_{\mathbf{M}}(z) - \overline{\mathbf{Q}}(z)) \rightarrow 0, \quad \text{a.s.}, \quad \forall z \in \mathcal{C}, \mathcal{C} \subset \mathbb{C}$$

this convergence implies that the Stieltjes transform of  $\mu_{\mathbf{M}}$  "converges" in the sense that

$$m_{\mu_{\mathbf{M}}}(z) - \bar{m}_n(z) \rightarrow 0$$

where  $\bar{m}_n(z) = \frac{1}{n} \operatorname{tr} \overline{\mathbf{Q}}(z)$ .

### Definition 23.3.2 (Matrix Equivalents)

For  $\mathbf{x}, \mathbf{Y} \in \mathbb{R}^{n \times n}$  two random or deterministic matrices, we write

$$\mathbf{x} \leftrightarrow \mathbf{Y} \quad (23.11)$$

if, for all  $\mathbf{A} \in \mathbb{R}^{n \times n}$  and  $\mathbf{a}, \mathbf{b} \in \mathbb{R}^n$  of unit norms (respectively, operator and Euclidean), we have the simultaneous results

$$\frac{1}{n} \operatorname{tr} \mathbf{A}(\mathbf{x} - \mathbf{Y}) \rightarrow 0, \quad \mathbf{a}'(\mathbf{x} - \mathbf{Y})\mathbf{b} \rightarrow 0, \quad \|\mathbb{E}[\mathbf{x} - \mathbf{Y}]\| \rightarrow 0$$

where, for random quantities, the convergence is either in probability or almost sure.

## 23.4 Resolvent and Perturbation Identities

### Lemma 23.4.1 (Resolvent Identity)

For invertible matrices  $\mathbf{A}$  and  $\mathbf{B}$ , we have

$$\mathbf{A}^{-1} - \mathbf{B}^{-1} = \mathbf{A}^{-1} (\mathbf{B} - \mathbf{A}) \mathbf{B}^{-1} \quad (23.12)$$

### Lemma 23.4.2 (Sherman-Morrison)

For  $\mathbf{A} \in \mathbb{R}^{n \times n}$  invertible and  $\mathbf{u}, \mathbf{v} \in \mathbb{R}^n$ , then  $\mathbf{A} + \mathbf{u}\mathbf{v}'$  is invertible if and only if  $1 + \mathbf{v}'\mathbf{A}^{-1}\mathbf{u} \neq 0$  and

$$(\mathbf{A} + \mathbf{u}\mathbf{v}')^{-1} = \mathbf{A}^{-1} - \frac{\mathbf{A}^{-1}\mathbf{u}\mathbf{v}'\mathbf{A}^{-1}}{1 + \mathbf{v}'\mathbf{A}^{-1}\mathbf{u}} \quad (23.13)$$

or,

$$(\mathbf{A} + \mathbf{u}\mathbf{v}')^{-1} \mathbf{u} = \frac{\mathbf{A}^{-1}\mathbf{u}}{1 + \mathbf{v}'\mathbf{A}^{-1}\mathbf{u}} \quad (23.14)$$

**Lemma 23.4.3 (Quadratic-form-close-to-the-trace)**

Let  $\mathbf{x} \in \mathbb{R}^p$  have iid entries of zero mean, unit variance and  $\mathbb{E}[|x_i|^K] \leq \nu_K$  for some  $K \geq 1$ . Then for  $\mathbf{A} \in \mathbb{R}^{p \times p}$  and  $k \geq 1$

$$\mathbb{E} \left[ \left| \mathbf{x}^\top \mathbf{A} \mathbf{x} - \text{tr } \mathbf{A} \right|^k \right] \leq C_k \left[ (\nu_4 \text{tr}(\mathbf{A} \mathbf{A}'))^{k/2} + \nu_{2k} \text{tr}(\mathbf{A} \mathbf{A}')^{k/2} \right]$$

for some constant  $C_k > 0$  independent of  $p$ . In particular, if  $\|\mathbf{A}\| \leq 1$  and the entries of  $\mathbf{x}$  have bounded eighth-order moment,

$$\mathbb{E} \left[ \left( \mathbf{x}^\top \mathbf{A} \mathbf{x} - \text{tr } \mathbf{A} \right)^4 \right] \leq Cp^2$$

for some  $C > 0$  independent of  $p$ , and consequently, as  $p \rightarrow \infty$ ,

$$\frac{1}{p} \mathbf{x}^\top \mathbf{A} \mathbf{x} - \frac{1}{p} \text{tr } \mathbf{A} \xrightarrow{\text{a.s.}} 0$$

# Chapter 24

## Wigner Matrix

# Chapter 25

## Sample Covariance Matrix

Suppose  $\{\mathbf{x}\}$  be a sequence of random vectors defined in  $\mathbb{R}^n$ , and  $(x_i)_{1 \leq i \leq n}$  be the components of the random vector  $\mathbf{x}$ , such that

$$E(\mathbf{x}) = 0, \quad E(\mathbf{x} \otimes \mathbf{x}) = \mathbf{I}_n$$

where  $\mathbf{x}$  is also called **isotropic** random vector.

Suppose  $\{m_n\}$  be a sequence defined in  $\mathbb{N}$  such that

$$0 < \underline{\rho} := \liminf_{n \rightarrow \infty} \frac{n}{m_n} \leq \limsup_{n \rightarrow \infty} \frac{n}{m_n} =: \bar{\rho} < \infty$$

Let  $\mathbf{x}_1, \dots, \mathbf{x}_{m_n}$  be iid copies of  $\mathbf{x}$ , and  $\mathbb{X}$  be the  $m_n \times n$  random matrix with iid rows  $\mathbf{x}_1, \dots, \mathbf{x}_{m_n}$ , and their empirical covariance matrix is

$$\hat{\Sigma} := \frac{1}{m_n} \sum_{i=1}^{m_n} \mathbf{x}_i \otimes \mathbf{x}_i = \frac{1}{m_n} \mathbb{X}' \mathbb{X}$$

which is a  $n \times n$  symmetric positive semidefinite random matrix, and

$$E(\hat{\Sigma}) = \mathbb{E}(\mathbf{x} \otimes \mathbf{x}) = \mathbf{I}_n$$

For convenience, we define the random matrix

$$\mathbf{A} := m_n \hat{\Sigma} = \mathbb{X}' \mathbb{X} = \sum_{i=1}^{m_n} \mathbf{x}_i \otimes \mathbf{x}_i$$

## 25.1 Eigenvalues and Singular Values

### Theorem 25.1.1

The eigenvalues of  $\mathbf{A}$  are squares of the singular values of  $\mathbb{X}$ , in particular

$$\lambda_{\max}(\mathbf{A}) = s_{\max}(\mathbb{X})^2 = \max_{\|\mathbf{x}\|=1} \|\mathbb{X}\mathbf{x}\|^2 = \|\mathbb{X}\|_2^2$$

if  $m_n \geq n$ , then

$$\lambda_{\min}(\mathbf{A}) = s_{\min}(\mathbb{X})^2 = \min_{\|\mathbf{x}\|=1} \|\mathbb{X}\mathbf{x}\|^2 = \|\mathbb{X}^{-1}\|_2^{-2}$$

*Proof.*

□

## 25.2 Laguerre Orthogonal Ensemble

### Definition 25.2.1 (Wishart Distribution)

Suppose  $\mathbb{X}$  be a  $p \times n$  matrix, each column of which is independently drawn from a  $p$ -variate normal distribution with zero means:

$$\mathbf{x}_i = (x_i^1, \dots, x_i^p)' \sim N_p(0, \Sigma)$$

Then the Wishart distribution is the probability distribution of the  $p \times p$  random matrix,

$$\mathbf{M} = \mathbb{X}'\mathbb{X} = \sum_{i=1}^n \mathbf{x}_i \mathbf{x}_i^\top \quad (25.1)$$

and which can be denoted by

$$\mathbf{M} \sim W_p(\Sigma, n)$$

If  $p = \Sigma = 1$ , then this distribution is a chi-squared distribution with  $n$  degrees of freedom.

**Theorem 25.2.1**

If  $n \geq p$ , the probability density function of  $\mathbf{M}$  is

$$f(\mathbf{M}) = \frac{1}{2^{np/2} [\det(\boldsymbol{\Sigma})]^{n/2} \Gamma_p\left(\frac{n}{2}\right)} \det(\mathbf{M})^{(n-p-1)/2} \exp\left[-\frac{1}{2} \text{tr}(\boldsymbol{\Sigma}^{-1} \mathbf{M})\right] \quad (25.2)$$

concerning the Lebesgue measure on the cone of symmetric positive definite matrices. Here,  $\Gamma_p$  is the multivariate gamma function defined as

$$\Gamma_p\left(\frac{n}{2}\right) = \pi^{p(p-1)/4} \prod_{j=1}^p \Gamma\left(\frac{n}{2} - \frac{j-1}{2}\right)$$

*Remark.* Especially, if the random variables  $(x_i)_{1 \leq i \leq n}$  are iid standard Gaussians, then the distribution of the random matrix  $\hat{\boldsymbol{\Sigma}}$  can be derived from the Wishart distribution. The probability density function of  $\hat{\boldsymbol{\Sigma}}$  can be derived from (25.2), since

$$\mathbf{A} \sim W_n(\mathbf{I}_n, m_n), \quad \det(\hat{\boldsymbol{\Sigma}}) = m_n^{-n} \det(\mathbf{A}), \quad \text{tr}(\hat{\boldsymbol{\Sigma}}) = m_n^{-1} \text{tr}(\mathbf{A})$$

thus,

$$f(\hat{\boldsymbol{\Sigma}}) = \frac{m_n^{-n(m_n-n-1)/2+1}}{2^{m_n n/2} \Gamma_n\left(\frac{m_n}{2}\right)} \det(\hat{\boldsymbol{\Sigma}})^{(m_n-n-1)/2} \exp\left[-\frac{m_n}{2} \text{tr}(\hat{\boldsymbol{\Sigma}})\right] \quad (25.3)$$

**Theorem 25.2.2**

If the random variables  $(x_i)_{1 \leq i \leq n}$  are iid standard Gaussians, the joint probability density function of eigenvalues of  $\hat{\boldsymbol{\Sigma}}$  is

$$p(\boldsymbol{\Lambda}) = \tilde{Q}_{m_n, n}^{-1} \exp\left(-\frac{m_n}{2} \sum_{k=1}^n \lambda_k\right) \prod_{k=1}^n \lambda_k^{(m_n-n-1)/2} \prod_{i < j} |\lambda_i - \lambda_j| \quad (25.4)$$

where

$$0 \leq \lambda_1 \leq \dots \leq \lambda_n < \infty$$

and  $\tilde{Q}_{m_n, n}$  is the normalization constant.

*Proof.* First, we will give the characteristic function of  $\hat{\boldsymbol{\Sigma}}$ , i.e.,

$$\varphi_{\hat{\boldsymbol{\Sigma}}}(\mathbf{P}) = E \left[ \exp \left( i \sum_{1 \leq i \leq j \leq n} P_{ij} \hat{\boldsymbol{\Sigma}}_{ji} \right) \right] = E \left[ \exp \left( i \text{tr}(\mathbf{P} \hat{\boldsymbol{\Sigma}}) \right) \right]$$

where  $\{P_{ij}\}_{1 \leq i \leq j \leq n} \in \mathbb{R}^{(n+1)n/2}$  and  $\mathbf{P}$  is a real symmetric matrix, that

$$\mathbf{P} = \left\{ \hat{P}_{ij}, \hat{P}_{ij} = \hat{P}_{ji} \right\}_{i,j=1}^n, \quad \hat{P}_{ij} = \begin{cases} P_{ii}, & i = j \\ P_{ij}/2, & i < j \end{cases}$$

Thus, we have

$$\begin{aligned} &= \int_{\mathbb{R}^{m_n \times n}} \exp \left( i \operatorname{tr} (\mathbf{P} \hat{\Sigma}) \right) \cdot (2\pi)^{-m_n n/2} \exp \left( -\frac{1}{2} \sum_{k=1}^{m_n} \sum_{i=1}^n \left( x_i^{(k)} \right)^2 \right) \prod_{k=1}^{m_n} \prod_{i=1}^n dx_i^{(k)} \\ &= \int_{\mathbb{R}^{m_n \times n}} (2\pi)^{-m_n n/2} \exp \left( -\frac{1}{2} \sum_{k=1}^{m_n} \sum_{i=1}^n \sum_{j=1}^n \mathbf{Q}_{ij} x_i^{(k)} x_j^{(k)} \right) \prod_{k=1}^{m_n} \prod_{i=1}^n dx_i^{(k)} \end{aligned}$$

where

$$\mathbf{Q} = \mathbf{I}_n - \frac{2i}{m_n} \mathbf{P}$$

Since  $(x_i^{(k)})_{1 \leq i \leq n}$  are iid standard Gaussians,

$$\begin{aligned} &= \left[ \int_{\mathbb{R}^n} (2\pi)^{-n/2} \exp \left( -\frac{1}{2} \sum_{i=1}^n \sum_{j=1}^n \mathbf{Q}_{ij} x_i x_j \right) \prod_{i=1}^n dx_i \right]^{m_n} \\ &= \left[ \int_{\mathbb{R}^n} (2\pi)^{-n/2} \exp \left( -\frac{1}{2} \mathbf{x}^\top \mathbf{Q} \mathbf{x} \right) d\mathbf{x} \right]^{m_n} \\ &= \left[ \det(\mathbf{Q})^{-\frac{1}{2}} \int_{\mathbb{R}^n} (2\pi)^{-n/2} \exp \left( -\frac{1}{2} (\mathbf{Q}^{\frac{1}{2}} \mathbf{x})' (\mathbf{Q}^{\frac{1}{2}} \mathbf{x}) \right) d\mathbf{Q}^{\frac{1}{2}} \mathbf{x} \right]^{m_n} \\ &= [\det(\mathbf{Q})]^{-m_n/2} \end{aligned}$$

thus,

$$[\det(\mathbf{Q})]^{-m_n/2} = \left[ \det \left( \mathbf{I}_n - \frac{2i}{m_n} \mathbf{P} \right) \right]^{-m_n/2} = \prod_{k=1}^n \left( 1 - \frac{2i}{m_n} p_k \right)^{-m_n/2} \quad (25.5)$$

where  $\{p_k\}_{k=1}^n$  are the eigenvalues of  $\mathbf{P}$ .

Then, we will show that the characteristic function of (25.4) coincides with the above function. By the Wishart distribution, the probability density of the real symmetric and positive definite random matrix  $\hat{\Sigma}$  is

$$\tilde{Q}_{m_n, n}^{-1} \exp \left[ -\frac{m_n}{2} \operatorname{tr} (\hat{\Sigma}) \right] [\det(\hat{\Sigma})]^{(m_n - n - 1)/2} d\hat{\Sigma} \quad (25.6)$$

where  $\tilde{Q}_{m_n, n}$  is the normalization constant. Then, the characteristic function of (25.6), i.e.,

$$\tilde{Q}_{m_n, n}^{-1} \int_{S_n^+} \exp \left[ i \operatorname{tr} (\mathbf{P} \hat{\Sigma}) - \frac{m_n}{2} \operatorname{tr} (\hat{\Sigma}) \right] [\det(\hat{\Sigma})]^{(m_n - n - 1)/2} d\hat{\Sigma}$$



where the integration is over the set  $\mathcal{S}_n^+$  of  $n \times n$  real symmetric and positive definite matrices. Since

$$\sum_{k=1}^n \lambda_k = \text{tr}(\widehat{\Sigma}), \quad \prod_{k=1}^n \lambda_k^{(m_n-n-1)/2} = [\det(\widehat{\Sigma})]^{(m_n-n-1)/2}$$

and

$$d\widehat{\Sigma} = \prod_{i < j} |\lambda_i - \lambda_j| d\mathbf{\Lambda} H_1(dO)$$

where  $H_1$  is the normalized Haar measure of  $O(n)$ , and the integration over  $\mathbf{\Lambda}$  and  $O \in O(n)$  are independent. Since the orthogonal invariance of the density of (25.6), and the characteristic function is

$$Q_{m_n, n}^{-1} \int_{(\mathbb{R}_+)^n} \exp \left[ \sum_{k=1}^n \left( \nu p_k - \frac{m_n}{2} \right) \lambda_k \right] \prod_{k=1}^n \lambda_k^{(m_n-n-1)/2} \prod_{i < j} |\lambda_i - \lambda_j| d\mathbf{\Lambda} \quad (25.7)$$

where  $Q_{m_n, n} = m_n! \tilde{Q}_{m_n, n}$ .

If we viewed (25.5) and (25.7) as the function of  $\{p_k\}_{k=1}^n \in \mathbb{R}^n$ , then they can be **analytic continuation** to the domain

$$\{p_k + \nu p'_k, p'_k \geq 0\}_{k=1}^n$$

If we replace  $\{p_k\}_{k=1}^n$  by  $\{\nu p'_k, p'_k \geq 0\}_{k=1}^n$  on (25.5) since this is a set of the uniqueness of both (25.5) and `eqrefeq:characteristic-function-wishart` analytic functions, we have

$$Q_{m_n, n}^{-1} \int_{(\mathbb{R}_+)^n} \exp \left[ -\frac{m_n}{2} \sum_{k=1}^n q_k \lambda_k \right] \prod_{k=1}^n \lambda_k^{(m_n-n-1)/2} \prod_{i < j} |\lambda_i - \lambda_j| d\mathbf{\Lambda}$$

where  $q_k = 1 + \frac{2p'_k}{m_n} \geq 1, k = 1, \dots, n$ , and since

$$\forall i, j \quad \frac{q_i}{q_j} = \frac{1 + \frac{2p'_i}{m_n}}{1 + \frac{2p'_j}{m_n}} \rightarrow 1, \text{ as } m_n \rightarrow \infty$$

we have

$$\prod_{i < j} |q_i \lambda_i - q_j \lambda_j| = \prod_{i < j} q_i \left| \lambda_i - \frac{q_j}{q_i} \lambda_j \right| \rightarrow \prod_{k=1}^n q_k^{(n-1)/2} \prod_{i < j} |\lambda_i - \lambda_j|, \text{ as } m_n \rightarrow \infty$$

thus,

$$\prod_{k=1}^n q_k^{-m_n/2} \cdot Q_{m_n, n}^{-1} \int_{(\mathbb{R}_+)^n} \exp \left[ -\frac{m_n}{2} \sum_{k=1}^n q_k \lambda_k \right] \prod_{k=1}^n (q_k \lambda_k)^{(m_n-n-1)/2} \prod_{i < j} |q_i \lambda_i - q_j \lambda_j| d\mathbf{q} d\mathbf{\Lambda}$$

Since

$$\forall k \quad q_k \lambda_k \rightarrow \lambda_k, \text{ as } m_n \rightarrow \infty$$

we can "lifting" from  $\{\lambda_k\}_{k=1}^n$  to  $\mathcal{S}_n^+$  bring the integral to

$$\prod_{k=1}^n \left(1 + \frac{2p'_k}{m_n}\right)^{-m_n/2} \tilde{Q}_n^{-1} \int_{\mathcal{S}_n^+} \exp \left[ -\frac{m_n}{2} \text{tr}(\hat{\Sigma}) \right] [\det(\hat{\Sigma})]^{(m_n-n-1)/2} d\hat{\Sigma}$$

The integral here is equal to  $\tilde{Q}_n$ , the normalization constant of the probability measure (25.6). If we replace  $\{ip'_k\}_{k=1}^n$  back by  $\{p_k\}_{k=1}^n$ , then the above expression is

$$\prod_{k=1}^n \left(1 - \frac{2p_k}{m_n}\right)^{-m_n/2}$$

which coincides with (25.5). Thus the probability law of the Wishart matrices of  $\Sigma$  given by (25.6) implies that the corresponding joint probability density of eigenvalues is given by (25.4) for  $\Sigma$ .  $\square$

### Definition 25.2.2 (Laguerre Orthogonal Ensemble)

For the  $n \times n$  Laguerre orthogonal ensembles of statistics, the joint probability density function of eigenvalues is for arbitrary parameter  $\beta > 0$  and  $\alpha > -\frac{2}{\beta}$ , is

$$p(\Lambda) = K_{\alpha, \beta} \exp \left( -\frac{\beta}{2} \sum_{k=1}^n \lambda_k \right) \prod_{k=1}^n \lambda_k^{\frac{\alpha\beta}{2}} \prod_{i < j} |\lambda_i - \lambda_j|^\beta \quad (25.8)$$

where

$$0 \leq \lambda_1 \leq \dots \leq \lambda_n < \infty$$

and  $K_{n, m}$  are normalization constant.

And Equation (25.8) can be written in the standard Boltzmann-Gibbs form, that,

$$p(\Lambda) \propto \exp[-\beta E(\Lambda)]$$

where

$$E(\Lambda) = \frac{1}{2} \sum_{k=1}^n (\lambda_k - \alpha \log \lambda_k) - \frac{1}{2} \sum_{i \neq j} |\lambda_i - \lambda_j| \quad (25.9)$$

*Remark.* For the (25.4), which can be written as (25.8) form, that,

$$p(\Lambda) \propto \exp[-\beta m_n E(\Lambda)]$$

where  $\beta = 1$  and

$$E(\Lambda) = \frac{m_n}{2} \sum_{k=1}^n \left[ \lambda_k - \left( \frac{m_n - n - 1}{m_n} \right) \log \lambda_k \right] - \frac{1}{2m_n} \sum_{i \neq j} |\lambda_i - \lambda_j|$$

## 25.3 Marčenko-Pastur Theorem

In this section, we will investigate the empirical spectral measure of  $\hat{\Sigma}$ , which converges to a nonrandom distribution — Marčenko-Pastur distribution. Before further proof, we will introduce some basic concepts and tools.

With the above tools, we can prove the Marčenko-Pastur Theorem. Here, we only suppose  $\mathbf{x}$  has some smooth tail condition.

### Theorem 25.3.1 (Marčenko-Pastur Theorem)

Consider the resolvent

$$\mathbf{Q}(z) = (\hat{\Sigma} - z\mathbf{I}_n)^{-1}$$

Then, if

$$\frac{n}{m_n} \rightarrow \rho \text{ with } \rho \in (0, \infty), \text{ as } n \rightarrow \infty$$

we have

$$\mathbf{Q}(z) \leftrightarrow \overline{\mathbf{Q}}(z), \quad \overline{\mathbf{Q}}(z) = m(z)\mathbf{I}_n$$

with  $(z, m(z))$  the unique solution in  $\mathcal{Z} \left( \mathbb{C} \setminus \left[ (1 - \sqrt{\rho})^2, (1 + \sqrt{\rho})^2 \right] \right)$  be

$$zcm^2(z) - (1 - c - z)m(z) + 1 = 0$$

where the function  $m(z)$  is the Stieltjes transform of the probability measure  $\mu$  given explicitly by

$$\mu(dx) = (1 - \rho^{-1})^+ \delta_0(x) + \frac{\sqrt{(x - a_-)^+ (a_+ - x)^+}}{2\pi\rho x} dx$$

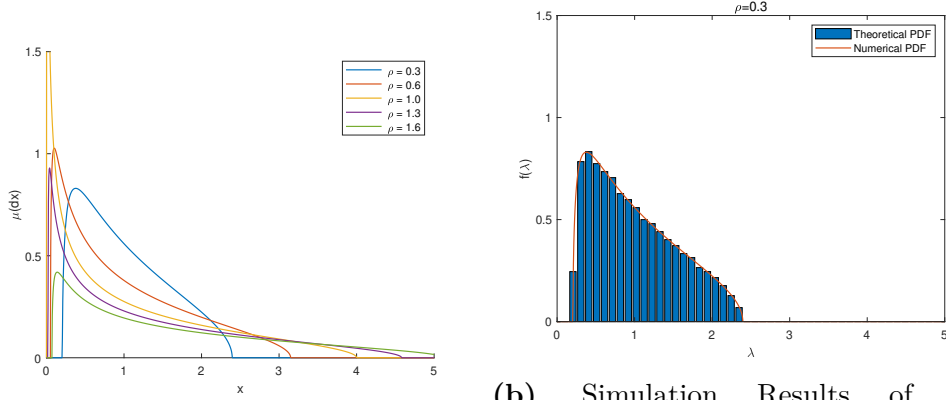
where  $a_{\pm} = (1 \pm \sqrt{\rho})^2$  and  $(x)^+ = \max(0, x)$ , and is known as the Marčenko-Pastur distribution. In particular, with probability one, the empirical spectral measure  $\mu_{\hat{\Sigma}}$  converges weakly to  $\mu$ .

*Proof.* (Intuitive Proof) Suppose  $\overline{\mathbf{Q}}(z) = \mathbf{F}(z)^{-1}$  for some matrix  $\mathbf{F}(z)$ . To prove  $\overline{\mathbf{Q}}(z)$  to be a deterministic equivalent for  $\mathbf{Q}(z)$ , particularly,

$$\frac{1}{n} \text{tr} \mathbf{A}(\mathbf{Q}(z) - \overline{\mathbf{Q}}(z)) \rightarrow 0 \quad \text{a.s.}$$

where  $\mathbf{A}$  is arbitrary, deterministic, and such that  $\|\mathbf{A}\| = 1$ . By Lemma 23.4.1, we have

$$\begin{aligned} \mathbf{Q}(z) - \overline{\mathbf{Q}}(z) &= \mathbf{Q}(z) \left( \mathbf{F}(z) + z\mathbf{I}_n - \hat{\Sigma} \right) \overline{\mathbf{Q}}(z) \\ &= \mathbf{Q}(z) \left( \mathbf{F}(z) + z\mathbf{I}_n - \frac{1}{m_n} \sum_{i=1}^{m_n} \mathbf{x}_i \mathbf{x}_i^\top \right) \overline{\mathbf{Q}}(z) \end{aligned}$$



(a) The Marčenko-Pastur Distribution for  $\rho = 0.3, 0.6, 1, 1.3, 1.6$   
 (b) Simulation Results of the Marčenko-Pastur Theorem When  $\rho = 0.3$

**Figure 25.1:** Illustrations of the Marčenko-Pastur Theorem

Thus, we turn to prove that,

$$\frac{1}{n} \operatorname{tr} \left[ (\mathbf{F}(z) + z\mathbf{I}_n) \overline{\mathbf{Q}}(z) \mathbf{A} \mathbf{Q}(z) \right] - \frac{1}{n} \cdot \frac{1}{m_n} \sum_{i=1}^{m_n} \mathbf{x}_i^\top \overline{\mathbf{Q}}(z) \mathbf{A} \mathbf{Q}(z) \mathbf{x}_i \rightarrow 0 \quad \text{a.s.}$$

By Lemma 23.4.2, we have

$$\mathbf{Q}(z) \mathbf{x}_i = \frac{\mathbf{Q}_{-i}(z) \mathbf{x}_i}{1 + \frac{1}{m_n} \mathbf{x}_i^\top \mathbf{Q}_{-i}(z) \mathbf{x}_i}$$

where

$$\mathbf{Q}_{-i}(z) = \left( \frac{1}{m_n} \sum_{j \neq i} \mathbf{x}_j \mathbf{x}_j^\top - z \mathbf{I}_n \right)^{-1}$$

is independent of  $\mathbf{x}_i$ . By Lemma 23.4.3, we have

$$\frac{1}{n} \mathbf{x}_i^\top \overline{\mathbf{Q}}(z) \mathbf{A} \mathbf{Q}(z) \mathbf{x}_i = \frac{\frac{1}{n} \mathbf{x}_i^\top \overline{\mathbf{Q}}(z) \mathbf{A} \mathbf{Q}_{-i}(z) \mathbf{x}_i}{1 + \frac{1}{m_n} \mathbf{x}_i^\top \mathbf{Q}_{-i}(z) \mathbf{x}_i} \simeq \frac{\frac{1}{n} \operatorname{tr} [\overline{\mathbf{Q}}(z) \mathbf{A} \mathbf{Q}_{-i}(z)]}{1 + \frac{1}{m_n} \operatorname{tr} [\mathbf{Q}_{-i}(z)]}$$

Hence, we need to prove the approximation that

$$\frac{1}{n} \operatorname{tr} \left[ (\mathbf{F}(z) + z\mathbf{I}_n) \overline{\mathbf{Q}}(z) \mathbf{A} \mathbf{Q}(z) \right] \simeq \frac{\frac{1}{n} \operatorname{tr} [\overline{\mathbf{Q}}(z) \mathbf{A} \mathbf{Q}(z)]}{1 + \frac{1}{m_n} \operatorname{tr} [\mathbf{Q}(z)]}$$

If  $\mathbf{F}(z)$  exist, for the approximation above to hold,  $\mathbf{F}(z)$  must be of the type

$$\mathbf{F}(z) \simeq \left( -z + \frac{1}{1 + \frac{1}{m_n} \operatorname{tr} \mathbf{Q}(z)} \right) \mathbf{I}_n$$

By Equation 23.8, we have,

$$m(z) \equiv \frac{1}{n} \operatorname{tr} [\bar{\mathbf{Q}}(z)] = \frac{1}{n} \operatorname{tr} [\mathbf{F}(z)^{-1}]$$

taking  $\mathbf{A} = \mathbf{I}_n$ , we have

$$\frac{1}{n} \operatorname{tr} [\mathbf{Q}(z)] \simeq \frac{1}{n} \operatorname{tr} [\bar{\mathbf{Q}}(z)] = m(z) = \frac{1}{-z + \frac{1}{1 + \frac{n}{m_n} \frac{1}{n} \operatorname{tr} [\mathbf{Q}(z)]}} \simeq \frac{1}{-z + \frac{1}{1 + \rho m(z)}}$$

As  $n, m_n \rightarrow \infty$ ,  $m(z)$  is solution to

$$m(z) = \frac{1}{-z + \frac{1}{1 + \rho m(z)}}$$

or equivalently

$$z \rho m^2(z) - (1 - \rho - z)m(z) + 1 = 0$$

This equation has two solutions defined via the two values of the complex square root function. Let

$$z = r e^{i\theta} \text{ where } r \geq 0, \theta \in [0, 2\pi) \Rightarrow \sqrt{z} \in \{\pm \sqrt{r} e^{i\theta/2}\}$$

and we can conclude that

$$m(z) = \frac{1 - \rho - z}{2\rho z} + \frac{\sqrt{((1 + \sqrt{\rho})^2 - z)((1 - \sqrt{\rho})^2 - z)}}{2\rho z}$$

only one of which is such that  $\Im[z]\Im[m(z)] > 0$  as imposed by the definition of Stieltjes transforms. By the inverse Stieltjes transform theorem, Theorem 23.2.1, we find that  $m(z)$  is the Stieltjes transform of the measure  $\mu$  with

$$\mu([a, b]) = \frac{1}{\pi} \lim_{\epsilon \downarrow 0} \int_a^b \Im[m(x + i\epsilon)] dx$$

for all continuity points  $a, b \in \mathbb{R}$  of  $\mu$ . This term under the square root in  $m(z)$  is negative only in the set

$$[(1 - \sqrt{\rho})^2, (1 + \sqrt{\rho})^2]$$

(and thus of non-real square root), the latter defines the support of the continuous part of the measure  $\mu$  with density

$$\frac{\sqrt{((1 + \sqrt{\rho})^2 - x)(x - (1 - \sqrt{\rho})^2)}}{2\rho\pi x}$$

at point  $x$  in the set. The case  $x = 0$  brings a discontinuity in  $\mu$  with weight equal to

$$\mu(\{0\}) = -\lim_{y \downarrow 0} \text{sym}(iy) = \frac{\rho - 1}{2\rho} \pm \frac{\rho - 1}{2\rho}$$

where the sign is established by a second order development of  $zm(z)$  in the neighborhood of zero: that is, "+" for  $c > 1$  inducing a mass  $1 - 1/\rho$  for  $p > n$ , or "-" for  $c < 1$  in which case  $\mu(\{0\}) = 0$  and  $\mu$  has no mass at zero.  $\square$

*Remark.* The asymptotic phenomenon holds not only in the Gaussian case, which also holds

1. if  $(x_i)_{1 \leq i \leq n}$  are iid with finite second moment.
2. if  $\mathbf{x}$  is isotropic and log-concave<sup>1</sup> random vector.

## 25.4 Limits of Extreme Eigenvalues

The weak convergence in Theorem 25.3.1 does not provide much information at the edge on the behavior of the extremal atoms, and what one can extract is that

$$\limsup_{n \rightarrow \infty} \lambda_{\min}(\hat{\Sigma}) \leq (1 - \sqrt{\rho})^2 \leq (1 + \sqrt{\rho})^2 \leq \liminf_{n \rightarrow \infty} \lambda_{\max}(\hat{\Sigma}) \quad \text{a.s.} \quad (25.10)$$

where the first inequality is considered only in the case where  $m_n \geq n$ .

The weak convergence above does not provide much information at the edge on the behavior of the extremal atoms. Now, we have more exact result, that if  $(X_{n,k})_{n \geq 1, 1 \leq k \leq n}$  are iid with finite fourth moment then,

$$(1 - \sqrt{\rho})^2 = \lim_{n \rightarrow \infty} \lambda_{\min}(\hat{\Sigma}) \leq \lim_{n \rightarrow \infty} \lambda_{\max}(\hat{\Sigma}) = (1 + \sqrt{\rho})^2 \quad \text{a.s.} \quad (25.11)$$

where the first inequality is considered only in the case where  $m_n \geq n$ .

*Remark.* The convergence of the smallest eigenvalue in the left-hand side of (25.11) holds if  $(x_i)_{1 \leq i \leq n}$  are iid with finite second moment.

---

<sup>1</sup>A probability measure  $\mu$  on  $\mathbb{R}^n$  with density  $\varphi$  is log-concave when  $\varphi = e^{-V}$  with  $V$  convex.

**Theorem 25.4.1**

If  $\bar{\rho} < 1$  (in particular  $m_n > n$  for  $n \gg 1$ ) and if the centered isotropic random vector  $\mathbf{x}$  is log-concave or if  $(x_i)_{1 \leq i \leq n}$  are iid then

$$\liminf_{n \rightarrow \infty} \frac{E(\lambda_{\min}(\mathbf{A}_n))}{(\sqrt{m_n} - \sqrt{n})^2} \geq 1 \quad (25.12)$$

If additionally  $\lim_{n \rightarrow \infty} \frac{n}{m_n} = \rho$  with  $\rho \in (0, 1)$ , in other words  $\underline{\rho} = \bar{\rho} \in (0, 1)$ , then

$$\lambda_{\min}(\hat{\Sigma}_n) \xrightarrow{p} (1 - \sqrt{\rho})^2 \text{ as } n \rightarrow \infty \quad (25.13)$$

*Proof.* □

**Theorem 25.4.2**

If the centered isotropic random vector  $\mathbf{x}$  is log-concave or if  $(x_i)_{1 \leq i \leq n}$  are iid with finite 4-th moment then

$$\limsup_{n \rightarrow \infty} \frac{E(\lambda_{\max}(\mathbf{A}_n))}{(\sqrt{m_n} + \sqrt{n})^2} \leq 1 \quad (25.14)$$

If additionally  $\lim_{n \rightarrow \infty} \frac{n}{m_n} = \rho$  with  $\rho \in (0, 1)$ , in other words  $\underline{\rho} = \bar{\rho} \in (0, 1)$ , then

$$\lambda_{\max}(\hat{\Sigma}_n) \xrightarrow{p} (1 + \sqrt{\rho})^2 \text{ as } n \rightarrow \infty \quad (25.15)$$

*Proof.* □

# **Part X**

## **Statistics Inference**



# Chapter 26

## Statistical Theory

### 26.1 Populations and Samples

### 26.2 Statistics

#### 26.2.1 Sufficient Statistics

##### Definition 26.2.1 (Sufficient Statistics)

A statistic  $T$  is said to be sufficient for  $X$ , or for the family  $\mathcal{P} = \{P_\theta, \theta \in \Omega\}$  of possible distributions of  $X$ , or for  $\theta$ , if the conditional distribution of  $X$  given  $T = t$  is independent of  $\theta$  for all  $t$ .

##### Theorem 26.2.1 (Fisher-Neyman Factorization Theorem)

If the probability density function is  $p_\theta(x)$ , then  $T$  is sufficient for  $\theta$  if and only if nonnegative functions  $g$  and  $h$  can be found such that

$$p_\theta(x) = h(x)g_\theta[T(x)].$$

*Proof.*

□

#### 26.2.2 Complete Statistics

##### Definition 26.2.2 (Complete Statistics)

A statistic  $T$  is said to be complete, if  $Eg(T) = 0$  for all  $\theta$  and some function  $g$  implies that  $P(g(T) = 0 \mid \theta) = 1$  for all  $\theta$ .

## 26.3 Estimators

### Definition 26.3.1 (Estimator)

An estimator is a real-valued function defined over the sample space, that is

$$\delta : \mathbf{X} \rightarrow \mathbb{R}. \quad (26.1)$$

It is used to estimate an estimand,  $\theta$ , a real-valued function of the parameter.

### Unbiasedness

#### Definition 26.3.2 (Unbiasedness)

An estimator  $\hat{\theta}$  of  $\theta$  is unbiased if

$$E\hat{\theta} = \theta, \quad \forall \theta \in \Theta. \quad (26.2)$$

*Remark.* • Unbiased estimators of  $\theta$  may not exist.

**Example** (Nonexistence of Unbiased Estimator).

### Consistency

#### Definition 26.3.3 (Consistency)

An estimator  $\hat{\theta}_n$  of  $\theta$  is consistent if

$$\lim_{n \rightarrow \infty} P(|\hat{\theta}_n - \theta| > \varepsilon) = 0, \quad \forall \varepsilon > 0, \quad (26.3)$$

that is,

$$\hat{\theta}_n \xrightarrow{p} \theta. \quad (26.4)$$

**Example** (Consistency of Sample Moments).

*Remark.* 1. Unbiased But Consistent  
2. Biased But Not Consistent

## Asymptotic Normality

### Definition 26.3.4 (Asymptotic Normality)

An estimator  $\hat{\theta}_n$  of  $\theta$  is asymptotic normality if

$$\sqrt{n} (\hat{\theta} - \theta) \xrightarrow{d} N(0, \sigma_\theta^2). \quad (26.5)$$

## Efficiency

### Definition 26.3.5 (Efficiency)

## Robustness

### Definition 26.3.6 (Robustness)

# Chapter 27

## Point Estimation

### 27.1 Maximum Likelihood Estimator

Suppose that  $\mathbf{x}_n = (x_1, \dots, x_n)$ , within a parametric family

$$p(X; \theta_0) \in \mathcal{P} = \{p(X; \theta) : \theta \in \Theta\}$$

The maximum likelihood estimate for observed  $\mathbf{x}_n$  is the value  $\theta \in \Theta$  which maximizes  $L_n(\theta) := p(\mathbf{x}_n; \theta)$ , i.e.,

$$\hat{\theta} = \max_{\theta \in \Theta} L_n(\theta). \quad (27.1)$$

In practice, it is often convenient to work with the natural logarithm of the likelihood function, called the log-likelihood:

$$\ell_n(\theta) := \log L_n(\theta)$$

Since the logarithm is a monotonic function, the maximum of  $\ell_n(\theta)$  occurs at the same value of  $\theta$  as does the maximum of  $L_n(\theta)$

#### 27.1.1 Consistency

To establish consistency, the following conditions are sufficient:

- (C1) Identification:  $\theta_0$  is identified in the sense that if  $\theta \neq \theta_0$  and  $\theta \in \Theta$ , then  $p(X; \theta) \neq p(X; \theta_0)$  with respect to the dominating measure  $\mu$ .
- (C2) The parameter space  $\Theta$  of the model is compact.
- (C3) The function  $\log p(X; \theta)$  is continuous in  $\theta$  for almost all values of  $x$ , i.e.,

$$P[\log p(X; \theta) \in C^0(\Theta)] = 1 \quad (27.2)$$

(C4) Dominance: there exists  $D(x)$  integrable with respect to the distribution  $p(X; \theta_0)$  such that  $|\log p(X; \theta)| < D(x)$  for all  $\theta \in \Theta$ .

**Lemma 27.1.1**

If  $\theta_0$  is identified and  $E_{\theta_0} [|\ln p(X; \theta)|] < \infty, \forall \theta \in \Theta$ , then  $\ell(\theta)$  is uniquely maximized at  $\theta = \theta_0$ .

*Proof.* By the strict version of Jensen's inequality, with  $\theta \neq \theta_0$ ,

$$\begin{aligned} \ell(\theta_0) - \ell(\theta) &= \mathbb{E}_{\theta_0} \left\{ -\ln \left[ \frac{p(z | \theta)}{p(z | \theta_0)} \right] \right\} > -\ln \mathbb{E}_{\theta_0} \left[ \frac{p(z | \theta)}{p(z | \theta_0)} \right] \\ &= -\ln \left[ \int f(z | \theta) dz \right] = 0 \end{aligned}$$

□

**Theorem 27.1.1 (Consistency of MLE)**

Under the Assumption (1)-(4), we have

$$\hat{\theta} \xrightarrow{p} \theta_0 \quad (27.3)$$

*Proof.* Suppose

$$\Theta(\epsilon) = \{\theta : \|\theta - \theta_0\| < \epsilon\}, \quad \forall \epsilon > 0$$

Since  $Q_0(\theta)$  is a continuous function, thus

$$\theta^* := \sup_{\theta \in \Theta \cap \Theta(\epsilon)^C} \{\ell(\theta)\}$$

is achieved for a  $\theta$  in the compact set  $\theta \in \Theta \cap \Theta(\epsilon)^C$  (For open set  $\Theta(\epsilon)$ ,  $\Theta \cap \Theta(\epsilon)^C$  is a compact set). And  $\theta_0$  is the unique maximized,

$$\exists \delta > 0, \quad \ell(\theta_0) - \ell(\theta^*) = \delta$$

1. For  $\theta \in \Theta \cap \Theta(\epsilon)^C$ . suppose

$$A_n = \left\{ \sup_{\theta \in \Theta \cap \Theta(\epsilon)^C} |\hat{\ell}(\theta; \mathbf{X}_n) - \ell(\theta)| < \frac{\delta}{2} \right\}$$

then,

$$A_n \implies \hat{\ell}(\theta; \mathbf{X}_n) < \ell(\theta) + \frac{\delta}{2} \leq \ell(\theta^*) + \frac{\delta}{2} = \ell(\theta_0) - \frac{\delta}{2}$$

2. For  $\theta \in \Theta(\epsilon)$ , suppose

$$B_n = \left\{ \sup_{\theta \in \Theta(\epsilon)} |\hat{\ell}(\theta) - \ell(\theta)| < \frac{\delta}{2} \right\}$$

then

$$B_n \implies \forall \theta \in \Theta(\epsilon), \hat{\ell}(\theta) > \ell(\theta) - \frac{\delta}{2}$$

By the uniform law of large numbers, the dominance condition together with continuity establishes the uniform convergence in the probability of the log-likelihood:

$$\sup_{\theta \in \Theta} |\hat{\ell}(\theta) - \ell(\theta)| \xrightarrow{P} 0$$

Thus, we can conclude that

$$P(A_n \cap B_n) \rightarrow 1$$

Within the definition

$$\hat{\theta} = \max_{\theta \in \Theta} \hat{\ell}(\theta)$$

we have,

$$A_n \cap B_n \implies \hat{\theta} \in \Theta(\epsilon)$$

Hence,

$$\forall \epsilon > 0, P[\hat{\theta} \in \Theta(\epsilon)] \rightarrow 1 \implies \hat{\theta} \xrightarrow{P} \theta_0$$

□

## 27.1.2 Fisher Information

### Definition 27.1.1 (Fisher Information)

The Fisher information of a random variable  $X$  with probability density function  $p(X; \theta)$  is defined as

$$I(\theta) = \mathbb{E} \left[ \left( \frac{\partial}{\partial \theta} \ln p(X; \theta) \right)^2 \right]. \quad (27.4)$$

Alternatively, the Fisher information can be expressed as

$$I(\theta) = -\mathbb{E} \left[ \frac{\partial^2}{\partial \theta^2} \ln p(X; \theta) \right] \quad (27.5)$$

### 27.1.3 Asymptotic Normality

(C5) The information matrix  $I(\boldsymbol{\theta})$  is positive definite.

(C6)  $\left\| \frac{\partial^2 \log p(X; \boldsymbol{\theta})}{\partial \boldsymbol{\theta} \partial \boldsymbol{\theta}^\top} \right\| \leq M(x)$  for all  $\boldsymbol{\theta} \in \Theta$  and  $\mathbb{E}_{\boldsymbol{\theta}_0} M(x) < \infty$ .

*Proof.* Since the MLE is the maximizer of the log-likelihood function, the score function evaluated at the MLE is zero, i.e.,  $\ell'_n(\hat{\boldsymbol{\theta}}) = \mathbf{0}$ . By the Taylor expansion of the score function around  $\boldsymbol{\theta}_0$ , we have

$$\mathbf{0} = \ell'_n(\hat{\boldsymbol{\theta}}) = \ell'_n(\boldsymbol{\theta}_0) + (\hat{\boldsymbol{\theta}} - \boldsymbol{\theta}_0) \ell''_n(\tilde{\boldsymbol{\theta}})$$

where  $\tilde{\boldsymbol{\theta}}$  lies between  $\hat{\boldsymbol{\theta}}$  and  $\boldsymbol{\theta}_0$ . Define  $J_n(\boldsymbol{\theta}) = -\frac{1}{n} \ell''_n(\boldsymbol{\theta})$ . Then we have

$$\sqrt{n}(\hat{\boldsymbol{\theta}} - \boldsymbol{\theta}_0) = J_n^{-1}(\tilde{\boldsymbol{\theta}}) n^{-1/2} \ell'_n(\boldsymbol{\theta}_0).$$

Then we need to show that: 1)  $n^{-1/2} \ell'_n(\boldsymbol{\theta}_0) \xrightarrow{d} \mathcal{N}(\mathbf{0}, I(\boldsymbol{\theta}_0))$ ; 2)  $J_n(\tilde{\boldsymbol{\theta}}) \xrightarrow{p} I(\boldsymbol{\theta}_0)$ .

For the first term, we have

$$n^{-1/2} \ell'_n(\boldsymbol{\theta}_0) = n^{-1/2} \sum_{i=1}^n \ell'(X_i; \boldsymbol{\theta}_0) = \sqrt{n} \frac{1}{n} \sum_{i=1}^n \frac{\partial \log p(X_i; \boldsymbol{\theta}_0)}{\partial \boldsymbol{\theta}},$$

thus, by the central limit theorem, we have

$$\sqrt{n} \frac{1}{n} \sum_{i=1}^n \frac{\partial \log p(X_i; \boldsymbol{\theta}_0)}{\partial \boldsymbol{\theta}} \xrightarrow{d} \mathcal{N}(\mathbf{0}, I(\boldsymbol{\theta}_0)).$$

For the second term, denote  $I_0^*(\boldsymbol{\theta}) = \mathbb{E}_{\boldsymbol{\theta}_0} \left[ -\frac{\partial^2 \log p(X; \boldsymbol{\theta})}{\partial \boldsymbol{\theta} \partial \boldsymbol{\theta}^\top} \right]$ , by the triangle inequality, we have

$$\|J_n(\tilde{\boldsymbol{\theta}}) - I(\boldsymbol{\theta}_0)\| \leq \|J_n(\tilde{\boldsymbol{\theta}}) - I_0^*(\tilde{\boldsymbol{\theta}})\| + \|I_0^*(\tilde{\boldsymbol{\theta}}) - I(\boldsymbol{\theta}_0)\|.$$

According to Lemma, we have

$$\|J_n(\tilde{\boldsymbol{\theta}}) - I_0^*(\tilde{\boldsymbol{\theta}})\| \leq \sup_{\boldsymbol{\theta} \in \Theta} \|J_n(\boldsymbol{\theta}) - I_0^*(\boldsymbol{\theta})\| \xrightarrow{p} 0.$$

Since  $I_0^*(\boldsymbol{\theta})$  is continuous in  $\boldsymbol{\theta}$ , and  $\tilde{\boldsymbol{\theta}} \xrightarrow{p} \boldsymbol{\theta}_0$ , by the continuous mapping theorem, we have

$$I_0^*(\tilde{\boldsymbol{\theta}}) \xrightarrow{p} I(\boldsymbol{\theta}_0).$$

Combining the above results, we have

$$J_n(\tilde{\boldsymbol{\theta}}) \xrightarrow{p} I(\boldsymbol{\theta}_0).$$

which completes the proof. □

### 27.1.4 Efficiency

## 27.2 Modified Likelihood Estimator

Seek a modified likelihood function that depends on as few of the nuisance parameters as possible while sacrificing as little information as possible.

### 27.2.1 Marginal Likelihood

### 27.2.2 Conditional Likelihood

Let  $\boldsymbol{\theta} = (\boldsymbol{\varphi}, \boldsymbol{\Lambda})$ , where  $\boldsymbol{\varphi}$  is the parameter vector of interest and  $\boldsymbol{\Lambda}$  is a vector of nuisance parameters. The conditional likelihood can be obtained as follows:

1. Find the complete sufficient statistic  $S_{\boldsymbol{\Lambda}}$ , respectively for  $\boldsymbol{\Lambda}$ .
2. Construct the conditional log-likelihood

$$\ell_c = \ln(f_{Y|S_{\boldsymbol{\Lambda}}}) \quad (27.6)$$

where  $f_{Y|S_{\boldsymbol{\Lambda}}}$  is the conditional distribution of the response  $Y$  given  $S_{\boldsymbol{\Lambda}}$ .

*Remark.* Two cases might occur, that, for fixed  $\boldsymbol{\varphi}_0$ ,  $S_{\boldsymbol{\Lambda}}(\boldsymbol{\varphi}_0)$  depends on  $\boldsymbol{\varphi}_0$ ; or  $S_{\boldsymbol{\Lambda}}(\boldsymbol{\varphi}_0) = S_{\boldsymbol{\Lambda}}$  is independent of  $\boldsymbol{\varphi}_0$ .

1. Independent:
2. Dependent:

Suppose that the log-likelihood for  $\boldsymbol{\theta} = (\boldsymbol{\varphi}, \boldsymbol{\Lambda})$  can be written in the exponential family form

$$\ell(\boldsymbol{\theta}, \mathbf{y}) = \boldsymbol{\theta}'\mathbf{s} - b(\boldsymbol{\theta}) \quad (27.7)$$

Also, suppose  $\ell(\boldsymbol{\theta}, \mathbf{y})$  has a decomposition of the form

$$\ell(\boldsymbol{\theta}, \mathbf{y}) = \boldsymbol{\varphi}'\mathbf{s}_1 + \boldsymbol{\Lambda}'\mathbf{s}_2 - b(\boldsymbol{\varphi}, \boldsymbol{\Lambda}) \quad (27.8)$$

*Remark.* The above decomposition can be achieved only if  $\boldsymbol{\varphi}$  is a linear function of  $\boldsymbol{\theta}$ . The choice of nuisance parameter  $\lambda$  is arbitrary and the inferences regarding  $\boldsymbol{\varphi}$  should be unaffected by the parameterization chosen for  $\lambda$ .

The conditional likelihood of the data  $\mathbf{Y}$  given  $\mathbf{s}_2$  is

$$\ell(\boldsymbol{\varphi} | \mathbf{s}_2) = \boldsymbol{\varphi}'\mathbf{s}_1 - b^*(\boldsymbol{\varphi}, \boldsymbol{\Lambda}) \quad (27.9)$$

which is independent of the nuisance parameter and may be used for inferences regarding  $\boldsymbol{\varphi}$ .

**Example.**  $Y_1 \sim P(\mu_1), Y_2 \sim P(\mu_2)$  are independent. Suppose  $\varphi = \ln\left(\frac{\mu_2}{\mu_1}\right) = \ln(\mu_2) - \ln(\mu_1)$  is the parameter of interest and the nuisance parameter is



1.  $\lambda_1 = \ln(\mu_1)$ .
- 2.

Then, give the conditional log-likelihood for different nuisance parameters.

*Proof.* 1. The log-likelihood function in the form of  $(\varphi, \lambda)$  is

$$\begin{aligned}
 \ell(\phi, \lambda_1) &\propto \ln \left[ e^{-(\mu_1 + \mu_2)} \mu_1^{y_1} \mu_2^{y_2} \right] \\
 &= -(\mu_1 + \mu_2) + y_1 \ln(\mu_1) + y_2 \ln(\mu_2) \\
 &= -\mu_1 \left( 1 + \frac{\mu_2}{\mu_1} \right) + y_1 \ln(\mu_1) + y_2 \ln(\mu_1) \\
 &\quad - y_2 [\ln(\mu_1) - \ln(\mu_2)] \\
 &= -e^{\lambda_1} (1 + e^\varphi) + (y_1 + y_2) \lambda_1 - y_2 \varphi \\
 &= s_1 \varphi + s_2 \lambda_1 - b(\varphi, \lambda_1)
 \end{aligned}$$

where  $s_1 = -y_2$ ,  $s_2 = y_1 + y_2$ ,  $b(\varphi, \lambda_1) = e^{\lambda_1} (1 + e^\varphi)$ .

Then, the conditional distribution of  $Y_1, Y_2$  given  $S_2 = Y_1 + Y_2$  is  $b\left(S_2, \frac{\mu_1}{\mu_1 + \mu_2}\right)$ , thus,

$$\begin{aligned}
 \ell(\varphi \mid S_2 = s_2) &\propto y_1 \ln \left( \frac{\mu_1}{\mu_1 + \mu_2} \right) + y_2 \ln \left( \frac{\mu_2}{\mu_1 + \mu_2} \right) \\
 &= y_1 \ln \left( \frac{\mu_1}{\mu_1 + \mu_2} \right) + y_2 \ln \left( \frac{\mu_1}{\mu_1 + \mu_2} \right) \\
 &\quad - y_2 \left[ \ln \left( \frac{\mu_1}{\mu_1 + \mu_2} \right) - \ln \left( \frac{\mu_2}{\mu_1 + \mu_2} \right) \right] \\
 &= (y_1 + y_2) \ln \left( \frac{1}{1 + e^\varphi} \right) - y_2 \varphi \\
 &= s_1 \varphi - b^*(\varphi, s_2)
 \end{aligned}$$

where  $b^*(\varphi, s_2) = -s_2 \ln \left( \frac{1}{1 + e^\varphi} \right)$ .

□

### 27.2.3 Profile Likelihood

### 27.2.4 Quasi Likelihood

## 27.3 Minimum-Variance Unbiased Estimator

### Definition 27.3.1 (UMVU Estimators)

An unbiased estimator  $\delta(\mathbf{X})$  of  $g(\theta)$  is the uniform minimum variance unbiased (UMVU) estimator of  $g(\theta)$  if

$$\text{Var}_\theta \delta(\mathbf{X}) \leq \text{Var}_\theta \delta'(\mathbf{X}), \quad \forall \theta \in \Theta, \quad (27.10)$$

where  $\delta'(\mathbf{X})$  is any other unbiased estimator of  $g(\theta)$ .

*Remark.* If there exists an unbiased estimator of  $g$ , the estimand  $g$  will be called  $U$ -estimable.

1. If  $T(\mathbf{X})$  is a complete sufficient statistic, estimator  $\delta(\mathbf{X})$  that only depends on  $T(\mathbf{X})$ , then for any  $U$ -estimable function  $g(\theta)$  with

$$E_\theta \delta(T(\mathbf{X})) = g(\theta), \quad \forall \theta \in \Theta, \quad (27.11)$$

hence,  $\delta(T(\mathbf{X}))$  is the unique UMVU estimator of  $g(\theta)$ .

2. If  $T(\mathbf{X})$  is a complete sufficient statistic and  $\delta(\mathbf{X})$  is any unbiased estimator of  $g(\theta)$ , then the UMVU estimator of  $g(\theta)$  can be obtained by

$$E[\delta(\mathbf{X}) \mid T(\mathbf{X})]. \quad (27.12)$$

**Example** (Estimating Polynomials of a Normal Variance). Let  $X_1, \dots, X_n$  be distributed with joint density

$$\frac{1}{(\sqrt{2\pi}\sigma)^n} \exp \left[ -\frac{1}{2\sigma^2} \sum (x_i - \xi)^2 \right]. \quad (27.13)$$

Discussing the UMVU estimators of  $\xi^r$ ,  $\sigma^r$ ,  $\xi/\sigma$ .

*Proof.* 1.  **$\sigma$  is known:**

Since  $\bar{X} = \frac{1}{n} \sum_{i=1}^n X_i$  is the complete sufficient statistic of  $X_i$ , and

$$E(\bar{X}) = \xi,$$

then the UMVU estimator of  $\xi$  is  $\bar{X}$ .

Therefore, the UMVU estimator of  $\xi^r$  is  $\bar{X}^r$  and the UMVU estimator of  $\xi/\sigma$  is  $\bar{X}/\sigma$ .

**2.  $\xi$  is known:**

Since  $s^r = \sum (x_i - \xi)^r$  is the complete sufficient statistic of  $X_i$ .

Assume

$$E \left[ \frac{s^r}{\sigma^r} \right] = \frac{1}{K_{n,r}},$$

where  $K_{n,r}$  is a constant depends on  $n, r$ .

Since  $s^2/\sigma^2 \sim \text{Ga}(n/2, 1/2) = \chi^2(n)$ , then

$$E \left[ \frac{s^r}{\sigma^r} \right] = E \left[ \left( \frac{s^2}{\sigma^2} \right)^{\frac{r}{2}} \right] = \int_0^\infty x^{\frac{r}{2}} \frac{1}{2^{\frac{n}{2}} \Gamma(\frac{n}{2})} x^{\frac{n}{2}-1} e^{-\frac{x}{2}} dx = \frac{\Gamma(\frac{n+r}{2})}{\Gamma(\frac{n}{2})} \cdot 2^{\frac{r}{2}}.$$

therefore,

$$K_{n,r} = \frac{\Gamma(\frac{n}{2})}{2^{\frac{r}{2}} \cdot \Gamma(\frac{n+r}{2})}.$$

Hence,

$$E[s^r K_{n,r}] = \sigma^r \text{ and } E[\xi s^{-1} K_{n,-1}] = \xi/\sigma,$$

which means the UMVU estimator of  $\sigma^r$  is  $s^r K_{n,r}$  and the UMVU estimator of  $\xi/\sigma$  is  $\xi s^{-1} K_{n,-1}$ .

**3. Both  $\xi$  and  $\sigma$  is unknown:**

Since  $(\bar{X}, s_x^r)$  are the complete sufficient statistic of  $X_i$ , where  $s_x^2 = \sum (x_i - \bar{X})^2$ .

Since  $s_x^2/\sigma^2 \sim \chi^2(n-1)$ , then

$$E \left[ \frac{s_x^r}{\sigma^r} \right] = \frac{1}{K_{n-1,r}}.$$

Hence,

$$E[s_x^r K_{n-1,r}] = \sigma^r,$$

which means the UMVU estimator of  $\sigma^r$  is  $s_x^r K_{n-1,r}$ , and

$$E(\bar{X}^r) = \xi^r,$$

which means the UMVU estimator of  $\xi^r$  is  $\bar{X}^r$ .

Since  $\bar{X}$  and  $s_x^r$  are independent, then

$$E[\bar{X} s_x^{-1} K_{n-1,-1}] = \xi/\sigma$$

which means the UMVU estimator of  $\xi/\sigma$  is  $\bar{X} s_x^{-1} K_{n-1,-1}$ .

□

**Example.** Let  $X_1, \dots, X_n$  be i.i.d sample from  $U(\theta_1 - \theta_2, \theta_1 + \theta_2)$ , where  $\theta_1 \in \mathbb{R}, \theta_2 \in \mathbb{R}^+$ . Discussing the UMVU estimators of  $\theta_1, \theta_2$ .

*Proof.* Let  $X_{(i)}$  be the  $i$ -th order statistic of  $X_i$ , then  $(X_{(1)}, X_{(n)})$  is the complete and sufficient statistic for  $(\theta_1, \theta_2)$ . Thus it suffices to find a function  $(X_{(1)}, X_{(n)})$ , which is unbiased of  $(\theta_1, \theta_2)$ .

Let

$$Y_i = \frac{X_i - (\theta_1 - \theta_2)}{2\theta_2} \sim U(0, 1),$$

and

$$Y_{(i)} = \frac{X_{(i)} - (\theta_1 - \theta_2)}{2\theta_2},$$

be the  $i$ -th order statistic of  $Y_i$ , then we got

$$\begin{aligned} E[X_{(1)}] &= 2\theta_2 E[Y_{(1)}] + (\theta_1 - \theta_2) \\ &= 2\theta_2 \int_0^1 ny(1-y)^{n-1} dy + (\theta_1 - \theta_2) \\ &= \theta_1 - \frac{3n+1}{n+1}\theta_2 \\ E[X_{(n)}] &= 2\theta_2 E[Y_{(n)}] + (\theta_1 - \theta_2) \\ &= 2\theta_2 \int_0^1 ny^n dy + (\theta_1 - \theta_2) \\ &= \theta_1 + \frac{n-1}{n+1}\theta_2 \end{aligned}$$

Thus,

$$\begin{aligned} \theta_1 &= E \left[ \frac{n-1}{4n} X_{(1)} + \frac{3n+1}{4n} X_{(n)} \right], \\ \theta_2 &= E \left[ -\frac{n+1}{4n} X_{(1)} + \frac{n+1}{4n} X_{(n)} \right], \end{aligned}$$

which means the UMVU estimator is

$$\hat{\theta}_1 = \frac{n-1}{4n} X_{(1)} + \frac{3n+1}{4n} X_{(n)}, \quad \hat{\theta}_2 = -\frac{n+1}{4n} X_{(1)} + \frac{n+1}{4n} X_{(n)}.$$

□

## 27.4 Accuracy of Estimators

**Example** (Normal Probability). Let  $X_1, \dots, X_n$  be iid as  $\mathcal{N}(\theta, 1)$  and consider the estimation of  $p = P(X_i \leq u) = \Phi(u - \theta)$ . The maximum likelihood estimator of  $p$  is  $\hat{p} = \Phi(u - \bar{X})$ , and we shall attempt to obtain large-sample approximations for the bias and variance of this estimator.

*Proof.* Since  $\bar{X} - \theta$  is likely to be small, it is natural to write

$$\Phi(u - \bar{X}) = \Phi[(u - \theta) - (\bar{X} - \theta)]$$

and to expand the right side about  $u - \theta$  by Taylor's theorem as

$$\begin{aligned} \Phi(u - \bar{X}) = & \Phi(u - \theta) - (\bar{X} - \theta)\phi(u - \theta) + \frac{1}{2}(\bar{X} - \theta)^2\phi'(u - \theta) \\ & - \frac{1}{6}(\bar{X} - \theta)^3\phi''(u - \theta) + \frac{1}{24}(\bar{X} - \theta)^4\phi'''(\xi), \end{aligned} \quad (27.14)$$

where  $\xi$  is a random quantity that lies between  $u - \theta$  and  $u - \bar{X}$ .

To calculate the bias, we take the expectation of (27.14), which yields

$$\mathbb{E}[\hat{p}] = p + \frac{1}{2n}\phi'(u - \theta) + \frac{1}{24}\mathbb{E}[(\bar{X} - \theta)^4\phi'''(\xi)].$$

Since the derivatives of  $\phi(x)$  all are of the form  $P(x)\phi(x)$ , where  $P(x)$  is a polynomial in  $x$  and are therefore all bounded. It follows that

$$|\mathbb{E}[(\bar{X} - \theta)^4\phi'''(\xi)]| < M\mathbb{E}(\bar{X} - \theta)^4 = 3M/n^2,$$

for some finite  $M$ . Using the fact that  $\phi'(x) = -x\phi(x)$ , we therefore find that

$$\mathbb{E}(\hat{p}) = p - \frac{1}{2n}(u - \theta)\phi(u - \theta) + O(1/n^2),$$

where the error term is uniformly  $O(1/n^2)$ . The estimator  $\delta$  therefore has a bias of order  $1/n$  which tends to zero as  $\theta \rightarrow \pm\infty$ .

In the same way, by Taylor's theorem, we can expand the square of the estimator as

$$\begin{aligned} \Phi^2(u - \bar{X}) = & \Phi^2(u - \theta) - 2(\bar{X} - \theta)\Phi(u - \theta)\phi(u - \theta) \\ & + (\bar{X} - \theta)^2[\phi^2(u - \theta) + \Phi(u - \theta)\phi'(u - \theta)] \\ & - \frac{1}{3}(\bar{X} - \theta)^3[2\phi(u - \theta)\phi'(u - \theta) + \Phi(u - \theta)\phi''(u - \theta) + \phi'^2(u - \theta)] \end{aligned}$$

where  $\xi$  lies between  $u - \theta$  and  $u - \bar{X}$ . Taking the expectation of this expansion, we find that

$$\mathbb{E}[\hat{p}^2] = p^2 + \frac{1}{n}\phi^2(u - \theta) - \frac{p}{n}(\mu - \theta)\phi(u - \theta),$$

and

$$[\mathbb{E}(\hat{p})]^2 = p^2 - \frac{p}{n}(\mu - \theta)\phi(u - \theta) + O(1/n^2).$$

Thus

$$\text{Var}(\hat{p}) = \mathbb{E}[\hat{p}^2] - [\mathbb{E}(\hat{p})]^2 = \frac{1}{n}\phi^2(u - \theta) + O\left(1/n^2\right).$$

and hence that

$$\text{Var}(\sqrt{n}\delta) \rightarrow \phi^2(u - \theta)$$

Since  $\sqrt{n}(\bar{X} - \theta) \xrightarrow{d} N(0, 1)$ , and  $f(\theta) = \Phi(\mu - \theta)$  is differentiable with  $f'(\theta) = -\phi(\mu - \theta)$  not equal to zero when  $\mu \neq \theta$ , then by the delta method (13.1.4), we have

$$\sqrt{n}(\hat{p} - p) \xrightarrow{d} N\left(0, \phi^2(u - \theta)\right),$$

the limit of the variance in this case is equal to the asymptotic variance.

It is interesting to see what happens if the expansion (27.14) is carried one step less far. Then

$$E[\Phi(u - \bar{X})] = p + \frac{1}{2n}\phi'(u - \theta) + \frac{1}{6}E[(\bar{X} - \theta)^3\phi''(\xi)].$$

Since the third derivative of  $\phi$  is bounded, the remainder now satisfies

$$\frac{1}{6}E\left|(\bar{X} - \theta)^3\phi'''(\xi)\right| < M'E|\bar{X} - \theta|^3 = O\left(\frac{1}{n^{3/2}}\right).$$

The conclusion is therefore weaker than before. □

# Chapter 28

## Interval Estimation

28.1 Confidence Interval

28.2 Pivot

28.3 Likelihood Interval

28.4 Prediction Interval

28.5 Tolerance Interval

# Chapter 29

## Testing Hypotheses

### 29.1 Testing Hypotheses

### 29.2 Parametric Tests

### 29.3 Specific Tests

#### 29.3.1 Goodness of Fit

Likelihood-Ratio Test

#### 29.3.2 Rank statistics



# Chapter 30

## Bayesian Inference

### 30.1 Bayes Estimator

We shall look for some estimators that make the risk function  $R(\theta, \delta)$  small in some overall sense. There are two ways to solve it: minimize the average risk, and minimize the maximum risk.

This chapter will discuss the first method, also known as, Bayes Estimator.

#### Definition 30.1.1 (Bayes Estimator)

The Bayes Estimator  $\delta$  with respect to  $\Lambda$  is minimizing the Bayes Risk of  $\delta$

$$r(\Lambda, \delta) = \int R(\theta, \delta) d\Lambda(\theta) \quad (30.1)$$

where  $\Lambda$  is the probability distribution.

In Bayesian arguments, it is important to keep track of which variables are being conditioned. Hence, the notations are as follows:

- The density of  $X$  will be denoted by  $X \sim f(x | \theta)$ .
- The prior distribution will be denoted by  $\Pi \sim \pi(\theta | \lambda)$  or  $\Lambda \sim \gamma(\lambda)$ , where  $\lambda$  is another parameter (sometimes called a hyperparameter).
- The posterior distribution, which calculates the conditional distributions as that of  $\theta$  given  $x$  and  $\lambda$ , or  $\lambda$  given  $x$ , which is denoted by  $\Pi \sim \pi(\theta | x, \lambda)$  or  $\Lambda \sim \gamma(\lambda | x)$ , that is

$$\pi(\theta | x, \lambda) = \frac{f(x | \theta) \pi(\theta | \lambda)}{m(x | \lambda)}, \quad (30.2)$$

where marginal distributions  $m(x | \lambda) = \int f(x | \theta) \pi(\theta | \lambda) d\theta$ .

**Theorem 30.1.1**

Let  $\Theta$  have distribution  $\Lambda$ , and given  $\Theta = \theta$ , let  $X$  have distribution  $P_\theta$ . Suppose, the following assumptions hold for the problem of estimating  $g(\Theta)$  with non-negative loss function  $L(\theta, d)$ ,

- There exists an estimator  $\delta_0$  with finite risk.
- For almost all  $x$ , there exists a value  $\delta_\Lambda(x)$  minimizing

$$E\{L[\Theta, \delta(x)] \mid X = x\}. \quad (30.3)$$

Then,  $\delta_\Lambda(x)$  is a Bayes Estimator.

*Remark.* Improper prior

**Corollary 30.1.1**

Suppose the assumptions of Theorem 30.1.1 hold.

1. If  $L(\theta, d) = [d - g(\theta)]^2$ , then

$$\delta_\Lambda(x) = E[g(\Theta) \mid x]. \quad (30.4)$$

2. If  $L(\theta, d) = w(\theta) [d - g(\theta)]^2$ , then

$$\delta_\Lambda(x) = \frac{E[w(\theta) g(\Theta) \mid x]}{E[w(\theta) \mid x]}. \quad (30.5)$$

3. If  $L(\theta, d) = |d - g(\theta)|$ , then  $\delta_\Lambda(x)$  is any median of the conditional distribution of  $\Theta$  given  $x$ .
4. If

$$L(\theta, d) = \begin{cases} 0 & \text{when } |d - \theta| \leq c \\ 1 & \text{when } |d - \theta| > c \end{cases},$$

then  $\delta_\Lambda(x)$  is the midpoint of the interval  $I$  of length  $2c$  which maximizes  $P(\Theta \in I \mid x)$ .

*Proof.*

□

**Theorem 30.1.2**

Necessary condition for Bayes Estimator

Methodologies have been developed to deal with the difficulty which sometimes incorporates frequentist measures to assess the choice of  $\Lambda$ .

- Empirical Bayes.
- Hierarchical Bayes.
- Robust Bayes.

- Objective Bayes.

### 30.1.1 Single-Prior Bayes

The Single-Prior Bayes model in a general form as

$$\begin{aligned} X | \theta &\sim f(x | \theta), \\ \Theta | \gamma &\sim \pi(\theta | \lambda), \end{aligned} \quad (30.6)$$

where we assume that the functional form of the prior and the value of  $\lambda$  is known (we will write it as  $\gamma = \gamma_0$ ).

Given a loss function  $L(\theta, d)$ , we would then determine the estimator that minimizes

$$\int L(\theta, d(x)) \pi(\theta | x) d\theta, \quad (30.7)$$

where  $\pi(\theta | x)$  is posterior distribution given by

$$\pi(\theta | x) = \frac{f(x | \theta) \pi(\theta | \gamma_0)}{\int f(x | \theta) \pi(\theta | \gamma_0) d\theta}.$$

In general, this Bayes estimator under squared error loss is given by

$$E(\Theta | x) = \frac{\int \theta f(x | \theta) \pi(\theta | \gamma_0) d\theta}{\int f(x | \theta) \pi(\theta | \gamma_0) d\theta}. \quad (30.8)$$

**Example.** Consider

$$\begin{aligned} X_i &\stackrel{\text{i.i.d}}{\sim} N(\mu, \Gamma^{-1}), \quad i = 1, 2, \dots, n \\ \mu &\sim N(0, 1), \\ \Gamma &\sim \text{Gamma}(2, 1), \end{aligned}$$

calculate the Single-Prior Bayes estimator under squared error loss.

*Proof.*

$$\begin{aligned} p(\mathbf{X} | \mu, \Gamma) &= \Gamma^n (2\pi)^{-\frac{n}{2}} \exp \left[ -2\Gamma^2 \sum_{i=1}^n (x_i - \mu)^2 \right], \\ p(\mu) &= \frac{1}{\sqrt{2\pi}} \exp \left( -\frac{\mu^2}{2} \right), \\ p(\Gamma) &= \frac{1}{\Gamma(2)} \Gamma \exp(-\Gamma). \end{aligned}$$

Therefore,

$$h(\mathbf{X}, \mu, \Gamma) = C \Gamma^n \exp \left[ -2\Gamma^2 \sum_{i=1}^n (x_i - \mu)^2 \right] \exp \left( -\frac{\mu^2}{2} \right) \Gamma \exp(-\Gamma),$$

where  $C = \frac{(2\pi)^{-\frac{n+1}{2}}}{\Gamma(2)}$ .  
 For  $\mu$ , we have

$$\pi(\mu \mid \mathbf{X}, \Gamma) = \frac{h(\mathbf{X}, \mu, \Gamma)}{p(\mu \mid \mathbf{X})}$$

□

For exponential families

### Theorem 30.1.3

## 30.1.2 Hierarchical Bayes

In a Hierarchical Bayes model, rather than specifying the prior distribution as a single function, we specify it in a **hierarchy**. Thus, the Hierarchical Bayes model in a general form as

$$\begin{aligned} X \mid \theta &\sim f(x \mid \theta), \\ \Theta \mid \gamma &\sim \pi(\theta \mid \lambda), \\ \Gamma &\sim \psi(\gamma), \end{aligned} \tag{30.9}$$

where we assume that  $\psi(\cdot)$  is known and not dependent on any other unknown hyperparameters.

*Remark.* We can continue this hierarchical modeling and add more stages to the model, but this is not then done in practice.

Given a loss function  $L(\theta, d)$ , we would then determine the estimator that minimizes

$$\int L(\theta, d(x)) \pi(\theta \mid x) d\theta, \tag{30.10}$$

where  $\pi(\theta \mid x)$  is posterior distribution given by

$$\pi(\theta \mid x) = \frac{\int f(x \mid \theta) \pi(\theta \mid \gamma) \psi(\gamma) d\gamma}{\int \int f(x \mid \theta) \pi(\theta \mid \gamma) \psi(\gamma) d\theta d\gamma}.$$

*Remark.* The posterior distribution can also be written as

$$\pi(\theta \mid x) = \int \pi(\theta \mid x, \gamma) \pi(\gamma \mid x) d\gamma,$$

where  $\pi(\gamma \mid x)$  is the posterior distribution of  $\Gamma$ , unconditional on  $\theta$ . The equation 30.10 can be written as

$$\int L(\theta, d(x)) \pi(\theta \mid x) d\theta = \int \left[ \int L(\theta, d(x)) \pi(\theta \mid x, \gamma) d\theta \right] \pi(\gamma \mid x) d\gamma.$$

which shows that **the Hierarchical Bayes estimator can be thought of as a mixture of Single-Prior estimators.**

**Example** (Poisson Hierarchy). Consider

$$\begin{aligned} X_i | \lambda &\stackrel{\text{i.i.d.}}{\sim} \text{Poisson}(\lambda), \quad i = 1, 2, \dots, n \\ \lambda | b &\sim \text{Gamma}(a, b), \text{ a known,} \\ \frac{1}{b} &\sim \text{Gamma}(k, \tau), \end{aligned} \tag{30.11}$$

calculate the Hierarchical Bayes estimator under squared error loss.

**Theorem 30.1.4**

For the Hierarchical Bayes model (30.9),

$$K[\pi(\lambda | x), \psi(\lambda)] < K[\pi(\theta | x), \pi(\theta)], \tag{30.12}$$

where  $K$  is the Kullback-Leibler information for discrimination between two densities.

*Proof.*

□

*Remark.*

### 30.1.3 Empirical Bayes

### 30.1.4 Bayes Prediction

# Chapter 31

## Nonparametric Statistics

### 31.1 Probability Distribution

#### 31.1.1 Cumulative Distribution Function

Let  $X_1, \dots, X_n \sim F$  where  $F(x) = \mathbb{P}(X \leq x)$  is a distribution function on the real line.

##### Definition 31.1.1 (Empirical Cumulative Distribution Function)

The empirical cumulative distribution function  $\hat{F}_n$  is the CDF that puts mass  $1/n$  at each data point  $X_i$ , that,

$$\hat{F}_n(x) = \frac{1}{n} \sum_{i=1}^n I(X_i \leq x) \quad (31.1)$$

Denote  $Y = \sum_{i=1}^n I(X_i \leq x)$ , then  $Y$  is a binomial random variable with parameters  $n$  and  $F(x)$ , that is,  $Y \sim b(n, F(x))$ . Therefore, we have

$$E[\hat{F}_n(x)] = F(x), \quad \text{Var}[\hat{F}_n(x)] = \frac{1}{n} F(x)[1 - F(x)]$$

so that  $\hat{F}_n(a)$  is unbiased and its variance is of order  $1/n$ . In addition, it follows from CLT (15.1.2) that

$$\sqrt{n} [\hat{F}_n(x) - F(x)] \xrightarrow{d} \mathcal{N}(0, F(x)[1 - F(x)]).$$

Thus,  $\hat{F}_n(x)$  is a consistent estimator of  $F(x)$  for each fixed  $x$ .

However, a much stronger consistency property can be asserted if the difference between  $\hat{F}_n(x)$  and  $F(x)$  is considered not only for a fixed  $x$  but simultaneously

for all  $x$ , namely

$$D_n = \sup_x \left| \hat{F}_n(x) - F(x) \right| \xrightarrow{P} 0, \quad n \rightarrow \infty.$$

For a still stronger result, see, for example, Serfling (1980, Section 2.1.4) or Billingsley (1986, Theorem 20.6).

### 31.1.2 Probability Density Function

**Histogram Estimator** Since

$$f(x) = \lim_{h \rightarrow 0} \frac{F(x+h) - F(x-h)}{2h},$$

one might consider the estimator

$$\hat{f}_n(x) = \frac{\hat{F}_n(x+h) - \hat{F}_n(x-h)}{2h}.$$

and hope that with  $h = h_n$  tending to 0 as  $n \rightarrow \infty$ , the estimator  $\hat{f}_n(x)$ , the so-called Rosenblatt estimator, will be consistent for  $f(x)$ .

*Proof.* The basic properties of  $\hat{p}_n$  are obtained from the fact that

$$n \left[ \hat{F}_n(x+h_n) - \hat{F}_n(x-h_n) \right] \sim b(n, p),$$

where  $p = F(x+h_n) - F(x-h_n)$ . It follows that

$$\mathbb{E} \left[ \hat{f}_n(x) \right] = \frac{F(x+h_n) - F(x-h_n)}{2h_n}.$$

The bias is therefore

$$b(x) = \frac{F(x+h_n) - F(x-h_n)}{2h_n} - \lim_{h_n \rightarrow 0} \frac{F(x+h_n) - F(x-h_n)}{2h_n}$$

which tends to be zero, provided  $h_n \rightarrow 0$ , as  $n \rightarrow \infty$ .

Similarly, the variance of  $\hat{f}_n(x)$  is

$$\text{Var} \left[ \hat{f}_n(x) \right] = \frac{p(1-p)}{4nh_n^2}.$$

As  $h_n \rightarrow 0$ , the value of  $p$

$$p_n = F(x+h_n) - F(x-h_n) \rightarrow 0.$$

thus,

$$\text{Var} [\hat{f}_n(x)] \sim \frac{p_n}{2h_n} \cdot \frac{1}{2nh_n}.$$

Since the first factor on the right side tends to  $f(x) > 0$ ,  $\text{Var} [\hat{f}_n(x)] \rightarrow 0$  as  $h_n \rightarrow 0$  if in addition  $nh_n \rightarrow \infty$ , that is, if  $h_n$  tends to 0 more slowly than  $1/n$  or, equivalently, if  $\frac{1}{n} = o(h_n)$ . From these results, we immediately obtain sufficient conditions for the consistency of  $\hat{f}_n(x)$ .  $\square$

It is interesting to note that  $\hat{f}_n(x)$  is itself a probability density. Since it is non-negative, one only needs to show that  $\int_{-\infty}^{\infty} \hat{f}_n(x) dx = 1$ . This is easily seen by writing

$$\hat{f}_n(x) = \frac{1}{2nh} \sum_{i=1}^n I(x - h_n \leq X_i \leq x + h_n)$$

Then

$$\int \hat{f}_n(x) dx = \frac{1}{2nh} \sum_{i=1}^n \int_{-\infty}^{\infty} I_i(x) dx = \frac{1}{2nh} \sum_{i=1}^n \int_{x_j-h}^{x_j+h} dx = 1.$$

*Remark.* Although  $\hat{f}_n(x)$  is consistent for estimating  $f(x)$  when the  $h_n$ 's satisfy  $h_n \rightarrow 0$  and  $nh_n \rightarrow \infty$ , note that  $\hat{f}_n(x)$  as an estimator of  $f(x)$  involves two approximations:  $f(x)$  by  $[F(x + h_n) - F(x - h_n)]/2h_n$  and the latter by  $\hat{f}_n(x)$ . As a result, the estimator turns out to be less accurate than one might have hoped. Besides, the estimator  $\hat{f}_n(x)$ , although a density, is a step function with discontinuities at every point  $x_i \pm h_n, i = 1, \dots, n$ . If we assume the true  $f(x)$  to be a smooth density, we may prefer its estimator also to be smooth.

**Kernel Density Estimation** We can rewrite it in the form:

$$\hat{f}_n(x) = \frac{1}{2nh_n} \sum_{i=1}^n I(x - h_n < X_i \leq x + h_n) = \frac{1}{nh_n} \sum_{i=1}^n K_0\left(\frac{X_i - x}{h}\right),$$

where  $K_0(u) = \frac{1}{2}I(-1 < u \leq 1)$ . A simple generalization of the Rosenblatt estimator is given by

$$\hat{f}_n(x) = \frac{1}{nh_n} \sum_{i=1}^n K\left(\frac{X_i - x}{h_n}\right), \quad (31.2)$$

where  $K : \mathbb{R} \rightarrow \mathbb{R}$  is an integrable function satisfying  $\int K(u) d\mu = 1$ . Such a function  $K$  is called a kernel and the parameter  $h_n$  is called a bandwidth of the estimator (31.2). The function  $x \mapsto \hat{f}_n(x)$  is called the kernel density estimator or the Parzen-Rosenblatt estimator. In addition, we shall restrict attention to kernel  $K$  that are symmetric about 0.



In generalization of (31.2), note that  $\hat{f}_n$  is a probability density since it is non-negative and since

$$\int \hat{f}_n(x) dx = \frac{1}{n} \sum_{i=1}^n \int K(\mu - x_i) d\mu = \frac{1}{n} \sum_{i=1}^n \int K(\mu) d\mu = 1.$$

### Theorem 31.1.1

Let the density  $f$  satisfies  $f(x) \leq f_{\max} < \infty$  for all  $x$  and let  $K$  be a kernel that  $\int K^2(\mu) d\mu < \infty$ . Then for any sequence  $h_n, n = 1, 2, \dots$ ,

$$\text{Var} [\hat{f}_n(x)] = \frac{f_{\max}}{nh_n} \int K^2(\mu) d\mu + o\left(\frac{1}{nh_n}\right).$$

*Proof.* Denote

$$\eta_i(x) = K\left(\frac{x - X_i}{h}\right) - \mathbb{E}\left[K\left(\frac{x - X_i}{h}\right)\right].$$

The random variables  $\eta_i(x)$  are i.i.d with zero mean and variance

$$\begin{aligned} \mathbb{E} [\eta_i^2(x)] &\leq \mathbb{E} \left[ K^2\left(\frac{x - X_i}{h}\right) \right] \\ &= \int K^2\left(\frac{x - z}{h}\right) f(z) dz \leq f_{\max} h \int K^2(\mu) d\mu. \end{aligned}$$

Then, we have

$$\text{Var} [\hat{f}_n(x)] = \frac{1}{n^2 h^2} \sum_{i=1}^n \mathbb{E} [\eta_i^2(x)] = \frac{1}{nh^2} \mathbb{E} [\eta_1^2(x)] \leq \frac{f_{\max}}{nh} \int K^2(\mu) d\mu.$$

□

### Theorem 31.1.2 (Bias of Kernel Density Estimator)

Let  $f$  be three times differentiable with a bounded third derivative in a neighborhood of  $y$  and let  $K$  be a kernel symmetric about 0, with

$$\int K^2(\mu) d\mu < \infty, \quad \int \mu^2 K(\mu) d\mu = \tau^2 < \infty, \quad \int |\mu|^3 K(\mu) d\mu < \infty.$$

Then for any sequence  $h_n, n = 1, 2, \dots$ ,

$$\text{Bias} [\hat{f}_n(x)] = \frac{1}{2} h_n^2 f''(x) \tau^2 + o(h_n^2).$$

*Proof.* Suppressing the subscript  $n$ , we find for the bias of  $\hat{f}_n(x)$ ,

$$\begin{aligned} \text{Bias} [\hat{f}_n(x)] &= E \left[ \frac{1}{n} \sum_{i=1}^n \frac{1}{h} K \left( \frac{x - X_i}{h} \right) \right] - f(x) \\ &= \int \frac{1}{h} K \left( \frac{x - X}{h} \right) f(X) dX - f(x) \\ &= \int K(\mu) [f(x - h\mu) - f(x)] d\mu. \end{aligned}$$

By the Taylor expansion of  $f(x - h\mu)$  about  $x$ , we have

$$f(x - h\mu) = f(x) + h\mu f'(x) + \frac{1}{2}(h\mu)^2 f''(x) + \frac{1}{6}(h\mu)^3 f'''(\xi),$$

where  $\xi$  lies between  $x$  and  $x - h\mu$ . Using the fact that  $\int \mu K(\mu) d\mu = 0$ , we therefore have

$$\text{Bias} [\hat{f}_n(x)] = \frac{1}{2} h^2 f''(x) \int \mu^2 K(\mu) d\mu + R_n,$$

where, with  $|f'''(z)| \leq M$ ,

$$|R_n| \leq \frac{Mh^3}{6} \int |z|^3 K(z) dz = o(h^2),$$

which proves the first part of the theorem.  $\square$

## 31.2 Kernel Methods

### 31.2.1 Positive Definite Kernels

#### Definition 31.2.1 (Positive Definite Kernel)

Let  $\mathcal{X}$  be a set, a function  $K : \mathcal{X} \times \mathcal{X} \rightarrow \mathbb{R}$  is called a positive definite kernel on  $\mathcal{X}$  if and only if it is

1. symmetric, that is,

$$K(\mathbf{x}, \mathbf{x}') = K(\mathbf{x}', \mathbf{x}), \quad \forall \mathbf{x}, \mathbf{x}' \in \mathcal{X} \quad (31.3)$$

2. positive definite, that is,

$$\sum_{i=1}^n \sum_{j=1}^n c_i c_j K(\mathbf{x}_i, \mathbf{x}_j) \geq 0, \quad (31.4)$$

holds for any  $x_1, \dots, x_n \in \mathcal{X}$ , given  $n \in \mathbb{N}$ ,  $c_1, \dots, c_n \in \mathbb{R}$ .

## Construction of the Reproducing Kernel Hilbert Space

**Theorem 31.2.1 (Morse-Aronszajn's Theorem)**

For any set  $\mathcal{X}$ , suppose  $K : \mathcal{X} \times \mathcal{X} \rightarrow \mathbb{R}$  is positive definite, then there is a unique RKHS  $\mathcal{H} \subset \mathbb{R}^{\mathcal{X}}$  with reproducing kernel  $K$ .

*Proof.* 1. How to build a valid pre-RKHS  $\mathcal{H}_0$ ?

Consider the vector space  $\mathcal{H}_0 \subset \mathcal{R}^{\mathcal{X}}$  spanned by the functions  $\{K(\cdot, \mathbf{x})\}_{\mathbf{x} \in \mathcal{X}}$ .

For any  $f, g \in \mathcal{H}_0$ , suppose

$$f = \sum_{i=1}^m a_i K(\cdot, \mathbf{x}_i), \quad g = \sum_{j=1}^n b_j K(\cdot, \mathbf{y}_j)$$

and let the inner product of  $\mathcal{H}_0$  be

$$\langle f, g \rangle = \sum_{i=1}^m \sum_{j=1}^n a_i b_j K(\mathbf{x}_i, \mathbf{y}_j) \quad (31.5)$$

Let  $\mathbf{x} \in \mathcal{X}$ ,

$$\langle f, K(\cdot, \mathbf{x}) \rangle_{\mathcal{H}_0} = \sum_{i=1}^m a_i K(\mathbf{x}, \mathbf{x}_i) = f(\mathbf{x})$$

And, we also have

$$\langle f, g \rangle_{\mathcal{H}_0} = \sum_{i=1}^m a_i g(\mathbf{x}_i) = \sum_{j=1}^n b_j f(\mathbf{y}_j)$$

Suppose

$$f = \sum_{i=1}^m a_i K(\cdot, \mathbf{x}_i), \quad g = \sum_{j=1}^n b_j K(\cdot, \mathbf{y}_j), \quad h = \sum_{k=1}^p c_k K(\cdot, \mathbf{z}_k)$$

(a) Linearity: For any  $\alpha, \beta \in \mathbb{R}$ ,  $\langle \alpha f + \beta g, h \rangle_{\mathcal{H}_0} = \alpha \langle f, h \rangle_{\mathcal{H}_0} + \beta \langle g, h \rangle_{\mathcal{H}_0}$ .

$$\begin{aligned} \langle \alpha f + \beta g, h \rangle_{\mathcal{H}_0} &= \left[ \alpha \sum_{i=1}^m a_i K(\cdot, \mathbf{x}_i) + \beta \sum_{j=1}^n b_j K(\cdot, \mathbf{y}_j) \right] \cdot \sum_{k=1}^p c_k K(\cdot, \mathbf{z}_k) \\ &= \alpha \sum_{i=1}^m \sum_{k=1}^p a_i c_k K(\mathbf{x}_i, \mathbf{z}_k) + \beta \sum_{j=1}^n \sum_{k=1}^p b_j c_k K(\mathbf{y}_j, \mathbf{z}_k) \\ &= \alpha \langle f, h \rangle_{\mathcal{H}_0} + \beta \langle g, h \rangle_{\mathcal{H}_0} \end{aligned}$$

(b) Conjugate Symmetry:  $\langle f, g \rangle_{\mathcal{H}_0} = \langle g, f \rangle_{\mathcal{H}_0}$ .

$$\begin{aligned} \langle f, g \rangle_{\mathcal{H}_0} &= \sum_{i=1}^m \sum_{j=1}^n a_i b_j K(\mathbf{x}_i, \mathbf{y}_j) = \sum_{j=1}^n \sum_{i=1}^m b_j a_i K(\mathbf{y}_j, \mathbf{x}_i) \\ &= \langle g, f \rangle_{\mathcal{H}_0} \end{aligned}$$

- (c) Positive Definiteness:  $\langle f, f \rangle_{\mathcal{H}_0} \geq 0$  and  $\langle f, f \rangle_{\mathcal{H}_0} = 0$  if and only if  $f = 0$ .  
By positive definiteness of  $K$ , we have:

$$\langle f, f \rangle_{\mathcal{H}_0} = \|f\|_{\mathcal{H}_0}^2 = \sum_{i=1}^m \sum_{j=1}^m a_i a_j K(\mathbf{x}_i, \mathbf{x}_j) \geq 0$$

As for,  $\langle f, f \rangle_{\mathcal{H}_0} = 0$  if and only if  $f = 0$ , we have,  
" $\Rightarrow$ " If  $f = 0$ , that is  $f = \sum_{i=1}^m a_i K(\cdot, \mathbf{x}_i) = 0$ , we have

$$\langle f, f \rangle_{\mathcal{H}_0} = \sum_{i=1}^m a_i f = 0$$

" $\Leftarrow$ " For  $\forall \mathbf{x} \in \mathcal{X}$ , by Cauchy-Schwarz Inequality, we have,

$$|f(\mathbf{x})| = |\langle f, K(\cdot, \mathbf{x}) \rangle_{\mathcal{H}_0}| \leq \|f\|_{\mathcal{H}_0} \cdot K(\mathbf{x}, \mathbf{x})^{\frac{1}{2}}$$

therefore, if  $\|f\|_{\mathcal{H}_0} = 0$ , then  $f = 0$

Hence, the definition in equation (31.5) is a valid inner product, which is a valid pre-RKHS  $\mathcal{H}_0$ .

□

## Examples of Kernels

**Example** (Gaussian Kernel).

$$K(\mathbf{x}, \mathbf{y}) = \exp\left(-\frac{\|\mathbf{x} - \mathbf{y}\|^2}{2\sigma^2}\right), \quad \mathbf{x}, \mathbf{y} \in \mathbb{R}^d \quad (31.6)$$

*Proof.* 1. It is obvious that  $K(\mathbf{x}, \mathbf{y})$  is symmetric, we only need to show  $K(\mathbf{x}, \mathbf{y})$  is positive definite.

$$\begin{aligned} K(\mathbf{x}, \mathbf{y}) &= \exp\left(-\frac{\|\mathbf{x} - \mathbf{y}\|^2}{2\sigma^2}\right) \\ &= \exp\left(-\frac{1}{2\sigma^2}\|\mathbf{x}\|^2\right) \cdot \exp\left(\frac{1}{\sigma^2}\langle \mathbf{x}, \mathbf{y} \rangle\right) \cdot \exp\left(-\frac{1}{2\sigma^2}\|\mathbf{y}\|^2\right) \end{aligned}$$

By the Taylor expansion of the exponential function, that

$$\exp\left(\frac{x}{\sigma^2}\right) = \sum_{n=0}^{+\infty} \left\{ \frac{x^n}{\sigma^{2n} \cdot n!} \right\}$$

Hence,

$$\exp\left(\frac{1}{\sigma^2}\langle \mathbf{x}, \mathbf{y} \rangle\right) = \sum_{n=0}^{+\infty} \left\{ \frac{\langle \mathbf{x}, \mathbf{y} \rangle^n}{\sigma^{2n} \cdot n!} \right\}$$

By the Multinomial Theorem, we have

$$\begin{aligned} \langle \mathbf{x}, \mathbf{y} \rangle^n &= \left( \sum_{i=1}^d x_i y_i \right)^n = \sum_{k_1+k_2+\dots+k_d=n} \left[ \binom{n}{k_1, k_2, \dots, k_d} \prod_{i=1}^d (x_i y_i)^{k_i} \right] \\ &= \sum_{k_1+k_2+\dots+k_d=n} \left[ \binom{n}{k_1, k_2, \dots, k_d}^{\frac{1}{2}} \prod_{i=1}^d x_i^{k_i} \cdot \binom{n}{k_1, k_2, \dots, k_d}^{\frac{1}{2}} \prod_{i=1}^d y_i^{k_i} \right] \end{aligned}$$

Therefore,

$$\begin{aligned} K(\mathbf{x}, \mathbf{y}) &= \exp \left( -\frac{\|\mathbf{x} - \mathbf{y}\|^2}{2\sigma^2} \right) = \exp \left( -\frac{\|\mathbf{x}\|^2}{2\sigma^2} \right) \cdot \exp \left( -\frac{\|\mathbf{y}\|^2}{2\sigma^2} \right) \cdot \sum_{n=0}^{+\infty} \left\{ \frac{\langle \mathbf{x}, \mathbf{y} \rangle^n}{\sigma^{2n} \cdot n!} \right\} \\ &= \sum_{n=0}^{+\infty} \frac{\exp \left( -\frac{\|\mathbf{x}\|^2}{2\sigma^2} \right)}{\sigma^n \cdot \sqrt{n!}} \cdot \frac{\exp \left( -\frac{\|\mathbf{y}\|^2}{2\sigma^2} \right)}{\sigma^n \cdot \sqrt{n!}} \cdot \langle \mathbf{x}, \mathbf{y} \rangle^n \end{aligned}$$

Let

$$c_{\sigma,n}(\mathbf{x}) = \frac{\exp \left( -\frac{\|\mathbf{x}\|^2}{2\sigma^2} \right)}{\sigma^n \cdot \sqrt{n!}}, \quad f_{n,\mathbf{k}}(\mathbf{x}) = \binom{n}{k_1, k_2, \dots, k_d}^{\frac{1}{2}} \prod_{i=1}^d x_i^{k_i}$$

then,

$$\begin{aligned} K(\mathbf{x}, \mathbf{y}) &= \sum_{n=0}^{+\infty} \sum_{k_1+k_2+\dots+k_d=n} c_{\sigma,n}(\mathbf{x}) f_{n,\mathbf{k}}(\mathbf{x}) \cdot c_{\sigma,n}(\mathbf{y}) f_{n,\mathbf{k}}(\mathbf{y}) \\ &= \langle \Phi(\mathbf{x}), \Phi(\mathbf{y}) \rangle \end{aligned}$$

where  $\Phi(\mathbf{x})_{\sigma,n,\mathbf{k}} = c_{\sigma,n}(\mathbf{x}) f_{n,\mathbf{k}}(\mathbf{x})$ .

$$\begin{aligned} \sum_{i=1}^n \sum_{j=1}^n c_i c_j K(\mathbf{x}_i, \mathbf{x}_j) &= \sum_{i=1}^n \sum_{j=1}^n c_i c_j \langle \Phi(\mathbf{x}_i), \Phi(\mathbf{x}_j) \rangle \\ &= \left\langle \sum_{i=1}^n c_i \Phi(\mathbf{x}_i), \sum_{i=1}^n c_i \Phi(\mathbf{x}_i) \right\rangle \geq 0 \end{aligned}$$

for any  $x_1, \dots, x_n \in \mathcal{X}$ , given  $n \in \mathbb{N}, c_1, \dots, c_n \in \mathbb{R}$ , i.e.,  $K(\mathbf{x}, \mathbf{y})$  is positive definite.

□

# Chapter 32

## Minimax Theory

### 32.1 Fano's Inequality

Let  $X \sim P_\theta$ ,  $\theta \in \Theta_0 \subset \Theta$ , in which  $\Theta_0$  are assumed to be finite, e.g.,  $\{\theta_1, \dots, \theta_M\}$ , and  $\theta$  uniformly distributed on  $\Theta_0$ . Let  $\hat{\theta}$  be an estimator of  $\theta$  based on  $X$ , Then

$$P(\theta \neq \hat{\theta}) = \frac{1}{M} \sum_{i=1}^M P_{\theta_i}(\hat{\theta} \neq \theta_i) \geq 1 - \frac{I(\theta, X) + \log 2}{\log M}. \quad (32.1)$$

Question: How to upper bound  $I(\theta, X)$ ?

There are various ways to do that, one earlier bound is

$$I(\theta, X) \leq \frac{1}{M^2} \sum_{i=1}^M \sum_{j=1}^M D_{KL}(P_{\theta_i} \| \theta_j). \quad (32.2)$$

To use such a bound needs to be very careful with the construction of  $\Theta_0$ . Alternatively, insight from the information theory may provide another method to do it in a way that takes advantage of known metric entropy ???.

Typically,  $\Theta_0$  is a subset of  $\Theta$ .

Original problem  $X \sim P_\theta$ ,  $\theta \in \Theta$ . Here  $\theta$  can be a finite dimension or infinite dimension (e.g.,  $\theta = f(x)$ , pdf of  $X$ ).

Let  $\pi(\theta)$  be a prior distribution on  $\Theta$ . To apply Fano's Inequality, we need to choose  $\Theta$  finite. Regardless, we want a general bound on  $I(\theta; X)$ , in which  $\theta \sim \pi$ .

Recall

$$H(X) = \begin{cases} -\sum_i p_i \log p_i & \text{discrete} \\ -\int p(x) \log p(x) dx & \text{continuous} \end{cases} \quad (32.3)$$

Given the ?? Prof P. Shannon ?? achieves (within one bit) the lower bound on the expected code length of any prefix code, with code length  $\log \frac{1}{p_i}$  (ignoring rounding up), which is  $H(X)$ .

If we mistakenly use  $q$  to code, then the expected extra bits are (also called redundancy)

$$\sum_i p_i \log \frac{1}{q_i} - \sum_i p_i \log \frac{1}{p_i} = \sum_i p_i \log \frac{p_i}{q_i} \geq 0. \quad (32.4)$$

Bayes misfu?? mis?? the Bayes Redundancy. Let  $q$  be any pdf, redundancy at  $\theta$  is

$$\int f(x, \theta) \log \frac{f(x, \theta)}{q(x)} dx, \quad (32.5)$$

where  $f(x, \theta)$  is pdf of  $X$ . Bayes redundancy is

$$\begin{aligned} & \int_{\Theta} \left( \int f(x, \theta) \log \frac{f(x, \theta)}{q(x)} dx \right) \cdot \pi(\theta) d\theta \\ &= \int \int_{\Theta} f(x, \theta) \log \frac{\pi(\theta) f(x, \theta)}{\pi(\theta) q(x)} \cdot \pi(\theta) d\theta dx, \end{aligned} \quad (32.6)$$

Let  $q^*(x) = \int_{\Theta} f(x, \theta) \pi(\theta) d\theta$ ,

$$\begin{aligned} &= \int \int_{\Theta} f(x, \theta) \log \frac{\pi(\theta) f(x, \theta)}{\pi(\theta) q^*(x)} \cdot \pi(\theta) d\theta dx \\ &\quad + \int \int_{\Theta} f(x, \theta) \log \frac{q^*(x)}{q(x)} \pi(\theta) d\theta dx, \end{aligned} \quad (32.7)$$

in which, the first part is the Bayes redundancy of  $q^*$ , and the second part

$$\begin{aligned} &= \int \log \frac{q^*(x)}{q(x)} \left( \int_{\Theta} f(x, \theta) \pi(\theta) d\theta \right) dx \\ &= \int \log \frac{q^*(x)}{q(x)} q^*(x) dx \geq 0. \end{aligned} \quad (32.8)$$

Thus Bayes redundancy of  $q^*$  is  $I(\theta; X)$ .

Our approach is to provide a sensible upper bound on  $I(\theta; X)$ , that is not specific to the choice of  $\Theta_0$ . Rather, it reflects the native of the  $\{p_{\theta}, \theta \in \Theta\}$  (or a subset of it) more.

Suppose we have i.i.d observations  $X_1, X_2, \dots, X_n \sim p_{\theta}$ , let

$$d_k(p, q) = \sqrt{D(p||q)} = \sqrt{\int p(x) \log \frac{p(x)}{q(x)} dx}, \quad (32.9)$$

which is not a metric. Let  $G_{\varepsilon}$  be an  $\varepsilon$ -cover of the family  $\{p_{\theta}, \theta \in \Theta\}$ , i.e.,

$$\forall \theta \in \Theta, \exists \theta' \in G_{\varepsilon}, \text{ s.t. } D(p_{\theta}||p_{\theta'}) \leq \varepsilon^2. \quad (32.10)$$

Let  $M = |G_\varepsilon|$ ,  $q(x_1, \dots, x_n) = \frac{1}{M} \sum_{i=1}^M p_{\theta_i}^n$  (centroid) and  $p_{\theta_i}^n = p_{\theta_i}(x_1) \dots p_{\theta_i}(x_n)$ . Then,

$$\begin{aligned}
 D_{\text{KL}}(p_\theta^n \| q) &= \int p_\theta^n \log \frac{p_\theta^n}{\frac{1}{M} \sum_{i=1}^M p_{\theta_i}^n} dx \\
 &\leq \int p_\theta^n \log \frac{p_\theta^n}{\frac{1}{M} p_{\theta_i}^n} dx \\
 &= \log M + \inf_i \int p_\theta^n \log \frac{p_\theta^n}{p_{\theta_i}^n} dx \\
 &= \log M + \inf_i D_{\text{KL}}(p_\theta \| p_{\theta_i}) \\
 &\leq \log M + n\varepsilon^2.
 \end{aligned} \tag{32.11}$$

This holds for all  $\theta \in \Theta$ . If we have a subset  $\Theta$  and for any prior  $\pi$ , we have

$$I(\theta; X^n) \leq \log |G_\varepsilon| + n\varepsilon^2. \tag{32.12}$$

$$\begin{aligned}
 D_{\text{KL}}(p_{\theta_i}^n \| q) &\leq \log M, \quad 1 \leq i \leq M, \\
 D_{\text{KL}}(p_\theta^n \| q) &\leq \log M + n\varepsilon^2, \quad \forall \theta \in \Theta.
 \end{aligned}$$

If  $D_{\text{KL}}(p_{\theta_i} \| p_{\theta_j}) \geq \eta^2 > 0$ , then  $D_{\text{KL}}(p_{\theta_i}^n \| p_{\theta_j}^n) \geq n\eta^2$ . But with the centroid density  $q$ , we have

$$D_{\text{KL}}(p_{\theta_i}^n \| q) \leq \log M.$$

Then, we have the interesting situation that  $D_{\text{KL}}(p_{\theta_i}^n \| p_{\theta_j}^n)$  are very large for  $i \neq j$ , yet  $D_{\text{KL}}(p_{\theta_i}^n \| q)$  are small. Relaxing speaking!

## 32.2 Minimax Rate

Consider a loss function  $l(\theta, \theta')$



**Theorem 32.2.1**

Suppose on a finite set  $\Theta_0 \subset \Theta$ , we have

$$\min_{\theta_i \neq \theta_j} l(\theta_i, \theta_j) \geq \Delta > 0, \quad (32.13)$$

for any  $\theta_i \neq \theta_j \in \Theta_0$  and  $\theta \in \Theta$ , we have

$$l(\theta_i, \theta) + l(\theta_j, \theta) \geq c\Delta, \quad (32.14)$$

for some constant  $c > 0$ . Then

$$\inf_{\hat{\theta}} \sup_{\theta \in \Theta} \mathbb{E}_{\theta} l(\theta, \hat{\theta}) \geq \frac{c\Delta}{2} \left( 1 - \frac{V_k(\varepsilon) + n\varepsilon^2 + \log 2}{\log |\Theta_0|} \right), \quad (32.15)$$

where  $V_k(\varepsilon)$  is the covering entropy of  $\{p_{\theta}, \theta \in \Theta\}$  under  $d_{\text{KL}}$ .

For  $f, g$  with different supports, we possibly have  $D(f\|g) = \infty$ , suppose density are w.r.t a probability measure  $\mu$ . Given original observations  $X_1, \dots, X_n \sim f \in \mathcal{F}$ , let  $Y_1, \dots, Y_n$  be i.i.d uniform w.r.t  $\mu$ , and  $V_1, \dots, V_n$  be coin flips. Suppose

$$Z_i = \begin{cases} X_i, & \text{if } V_i = 1, \\ Y_i, & \text{if } V_i = 0. \end{cases}$$

Then,  $Z_1, \dots, Z_n \sim \frac{f}{2} + \frac{1}{2}$ .

*Proof.* Let

$$\tilde{\theta} = \arg \min_{\theta_i \in \Theta_0} l(\theta_i, \hat{\theta}).$$

Then, we have  $\theta \neq \tilde{\theta}$ , we know

$$l(\theta, \hat{\theta}) \geq l(\tilde{\theta}, \hat{\theta}),$$

and

$$l(\theta, \hat{\theta}) + l(\tilde{\theta}, \hat{\theta}) \geq c\Delta.$$

Consequently,  $l(\theta, \hat{\theta}) \geq \frac{c\Delta}{2}$ . Thus,

$$\inf_{\hat{\theta}} \sup_{\theta \in \Theta} \mathbb{E}_{\theta} l(\theta, \hat{\theta}) \geq \inf_{\hat{\theta}} \sup_{\theta \in \Theta_0} \mathbb{E}_{\theta} l(\theta, \hat{\theta}).$$

□

### 32.3 Applications

1. Consider a fixed design regression with independent observations:

$$Y_i = u(X_i) + e_i, \quad 1 \leq i \leq n,$$

where  $x_i = i/n$ ,  $e_i \sim N(0, 1)$  and  $\mathcal{U}$  consists of all functions  $g$  on  $[0, 1]$  that are uniformly bounded between  $-A$  and  $A$  for some positive constant  $A$  and  $|g(x) - g(y)| \leq L|x - y|$  for some constant  $L > 0$ . The loss function of interest is  $\ell(u, \hat{u}) = \int_0^1 (u(x) - \hat{u}(x))^2 dx$ . Show that

$$\inf_{\hat{u}} \sup_{u \in \mathcal{U}} E_u \ell(u, \hat{u}) \asymp n^{-2/3}.$$

Note that some results given in the lectures on regression may not be directly applicable because here we deal with a fixed design. You may consider a piecewise constant estimator for upper bounding and you may use the fact that  $\mathcal{U}$  has  $L_2$  metric entropy order  $1/\epsilon$ .

*Proof.* Since  $X$  are fixed design, then  $y_i \sim N(u(x_i), 1)$ ,  $i = 1, 2, \dots, n$ , which are independent but no longer iid. Let  $P_u$  denotes the distribution of  $Y$  with regression function  $u$ . Thus,

$$D_{\text{KL}}(p_u^n \| p_v^n) = \frac{1}{2} \sum_{i=1}^n [u(x_i) - v(x_i)]^2,$$

where  $p_u^n(\mathbf{y}) = \prod_{i=1}^n p_u(y_i)$ , and  $p_v^n(\mathbf{y}) = \prod_{i=1}^n p_v(y_i)$ .

**Lower Bound:**

□

2. Consider the collection  $\mathcal{A} = \{(a, b) : -\infty < a < b < \infty\}$  comprising sets in the real number line. Show its VC dimension is 2.

*Proof.* Since we have  $S_{\mathcal{A}}(n) = \frac{n(n+1)}{2} + 1$ , it follows that the VC dimension of  $\mathcal{A}$  is 2. □

# Chapter 33

## Multivariate Extensions

### 33.1 Applications

#### 33.1.1 Mean vector

Let  $\mathbf{X}_i$ ,  $i = 1, \dots, n$  be drawn from a  $p$ -dimensional distribution with mean vector  $\boldsymbol{\mu}$  and covariance matrix  $\boldsymbol{\Sigma}$ . The sample mean vector is given by

$$\bar{\mathbf{X}} = \frac{1}{n} \sum_{i=1}^n \mathbf{X}_i.$$

By the multivariate central limit theorem, we have

$$\sqrt{n}(\bar{\mathbf{X}} - \boldsymbol{\mu}) \xrightarrow{d} \mathcal{N}(\mathbf{0}, \boldsymbol{\Sigma}).$$

Applying the continuous mapping theorem and Theorem 16.1.1, we obtain

$$n(\bar{\mathbf{X}} - \boldsymbol{\mu})^\top \boldsymbol{\Sigma}^{-1}(\bar{\mathbf{X}} - \boldsymbol{\mu}) \sim \chi_p^2.$$

If  $\boldsymbol{\Sigma}$  is known, then the confidence region for  $\boldsymbol{\mu}$  with confidence level  $1 - \alpha$  is given by

$$n(\bar{\mathbf{X}} - \boldsymbol{\mu})^\top \boldsymbol{\Sigma}^{-1}(\bar{\mathbf{X}} - \boldsymbol{\mu}) \leq C_p,$$

where

$$\int_0^{C_p} \chi_p^2(t) dt = 1 - \alpha.$$

In applications where  $\boldsymbol{\Sigma}$  is unknown, we can use the sample covariance matrix  $\hat{\boldsymbol{\Sigma}}$  to replace  $\boldsymbol{\Sigma}$ , and  $\hat{\boldsymbol{\Sigma}}^{-1}$  is a consistent estimator of  $\boldsymbol{\Sigma}^{-1}$ . This leads to the following confidence region for  $\boldsymbol{\mu}$ :

$$n(\bar{\mathbf{X}} - \boldsymbol{\mu})^\top \hat{\boldsymbol{\Sigma}}^{-1}(\bar{\mathbf{X}} - \boldsymbol{\mu}) \leq C_p,$$

which provides asymptotically valid confidence regions for  $\boldsymbol{\mu}$  at level  $1 - \alpha$  for any non-singular  $\boldsymbol{\Sigma}$  and fixed shape of distribution  $F$ .

*Proof.* To show that

$$\Pr \left( n(\bar{\mathbf{X}} - \boldsymbol{\mu})^\top \boldsymbol{\Sigma}^{-1} (\bar{\mathbf{X}} - \boldsymbol{\mu}) \leq C_p \right) \rightarrow 1 - \alpha, \text{ as } n \rightarrow \infty.$$

□

*Remark.* If  $\mathbf{X}_i$ ,  $i = 1, \dots, n$  are drawn from a  $p$ -dimensional multivariate normal distribution, the exact distribution of  $n(\bar{\mathbf{X}} - \boldsymbol{\mu})^\top \boldsymbol{\Sigma}^{-1} (\bar{\mathbf{X}} - \boldsymbol{\mu})$  is a Hotelling's  $T^2$  distribution with parameters  $p$  and  $n - 1$ .

We next consider the power of

### 33.1.2 Difference of two mean vectors

$$T_{mn} = [(\boldsymbol{\eta} - \boldsymbol{\xi}) - (\bar{\mathbf{Y}} - \bar{\mathbf{X}})]^\top \left( \frac{1}{m} \hat{\boldsymbol{\Sigma}} + \frac{1}{n} \hat{\boldsymbol{\Gamma}} \right)^{-1} [(\boldsymbol{\eta} - \boldsymbol{\xi}) - (\bar{\mathbf{Y}} - \bar{\mathbf{X}})] \leq C_p,$$

Univariate case:

$$T_{mn} = \frac{(\bar{Y} - \bar{X}) / \sqrt{\frac{1}{m} + \frac{1}{n}}}{\sqrt{\left[ \sum_i (X_i - \bar{X})^2 + \sum_j (Y_j - \bar{Y})^2 \right] / (m + n - 2)}}$$

Since

$$\frac{(\bar{Y} - \bar{X}) - (\eta - \xi)}{\sqrt{\frac{\sigma^2}{m} + \frac{\gamma^2}{n}}} \sim \mathcal{N}(0, 1)$$

and

$$\frac{\sum_i (X_i - \bar{X})^2}{\sigma^2} + \frac{\sum_j (Y_j - \bar{Y})^2}{\gamma^2} \sim \chi_{m+n-2}^2,$$

then, we have

$$T'_{mn} = \frac{\frac{(\bar{Y} - \bar{X}) - (\eta - \xi)}{\sqrt{\frac{\sigma^2}{m} + \frac{\gamma^2}{n}}}}{\sqrt{\left( \frac{\sum_i (X_i - \bar{X})^2}{\sigma^2} + \frac{\sum_j (Y_j - \bar{Y})^2}{\gamma^2} \right) / (m + n - 2)}} \sim t_{m+n-2}.$$

Under the null hypothesis, we have

$$T'_{mn} = T_{mn} \cdot \sqrt{\frac{\sum_i (X_i - \bar{X})^2 + \sum_j (Y_j - \bar{Y})^2}{\gamma^2 \sum_i (X_i - \bar{X})^2 + \sigma^2 \sum_j (Y_j - \bar{Y})^2}} \cdot \sqrt{\frac{(m+n)\sigma^2\gamma^2}{n\sigma^2 + m\gamma^2}}.$$

It is easy to see that

$$\begin{aligned} \sum_i (X_i - \bar{X})^2 + \sum_j (Y_j - \bar{Y})^2 &\rightarrow_p m\sigma^2 + n\gamma^2, \\ \gamma^2 \sum_i (X_i - \bar{X})^2 + \sigma^2 \sum_j (Y_j - \bar{Y})^2 &\rightarrow_p (m+n)\sigma^2\gamma^2 > 0, \end{aligned}$$

and by the continuous mapping theorem with the fact that  $m/(m+n) \rightarrow \rho$ , we have

$$\sqrt{\frac{\sum_i (X_i - \bar{X})^2 + \sum_j (Y_j - \bar{Y})^2}{\gamma^2 \sum_i (X_i - \bar{X})^2 + \sigma^2 \sum_j (Y_j - \bar{Y})^2}} \cdot \sqrt{\frac{(m+n)\sigma^2\gamma^2}{n\sigma^2 + m\gamma^2}} \rightarrow_p \sqrt{\frac{\rho\sigma^2 + (1-\rho)\gamma^2}{(1-\rho)\sigma^2 + \rho\gamma^2}}.$$

*Proof.*

$$\Pr(T_{mn} \leq C_k) \rightarrow \gamma, \text{ as } m, n \rightarrow \infty$$

For convenience, suppose  $m < n$ , then we define

$$\mathbf{Z}_i = \mathbf{X}_i - \sqrt{\frac{m}{n}} Y_i + \frac{m}{\sqrt{mn}}$$

□

### 33.1.3 Simple Linear Regression

Suppose

$$\mathbf{X}_i = \boldsymbol{\alpha} + \mathbf{v}_i \boldsymbol{\beta} + \boldsymbol{\varepsilon}_i, \quad i = 1, \dots, n$$

where  $\mathbf{v}_i$  is a  $p$ -dimensional vector of known constants, and  $\boldsymbol{\varepsilon}_i$  is a  $p$ -dimensional random vector with mean  $\mathbf{0}$  and covariance matrix  $\boldsymbol{\Sigma}$ . The least squares estimator of  $\boldsymbol{\alpha}$  and  $\boldsymbol{\beta}$  are given by

$$\hat{\beta}_j = \frac{\sum_i (X_{i,j} - \bar{X}_j)(v_{i,j} - \bar{v}_j)}{\sum_i (v_{i,j} - \bar{v}_j)^2}, \quad \hat{\alpha}_j = \bar{X}_j - \hat{\beta}_j \bar{v}_j.$$

where  $\bar{\mathbf{X}} = \frac{1}{n} \sum_i \mathbf{X}_i$  and  $\bar{\mathbf{v}} = \frac{1}{n} \sum_i \mathbf{v}_i$ .

Denote

$$d_{n,j}^{(i)} = \frac{v_{i,j} - \bar{v}_j}{\sqrt{\sum_i (v_{i,j} - \bar{v}_j)^2}}.$$

We can rewrite the least squares estimator as

$$\hat{\beta}_j - \beta_j = \frac{\sum_i d_{n,j}^{(i)} [(X_{i,j} - \bar{X}_j) - \mathbb{E}(X_{i,j} - \bar{X}_j)]}{\sqrt{\sum_i (v_{i,j} - \bar{v}_j)^2}},$$

By Theorem, we have

$$\left( \sqrt{\sum_i (v_{i,1} - \bar{v}_1)^2} (\hat{\beta}_1 - \beta_1), \dots, \sqrt{\sum_i (v_{i,p} - \bar{v}_p)^2} (\hat{\beta}_p - \beta_p) \right)^\top \xrightarrow{d} \mathcal{N}(\mathbf{0}, \Sigma).$$

### 33.1.4 Multinomial One-Sample Test

Consider a sequence of  $n$  independent multinomial trials with  $k + 1$  categories, where the probability of the  $i$ -th category is  $p_i$ ,  $i = 1, \dots, k + 1$ .

Let us now consider testing the hypothesis

$$H_0 : p_i = p_i^{(0)}, \quad i = 1, \dots, k + 1$$

against the alternatives that  $p_i \neq p_i^{(0)}$  for at least some  $i$ . The standard test for this problem is Pearson's  $\chi^2$ -test, which rejects  $H_0$  when

$$Q = n \sum_{i=1}^{k+1} \left( \frac{Y_i}{n} - p_i^{(0)} \right)^2 / p_i^{(0)} \geq C_k,$$

The asymptotic distribution of  $Q$  is  $\chi_k^2$  under  $H_0$ .

*Proof.* It follows from (5.4.14) and Theorem 16.1.1 that

$$n \sum_{i=1}^k \sum_{j=1}^k a_{ij} \left( \frac{Y_i}{n} - p_i^{(0)} \right) \left( \frac{Y_j}{n} - p_j^{(0)} \right) \xrightarrow{d} \chi_k^2, \quad (33.1)$$

where

$$a_{ij} = \begin{cases} \frac{1}{p_i^{(0)}} + \frac{1}{p_{k+1}^{(0)}} & \text{if } i = j, \\ \frac{1}{p_{k+1}^{(0)}} & \text{if } i \neq j. \end{cases}$$

The left side of (33.1) can be written as

$$n \sum_{i=1}^k \frac{1}{p_i^{(0)}} \left( \frac{Y_i}{n} - p_i^{(0)} \right)^2 + \frac{n}{p_{k+1}^{(0)}} \sum_{i=1}^k \sum_{j=1}^k \left( \frac{Y_i}{n} - p_i^{(0)} \right) \left( \frac{Y_j}{n} - p_j^{(0)} \right),$$

The last term is equal to

$$n \left[ \sum_{i=1}^k \left( \frac{Y_i}{n} - p_i^{(0)} \right) \right]^2 / p_{k+1}^{(0)} = n \left( \frac{Y_{k+1}}{n} - p_{k+1}^{(0)} \right)^2 / p_{k+1}^{(0)},$$

and completes the proof.  $\square$

### 33.1.5 Contingency Table

# **Part XI**

## **Computational Statistics**

# Chapter 34

## Random Generator

Random number generator is a key component in computational statistics. It is used to generate random numbers from a given probability distribution. In this chapter, we will introduce some basic concepts and algorithms of random number generation.

### 34.1 Uniform Random Number Generation

### 34.2 Non-uniform Random Number Generation

For non-uniform random number generation

#### 34.2.1 Inversion Method

##### Theorem 34.2.1

Let  $X$  be a random variable with cumulative distribution function  $F(x)$ , then  $F(x)$  is a non-decreasing function and  $F(x) \in [0, 1]$ . Let  $U \sim \mathcal{U}(0, 1)$  be a random variable with uniform distribution on  $(0, 1)$ , then

$$F^{-1}(U) \sim F. \quad (34.1)$$

*Proof.* Since  $F(x)$  is a non-decreasing function, it is invertible. Let  $Y = F^{-1}(U)$ , then

$$\Pr(Y \leq y) = \Pr(F^{-1}(U) \leq y) = \Pr(U \leq F(y)) = F(y).$$

□



**Example** (Normal Distribution). Let  $X \sim \mathcal{N}(0, 1)$  be a random variable with standard normal distribution, then the cumulative distribution function of  $X$  is

$$F(x) = \frac{1}{\sqrt{2\pi}} \int_{-\infty}^x \exp\left(-\frac{t^2}{2}\right) dt. \quad (34.2)$$

Since there is no closed form of the inverse of  $F(x)$ , we can use the approximation form:

$$F^{-1}(u) \approx t - \frac{a_0 - a_1 t}{1 + b_1 t + b_2 t^2}, \quad (34.3)$$

where  $t = \sqrt{-2 \log u}$ .

### 34.2.2 Rejection Sampling Method

The inversion method is a general method to generate random numbers from a given probability distribution. However, it is not always easy to find the inverse of the cumulative distribution function. In this case, we can use the rejection sampling method.

Suppose we want to generate random numbers from a probability distribution  $f(x)$ , which is not easy to sample from. We can find a proposal distribution  $g(x)$ , which is easy to sample from and satisfies

$$\exists M > 0, \quad f(x) \leq M g(x). \quad (34.4)$$

Then the rejection sampling method is as follows:

---

**Algorithm 2:** Rejection Sampling Method

---

**Input:** Proposal distribution  $g(x)$ , constant  $M$

1 Draw a sample  $x \sim g(x)$  and  $u \sim \mathcal{U}(0, 1)$ ;

2 **if**  $u \leq \frac{f(x)}{Mg(x)}$  **then**

3     Accept  $x$ ;

4 **else**

5     Reject  $x$  and go to step 1;

**Output:** Sample  $x$

---

#### Theorem 34.2.2

The rejection sampling method generates a sample  $x$  from the probability distribution  $f(x)$ .

*Proof.* Let  $I = 1$  if  $x$  is accepted and  $I = 0$  if  $x$  is rejected. Then the probability of accepting  $x$  given  $x$  is

$$\Pr(I = 1 \mid x) = \Pr(u \leq \frac{f(x)}{Mg(x)}) = \frac{f(x)}{Mg(x)}.$$

Thus, the probability of accepting  $x$  is

$$\Pr(x \mid I = 1) = \frac{\Pr(x, I = 1)}{\Pr(I = 1)} = \frac{\Pr(x) \Pr(I = 1 \mid x)}{\int \Pr(x) \Pr(I = 1 \mid x) \mathrm{d}x} = \frac{f(x)/M}{\int f(x)/M \mathrm{d}x} = f(x).$$

□

### 34.3 Markov Chain Monte Carlo

Markov Chain Monte Carlo (MCMC) is a class of algorithms for sampling from a probability distribution based on constructing a Markov chain that has the desired distribution as its equilibrium distribution. The state of the chain after several steps is then used as a sample of the desired distribution. The quality of the sample improves as a function of the number of steps.

#### 34.3.1 Metropolis-Hastings Sampling

We want to sample from a distribution  $\pi(x)$ , where  $x \in \mathcal{X}$ .

---

**Algorithm 3:** Random Walk Metropolis-Hastings Sampling

---

**Input:** Initial state  $x^{(0)}$ , number of iterations  $N$

```

1 for  $i = 1, \dots, N$  do
2   Sample  $y \sim \mathcal{N}(x^{(i-1)}, \Sigma)$ ;
3   Compute  $\alpha = \min \left\{ 1, \frac{\pi(y)}{\pi(x^{(i-1)})} \right\}$ ;
4   Sample  $u \sim \mathcal{U}(0, 1)$ ;
5   if  $u < \alpha$  then
6      $x^{(i)} = y$ ;
7   else
8      $x^{(i)} = x^{(i-1)}$ ;

```

**Output:** Samples  $x^{(1)}, \dots, x^{(N)}$

---

#### Random Walk Metropolis-Hastings Sampling

*Remark.* It is usually difficult to sample from a high-dimensional distribution due to the rejection progressively increasing with the dimensionality.

### 34.3.2 Gibbs Sampling

Gibbs sampling is a special case of Metropolis-Hastings sampling. It is applicable when the joint distribution is not known explicitly or is difficult to sample from directly, but the conditional distribution of each variable is known and easier to sample from. Gibbs sampling only attempts transitions in the coordinate axis direction, using the conditional distribution of the current point to determine the next step's proposal distribution. All proposal samples are accepted without rejection, resulting in potentially higher efficiency.

Let  $\mathbf{x} = (x_1, \dots, x_d)^\top$  be a random vector with joint distribution  $\pi(\mathbf{x})$ , and let  $\pi(x_i|\mathbf{x}_{-i})$  be the conditional distribution of  $x_i$  given  $\mathbf{x}_{-i} = (x_1, \dots, x_{i-1}, x_{i+1}, \dots, x_d)^\top$ . The Gibbs sampling algorithm is as follows:

---

**Algorithm 4:** Gibbs Sampling

---

**Input:** Initial state  $\mathbf{x}^{(0)}$ , number of iterations  $N$ , burn-in period  $B$

```

1 for  $i = 1, \dots, N$  do
2   for  $j = 1, \dots, d$  do
3      $\lfloor$  Sample  $x_j^{(i)} \sim \pi(x_j|x_1^{(i)}, \dots, x_{j-1}^{(i)}, x_{j+1}^{(i-1)}, \dots, x_d^{(i-1)})$ ;
4    $\rfloor$   $\mathbf{x}^{(i)} = (x_1^{(i)}, \dots, x_d^{(i)})^\top$ ;
```

**Output:** Samples  $\mathbf{x}^{(B+1)}, \dots, \mathbf{x}^{(N)}$

---

# Chapter 35

## Monte Carlo Integration

Suppose we want to estimate the expectation of a function  $h(x)$  for a probability distribution  $\pi(x)$ , i.e., we want to estimate

$$\mu = \mathbb{E}_\pi[h(x)] = \int h(x)\pi(x) \, dx. \quad (35.1)$$

### 35.1 Monte Carlo Integration

If we can sample from  $\pi(x)$ , then we can use the Monte Carlo integration method as follows:

$$\hat{\mu} = \frac{1}{N} \sum_{i=1}^N h(x_i), \quad (35.2)$$

where  $x_i \sim \pi(x)$ ,  $i = 1, \dots, N$ .

### 35.2 Importance Sampling

Importance sampling is a variance reduction technique that can be used when sampling from a distribution  $\pi(x)$  is difficult, but sampling from a distribution  $g(x)$  is easy. The idea is to sample from  $g(x)$  and then reweight the samples so that they are distributed according to  $\pi(x)$ .

If the probability distribution  $\pi(x)$  is difficult to sample from, we can find a proposal distribution  $g(x)$ . Then the Importance sampling method is as follows:

---

**Algorithm 5:** Importance Sampling Method

---

**Input:** Proposal distribution  $g(x)$ , number of samples  $N$ 

- 1 **for**  $i = 1, \dots, N$  **do**
- 2     Draw a sample  $x_i \sim g(x)$ ;
- 3     Compute  $w_i = \frac{\pi(x_i)}{g(x_i)}$ ;
- 4 Calculate  $\hat{\mu} = \frac{1}{N} \sum_{i=1}^N w_i h(x_i)$ ;

**Output:** Estimate  $\hat{\mu}$ 

---

**Normalized Importance Sampling** If we do not know the normalization constant of  $\pi(x)$ , we can use the normalized importance sampling method as follows:

$$\hat{\mu} = \frac{\sum_{i=1}^N w_i h(x_i)}{\sum_{i=1}^N w_i}. \quad (35.3)$$

# Chapter 36

## Bootstrap

Bootstrap is a statistical method for estimating the sampling distribution of an estimator by sampling with replacement from the original sample, most often to derive robust estimates of standard errors and confidence intervals of a population parameter like a mean, median, proportion, odds ratio, correlation coefficient, or regression coefficient.

### 36.1 Bootstrap Principle

Suppose the i.i.d samples  $\{\mathbf{x}_1, \mathbf{x}_2, \dots, \mathbf{x}_n\}$  from an unknown probability distribution  $F$  on some probability space  $\mathcal{X}$ . Let  $\hat{\theta}_n$  be an estimator of  $\theta$  based on the sample  $\{\mathbf{x}_1, \mathbf{x}_2, \dots, \mathbf{x}_n\}$ , that is,

$$\hat{\theta}_n = s(\mathbf{x}_1, \mathbf{x}_2, \dots, \mathbf{x}_n),$$

where  $s(\cdot)$  is some algorithm.

The bootstrap principle is to use the sample  $\{\mathbf{x}_1, \mathbf{x}_2, \dots, \mathbf{x}_n\}$  to estimate the sampling distribution of  $\hat{\theta}_n$ .

**Nonparametric Bootstrap** In the  $b$ -th bootstrap replicate, we sample with replacement from the original sample  $\{\mathbf{x}_1, \mathbf{x}_2, \dots, \mathbf{x}_n\}$  to get  $\{\mathbf{x}_1^{*b}, \mathbf{x}_2^{*b}, \dots, \mathbf{x}_n^{*b}\}$ , and then compute the bootstrap estimate of  $\hat{\theta}_n$  as

$$\hat{\theta}_n^{*b} = s(\mathbf{x}_1^{*b}, \mathbf{x}_2^{*b}, \dots, \mathbf{x}_n^{*b}). \quad (36.1)$$

**Parametric Bootstrap**

**Bayesian Bootstrap**

**Smooth Bootstrap****Block Bootstrap****36.2 Standard Error Estimation**

The bootstrap estimate of the standard error of  $\hat{\theta}_n$  is

$$\widehat{\text{se}}_{\text{boot}} = \sqrt{\frac{1}{B-1} \sum_{b=1}^B (\hat{\theta}_n^{*b} - \hat{\theta}_n^*)^2}, \quad (36.2)$$

where  $\hat{\theta}_n^{*b}$  is the  $b$ -th bootstrap replicate of  $\hat{\theta}_n$  and  $\hat{\theta}_n^*$  is the bootstrap estimate of  $\hat{\theta}_n$ , that is,

$$\hat{\theta}_n^* = \frac{1}{B} \sum_{b=1}^B \hat{\theta}_n^{*b}. \quad (36.3)$$

**36.3 Bias Estimation**

The bootstrap estimate of the bias of  $\hat{\theta}_n$  is

$$\widehat{\text{Bias}}_{\text{boot}} = \hat{\theta}_n^* - \hat{\theta}_n. \quad (36.4)$$

*Remark.* The bias-corrected bootstrap estimate of  $\hat{\theta}_n$  is

$$\hat{\theta}_n^* = \hat{\theta}_n - \widehat{\text{Bias}}_{\text{boot}} = 2\hat{\theta}_n - \hat{\theta}_n^*. \quad (36.5)$$

**36.4 Confidence Interval Estimation**

**Percentile Confidence Interval** The  $1 - \alpha$  percentile confidence interval of  $\hat{\theta}_n$  is

$$\left[ \hat{\theta}_n^*(\alpha/2), \hat{\theta}_n^*(1 - \alpha/2) \right], \quad (36.6)$$

where  $\hat{\theta}_n^*(\alpha/2)$  is the  $\alpha/2$ -th percentile of the bootstrap distribution of  $\hat{\theta}_n$  and  $\hat{\theta}_n^*(1 - \alpha/2)$  is the  $(1 - \alpha/2)$ -th percentile of the bootstrap distribution of  $\hat{\theta}_n$ .

**Bootstrap-t Confidence Interval** The  $1 - \alpha$  bootstrap-t confidence interval of  $\hat{\theta}_n$  is

$$\left[ \hat{\theta}_n - t_{n-1}^*(1 - \alpha/2) \widehat{\text{se}}_{\text{boot}}, \hat{\theta}_n - t_{n-1}^*(\alpha/2) \widehat{\text{se}}_{\text{boot}} \right], \quad (36.7)$$

where  $t_{n-1}^*(1 - \alpha/2)$  is the  $(1 - \alpha/2)$ -th percentile of the bootstrap distribution of  $t_{n-1}$  and  $t_{n-1}^*(\alpha/2)$  is the  $\alpha/2$ -th percentile of the bootstrap distribution of  $t_{n-1}$ .

**Bias-corrected and accelerated (BCa) Confidence Interval** The  $1 - \alpha$  bias-corrected and accelerated (BCa) confidence interval of  $\hat{\theta}_n$  is

$$\left[ \hat{\theta}_n^* (\alpha_{\text{BCa}}/2), \hat{\theta}_n^* (1 - \alpha_{\text{BCa}}/2) \right], \quad (36.8)$$

where  $\alpha_{\text{BCa}}$  is the percentile of the bootstrap distribution of  $\hat{\theta}_n$  that satisfies

$$\alpha_{\text{BCa}} = \Phi \left( z_0 + \frac{z_0 + z_\alpha - \widehat{\text{Bias}}_{\text{boot}}}{1 - \hat{a}} \right), \quad (36.9)$$

where  $\Phi(\cdot)$  is the cumulative distribution function of the standard normal distribution,  $z_0$  is the  $1 - \alpha/2$  percentile of the standard normal distribution,  $z_\alpha$  is the  $\alpha/2$  percentile of the standard normal distribution, and  $\hat{a}$  is the percentile of the bootstrap distribution of  $\hat{\theta}_n$  that satisfies

$$\hat{a} = \frac{\sum_{b=1}^B (\hat{\theta}_n^{*b} - \hat{\theta}_n^*)^3}{6 \left[ \sum_{b=1}^B (\hat{\theta}_n^{*b} - \hat{\theta}_n^*)^2 \right]^{3/2}}. \quad (36.10)$$

## 36.5 Hypothesis Testing

**One-sample Hypothesis Testing**

**Two-sample Hypothesis Testing**

## 36.6 Jackknife

Let  $\mathbf{x}_{-i}$  be the sample with  $x_i$  removed,  $\mathbf{x}_{-i} = (x_1, x_2, \dots, x_{i-1}, x_{i+1}, \dots, x_n)^\top$ , and denote the corresponding value of the statistic of interest as

$$\hat{\theta}_{-i} = s(\mathbf{x}_{-i}) \quad (36.11)$$

**Bias of  $\hat{\theta}$**  For almost all reasonable and practical estimates, we have

$$\text{Bias}(\hat{\theta}_n) = \mathbb{E}(\hat{\theta}_n) - \theta \rightarrow 0, \quad n \rightarrow \infty$$

Then, it is reasonable to assume a power series of the type

$$\mathbb{E}(\hat{\theta}_n) = \theta + \frac{a_1}{n} + \frac{a_2}{n^2} + \frac{a_3}{n^3} + \dots$$



with some coefficients  $\{a_k\}$ . And we have

$$\mathbb{E}(\hat{\theta}_{-i}) = \mathbb{E}(\hat{\theta}_{n-1}) = \theta + \frac{a_1}{n-1} + \frac{a_2}{(n-1)^2} + \frac{a_3}{(n-1)^3} + \dots$$

For the sake of a smaller variance, we average all such estimates and let

$$\hat{\theta}_{(\cdot)} = \frac{1}{n} \sum_{i=1}^n \hat{\theta}_{-i}$$

thus,

$$\mathbb{E}(\hat{\theta}_{(\cdot)}) = \theta + \frac{a_1}{n-1} + \frac{a_2}{(n-1)^2} + \frac{a_3}{(n-1)^3} + \dots$$

Thus, we have

$$(n-1)\mathbb{E}[\hat{\theta}_{(\cdot)} - \hat{\theta}_n] = \frac{a_1}{n} + \frac{a_2}{n^2} + \frac{a_3}{n^3} + \dots = \text{Bias}(\hat{\theta})$$

Hence, we can get the jackknife estimate bias for  $\hat{\theta}$  be

$$\widehat{\text{Bias}}_{\text{jack}} = (n-1)(\hat{\theta}_{(\cdot)} - \hat{\theta}_n) \quad (36.12)$$

*Remark.* It is easy to combine the averaged Jackknife estimator  $\hat{\theta}_{-i}$  with the original  $\hat{\theta}$ , to kill the main term in the bias of  $\hat{\theta}$ , thus,

$$\begin{aligned} \mathbb{E}[n\hat{\theta}_n - (n-1)\hat{\theta}_{(\cdot)}] &= [n\theta - (n-1)\theta] + [a_1 - a_1] + \left[\frac{a_2}{n} - \frac{a_2}{n-1}\right] + \dots \\ &= \theta + \frac{a_2}{n(n-1)} + \dots = \theta + \frac{a_2}{n^2} + O(n^{-3}) \end{aligned}$$

This removes the bias in the special case that the bias is  $O(n^{-1})$  and removes it to  $O(n^{-2})$  in other cases.

**Variance of  $\hat{\theta}$**  The jackknife estimate of variance for  $\hat{\theta}$  is

$$\widehat{\text{Var}}_{\text{jack}} = \frac{n-1}{n} \sum_{i=1}^n (\hat{\theta}_{-i} - \hat{\theta}_{(\cdot)})^2, \quad \text{where } \hat{\theta}_{(\cdot)} = \frac{1}{n} \sum_{i=1}^n \hat{\theta}_{-i} \quad (36.13)$$

The jackknife method of estimation can fail if the statistic  $\hat{\theta}_{\text{jack}}$  is not smooth. Smoothness implies that relatively small changes to data values will cause only a small change in the statistic.

**Example** (Sample Mean).

**Example** (Sample Correlation Coefficient).

# **Part XII**

## **Regression Analysis**

## Chapter 37

# Linear Regression

# Chapter 38

## Generalized Linear Model

### 38.1 Introduction

Suppose the response  $Y$  has a distribution in the exponential family

$$f(y | \theta, \phi) = \exp \left[ \frac{y\theta - b(\theta)}{a(\phi)} + c(y, \phi) \right]$$

with link function  $g$ , such that,

$$E(Y | \mathbf{x}) = \mu = g^{-1}(\eta), \quad \eta = \mathbf{x}^\top \boldsymbol{\beta} \quad (38.1)$$

where the link function provides the relationship between the linear predictor and the mean of the distribution function. If  $\eta = \theta$ , the link function is called **canonical link function**.

*Remark.* A generalized linear model (GLM) is a flexible generalization of ordinary linear regression that allows for the response variable to have an error distribution other than the normal distribution.

**Table 38.1:** Commonly Used Link Functions

Distribution	Support of Distribution	Link Function $g(\mu)$	Mean Function $g^{-1}(\eta)$
Normal	real: $(-\infty, +\infty)$	$\mu$	$\eta$
Bernoulli	integer: $\{0, 1\}$	$\log \left( \frac{\mu}{1-\mu} \right)$	$\frac{1}{1+\exp(-\eta)}$
Poisson	integer: $0, 1, 2, \dots$	$\log(\mu)$	$\exp(\eta)$

**Maximum Likelihood** Suppose the log-likelihood function be

$$\ell(\boldsymbol{\beta} \mid \mathbf{x}, y) = \log [f(y \mid \theta, \phi)] = \log [f(y \mid g^{-1}(\eta), \phi)] \quad (38.2)$$

where  $g$  is the canonical link function and  $\eta = \mathbf{x}^\top \boldsymbol{\beta}$ .

Let

$$U(\boldsymbol{\beta}) = \frac{\partial \ell(\boldsymbol{\beta})}{\partial \boldsymbol{\beta}}, \quad A(\boldsymbol{\beta}) = -\frac{\partial^2 \ell(\boldsymbol{\beta})}{\partial \boldsymbol{\beta}' \partial \boldsymbol{\beta}}$$

be the score function and observed information matrix.

If  $\hat{\boldsymbol{\beta}}$  is the maximum likelihood estimate, then

$$U(\hat{\boldsymbol{\beta}}) = \mathbf{0}$$

By the mean value theorem,

$$\begin{aligned} U(\hat{\boldsymbol{\beta}}) - U(\boldsymbol{\beta}_0) &= \frac{\partial U(\boldsymbol{\beta}^*)}{\partial \boldsymbol{\beta}} (\hat{\boldsymbol{\beta}} - \boldsymbol{\beta}_0) \\ \Rightarrow -U(\boldsymbol{\beta}_0) &= -A(\boldsymbol{\beta}^*) (\hat{\boldsymbol{\beta}} - \boldsymbol{\beta}_0) \end{aligned}$$

where  $\boldsymbol{\beta}^* \in [\boldsymbol{\beta}_0, \hat{\boldsymbol{\beta}}]$ . Thus,

$$\hat{\boldsymbol{\beta}} = \boldsymbol{\beta}_0 + A^{-1}(\boldsymbol{\beta}^*) U(\boldsymbol{\beta}_0)$$

Suppose  $\hat{\boldsymbol{\beta}}_t, \hat{\boldsymbol{\beta}}_{t+1}$  be the maximum likelihood estimate at the  $t$ -th and  $(t+1)$ -th iterations, respectively. Two algorithms can be used to obtain the maximum likelihood estimate  $\hat{\boldsymbol{\beta}}$ .

1. Newton-Raphson Method:

$$\hat{\boldsymbol{\beta}}_{t+1} = \hat{\boldsymbol{\beta}}_t + A^{-1}(\hat{\boldsymbol{\beta}}_t) U(\hat{\boldsymbol{\beta}}_t) \Leftrightarrow A(\hat{\boldsymbol{\beta}}_t) \hat{\boldsymbol{\beta}}_{t+1} = A(\hat{\boldsymbol{\beta}}_t) \hat{\boldsymbol{\beta}}_t + U(\hat{\boldsymbol{\beta}}_t) \quad (38.3)$$

where

$$U(\boldsymbol{\beta}) = \frac{\partial \ell(\boldsymbol{\beta})}{\partial \boldsymbol{\beta}} \quad (38.4)$$

is the score function and

$$A(\boldsymbol{\beta}) = -\frac{\partial^2 \ell(\boldsymbol{\beta})}{\partial \boldsymbol{\beta}' \partial \boldsymbol{\beta}} \quad (38.5)$$

is the observed information matrix.

2. Fisher Scoring Method:

$$\hat{\boldsymbol{\beta}}_{t+1} = \hat{\boldsymbol{\beta}}_t + I^{-1}(\hat{\boldsymbol{\beta}}_t) U(\hat{\boldsymbol{\beta}}_t) \Leftrightarrow I(\hat{\boldsymbol{\beta}}_t) \hat{\boldsymbol{\beta}}_{t+1} = I(\hat{\boldsymbol{\beta}}_t) \hat{\boldsymbol{\beta}}_t + U(\hat{\boldsymbol{\beta}}_t) \quad (38.6)$$

where  $U(\boldsymbol{\beta})$  is the score function and

$$I(\boldsymbol{\beta}) = E[A(\boldsymbol{\beta})] = -E\left[\frac{\partial^2 \ell(\boldsymbol{\beta})}{\partial \boldsymbol{\beta}' \partial \boldsymbol{\beta}}\right] \quad (38.7)$$

is the Fisher information matrix.

## Bayesian Methods

## 38.2 Binary Data

Suppose

$$Y \sim b(m, \pi), \quad i = 1, 2, \dots, n \quad (38.8)$$

with link function

$$\eta = g(\pi) = \log\left(\frac{\pi}{1 - \pi}\right) = \mathbf{x}^\top \boldsymbol{\beta} \quad (38.9)$$

*Remark.*

The likelihood function is

$$f(\boldsymbol{\pi} \mid \mathbf{x}, \mathbf{y}) = \prod_{i=1}^n \binom{m_i}{y_i} \pi_i^{y_i} (1 - \pi_i)^{m_i - y_i} \quad (38.10)$$

and the log-likelihood function is

$$\begin{aligned} \ell(\boldsymbol{\beta}) &= \log[f(\boldsymbol{\pi} \mid \mathbf{x}, \mathbf{y})] = \sum_{i=1}^n \ell_i(\boldsymbol{\beta}) \\ &= \sum_{i=1}^n \left\{ \log \left[ \binom{m_i}{y_i} \right] + y_i \log(\pi_i) + (m_i - y_i) \log(1 - \pi_i) \right\} \\ &= \sum_{i=1}^n \left[ y_i \log\left(\frac{\pi_i}{1 - \pi_i}\right) + m_i \log(1 - \pi_i) \right] + \sum_{i=1}^n \log \left[ \binom{m_i}{y_i} \right] \end{aligned} \quad (38.11)$$

where

$$\pi_i = \frac{\exp(\mathbf{x}_i^\top \boldsymbol{\beta})}{1 + \exp(\mathbf{x}_i^\top \boldsymbol{\beta})} \quad (38.12)$$

Thus,

$$\begin{aligned} U_r(\boldsymbol{\beta}) &= \sum_{i=1}^n (y_i - m_i \pi_i) x_{ir} \\ I_{sr}(\boldsymbol{\beta}) &= \sum_{i=1}^n m_i \pi_i (1 - \pi_i) x_{is} x_{ir} \end{aligned}$$

## 38.3 Polytomous Data

**Definition 38.3.1 (Polytomous Data)**

A response is polytomous if the response of an individual or item in a study is **restricted to one of a fixed set of possible values**.

*Remark.* There are two types of scales, pure scales and compound scales<sup>1</sup>. For pure scales, there are several types:

1. **Nominal Scale:** a scale used for labeling variables into distinct classifications and does not involve a quantitative value or order.
2. **Ordinal Scale:** a variable measurement scale used to simply depict the order of variables and not the difference between each of the variables.
3. **Interval Scale:** a numerical scale where the order of the variables is known as well as the difference between these variables.

Let the category probabilities given  $\mathbf{x}_i$  be

$$\pi_j(\mathbf{x}_i) = P(Y = y_j \mid \mathbf{x} = \mathbf{x}_i) \quad (38.13)$$

and the cumulative probabilities given  $\mathbf{x}_i$  be

$$r_j(\mathbf{x}_i) = P\left(Y \leq \sum_{r \leq j} y_r \mid \mathbf{x} = \mathbf{x}_i\right) \quad (38.14)$$

where  $i = 1, 2, \dots, n$ ,  $j = 1, 2, \dots, k$ .

Here, multinomial distribution is in many ways the most natural distribution to consider in the context of a polytomous response variable. The density function of the multinomial distribution is,

$$P(Y_1 = y_1, \dots, Y_k = y_k) = \begin{cases} \frac{m!}{y_1! \dots y_k!} \pi_1^{y_1} \dots \pi_k^{y_k}, & \sum_{i=1}^k y_i = m \\ 0 & \text{otherwise} \end{cases}$$

for non-negative integers  $y_1, \dots, y_k$ .

As for the link function, we have

### Nominal Scale

$$\pi_j(\mathbf{x}_i) = \frac{\exp[\eta_j(\mathbf{x}_i)]}{\sum_{j=1}^k \exp[\eta_j(\mathbf{x}_i)]} \quad (38.15)$$

where  $\eta_j(\mathbf{x}_i) = \eta_j(\mathbf{x}_0) + (\mathbf{x}_i - \mathbf{x}_0)' \boldsymbol{\beta}_j + \alpha_i$ .

### Ordinal Scale

1. Logistic Scale:

$$\log \left[ \frac{r_j(\mathbf{x}_i)}{1 - r_j(\mathbf{x}_i)} \right] = \theta_j - \mathbf{x}_i^\top \boldsymbol{\beta} \quad (38.16)$$

2. Complementary Log-Log Scale:

$$\log \{-\log [1 - r_j(\mathbf{x}_i)]\} = \theta_j - \mathbf{x}_i^\top \boldsymbol{\beta} \quad (38.17)$$

---

<sup>1</sup>A bivariate response with one response ordinal and the other continuous is an example of compound scales.

**Interval Scale** Suppose the  $j$ -th category exits a cardinal number or score,  $s_j$ , where the difference between scores is a measure of distance between or separation of categories.

1.

$$\log \left[ \frac{r_j(\mathbf{x}_i)}{1 - r_j(\mathbf{x}_i)} \right] = \varsigma_0 + \varsigma_1 \left( \frac{s_j + s_{j+1}}{2} \right) - \mathbf{x}_i^\top \boldsymbol{\beta} - \mathbf{x}_i^\top \boldsymbol{\xi} (c_j - \bar{c}) \quad (38.18)$$

where  $c_j = \frac{s_j + s_{j+1}}{2}$  or  $c_j = \text{logit} \left( \frac{s_j + s_{j+1}}{2} \right)$ .

2.

$$\pi_j(\mathbf{x}_i) = \frac{\exp[\eta_j(\mathbf{x}_i)]}{\sum_{j=1}^k \exp[\eta_j(\mathbf{x}_i)]} \quad (38.19)$$

where  $\eta_j(\mathbf{x}_i) = \eta_j + (\mathbf{x}_i^\top \boldsymbol{\beta}) s_j + \alpha_i$ .

3.

$$\sum_{j=1}^k \pi_j(\mathbf{x}_i) s_j = \mathbf{x}_i^\top \boldsymbol{\beta} \quad (38.20)$$

## 38.4 Count Data

Departures from the idealized Poisson model are to be expected. Therefore, we avoid the assumption of Poisson variation and assume only that

$$\text{Var}(Y) = \sigma^2 E(Y) \quad (38.21)$$

with link function

$$\log(\mu) = \eta = \mathbf{x}^\top \boldsymbol{\beta} \quad (38.22)$$

where  $\mu = E(Y | \mathbf{x})$ .

For the response in the Poisson distribution, i.e.

$$P(Y = y | \mu) = \frac{e^{-\mu} \mu^y}{y!}$$

and the log-likelihood function is

$$\ell(\boldsymbol{\beta}) \propto \sum_{i=1}^n (y_i \log(\mu_i) - \mu_i) \quad (38.23)$$

where  $\mu_i = E(Y | \mathbf{x} = \mathbf{x}_i)$ .



# Chapter 39

## Quantile Regression

### Definition 39.0.1 (Smooth Quantile Loss)

Consider the following smooth quantile loss with a convolution operation:

$$\rho_{\tau,h}(\mu) := \int \rho_{\tau}(v) \kappa_h(\nu - \mu) \, d\nu, \quad (39.1)$$

where  $\rho_{\tau}$  is the quantile loss function and  $\kappa_h$  is a kernel function with bandwidth  $h$ .

# Chapter 40

## Survival Analysis

### 40.1 General Formulation

#### Definition 40.1.1 (Survival Function)

The survival function<sup>a</sup> is defined to be

$$S(t) = P(T > t) = \int_t^\infty f(u) \, du = 1 - F(t). \quad (40.1)$$

where  $t$  is some specified time,  $T$  is a random variable denoting the time of death.

---

<sup>a</sup>The survival function is the probability that the time of death is later than some specified time  $t$ .

#### Definition 40.1.2 (Lifetime Distribution Function)

The lifetime distribution function is defined to be

$$F(t) = P(T \leq t) \quad (40.2)$$

If  $F$  is differentiable then the derivative, which is the density function of the lifetime distribution<sup>a</sup>, is defined to be

$$f(t) = F'(t) = \frac{d}{dt} F(t) \quad (40.3)$$

---

<sup>a</sup>The function  $f$  is sometimes called the event density; it is the rate of death or failure events per unit time.

**Definition 40.1.3 (Hazard Function)**

The Hazard function<sup>a</sup> is defined to be

$$\lambda(t) = \lim_{\varepsilon \rightarrow 0^+} \left[ \frac{P(t \leq T < t + \varepsilon \mid T \geq t)}{\varepsilon} \right] = \frac{f(t)}{S(t)} \quad (40.4)$$

<sup>a</sup>The Hazard function is the event rate at time  $t$  conditional on survival until time  $t$  or later (that is,  $T \geq t$ ).

**Property.** The relationship among  $\lambda(t)$ ,  $f(t)$ ,  $S(t)$ ,

1.

$$\lambda(t) = -\frac{d \log[S(t)]}{dt} \quad (40.5)$$

2.

$$S(t) = \exp \left[ -\int_0^t \lambda(x) dx \right] \quad (40.6)$$

3.

$$f(t) = \lambda(t) \exp \left[ -\int_0^t \lambda(x) dx \right] \quad (40.7)$$

*Proof.*

□

**Example** (Constant Hazards). Suppose

$$\lambda(t) = \lambda \quad (40.8)$$

then

$$\begin{aligned} S(t) &= \exp \left[ -\int_0^t \lambda(x) dx \right] = \exp \left[ -\int_0^t \lambda dx \right] = \exp(-\lambda t) \\ f(t) &= \lambda(t) \exp \left[ -\int_0^t \lambda(x) dx \right] = \lambda \exp \left[ -\int_0^t \lambda dx \right] = \lambda \exp(-\lambda t) \end{aligned}$$

which is the exponential distribution.

**Example** (Bathtub Hazards).

$$\lambda(t) = \alpha t + \frac{\beta}{1 + \gamma t} \quad (40.9)$$

## 40.2 Estimation of Survival Function

**Parametric Approach** Suppose  $t_1, t_2, \dots, t_n$  are failure times corresponding to censor indicators  $\delta_1, \delta_2, \dots, \delta_n$ . The likelihood function is

$$\begin{aligned} f(\boldsymbol{\theta} \mid \mathbf{t}, \boldsymbol{\delta}) &= \prod_{i=1}^n [f(t_i)]^{\delta_i} [S(t_i)]^{1-\delta_i} \\ &= \prod_{i=1}^n \left( \frac{f(t_i)}{S(t_i)} \right)^{\delta_i} S(t_i) \\ &= \prod_{i=1}^n [\lambda(t_i)]^{\delta_i} S(t_i) \end{aligned} \quad (40.10)$$

where  $\lambda(t), S(t)$  depends on some parameter  $\theta$ .

**Example.** Suppose  $\mathbf{T}$  have exponential density, that,

$$f(t) = \lambda e^{-\lambda t}, \quad S(t) = e^{-\lambda t}$$

Thus,

$$\begin{aligned} \ell(\lambda) &= \log[\ell(\theta)] = \sum_{i=1}^n [\delta_i \log(\lambda) - \lambda t_i] \\ &= \left( \sum_{i=1}^n \delta_i \right) \log(\lambda) - \lambda \left( \sum_{i=1}^n t_i \right) \end{aligned}$$

Hence,

$$\frac{\partial \ell(\lambda)}{\partial \lambda} = \frac{\sum_{i=1}^n \delta_i}{\lambda} - \sum_{i=1}^n t_i = 0 \Rightarrow \hat{\lambda} = \frac{\sum_{i=1}^n \delta_i}{\sum_{i=1}^n t_i}$$

**Nonparametric Approach** Then, for  $t_{(k)} \leq t < t_{(k+1)}$ ,

$$\begin{aligned} \hat{S}(t) &= \prod_{j=1}^k \left( \frac{n_j - d_j}{n_j} \right) \\ &= \left( 1 - \frac{d_1}{n_1} \right) \left( 1 - \frac{d_2}{n_2} \right) \cdots \left( 1 - \frac{d_k}{n_k} \right) \\ &\approx [1 - \hat{\lambda}(t_1)] [1 - \hat{\lambda}(t_2)] \cdots [1 - \hat{\lambda}(t_k)] \end{aligned} \quad (40.11)$$

where  $\hat{S}(t)$  is referred to as Kaplan-Meier estimate.

### 40.3 Proportional Hazards Model

Let  $t_1, t_2, \dots, t_n$  be the failure times associated with censor indicator  $\delta_1, \delta_2, \dots, \delta_n$  and the covariate vectors  $\mathbf{x}_i$ .

Further, let  $t_{(1)} \leq t_{(2)} \leq \dots \leq t_{(m)}$  be the ordered uncensored failure times corresponding to  $\delta_{(j)} = 1, j = 1, 2, \dots, m$ , and  $x_{(1)}, x_{(2)}, \dots, x_{(m)}$  are the associated covariate vectors. Note  $(j)$  represents the label for the individual who dies at  $t_{(j)}$ .

The proportional hazards model specifying the hazard at time  $t$  for an individual whose covariate vector is  $\mathbf{x}$  is given by

$$\lambda(t) = \lambda_0(t)e^{\mathbf{x}^\top \boldsymbol{\beta}} \quad (40.12)$$

where  $\lambda_0(t)$  is referred to as the baseline hazard function.

The exact likelihood function is

$$\ell[\boldsymbol{\beta}, \lambda_0(t)] = \prod_{i=1}^n [\lambda_i(t_i)]^{\delta_i} S(t_i) \quad (40.13)$$

depends on both the nonparametric function  $\lambda_0(t)$  and the parameter  $\boldsymbol{\beta}$ . Thus, it might be difficult to estimate  $\lambda_0(t)$  and  $\boldsymbol{\beta}$  simultaneously.

The partial likelihood function is

$$\ell_p(\boldsymbol{\beta}) = \prod_{j=1}^m \frac{e^{\mathbf{x}'_{(j)} \boldsymbol{\beta}}}{\sum_{l \in R(t_{(j)})} e^{\mathbf{x}'_l \boldsymbol{\beta}}} = \prod_{i=1}^n \left[ \frac{e^{\mathbf{x}'_i \boldsymbol{\beta}}}{\sum_{l \in R(t_i)} e^{\mathbf{x}'_l \boldsymbol{\beta}}} \right]^{\delta_i} \quad (40.14)$$

where  $R(t)$  is the set of individuals who are alive and uncensored at a time just before  $t_i$ , which is called the risk set.

# Chapter 41

## Nonparametric Regression

### 41.1 Uniform Stability of Regularized Kernel Model

#### 41.1.1 Introduction

Suppose  $\mathcal{X}$  be the input space,  $\mathcal{Y}$  be the output space,  $\mathcal{D}$  be some (almost) completely unknown probability distribution on  $\mathcal{X} \times \mathcal{Y}$ . Given the  $n$  i.i.d observed data, which sampled from an unknown distribution  $\mathcal{D}$ , that,

$$S := \{(\mathbf{x}_1, y_1), \dots, (\mathbf{x}_n, y_n)\} \subseteq \mathcal{D} \quad (41.1)$$

and the goal of us is to estimate the functional relationship between  $\mathcal{X}$  and  $\mathcal{Y}$ .

To formalize the problem, we now aim at finding a predictor function  $f^*$  among the function space  $\mathcal{F} := \{f : \mathcal{X} \rightarrow \mathcal{Y}\}$  based on the observed data  $S$ , which minimizes the true risk

$$R[f] := \mathbb{E}_{\mathcal{D}} [L(y, f(\mathbf{x}))] \quad (41.2)$$

where  $L : \mathcal{Y} \times \mathcal{Y} \rightarrow \mathbb{R}$  is an arbitrary convex loss function, typically assumed that the smaller  $L(y, f(\mathbf{x}))$  is, the better the approximation of  $y$  is. Thus, we are trying to find a predictor  $f^*$  with risk close to the optimal risk

$$R^* := \inf \{R[f] \mid f : \mathcal{X} \rightarrow \mathcal{Y}\} \quad (41.3)$$

Finding the predictor function  $f^*$  which minimizing the empirical risk

$$R_{\text{emp}}(f) = \frac{1}{n} \sum_{i=1}^n L(y_i, f(\mathbf{x}_i)) \quad (41.4)$$

is a natural thing for us to be trying to do. However, as it is known to all, just minimizing the empirical risk is suicidal, which almost certainly leads to overfitting. Minimizing  $R_{\text{emp}}$  only makes sense if we simultaneously somehow restrict ourselves

to the  $\mathcal{F}$ , which are of just the right level of complexity. One way to do this is by explicitly restricting the function space  $\mathcal{F}$  to a "simple" space, as in structural risk minimization, which is to introduce a penalty functional  $\Omega[f]$  that somehow measures the complexity of each function  $f \in \mathcal{F}$ , and to minimize the regularized risk

$$R_{\text{reg}}[f] = R_{\text{emp}}[f] + \Omega[f] \quad (41.5)$$

In this report, we restrict the predictor function  $f \in \mathcal{F}$  among the reproducing kernel Hilbert space  $\mathcal{H}$ , and the regularized risk has the form

$$R_{\text{reg}}[f] = \frac{1}{n} \sum_{i=1}^n L(y_i, f(\mathbf{x}_i)) + \frac{\lambda}{2} \|f\|_{\mathcal{H}}^2 \quad (41.6)$$

thus, we can estimate  $f^*$  by solving the following optimization problem

$$\hat{f} = \arg \min_{f \in \mathcal{H}} \frac{1}{n} \sum_{i=1}^n L(y_i, f(\mathbf{x}_i)) + \frac{\lambda}{2} \|f\|_{\mathcal{H}}^2 \quad (41.7)$$

where  $\lambda > 0$  is a regularization parameter to reduce the danger of overfitting. Since  $L(y, f(\mathbf{x}))$  is convex in  $f$ , the minimizer  $\hat{f}$  is uniquely determined and a simple gradient descent algorithm can be used to find  $\hat{f}$ . So the main focus of this report is to answer a remaining question we want to know, whether the risk  $R[\hat{f}]$  is close to the optimal risk  $R^*$ , which will influence the stability of our algorithm.

### 41.1.2 Some Notations and Concepts

Before getting into the formal discussion, we will introduce some notations and concepts.

- $S^{\setminus i} := S \setminus \{(\mathbf{x}_i, y_i)\}$  be the sample where the  $i$ -th observation is removed.
- $S^i := S^{\setminus i} \cup \{(\mathbf{x}, y)\}$  be the sample where the  $i$ -th observation is replaced by  $(\mathbf{x}, y)$ .

and let  $\hat{f}_{\setminus i}$  be the estimated result based on sample  $S^{\setminus i}$ ,  $\hat{f}_i$  based on sample  $S^i$  and  $\hat{f}$  based on sample  $S$ .

To quantify the stability of our algorithm, we will introduce one important concept — **Uniform Stability**.

#### Definition 41.1.1 (Uniform Stability)

The algorithm is uniformly  $\beta$ -stable with respect to the loss function  $L(y, f(\mathbf{x}))$ , if for all samples  $S := \{\mathbf{x}_i, y_i\}_{i=1}^n \subset \mathcal{D}$  and  $i \in [n]$ ,

$$\sup_{(x,y) \in \mathcal{D}} |L(y, \hat{f}(\mathbf{x})) - L(y, \hat{f}_{\setminus i}(\mathbf{x}))| \leq \beta \quad (41.8)$$

i.e. the algorithm is "stable" to remove a single sample at all points.

### 41.1.3 Uniform Stability of Regularized Kernel Model

Firstly, we will provide an auxiliary lemma.

#### Lemma 41.1.1 (Convex Functions and Derivatives)

For any differentiable convex function  $f : \mathbb{R} \rightarrow \mathbb{R}$  and any  $a, b \in \mathbb{R}$ , we have

$$[f'(a) - f'(b)](a - b) \geq 0 \quad (41.9)$$

*Proof.* Due to the convexity of  $f$  we know that  $f(a) + (b - a)f'(a) \leq f(b)$  and, likewise,  $f(b) + (a - b)f'(b) \leq f(a)$ . Summing up both inequalities and subtracting the terms in  $f(a)$  and  $f(b)$  proves (41.9).  $\square$

Then, we will show that the algorithm we studied in this paper satisfied the definition 41.1.1, and the corresponding value of  $\beta$  can be calculated.

#### Theorem 41.1.1 (Algorithmic Stability of Risk Minimizers (Bousquet and Elisseeff,

The algorithm that minimizes the regularized empirical risk in (41.6) has stability

$$\beta = \frac{2C^2\kappa^2}{n\lambda} \quad (41.10)$$

where  $\kappa$  is bound on  $\|k(x, \cdot)\| = \sqrt{k(x, x)}$ ,  $\|\cdot\|$  is the RKHS norm induced by  $k$ , and  $C$  is a bound on the Lipschitz constant of the loss function  $L(y, f(\mathbf{x}))$ , which can be viewed as a function of  $f$ .

*Remark.* We can see that the stability of the algorithm depends on the regularization constant via  $\frac{1}{\lambda n}$ , hence we may be able to afford to choose weaker regularization if the sample size  $n$  increases.

*Proof.* To distinguish between different training sets, we use  $R_{\text{reg}}[f, S]$  and  $R_{\text{reg}}[f, S^{\setminus i}]$  (and likewise  $R_{\text{emp}}[f, S]$ ) during the remainder of the proof.

Since  $\hat{f}$  minimizes  $R_{\text{reg}}[f, S]$ , that is, the **functional derivative** Stéphane Canu, 2014 of  $R_{\text{reg}}[f, S]$  at  $\hat{f}$  vanishes, and so does  $R_{\text{reg}}[f, S^{\setminus i}]$  at  $\hat{f}_{\setminus i}$ ,

$$\begin{aligned} \partial_f R_{\text{reg}}[\hat{f}, S] &= \partial_f R_{\text{emp}}[\hat{f}, S] + \lambda \hat{f} = 0 \\ \partial_f R_{\text{reg}}[\hat{f}_{\setminus i}, S^{\setminus i}] &= \partial_f R_{\text{emp}}[\hat{f}_{\setminus i}, S^{\setminus i}] + \lambda \hat{f}_{\setminus i} = 0 \end{aligned} \quad (41.11)$$

Next, we construct an auxiliary risk function  $\tilde{R}[f]$  by

$$\tilde{R}[f] := \langle \partial_f R_{\text{emp}}[\hat{f}, S] - \partial_f R_{\text{emp}}[\hat{f}_{\setminus i}, S^{\setminus i}], f - \hat{f}_{\setminus i} \rangle + \frac{\lambda}{2} \|f - \hat{f}_{\setminus i}\|_{\mathcal{H}}^2 \quad (41.12)$$



$\tilde{R}[f]$  is a convex function in  $f$  (the first term is linear, the second quadratic).

Additionally, by construction, we have

$$\tilde{R}[\hat{f}_{\setminus i}] = 0 \quad (41.13)$$

Furthermore, taking the functional derivative of  $\tilde{R}[f]$ , that,

$$\partial_f \tilde{R}[f] = \partial_f R_{\text{emp}} [\hat{f}, S] - \partial_f R_{\text{emp}} [\hat{f}_{\setminus i}, S^{\setminus i}] + \lambda (f - \hat{f}_{\setminus i}) = \partial_f R_{\text{emp}} [\hat{f}, S] + \lambda f \quad (41.14)$$

the functional derivative of  $\tilde{R}[f]$  (41.14) vanishes at  $f = \hat{f}$  due to (41.11), thus the minimum of  $\tilde{R}[f]$  is obtained for  $f = \hat{f}$ . Therefore, combined with  $\tilde{R}[\hat{f}_{\setminus i}] = 0$ , we can conclude that  $\tilde{R}[\hat{f}] \leq 0$ .

To obtain bounds on  $\|\hat{f} - \hat{f}_{\setminus i}\|$ , we have to get rid of some of the first terms in  $\tilde{R}[f]$ , since

$$\begin{aligned} & n \langle \partial_f R_{\text{emp}} [\hat{f}, S] - \partial_f R_{\text{emp}} [\hat{f}_{\setminus i}, S^{\setminus i}], \hat{f} - \hat{f}_{\setminus i} \rangle \\ &= \sum_{j \neq i} [L'(y_j, \hat{f}(\mathbf{x}_j)) - L'(y_j, \hat{f}_{\setminus i}(\mathbf{x}_j))] [\hat{f}(\mathbf{x}_j) - \hat{f}_{\setminus i}(\mathbf{x}_j)] \\ & \quad + L'(y_i, \hat{f}(\mathbf{x}_i)) [\hat{f}(\mathbf{x}_i) - \hat{f}_{\setminus i}(\mathbf{x}_i)] \\ & \geq L'(y_i, \hat{f}(\mathbf{x}_i)) [\hat{f}(\mathbf{x}_i) - \hat{f}_{\setminus i}(\mathbf{x}_i)] \end{aligned} \quad (41.15)$$

The first equation is since the functional derivative  $\partial_f(f) = k(\mathbf{x}, \cdot)$  and then collecting the common terms between  $R_{\text{emp}} [\hat{f}, S]$  and  $R_{\text{emp}} [\hat{f}_{\setminus i}, S^{\setminus i}]$ . And, as for the last inequation, we use lemma 41.1.1 applied to the loss function  $L(y, f(\mathbf{x}))$  which is a convex function of  $f(\mathbf{x})$ .

Combine the above result with the fact  $\tilde{R}[\hat{f}] \leq 0$ , we have

$$\langle \partial_f R_{\text{emp}} [\hat{f}, S] - \partial_f R_{\text{emp}} [\hat{f}_{\setminus i}, S^{\setminus i}], \hat{f} - \hat{f}_{\setminus i} \rangle + \frac{\lambda}{2} \|\hat{f} - \hat{f}_{\setminus i}\|_{\mathcal{H}}^2 \leq 0 \quad (41.16)$$

thus,

$$L'(y_i, \hat{f}(\mathbf{x}_i)) [\hat{f}(\mathbf{x}_i) - \hat{f}_{\setminus i}(\mathbf{x}_i)] + \frac{n\lambda}{2} \|\hat{f} - \hat{f}_{\setminus i}\|_{\mathcal{H}}^2 \leq 0 \quad (41.17)$$

and by the convexity of loss function  $L(y, f(\mathbf{x}))$ ,

$$\begin{aligned} \frac{n\lambda}{2} \|\hat{f} - \hat{f}_{\setminus i}\|_{\mathcal{H}}^2 & \leq L'(y_i, \hat{f}(\mathbf{x}_i)) [\hat{f}_{\setminus i}(\mathbf{x}_i) - \hat{f}(\mathbf{x}_i)] \\ & \leq L(y_i, \hat{f}(\mathbf{x}_i)) - L(y_i, \hat{f}_{\setminus i}(\mathbf{x}_i)) \\ & \leq |L(y_i, \hat{f}(\mathbf{x}_i)) - L(y_i, \hat{f}_{\setminus i}(\mathbf{x}_i))| \end{aligned} \quad (41.18)$$

By the Cauchy-Schwarz inequality, we can see that, for any  $f, f' \in \mathcal{H}$  and any  $\mathbf{x} \in \mathcal{X}$ ,

$$|f(\mathbf{x}) - f'(\mathbf{x})| = |\langle f - f', k(\mathbf{x}, \cdot) \rangle| \leq \|f - f'\|_{\mathcal{H}} \|k(\mathbf{x}, \cdot)\|_{\mathcal{H}} \leq \kappa \|f - f'\|_{\mathcal{H}} \quad (41.19)$$

and since  $L(y, f(\mathbf{x}))$  is Lipschitz continuous at  $\mathbf{x}_i$ , we have

$$|L(y, \hat{f}(\mathbf{x}_i)) - L(y, \hat{f}_i(\mathbf{x}_i))| \leq C |\hat{f}(\mathbf{x}_i) - \hat{f}_i(\mathbf{x}_i)| \leq C\kappa \|\hat{f} - \hat{f}_i\|_{\mathcal{H}} \quad (41.20)$$

Combine equation (41.18) and (41.20), we get

$$\frac{n\lambda}{2} \|\hat{f} - \hat{f}_i\|_{\mathcal{H}}^2 \leq C\kappa \|\hat{f} - \hat{f}_i\|_{\mathcal{H}} \quad (41.21)$$

thus,

$$\|\hat{f} - \hat{f}_i\|_{\mathcal{H}} \leq \frac{2C\kappa}{n\lambda} \quad (41.22)$$

Therefore, by the equation (41.20) for every  $\mathbf{x}$ , we have

$$|L(y, \hat{f}(\mathbf{x})) - L(y, \hat{f}_i(\mathbf{x}))| \leq C\kappa \|\hat{f} - \hat{f}_i\|_{\mathcal{H}} \leq \frac{2C^2\kappa^2}{n\lambda} \quad (41.23)$$

□

Within the uniform stability of our algorithm, we will also prove that the  $\beta$ -stable algorithm exhibits uniform convergence of the empirical risk  $R_{\text{emp}}[f]$  to the true risk  $R[f]$ .

#### Theorem 41.1.2 (McDiarmid's Bound (McDiarmid, 1989))

Suppose  $\xi_1, \dots, \xi_n$  be i.i.d real value random variables and assume that there exists a function  $g : \mathbb{R}^n \rightarrow \mathbb{R}$  with the property that for all  $i \in [n]$  and  $c_i > 0$ ,

$$\sup_{\xi_1, \dots, \xi_n, \xi'_i \in \mathbb{R}} |g(\xi_1, \dots, \xi_n) - g(\xi_1, \dots, \xi_{i-1}, \xi'_i, \xi_{i+1}, \dots, \xi_n)| \leq c_i \quad (41.24)$$

where  $\xi'_i$  is drawn from the same distribution as  $\xi_i$ . Then

$$\mathbb{P}\{|g(\xi_1, \dots, \xi_n) - \mathbb{E}[g(\xi_1, \dots, \xi_n)]| > \varepsilon\} \leq 2 \exp\left(-\frac{2\varepsilon^2}{\sum_{i=1}^n c_i^2}\right) \quad (41.25)$$

**Theorem 41.1.3 (Bousquet and Elisseeff (Bousquet and Elisseeff, 2001; Ofer Dekel**

Assume that we have a  $\beta$ -stable algorithm with the additional requirement that the loss function  $L(y, f(\mathbf{x})) \leq M$  for all  $(\mathbf{x}, y) \in \mathcal{D}$  and for all samples  $S := \{\mathbf{x}_i, y_i\}_{i=1}^n \subset \mathcal{D}$ . Then, for any  $n \geq 1$

$$\mathbb{P} \left\{ |R_{\text{emp}}[\hat{f}, S] - R[\hat{f}]| > \varepsilon + 2\beta \right\} \leq 2 \exp \left( -\frac{n\varepsilon^2}{2(n\beta + M)^2} \right) \quad (41.26)$$

*Proof.* Within the i.i.d assumption, we have

$$\mathbb{E}_{S \sim \mathcal{D}} [R_{\text{emp}}[\hat{f}]] = \frac{1}{n} \sum_{i=1}^n \mathbb{E}_{S \sim \mathcal{D}} [L(y_i, \hat{f}(\mathbf{x}_i))] = \mathbb{E}_{S \sim \mathcal{D}} [L(y_i, \hat{f}(\mathbf{x}_i))] \quad (41.27)$$

If we replace  $(\mathbf{x}_i, y_i)$  by  $(\mathbf{x}, y)$ , we can get

$$\mathbb{E}_{S \sim \mathcal{D}} [R_{\text{emp}}[\hat{f}]] = \mathbb{E}_{S, (\mathbf{x}, y) \sim \mathcal{D}} [L(y, \hat{f}(\mathbf{x}))] \quad (41.28)$$

and with the observation that

$$\mathbb{E}_{\mathcal{D}} [R[\hat{f}]] = \mathbb{E}_{S, (\mathbf{x}, y) \sim \mathcal{D}} [L(y, \hat{f}(\mathbf{x}))] \quad (41.29)$$

In order to bound on the expected difference between  $R_{\text{emp}}[\hat{f}, S]$  and  $R[\hat{f}]$ , which leads to

$$\begin{aligned} \mathbb{E}_{\mathcal{D}} [R_{\text{emp}}[\hat{f}, S] - R[\hat{f}]] &= \mathbb{E}_{S, (\mathbf{x}, y) \sim \mathcal{D}} [L(y, \hat{f}(\mathbf{x}))] - \mathbb{E}_{S, (\mathbf{x}, y) \sim \mathcal{D}} [L(y, \hat{f}(\mathbf{x}))] \\ &= \mathbb{E}_{S, (\mathbf{x}, y) \sim \mathcal{D}} [L(y, \hat{f}(\mathbf{x})) - L(y, \hat{f}(\mathbf{x}))] \leq 2\beta \end{aligned} \quad (41.30)$$

By the triangle inequality, we have

$$|R[\hat{f}] - R[\hat{f}_i]| \leq |R[\hat{f}] - R[\hat{f}_{\setminus i}]| + |R[\hat{f}_{\setminus i}] - R[\hat{f}_i]| \leq 2\beta \quad (41.31)$$

Also, we have

$$\begin{aligned} |R_{\text{emp}}[\hat{f}, S] - R_{\text{emp}}(\hat{f}_i, S^i)| &\leq \frac{1}{n} \sum_{j \neq i} |L(y_j, \hat{f}(\mathbf{x}_j)) - L(y_j, \hat{f}_i(\mathbf{x}_j))| \\ &\quad + \frac{1}{n} |L(y_i, \hat{f}(\mathbf{x}_i)) - L(y_i, \hat{f}_i(\mathbf{x}_i))| \\ &\leq \frac{n-1}{n} 2\beta + \frac{2M}{n} \leq 2\beta + \frac{2M}{n} \end{aligned} \quad (41.32)$$

and,

$$\begin{aligned} \left| \left[ R_{\text{emp}}[\hat{f}, S] - R[\hat{f}] \right] - \left[ R_{\text{emp}}(\hat{f}_i, S^i) - R[\hat{f}_i] \right] \right| &\leq \left| R_{\text{emp}}[\hat{f}, S] - R_{\text{emp}}(\hat{f}_i, S^i) \right| \\ &\quad + \left| R[\hat{f}] - R[\hat{f}_i] \right| \\ &\leq 4\beta + \frac{2M}{n} \end{aligned} \quad (41.33)$$

Thus, by the Theorem 41.1.2, we have  $c_i = 4\beta + \frac{2M}{n}$ , that,

$$\begin{aligned} \mathbb{P} \left\{ |R_{\text{emp}}[\hat{f}, S] - R[\hat{f}] - 2\beta| > \varepsilon \right\} &\leq \mathbb{P} \left\{ |R_{\text{emp}}[\hat{f}, S] - R[\hat{f}]| > \varepsilon + 2\beta \right\} \\ &\leq 2 \exp \left( -\frac{n\varepsilon^2}{2(2n\beta + M)^2} \right) \end{aligned} \quad (41.34)$$

□

Within the above two theorems, we can directly get the following practical consequence.

#### Corollary 41.1.1 (Uniform Convergence Bounds for RKHS)

The algorithm minimizing the regularized risk  $R_{\text{reg}}[f]$ , as in (41.6), and with the assumptions of Theorem 41.1.1 and 41.1.3, we obtain

$$\mathbb{P} \left\{ |R_{\text{emp}}[\hat{f}] - R[\hat{f}]| > \varepsilon + 2\beta \right\} \leq 2 \exp \left( -\frac{n}{2} \left( \frac{\varepsilon}{M} \right)^2 \left( 1 + \frac{4}{\lambda M} (C\kappa)^2 \right)^{-2} \right) \quad (41.35)$$

where

$$\beta = \frac{2C^2\kappa^2}{n\lambda}$$

*Remark.* For practical considerations, (41.35) may be very useful, even if the rates are not optimal, since the bound is predictive even for small sample sizes and moderate regularization strength. Still, we expect that the constants

## Comments

The idea of the discussion content was inspired by Hofmann, Schölkopf, and Smola, 2008, Section 3 review, and the report is organized following Schölkopf, 2002, Chapter 12 structure, and so are the main proof ideas.

## References

- Bousquet, Olivier and André Elisseeff (2001). “Algorithmic Stability and Generalization Performance”. In: *Advances in Neural Information Processing Systems*. Vol. 13. MIT Press. URL: <https://papers.nips.cc/paper/2000/hash/49ad23d1ec9fa4bd8d77d02681df5cfa-Abstract.html> (visited on 01/08/2022).
- (2002). “Stability and Generalization”. In: *Journal of Machine Learning Research* 2 (Mar), pp. 499–526. ISSN: ISSN 1533-7928. URL: <https://www.jmlr.org/papers/v2/bousquet02a.html> (visited on 01/08/2022).
- Hofmann, Thomas, Bernhard Schölkopf, and Alexander J. Smola (June 1, 2008). “Kernel Methods in Machine Learning”. In: *The Annals of Statistics* 36.3. ISSN: 0090-5364. DOI: [10.1214/009053607000000677](https://doi.org/10.1214/009053607000000677).
- McDiarmid, Colin (Aug. 3, 1989). “On the Method of Bounded Differences”. In: *Surveys in Combinatorics, 1989*. Ed. by J. Siemons. 1st ed. London Mathematical Society Lecture Note Series. Cambridge: Cambridge University Press, pp. 148–188. ISBN: 978-0-521-37823-9 978-1-107-35994-9. DOI: [10.1017/CB09781107359949.008](https://doi.org/10.1017/CB09781107359949.008).
- Ofer Dekel and Thach Nguyen (Aug. 3, 2011). “Algorithmic Stability”.
- Schölkopf, Bernhard (2002). *Learning with Kernels: Support Vector Machines, Regularization, Optimization, and Beyond*. Adaptive Computation and Machine Learning. Cambridge, Mass: MIT Press. ISBN: 978-0-262-19475-4 978-0-262-25693-3 978-0-585-47759-6.
- Stéphane Canu (Mar. 4, 2014). “Lecture 4: Kernels and Associated Functions” (Sao Paulo). URL: [https://cel.archives-ouvertes.fr/cel-01003007/file/Lecture4\\_Kernels\\_Functions\\_RKHS.pdf](https://cel.archives-ouvertes.fr/cel-01003007/file/Lecture4_Kernels_Functions_RKHS.pdf).

# Chapter 42

## High Dimensional Regression Analysis

### 42.1 Lasso for Linear Regression

If  $p > n$ , then the least squares estimator is not unique. In this case, we can use the lasso to select a unique estimator. The lasso estimator is defined as

$$\hat{\beta}_{\text{lasso}} := \arg \min_{\beta \in \mathbb{R}^p} \left\{ \|\mathbf{y} - \mathbf{X}\beta\|_2^2 / n + \lambda \|\beta\|_1 \right\}, \quad (42.1)$$

where  $\lambda > 0$  is a tuning parameter. In addition, the optimization problem (42.1) can be rewritten as

$$\hat{\beta}_{\text{lasso}} := \arg \min_{\beta \in \mathbb{R}^p: \|\beta\|_1 \leq R} \|\mathbf{y} - \mathbf{X}\beta\|_2^2 / n,$$

with a one-to-one correspondence between  $\lambda$  and  $R$ .

#### 42.1.1 Numerical Algorithms

The lasso estimator can be computed by the following numerical algorithms.

##### Cyclic Coordinate Descent

Cyclic coordinate descent is an iterative algorithm. At each iteration, we update one coordinate of  $\beta$  while fixing all other coordinates. The cyclic coordinate descent algorithm for the lasso is given in Algorithm 6.

**Algorithm 6:** Cyclic Coordinate Descent for the Lasso Estimator

---

**Input:** Data  $\{(\mathbf{X}_i, y_i)\}_{i=1}^n$ , tuning parameter  $\lambda > 0$ , initial value  $\boldsymbol{\beta}^{(0)} \in \mathbb{R}^p$ , and tolerance  $\epsilon > 0$ .

- 1  $\mathbf{r} \leftarrow \mathbf{y} - \mathbf{X}\boldsymbol{\beta}^{(0)}$ ;
- 2 **while**  $\|\boldsymbol{\beta}^{(t)} - \boldsymbol{\beta}^{(t-1)}\|_\infty > \epsilon$  **do**
- 3     **for**  $j = 1, \dots, p$  **do**
- 4          $\beta_j^{(t+1)} \leftarrow \frac{S_\lambda(\mathbf{X}_j^\top \mathbf{r}/n)}{\mathbf{X}_j^\top \mathbf{X}_j/n}$ ;
- 5          $\mathbf{r} \leftarrow \mathbf{r} + (\beta_j^{(t+1)} - \beta_j^{(t)}) \mathbf{X}_j$ ;

**Output:** Estimate  $\hat{\boldsymbol{\beta}}_{\text{lasso}}$ .

---

**42.1.2 Selection of the Tuning Parameter****K-Fold Cross-Validation****42.2 Theory for the Lasso**

Let  $\boldsymbol{\beta}^0$  be the true parameter vector, and let  $\hat{\boldsymbol{\beta}}$  be the lasso estimator. We assume that the design matrix  $\mathbf{X}$  is orthogonal, i.e.,  $\mathbf{X}^\top \mathbf{X} = \mathbf{I}_p$ , and that the noise vector  $\boldsymbol{\varepsilon}$  is a random vector with mean zero and covariance matrix  $\sigma^2 \mathbf{I}_n$ . We further denote  $S_0 := \{j : \beta_j^0 \neq 0\}$  as the active set, and  $s_0 := |S_0|$  as the cardinality of the active set. We also denote  $\phi_0 := \min_{j \in S_0} |\beta_j^0|$  as the minimum absolute value of the nonzero coefficients.

In this section, we will discuss two important properties of the lasso estimator:

1. Prediction Error:

$$\|\mathbf{X}(\hat{\boldsymbol{\beta}} - \boldsymbol{\beta}^0)\|_2^2 / n \lesssim \frac{\log p}{n} s_0, \quad \text{with } s_0 = o(n / \log p).$$

2.  $l_1$ -Error:

$$\|\hat{\boldsymbol{\beta}} - \boldsymbol{\beta}^0\|_1 \lesssim \sqrt{\frac{\log p}{n}} s_0, \quad \text{with } s_0 = o(\sqrt{n / \log p}).$$

**Lemma 42.2.1 (Basic Inequality)**

$$\|\mathbf{X}(\hat{\boldsymbol{\beta}} - \boldsymbol{\beta}^0)\|_2^2 / n + \lambda \|\hat{\boldsymbol{\beta}}\|_1 \leq 2\boldsymbol{\varepsilon}^\top \mathbf{X}(\hat{\boldsymbol{\beta}} - \boldsymbol{\beta}^0) / n + \lambda \|\boldsymbol{\beta}^0\|_1.$$

*Proof.* Since  $\hat{\boldsymbol{\beta}}$  is the minimizer of the objective function in (42.1), we have

$$\|\mathbf{y} - \mathbf{X}\hat{\boldsymbol{\beta}}\|_2^2 / n + \lambda \|\hat{\boldsymbol{\beta}}\|_1 \leq \|\mathbf{y} - \mathbf{X}\boldsymbol{\beta}^0\|_2^2 / n + \lambda \|\boldsymbol{\beta}^0\|_1.$$

By rearranging the terms, we have

$$\left\| \mathbf{X} (\hat{\boldsymbol{\beta}} - \boldsymbol{\beta}^0) \right\|_2^2 / n + \lambda \left\| \hat{\boldsymbol{\beta}} \right\|_1 \leq 2\boldsymbol{\varepsilon}^\top \mathbf{X} (\hat{\boldsymbol{\beta}} - \boldsymbol{\beta}^0) / n + \lambda \left\| \boldsymbol{\beta}^0 \right\|_1.$$

□

### Corollary 42.2.1

Assume that  $\hat{\sigma}_j^2 = 1$  for all  $j$  and that the compatibility condition holds for  $S_0$ , with  $\hat{\boldsymbol{\Sigma}}$  normalized in this way. For some  $t > 0$ , let the regularization parameter be

$$\lambda := 4\hat{\sigma} \sqrt{\frac{t^2 + 2 \log p}{n}},$$

where  $\hat{\sigma}^2$  is an estimator of the noise variance  $\sigma^2$ . Then with probability at least  $1 - \alpha$ , where

$$\alpha := 2 \exp \left[ -t^2/2 \right] + \Pr(\hat{\sigma} \leq \sigma),$$

we have

$$\left\| \mathbf{X} (\hat{\boldsymbol{\beta}} - \boldsymbol{\beta}^0) \right\|_2^2 / n + \lambda \left\| \hat{\boldsymbol{\beta}} - \boldsymbol{\beta}^0 \right\|_1 \leq 4\lambda^2 s_0 / \phi_0^2.$$

*Proof.* By Lemma 6.2, for

$$\mathcal{F} = \left\{ \max_{1 \leq j \leq p} 2|\boldsymbol{\varepsilon}^\top \mathbf{X}^{(j)}|/n \leq \lambda_0 \right\},$$

we have for all  $t > 0$ ,

$$\Pr(\mathcal{F}) = 1 - 2 \exp \left[ -t^2/2 \right], \quad \text{where} \quad \lambda_0 = 2\sigma \sqrt{\frac{t^2 + 2 \log p}{n}}.$$

As in the proof of corollary 6.2, if  $\hat{\sigma} > \sigma$ , then we have  $\lambda \geq 2\lambda_0$ , and

$$\Pr(\mathcal{F} \cap \{\hat{\sigma} > \sigma\}) = 1 - \Pr(\mathcal{F}^c \cup \{\hat{\sigma} \leq \sigma\}) \geq 1 - \Pr(\mathcal{F}^c) - \Pr(\hat{\sigma} \leq \sigma) = 1 - \alpha,$$

where  $\alpha = 2 \exp \left[ -t^2/2 \right] + \Pr(\hat{\sigma} \leq \sigma)$ . And since the compatibility condition holds, according to Theorem 6.1, we have

$$\left\| \mathbf{X} (\hat{\boldsymbol{\beta}} - \boldsymbol{\beta}^0) \right\|_2^2 / n + \lambda \left\| \hat{\boldsymbol{\beta}} - \boldsymbol{\beta}^0 \right\|_1 \leq 4\lambda^2 s_0 / \phi_0^2,$$

with probability at least  $1 - \alpha$ . □



## 42.3 Other Lasso-Type Estimators

### 42.3.1 Adaptive Lasso

The adaptive lasso estimator is defined as

$$\hat{\beta}_{\text{adaptive}} := \arg \min_{\beta \in \mathbb{R}^p} \left\{ \frac{1}{2n} \|\mathbf{y} - \mathbf{X}\beta\|_2^2 + \lambda \sum_{j=1}^p \frac{|\beta_j|}{\hat{\sigma}_j} \right\},$$

where  $\lambda > 0$  is a tuning parameter, and  $\hat{\sigma}_j$  is an estimator of  $\sigma_j$ .

### 42.3.2 Elastic Net

The elastic net estimator is defined as

$$\hat{\beta}_{\text{elastic}} := \arg \min_{\beta \in \mathbb{R}^p} \left\{ \frac{1}{2n} \|\mathbf{y} - \mathbf{X}\beta\|_2^2 + \lambda_1 \|\beta\|_1 + \lambda_2 \|\beta\|_2^2 \right\},$$

where  $\lambda_1, \lambda_2 > 0$  are tuning parameters.

### 42.3.3 Group Lasso

The group lasso estimator is defined as

$$\hat{\beta}_{\text{group}} := \arg \min_{\beta \in \mathbb{R}^p} \left\{ \frac{1}{2n} \|\mathbf{y} - \mathbf{X}\beta\|_2^2 + \lambda \sum_{g=1}^G \sqrt{p_g} \|\beta_g\|_2 \right\},$$

where  $\lambda > 0$  is a tuning parameter, and  $\beta_g$  is the subvector of  $\beta$  corresponding to the  $g$ th group.

### 42.3.4 Fused Lasso

The fused lasso estimator is defined as

$$\hat{\beta}_{\text{fused}} := \arg \min_{\beta \in \mathbb{R}^p} \left\{ \frac{1}{2n} \|\mathbf{y} - \mathbf{X}\beta\|_2^2 + \lambda \sum_{j=1}^{p-1} |\beta_{j+1} - \beta_j| \right\},$$

## 42.4 Nonconvex Penalties

The main drawback of the lasso estimator is that it is a biased estimator. To reduce the bias, we can use nonconvex penalties, such as the smoothly clipped absolute deviation (SCAD) penalty and the minimax concave penalty (MCP).

### 42.4.1 SCAD

The SCAD penalty is defined as

$$\psi_{\text{SCAD}}(\beta) := \begin{cases} \lambda\beta, & \text{if } |\beta| \leq \lambda, \\ \frac{(a\lambda - |\beta|)_+}{(a-1)\lambda}, & \text{if } \lambda < |\beta| \leq a\lambda, \\ \frac{1}{(a-1)\lambda^2} [(a+1)\lambda^2 - 2a\lambda|\beta| + \beta^2], & \text{if } a\lambda < |\beta|, \end{cases}$$

where  $a > 2$  is a constant.

### 42.4.2 MCP

The MCP penalty is defined as

$$\psi_{\text{MCP}}(\beta) := \begin{cases} \lambda\beta, & \text{if } |\beta| \leq \lambda, \\ \frac{|\beta|^2}{2(a-1)}, & \text{if } \lambda < |\beta| \leq a\lambda, \\ \frac{a\lambda|\beta| - \lambda^2/2}{a-1}, & \text{if } a\lambda < |\beta|, \end{cases}$$

where  $a > 1$  is a constant.

# Part XIII

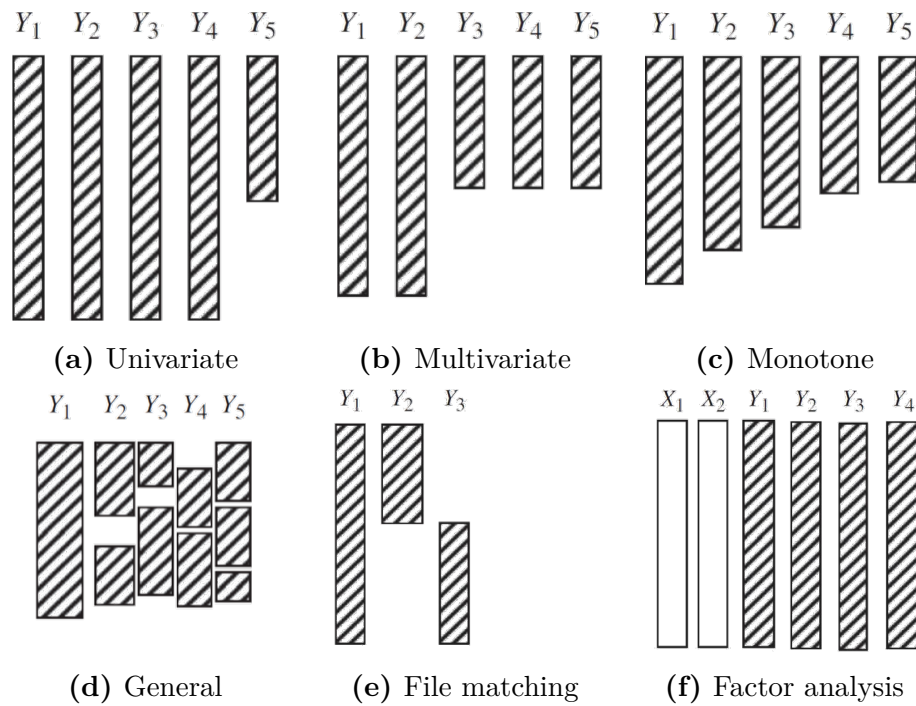
## Statistics Applications

# Chapter 43

## Missingness Data

### 43.1 The Problem of Missing Data

We are concerned with the problem of the analysis of such a data matrix when some of the entries in the matrix are not observed (Figure 43.1).



**Figure 43.1:** Examples of missingness patterns

Notations for missing data are as follows

- $Y = (y_{ij})$  denote the  $(n \times p)$  rectangular data matrix, where only a portion of  $Y$  are observed and  $y_{ij} = \star$  indicates this entry is missing;

- $M = (m_{ij})$  denote the *missingness indicator matrix* for  $y_{ij}$ , taking  $m_{ij} = 0$  for  $y_{ij}$  is observed, and  $m_{ij} = 1$  for  $y_{ij}$  is missing.
- In order to simplify, let  $y_i = (y_{i1}, y_{i2}, \dots, y_{ip})$ ,  $m_i = (m_{i1}, m_{i2}, \dots, m_{ip})$  and  $y_{(0)i}$  be the components of  $y_i$  that are observed for unit  $i$ ,  $y_{(1)i}$  be the components of  $y_i$  that are missing for unit  $i$ .

### 43.1.1 Missingness Mechanisms

The missingness mechanism is characterized by the conditional distribution of  $m_i$  given  $y_i$ , say

$$f_{M|Y}(m_i | y_i, \phi), \quad (43.1)$$

where  $\phi$  denotes unknown parameters.

#### Definition 43.1.1 (Missing Completely at Random, MCAR)

If missingness does not depend on the value of the data, missing or observed, i.e., if for all  $y_i$  and any distinct values  $y^*$  in the sample space of  $Y$ ,

$$f_{M|Y}(m_i | y_i, \phi) = f_{M|Y}(m_i | y^*, \phi), \quad (43.2)$$

then the data are called missing completely at random, MCAR.

#### Definition 43.1.2 (Missing at Random, MAR)

If missingness depends on  $y_i$  only through the observed components  $y_{(0)}$ , i.e., if for all  $y_i$  and any distinct values  $y_{(1)}^*$  in the sample space of  $y_{(1)}$ ,

$$f_{M|Y}(m_i | y_{(0)i}, y_{(1)i}, \phi) = f_{M|Y}(m_i | y_{(0)i}, y_{(1)}^*, \phi), \quad (43.3)$$

then the data are called missing at random, MAR.

#### Definition 43.1.3 (Missing Not at Random, MNAR)

If missingness depends on  $y_i$  the missing components  $y_{(1)}$ , i.e., if some  $y_i$  and some values  $y_{(1)}^*$  in the sample space of  $y_{(1)}$ ,

$$f_{M|Y}(m_i | y_{(0)i}, y_{(1)i}, \phi) \neq f_{M|Y}(m_i | y_{(0)i}, y_{(1)}^*, \phi), \quad (43.4)$$

then the data are called missing not at random, MNAR.

### 43.1.2 Commonly Used Methods for Missing Data

1. Complete-case Analysis: discard incompletely recorded units, only use the units with the complete data.
2. Weighting Procedures: randomization inferences from sample survey data without nonresponse commonly weight sampled units by their design weights.
3. Imputation Methods: impute the missing values, and the resultant completed data are analyzed by standard methods.
4. **Model-based Methods:** A broad class of procedures is generated by defining a model for the complete data and basing inferences on the likelihood or posterior distribution under that model, with parameters estimated by procedures such as maximum likelihood.
5. Hybrid Approaches: approaches based on estimating equations have been proposed that combine the aspects of modeling and weighting.

## 43.2 Likelihood-Based Inference with Missing Data

We can model the density of the joint distribution of  $Y$  and  $M$  using the "selection model" factorization

$$p(Y = y, M = m \mid \theta, \psi) = f_Y(y \mid \theta) f_{M|Y}(m \mid y, \psi),$$

where  $\theta$  is the parameter vector governing the data model, and  $\psi$  is the parameter vector governing the model for the missingness mechanism.

The full likelihood based on the observed values  $(y_{(0)}, m)$  and the assumed joint distribution model above is defined to be

$$L_{\text{full}}(\theta, \psi \mid y_{(0)}, m) = \int f_Y(y_{(0)}, y_{(1)} \mid \theta) f_{M|Y}(m \mid y_{(0)}, y_{(1)}, \psi) dy_{(1)} \quad (43.5)$$

The likelihood of  $\theta$  ignoring the missingness mechanism is defined to be

$$L_{\text{ign}}(\theta \mid y_{(0)}) = \int f_Y(y_{(0)}, y_{(1)} \mid \theta) dy_{(1)} \quad (43.6)$$

### 43.2.1 Ignorable Missingness Mechanism

#### Definition 43.2.1 (Ignorable missingness mechanism)

The missingness mechanism is called ignorable if for any given  $\tilde{m}$  and  $\tilde{y}_{(0)}$  the inferences for  $\theta$  based on the ignorable likelihood equation evaluated at  $m = \tilde{m}$  and  $\tilde{y}_0$  are the same as the full likelihood equation.

*Remark* (Another definition of ignorable missingness mechanism).

$$\frac{L_{\text{full}}(\theta, \psi \mid \tilde{y}_{(0)}, \tilde{m})}{L_{\text{full}}(\theta^*, \psi \mid \tilde{y}_{(0)}, \tilde{m})} = \frac{L_{\text{ign}}(\theta \mid \tilde{y}_{(0)})}{L_{\text{ign}}(\theta^* \mid \tilde{y}_{(0)})} \quad \forall \theta, \theta^*, \psi. \quad (43.7)$$

### Theorem 43.2.1

The missingness mechanism is ignorable for direct likelihood inference on  $(\tilde{m}, \tilde{y}_{(0)})$  if

1. Parameter distinctness: The parameters  $\theta$  and  $\psi$  are distinct, that is,  $\Omega_{\theta, \psi} = \Omega_{\theta} \times \Omega_{\psi}$ .
2. Factorization of the full likelihood: The full likelihood, with  $(y_0, m) = (\tilde{y}_0, \tilde{m})$  factors as

$$L_{\text{full}}(\theta, \psi \mid \tilde{y}_{(0)}, \tilde{m}) = L_{\text{ign}}(\theta \mid \tilde{y}_{(0)}) \times L_{\text{rest}}(\psi \mid \tilde{y}_{(0)}, \tilde{m}) \quad \forall \theta, \psi \in \Omega_{\theta, \psi} \quad (43.8)$$

### Corollary 43.2.1

If the missing data are MAR at  $(\tilde{m}, \tilde{y}_{(0)})$ , and  $\theta$  and  $\psi$  are distinct, the missingness mechanism is ignorable for likelihood inference.

*Proof.* Since,

$$f_{M|Y}(\tilde{m} \mid \tilde{y}_{(0)}, y_{(1)}, \psi) = f_{M|Y}(\tilde{m} \mid \tilde{y}_{(0)}, y_{(1)}^*, \psi) \quad \forall y_{(1)}, y_{(1)}^*, \psi \quad (43.9)$$

therefore,

$$\begin{aligned} f(\tilde{y}_{(0)}, \tilde{m} \mid \theta, \psi) &= f_{M|Y}(\tilde{m} \mid \tilde{y}_{(0)}, \psi) \times \int f_Y(\tilde{y}_{(0)}, y_{(1)} \mid \theta) dy_{(1)} \\ &= f_{M|Y}(\tilde{m} \mid \tilde{y}_{(0)}, \psi) \times f_Y(\tilde{y}_{(0)} \mid \theta) \end{aligned} \quad (43.10)$$

yields the factored likelihood equation 43.8.  $\square$

## **Ignorable Missingness Mechanism v.s. Nonignorable Missingness Mechanism**

**Example** (Exponential Sample). The joint density of  $n$  independent and identically distributed scalar units from the exponential distribution with mean  $\theta > 0$  is

$$f_Y(y \mid \theta) = \theta^{-n} \exp \left\{ -\sum_{i=1}^n \frac{y_i}{\theta} \right\}. \quad (43.11)$$

The log-likelihood function is

$$\ell_Y(\theta | y) = \ln \left\{ \theta^{-n} \exp \left( - \sum_{i=1}^n \frac{y_i}{\theta} \right) \right\} = -n \ln \theta - \sum_{i=1}^n \frac{y_i}{\theta}. \quad (43.12)$$

Differentiating to  $\theta$  gives the likelihood equation

$$-\frac{n}{\theta} + \sum_{i=1}^n \frac{y_i}{\theta^2} = 0. \quad (43.13)$$

Thus, we obtain the ML estimates

$$\hat{\theta} = \bar{y} = \frac{1}{n} \sum_{i=1}^n y_i. \quad (43.14)$$

**Example** (Incomplete Exponential Sample). Suppose we have an incomplete univariate exponential sample with  $y_{(0)} = (y_1, \dots, y_r)^\top$  observed and  $y_{(1)} = (y_{r+1}, \dots, y_n)^\top$  missing. Thus,  $m = (m_1, \dots, m_n)^\top$ , where  $m_i = 0, i = 1, \dots, r$  and  $m_i = 1, i = r+1, \dots, n$ .

The likelihood of the ignorable missingness mechanism is

$$L_{\text{ign}}(\theta | y_{(0)}) = \theta^{-r} \exp \left( - \sum_{i=1}^r \frac{y_i}{\theta} \right). \quad (43.15)$$

If each unit is observed with probability  $\psi$  that does not depend on  $Y$ , that is,

$$f_{M|Y}(m | y, \psi) = \frac{n!}{r!(n-r)!} \psi^r (1-\psi)^{n-r} \quad (43.16)$$

then,

$$f(y_{(0)}, m | \theta, \psi) = \frac{n!}{r!(n-r)!} \psi^r (1-\psi)^{n-r} \theta^{-r} \exp \left( - \sum_{i=1}^r \frac{y_i}{\theta} \right) \quad (43.17)$$

Because the missing data are MAR, if  $\psi$  and  $\theta$  are distinct, then likelihood-based inferences about  $\theta$  can be based on the ignorable likelihood, the ML estimate of  $\theta$  is

$$\hat{\theta} = \frac{1}{r} \sum_{i=1}^r y_i. \quad (43.18)$$

If each unit is observed only if values less than  $c$ , that is

$$f_{M|Y}(m | y, \psi) = \prod_{i=1}^n f(m_i | y_i, \psi), \quad (43.19)$$



where

$$f(m_i | y_i, \psi) = \begin{cases} 1, & y_i \geq c \\ 0, & \text{otherwise} \end{cases} \quad (43.20)$$

Hence,

$$\begin{aligned} L_{\text{full}}(\theta | y_{(0)}, m) &= \prod_{i=1}^r f_Y(y_i | \theta) \Pr(y_i < c | y_i, \theta) \times \prod_{i=r+1}^n \Pr(y_i \geq c | \theta) \\ &= \theta^{-r} \exp\left(-\sum_{i=1}^r \frac{y_i}{\theta}\right) \times \exp\left(-\frac{(n-r)c}{\theta}\right) \end{aligned} \quad (43.21)$$

Maximizing the above equation for  $\theta$  gives the ML estimate

$$\hat{\theta} = \frac{1}{r} \left[ \sum_{i=1}^r y_i + (n-r)c \right]. \quad (43.22)$$

The inflation of the sample mean in this expression reflects the censoring of the missing values.

### 43.2.2 Expectation-Maximization Algorithm

Let  $\theta^{(i)}$  be the current estimate of the parameter  $\theta$ . The E step of EM finds the expected complete-data loglikelihood if  $\theta$  were  $\theta^{(t)}$ :

$$Q(\theta | \theta^{(t)}) = \int \ell(\theta | Y_{(0)}, Y_{(1)}) f(Y_{(1)} | Y_{(0)}, \theta^{(t)}) dY_{(1)}. \quad (43.23)$$

The M step of EM determines  $\theta^{(t+1)}$  by maximizing this expected completedata loglikelihood:

$$Q(\theta^{(t+1)} | \theta^{(t)}) \geq Q(\theta | \theta^{(t)}), \quad \forall \theta. \quad (43.24)$$

Hence, the EM algorithm for likelihood-based inference with missing data is

1. replace missing values by estimated values
2. estimate parameters
3. re-estimate the missing values assuming the new parameter estimates are correct
4. re-estimate parameters, and so forth, iterating until apparent convergence

**Convergence Properties of EM Algorithm with Missing Data****Theorem 43.2.2**

Every GEM algorithm increases  $\ell(\theta | Y_{(0)})$  at each iteration, that is,

$$\ell(\theta^{(t+1)} | Y_{(0)}) \geq \ell(\theta^{(t)} | Y_{(0)}) \quad (43.25)$$

, with equality if and only if

$$Q(\theta^{(t+1)} | \theta^{(t)}) = Q(\theta^{(t)} | \theta^{(t)}) \quad (43.26)$$

*Proof.* The distribution of the complete data  $Y$  can be factored as follows:

$$f(Y | \theta) = f(Y_{(0)}, Y_{(1)} | \theta) = f(Y_{(0)} | \theta) f(Y_{(1)} | Y_{(0)}, \theta) \quad (43.27)$$

The corresponding decomposition of the log-likelihood is

$$\ell(\theta | Y) = \ell(\theta | Y_{(0)}, Y_{(1)}) = \ell(\theta | Y_{(0)}) + \ln f(Y_{(1)} | Y_{(0)}, \theta) \quad (43.28)$$

Let,

$$\ell(\theta | Y_{(0)}) = \ell(\theta | Y) - \ln f(Y_{(1)} | Y_{(0)}, \theta) \quad (43.29)$$

The expectation of both sides of the above equation over the distribution of the missing data  $Y_{(1)}$ , given the observed data  $Y_{(0)}$  and a current estimate of  $\theta$ , say  $\theta^{(t)}$ , is

$$\ell(\theta | Y_{(0)}) = Q(\theta | \theta^{(t)}) - H(\theta | \theta^{(t)}), \quad (43.30)$$

where

$$\begin{aligned} Q(\theta | \theta^{(t)}) &= \int \ell(\theta | Y_{(0)}, Y_{(1)}) f(Y_{(1)} | Y_{(0)}, \theta^{(t)}) dY_{(1)} \\ H(\theta | \theta^{(t)}) &= \int \ln f(Y_{(1)} | Y_{(0)}, \theta) f(Y_{(1)} | Y_{(0)}, \theta^{(t)}) dY_{(1)} \end{aligned} \quad (43.31)$$

Since,

$$\begin{aligned}
& H(\theta^{(t)}, \theta^{(t)}) - H(\theta, \theta^{(t)}) \\
&= \int \ln f(Y_{(1)} | Y_{(0)}, \theta^{(t)}) f(Y_{(1)} | Y_{(0)}, \theta^{(t)}) dY_{(1)} \\
&\quad - \int \ln f(Y_{(1)} | Y_{(0)}, \theta) f(Y_{(1)} | Y_{(0)}, \theta^{(t)}) dY_{(1)} \\
&= \int \ln \left[ \frac{f(Y_{(1)} | Y_{(0)}, \theta^{(t)})}{f(Y_{(1)} | Y_{(0)}, \theta)} \right] f(Y_{(1)} | Y_{(0)}, \theta^{(t)}) dY_{(1)} \\
&= \int -\ln \left[ \frac{f(Y_{(1)} | Y_{(0)}, \theta)}{f(Y_{(1)} | Y_{(0)}, \theta^{(t)})} \right] f(Y_{(1)} | Y_{(0)}, \theta^{(t)}) dY_{(1)} \\
&\geq -\ln \int \frac{f(Y_{(1)} | Y_{(0)}, \theta)}{f(Y_{(1)} | Y_{(0)}, \theta^{(t)})} f(Y_{(1)} | Y_{(0)}, \theta^{(t)}) dY_{(1)} = 0
\end{aligned} \tag{43.32}$$

Therefore,

$$H(\theta | \theta^{(t)}) \leq H(\theta^{(t)} | \theta^{(t)}) \tag{43.33}$$

Hence, the difference in values of  $\ell(\theta | Y_{(0)})$  at successive iterates is given by

$$\begin{aligned}
\ell(\theta^{(t+1)} | Y_{(0)}) - \ell(\theta^{(t)} | Y_{(0)}) &= [Q(\theta^{(t+1)} | \theta^{(t)}) - Q(\theta^{(t)} | \theta^{(t)})] \\
&\quad - [H(\theta^{(t+1)} | \theta^{(t)}) - H(\theta^{(t)} | \theta^{(t)})] \\
&\geq 0
\end{aligned} \tag{43.34}$$

□

### Theorem 43.2.3

Suppose a sequence of EM iterates is such that

1.  $D^{10}Q(\theta^{(t+1)} | \theta^{(t)}) = 0$ , where "D" here denotes derivative, and  $D^{10}$  means the derivative for the first argument, that is, define

$$D^{10}Q(\theta^{(t+1)} | \theta^{(t)}) = \left. \frac{\partial}{\partial \theta} Q(\theta | \theta^{(t)}) \right|_{\theta=\theta^{(t+1)}} = 0. \tag{43.35}$$

2.  $\theta^{(t)}$  converges to  $\theta^*$ .
3.  $f(Y_{(1)} | Y_{(0)}, \theta)$  is smooth in  $\theta$ , where smooth is defined in the proof.

Then

$$D\ell(\theta^* | Y_{(0)}) \equiv \left. \frac{\partial}{\partial \theta} \ell(\theta | Y_{(0)}) \right|_{\theta=\theta^*} = 0, \tag{43.36}$$

so that if the  $\theta^{(t)}$  converge, they converge to a stationary point.

*Proof.*

$$\begin{aligned}
 D\ell\left(\theta^{(t+1)} \mid Y_{(0)}\right) &= D^{10}Q\left(\theta^{(t+1)} \mid \theta^{(t)}\right) - D^{10}H\left(\theta^{(t+1)} \mid \theta^{(t)}\right) \\
 &= -D^{10}H\left(\theta^{(t+1)} \mid \theta^{(t)}\right) \\
 &= -\frac{\partial}{\partial\theta} \int \left[ \ln f\left(Y_{(1)} \mid Y_{(0)}, \theta\right) \right] f\left(Y_{(1)} \mid Y_{(0)}, \theta^{(t)}\right) dY_{(1)} \Bigg|_{\theta=\theta^{(t+1)}}
 \end{aligned} \tag{43.37}$$

which assumes sufficient smoothness to interchange the order of differentiation and integration,

$$\begin{aligned}
 &= -\int \frac{\partial}{\partial\theta} f\left(Y_{(1)} \mid Y_{(0)}, \theta\right) dY_{(1)} \Bigg|_{\theta=\theta^{(t+1)}} \\
 &= -\int \frac{\partial}{\partial\theta} f\left(Y_{(1)} \mid Y_{(0)}, \theta\right) dY_{(1)} \Bigg|_{\theta=\theta^{(t+1)}} = 0
 \end{aligned} \tag{43.38}$$

□

### Examples of EM Algorithm with Missing Data

**Example** (Multivariate Normal Sample). Let  $y = (y_{ij})$ , where  $i = 1, \dots, n$ ,  $j = 1, \dots, p$ , be a matrix representing an independent and identically distributed sample of  $n$  units from the multivariate normal distribution with mean vector  $\mu = (\mu_1, \dots, \mu_p)$  and covariance matrix  $\Sigma = (\sigma_{jk})$ ,  $j = 1, \dots, p$ ;  $k = 1, \dots, p$ . Thus,  $y_{ij}$  represents the value of the  $j$ th variable for the  $i$ th unit in the sample. The density of  $y$  is

$$f_Y(y \mid \mu, \Sigma) = (2\pi)^{-np/2} |\Sigma|^{-n/2} \exp \left\{ -\frac{1}{2} \sum_{i=1}^n (y_i - \mu) \Sigma^{-1} (y_i - \mu)^\top \right\}. \tag{43.39}$$

The loglikelihood of  $\theta = (\mu, \Sigma)$  is then

$$\ell_Y(\mu, \Sigma \mid y) = -\frac{n}{2} \ln |\Sigma| - \frac{1}{2} \sum_{i=1}^n (y_i - \mu) \Sigma^{-1} (y_i - \mu)^\top \tag{43.40}$$

Maximizing above equation with respect to  $\theta$  and  $\Sigma$  gives the ML estimate

$$\hat{\mu} = \bar{y}, \quad \hat{\Sigma} = \frac{n-1}{n} S, \tag{43.41}$$

where  $\bar{y} = (\bar{y}_1, \dots, \bar{y}_p)$  is the row vector of sample means, and  $S = (s_{jk})$  is the  $(p \times p)$  sample covariance matrix with  $(j, k)$  th element  $s_{jk} = \frac{1}{n-1} \sum_{i=1}^n (y_{ij} - \bar{y}_i) (y_{ik} - \bar{y}_k)$

**Example** (Incomplete Multivariate Normal Sample). Suppose  $Y = (Y_{(0)}, Y_{(1)})$ , where  $Y$  represents a random sample of size  $n$  on  $(Y_1, \dots, Y_p)$ ,  $Y_{(0)}$  the set of observed values, and  $Y_{(1)}$  the missing data. Also, let  $y_{(0),i}$  represent the set of variables with values observed for unit  $i$ ,  $i = 1, \dots, n$ .

The log-likelihood based on the observed data is then

$$\ell(\mu, \Sigma \mid Y_{(0)}) = -\frac{1}{2} \sum_{i=1}^n \ln |\Sigma_{(0),i}| - \frac{1}{2} \sum_{i=1}^n (y_{(0),i} - \mu_{(0),i})^\top \Sigma_{(0),i}^{-1} (y_{(0),i} - \mu_{(0),i}), \quad (43.42)$$

where  $\mu_{(0),i}$  and  $\Sigma_{(0),i}$  are the mean and covariance matrix of the observed components of  $Y$  for unit  $i$ .

The exponential family form of multivariate normal distribution with  $(\mu, \Sigma)$  is

$$f_Y(y \mid \mu, \Sigma) = (2\pi)^{-np/2} |\Lambda|^{n/2} \exp \left[ \eta^T \sum_{i=1}^n y_i - \frac{1}{2} \sum_{i=1}^n \text{tr}(\Lambda y_i y_i^T) - \frac{n}{2} \eta^T \Lambda \eta \right], \quad (43.43)$$

where  $\Lambda = \Sigma^{-1}$  and  $\eta = \Sigma^{-1}\mu$ . And

$$\ln f_Y(y \mid \mu, \Sigma) = -\frac{np}{2} \ln(2\pi) + \frac{n}{2} \ln |\Lambda| - \frac{n}{2} \eta^T \Lambda \eta + \eta^T \sum_{i=1}^n y_i - \frac{1}{2} \sum_{i=1}^n \text{tr}(\Lambda y_i y_i^T) \quad (43.44)$$

Hence,

$$\begin{aligned} Q(\theta \mid \theta^{(t)}) &= E_{Y_{(0)}, \theta^{(t)}} [\ell(\theta \mid Y_{(0)}, Y_{(1)})] \\ &= -\frac{np}{2} \ln(2\pi) + \frac{n}{2} \ln |\Lambda| - \frac{n}{2} \eta^T \Lambda \eta \\ &\quad + \eta^T E \left( \sum_{i=1}^n y_i \right) - \frac{1}{2} \sum_{i=1}^n \text{tr}(\Lambda E(y_i y_i^T)) \end{aligned} \quad (43.45)$$

Therefore, the EM algorithm for the incomplete multivariate normal sample is,

- E-step:

$$\begin{aligned} E \left( \sum_{i=1}^n y_{ij} \mid Y_{(0)}, \theta^{(t)} \right) &= \sum_{i=1}^n y_{ij}^{(t+1)}, \quad j = 1, \dots, p \\ E \left( \sum_{i=1}^n y_{ij} y_{ik} \mid Y_{(0)}, \theta^{(t)} \right) &= \sum_{i=1}^n (y_{ij}^{(t+1)} y_{ik}^{(t+1)} + c_{jki}^{(t+1)}), \quad j, k = 1, \dots, p \end{aligned} \quad (43.46)$$

where

$$\begin{aligned} y_{ij}^{(t+1)} &= \begin{cases} y_{ij}, & \text{if } y_{ij} \text{ is observed} \\ E(y_{ij} \mid y_{(0),i}, \theta^{(t)}), & \text{if } y_{ij} \text{ is missing} \end{cases} \\ c_{jki}^{(t+1)} &= \begin{cases} 0, & \text{if } y_{ij} \text{ or } y_{ik} \text{ is observed} \\ \text{Cov}(y_{ij}, y_{ik} \mid y_{(0),i}, \theta^{(t)}), & \text{if } y_{ij} \text{ and } y_{ik} \text{ are missing} \end{cases} \end{aligned} \quad (43.47)$$

- M-step:

$$\begin{aligned}
\mu_j^{(t+1)} &= n^{-1} \sum_{i=1}^n y_{ij}^{(t+1)}, \quad j = 1, \dots, p \\
\sigma_{jk}^{(t+1)} &= n^{-1} E \left( \sum_{i=1}^n y_{ij} y_{ik} \mid Y_{(0)}, \theta^{(t)} \right) - \mu_j^{(t+1)} \mu_k^{(t+1)} \\
&= n^{-1} \sum_{i=1}^n \left[ \left( y_{ij}^{(t+1)} - \mu_j^{(t+1)} \right) \left( y_{ik}^{(t+1)} - \mu_k^{(t+1)} \right) + c_{jki}^{(t+1)} \right], \quad j, k = 1, \dots, p
\end{aligned} \tag{43.48}$$

**Example** (Missing Outcomes in Multiple Linear Regression). Suppose a scalar outcome variable  $Y$  is regressed on  $p$  predictor variables  $X_1, \dots, X_p$ ,  $y_i, i = 1, \dots, m$  are missing, where

$$\begin{aligned}
E(Y \mid X_1, \dots, X_p) &= \beta_0 + \sum_{j=1}^p \beta_j X_j \\
\text{Var}(Y \mid X_1, \dots, X_p) &= \sigma^2
\end{aligned} \tag{43.49}$$

We assume the joint distribution of the data (including outcomes and predictors) is multivariate normal with

$$\begin{aligned}
\mu &= (\mu_1, \dots, \mu_p, \mu_y) \\
\Sigma &= \begin{pmatrix} \Sigma_{xx} & \Sigma_{xy} \\ \Sigma_{yx} & \sigma_{yy} \end{pmatrix}
\end{aligned} \tag{43.50}$$

, and that the missing data mechanism is ignorable.

Standard regression theory gives

$$\begin{aligned}
\beta &= \Sigma_{yx} \Sigma_{xx}^{-1}; \quad \beta_0 = \mu_y - \sum_{j=1}^p \beta_j \mu_j; \\
\sigma^2 &= \sigma_{yy} - \Sigma_{yx} \Sigma_{xx}^{-1} \Sigma_{xy}
\end{aligned} \tag{43.51}$$

The loglikelihood based on the observed data of  $\theta = (\beta, \sigma^2)$ , where  $\beta = (\beta_0, \beta_1, \dots, \beta_p)$ , given observed data  $\{(x_i, y_i), i = 1, \dots, n\}$  is

$$\ell(\beta, \sigma^2 \mid X, Y_{(0)}) = -\frac{n-m}{2} \ln \sigma^2 - \frac{1}{2} \sum_{i=m+1}^n \left( y_i - \beta_0 - \sum_{j=1}^p x_{ij} \beta_j \right)^2, \tag{43.52}$$

where only using the observed data.

EM algorithms can be applied to all observations and will obtain iteratively the same ML estimates as would have been obtained noniteratively using only the complete observations.

- E-step:

$$\begin{aligned} E(y_i | X, Y_{(0)}, \theta^{(t)}) &= \begin{cases} y_i, & \text{if } y_i \text{ is observed} \\ \beta^{(t)} \tilde{x}_i^\top & \text{if } y_i \text{ is missing} \end{cases} \\ E(y_i^2 | X, Y_{(0)}, \theta^{(t)}) &= \begin{cases} y_i^2, & \text{if } y_i \text{ is observed} \\ (\beta^{(t)} \tilde{x}_i^\top)^2 + \sigma^{(t)2}, & \text{if } y_i \text{ is missing} \end{cases}, \end{aligned} \quad (43.53)$$

where  $\tilde{x}_i = (1, x_i)$ .

- M-step:

$$\begin{aligned} \beta^{(t+1)} &= (\mathbf{X}^\top X)^{-1} \mathbf{X}^\top Y^{(t+1)} \\ \sigma^{(t+1)2} &= n^{-1} \left[ \sum_{i=m+1}^n (y_i - \beta^{(t)} x_i)^2 + m \sigma^{(t)2} \right], \end{aligned} \quad (43.54)$$

where  $X = (1, X_1, X_2, \dots, X_p)$

**Example** (Finite Mixture Linear Regression).

### 43.3 Missing Not At Random Models

Here, we based on

$$L_{\text{full}}(\theta, \psi | Y_{(0)}, X, M) \propto f(Y_{(0)}, M | X, \theta, \psi) \quad (43.55)$$

regarded as a function of the parameters  $\theta, \psi$  for fixed observed data  $Y_{(0)}$  and missingness pattern  $M$ ; here  $f(Y_{(0)}, M | X, \theta, \psi)$  is obtained by integrating  $Y_{(1)}$  out of the joint density  $f(Y, M | X, \theta, \psi)$  based on a joint model for  $Y$  and  $M$  given  $X$ .

The EM algorithm has the following form for MNAR selection models are as followed,

- E-step:

$$\begin{aligned} Q(\theta, \psi | \theta^{(t)}, \psi^{(t)}) &= \int \ell(\theta, \psi | X, Y_{(0)}, Y_{(1)}, M) \\ &\quad \cdot f(Y_{(1)} | X, Y_{(0)}, M, \theta = \theta^{(t)}, \psi = \psi^{(t)}) dY_{(1)} \end{aligned} \quad (43.56)$$

- M-step:

$$Q(\theta^{(t+1)}, \psi^{(t+1)} | \theta^{(t)}, \psi^{(t)}) \geq Q(\theta, \psi | \theta^{(t)}, \psi^{(t)}) \quad \forall \theta, \psi \quad (43.57)$$

### 43.3.1 Normal Models for MNAR Missing Data

1. Follow up a sample of nonrespondents and incorporate this information into the main analysis.
2. Adopt a Bayesian approach, assigning the parameters prior distributions. Bayesian inference does not generally require that the data provide information for all the parameters, although inferences tend to be sensitive to the choice of prior distribution.
3. Impose additional restrictions on model parameters.
4. Conduct analysis to assess the sensitivity of inferences for quantities of interest to different choices of the values of parameters poorly estimated from the data.
5. Selectively discard data to avoid modeling the missingness mechanism.



# Chapter 44

## Treatment-effects Analysis

### 44.1 Evaluations

#### 44.1.1 Average Treatment Effect

**Definition 44.1.1 (Average Treatment Effect)**

$$\mathbb{E}(Y_1 - Y_2) \tag{44.1}$$

#### 44.1.2 Mann-Whitney Statistic

**Definition 44.1.2 (Mann-Whitney Statistic)**

$$\Pr(Y_1 < Y_2) \tag{44.2}$$

#### 44.1.3 Distribution-type Index

**Definition 44.1.3 (Distribution-type Index)**

$$F(x) := \Pr(Y_1 - Y_2 = x) \tag{44.3}$$

# Chapter 45

## Graphical Lasso

Before deriving the logarithmic likelihood function, we first introduce the concept of sample covariance matrix, inverse covariance matrix, and probability density function of multivariate normal distribution. Let  $X$  be a  $p$ -dimensional random vector with mean  $\mu$  and covariance matrix  $\Sigma$ , then it follows multivariate normal distribution (also known as normal distribution):

$$X \sim N(\mu, \Sigma)$$

Its probability density function is:

$$f_X(x) = \frac{1}{(2\pi)^{p/2}|\Sigma|^{1/2}} \exp\left(-\frac{1}{2}(x - \mu)^\top \Sigma^{-1}(x - \mu)\right)$$

where  $|\Sigma|$  represents the determinant of the covariance matrix. The sample covariance matrix  $S$  can be calculated from sample data, and its specific expression is:

$$S = \frac{1}{n} \sum_{i=1}^n (x_i - \bar{x})(x_i - \bar{x})^\top$$

where  $\bar{x} = \frac{1}{n} \sum_{i=1}^n x_i$  represents the sample mean. The determinant  $|\Theta|$  of the inverse covariance matrix  $\Theta$  represents the independence between variables. L1 regularization can make it equal to zero, so that sparse estimation results can be obtained. The inverse covariance matrix is the precision matrix of a specific multivariate normal distribution, and it is defined as:

$$\Theta = (\theta_{ij}) = \Sigma^{-1}$$

The derivation of the logarithmic likelihood function models the covariance matrix and inverse covariance matrix. The likelihood function is as follows:

$$\begin{aligned}
L(\Theta|x) &= \log P(x|\Theta) \\
&= \log \prod_{i=1}^n P(x_i|\Theta) \\
&= \sum_{i=1}^n \log P(x_i|\Theta) \\
&= \sum_{i=1}^n \left( -\frac{p}{2} \log(2\pi) - \frac{1}{2} \log |\Sigma| - \frac{1}{2} (x_i - \mu)^\top \Sigma^{-1} (x_i - \mu) \right) \\
&= -\frac{np}{2} \log 2\pi - \frac{n}{2} \log |\Sigma| - \frac{1}{2} \sum_{i=1}^n (x_i - \mu)^\top \Sigma^{-1} (x_i - \mu) \\
&= -\frac{np}{2} \log 2\pi - \frac{n}{2} \log |\Theta| - \frac{1}{2} \sum_{i=1}^n (x_i - \mu)^\top \Theta (x_i - \mu)
\end{aligned}$$

Taking the negative value of  $L(\Theta|x)$  gives the logarithmic likelihood function, which is:

$$\begin{aligned}
\log L(\Theta|x) &= -L(\Theta|x) \\
&= \frac{np}{2} \log 2\pi + \frac{n}{2} \log |\Theta| + \frac{1}{2} \sum_{i=1}^n (x_i - \mu)^\top \Theta (x_i - \mu)
\end{aligned}$$

where  $\mu$  is the sample mean and  $\Sigma$  is obtained from  $\Theta$ . The Graphical Lasso loss function obtained by regularization combines this logarithmic likelihood term with the L1 regularization term.

# Chapter 46

## Semi-supervised Learning

### 46.1 Assumptions

1. **Smoothness assumption:**
2. **Low-density assumption:**
3. **Manifold assumption:**

*Remark.* For regression problems, the low-density assumption does not hold, since the decision boundary does not exist.

VAT (Virtual Adversarial Training):

$$\mathbb{E}_{(\mathbf{x}, y) \sim p_l(\mathbf{x}, y)} \ell(\mathbf{x}, y) + \lambda \cdot \mathbb{E}_{\mathbf{x} \sim p_u(\mathbf{x})} \mathbb{E}_{\tilde{\mathbf{x}} \sim q(\tilde{\mathbf{x}} | \mathbf{x})} D_{\text{KL}}(p_\theta(y | \mathbf{x}) \| p_\theta(y | \tilde{\mathbf{x}}))$$

**Covariance shift** may be a point?

# **Part XIV**

## **Machine Learning**

# Chapter 47

## Support Vector Machine

### Theorem 47.0.1

The minimizer of

$$f^*(X) = \arg \min_f \mathbb{E} \left\{ [1 - f(X)Y]_+ \mid X = x \right\}$$

is the sign of  $f(x) = \log \frac{p(x)}{1-p(x)}$ , i.e.,  $\text{sgn} \left[ p(x) - \frac{1}{2} \right]$ .

*Proof.* For the hinge loss function,

$$\begin{aligned} & E \left\{ [1 - Yg(X)]_+ \mid X = x \right\} \\ &= [1 - g(x)]_+ P(Y = 1 \mid X = x) + [1 + g(x)]_+ P(Y = -1 \mid X = x) \\ &= [1 - g(x)]_+ p(x) + [1 + g(x)]_+ [1 - p(x)] \\ &= \begin{cases} [1 - g(x)] p(x), & g(x) < -1 \\ 1 + [1 - 2p(x)] g(x), & -1 \leq g(x) \leq 1 \\ [1 + g(x)] [1 - p(x)], & g(x) > 1 \end{cases} \end{aligned}$$

When  $g(x) < -1$ ,

$$\arg \min_g E \left\{ [1 - Yg(X)]_+ \mid X = x \right\} = \arg \min_g [1 - g(x)] p(x) = -1$$

When  $g(x) > 1$ ,

$$\arg \min_g E \left\{ [1 - Yg(X)]_+ \mid X = x \right\} = \arg \min_g [1 + g(x)] [1 - p(x)] = 1$$

When  $-1 \leq g(x) \leq 1$ ,

$$\begin{aligned} & \arg \min_g E \left\{ [1 - Yg(X)]_+ \mid X = x \right\} \\ &= \arg \min_g \{ 1 + [1 - 2p(x)]g(x) \} \\ &= \begin{cases} -1, & p(x) < \frac{1}{2} \\ 0, & p(x) = \frac{1}{2} \\ 1, & p(x) > \frac{1}{2} \end{cases} \end{aligned}$$

Thus, for the  $g(x) \in [-1, 1]$  the minimizer of  $\arg \min_g E \left\{ [1 - Yg(X)]_+ \mid X = x \right\}$  is the sign of  $p(x) - \frac{1}{2}$ , that is the sign of  $f(x) = \log \frac{p(x)}{1-p(x)}$   $\square$

## Chapter 48

# Linear Discriminant Analysis



## Chapter 49

### K-Nearest Neighbor

# Chapter 50

## Decision Tree

# Chapter 51

## Kalman Filter

# **Part XV**

## **Deep Learning**

## Chapter 52

# Mixture of Experts

# **Part XVI**

## **Generative Models**

# Chapter 53

## Diffusion Model

### 53.1 Introduction

#### 53.1.1 Denoising Diffusion Probabilistic Model

Ho, Jain, and Abbeel ([2020](#))

### 53.2 Effort in Inference

Due to the sequential nature of the sampling process, the diffusion model is computationally expensive. Efforts have been made to accelerate this process, by resorting to:

**Higher-order Numerical Schemes** Dockhorn, Vahdat, and Kreis ([2022](#))

**High Resolution Image Generation** lower dimensional latent space ([rombach2022higha](#))

## References

- Dockhorn, Tim, Arash Vahdat, and Karsten Kreis (Oct. 2022). *GENIE: Higher-Order Denoising Diffusion Solvers*. DOI: [10.48550/arXiv.2210.05475](#). arXiv: [2210.05475](#). (Visited on 10/29/2024).
- Ho, Jonathan, Ajay Jain, and Pieter Abbeel (2020). “Denoising Diffusion Probabilistic Models”. In: *Advances in Neural Information Processing Systems*. Vol. 33. Curran Associates, Inc., pp. 6840–6851. (Visited on 11/20/2023).