

MATH2105: Lecture Note on Stochastic Processes

Shixiao W. Jiang

Institute of Mathematical Sciences, ShanghaiTech University, Shanghai 201210, China

jiangshx@shanghaitech.edu.cn

2024 年 5 月 1 日

Contents

1 Review of Probability Theory	5
1.1 Probability Space	5
1.1.1 Sample space	5
1.1.2 Events	5
1.1.3 Probability space	6
1.2 Conditional Prob.	6
1.3 Independent events	7
1.4 Bayes' formula	7
1.5 Discrete random variables	8
1.5.1 the Bernoulli random vari.	8
1.5.2 the Binomial random vari.	8
1.5.3 the geometric random vari.	9
1.5.4 the Poisson random vari.	9
1.6 Cont. Random vari.	10
1.6.1 Uniform random variable	11
1.6.2 Exponential random variable	11
1.6.3 Gamma ranodm variable	12
1.6.4 Normal random variable	12
1.6.5 Inverse Gamma Random Variable	12
1.6.6 History of Normal distribution	13
1.7 Jointly distributed random variables	13
1.7.1 independent random variables	13
1.7.2 Covariance and Variance of Sums of Random Variables	14
1.7.3 Sum of two independent variables	16
1.8 Moment Generating Functions	16
1.9 Limit Theorems	18
2 Conditional Probability	21
2.1 The Discrete Case	21
2.2 The Continuous Case	22
2.3 Computing Expectatoins by Conditioning	23
2.4 Computing Probabilities by Conditioning	27

2.5	Computing Conditional Expectation and Conditional Probability by Conditioning	29
3	Markov Chain	32
3.1	Introduction	32
3.2	Chapman-Kolmogorov Equations	35
3.3	Unconditional Distribution of the State	36
3.4	Classification of States	37
3.4.1	recurrent and transient states and classes	39
3.4.2	Periodicity	44
3.4.3	positive and null recurrent states and classes	45
3.4.4	ergodicity	46
3.4.5	Classification	47
3.5	Stationary Distributions, Limiting Probabilities, Limiting Distributions	49
3.5.1	Stationary Distributions	50
3.5.2	The limiting behavior of transition probability matrix	54
3.5.3	limiting probabilities and limiting distributions	54
3.5.4	Examples for stationary distributions and limiting probabilities	55
3.6	Mean Time Spent in Transient States	57
3.7	Time Reversible Markov Chains	62
3.8	Hidden Markov Model	66
3.8.1	Introduction to HMM	66
3.8.2	Key ingredients to HMM	68
3.8.3	Three Basic Problems for HMMs	68
3.8.4	Forward and Backward Approaches for Problem 1	69
4	Markov Chain Monte Carlo (MCMC)	71
4.1	Monte Carlo Methods	71
4.1.1	deterministic vs. stochastic	71
4.1.2	background	71
4.1.3	difference bw some terminology	72
4.1.4	statistical mechanics	72
4.1.5	applications	72
4.2	Generation of Random Numbers	75
4.2.1	generation of random numbers from $U(0, 1)$	75
4.2.2	generation of random numbers with general distributions	76
4.2.3	technique for reducing the variance	78
4.3	Introduction to Markov Chain Monte Carlo (MCMC)	81
4.3.1	Historical development of MCMC and its connection to statistical mechanics	81
4.3.2	Metropolis-Hastings Algorithm	82
4.3.3	Gibbs sampling	87
4.4	Parameter estimation problems	88
4.4.1	frequentist statistics vs. Bayesian statistics	88

4.4.2	The Bayesian approach	89
4.4.3	Applications of Bayesian statistics	91
4.4.4	Beta-Bernoulli model	92
4.4.5	Conjugate priors	94
4.4.6	Exponential families	96
5	Poisson Process	97
5.1	Introduction	97
5.2	The Exponential Distribution	97
5.2.1	Definition	97
5.2.2	Properties of the Exponential Distribution	98
5.3	The Poisson Process	101
5.3.1	Counting Processes	101
5.3.2	Definition of the Poisson Process	102
5.3.3	Interarrival and Waiting Time Distributions	104
5.3.4	Further Properties of Poisson Processes	106
5.3.5	Conditional Distribution of the Arrival Times	109
5.3.6	Applications of Central Limit Theorem	110
5.4	Generalization of the Poisson Process	111
5.4.1	Nonhomogeneous Poisson Process	111
5.4.2	Compound Poisson Process	111
5.5	Parameter Estimation for Poisson Processes	113
5.5.1	Estimate the rate λ using $N(t)$	113
5.5.2	Estimate the rate λ using S_n	114
6	Continuous-Time Markov Chains	115
6.1	Continuous-Time Markov Chains	115
6.2	Birth and Death Processes	116
6.3	The Transition Probability Function $P_{ij}(t)$	119
6.3.1	transition probabilities	119
6.3.2	instantaneous transition rates and Chapman-Kolmogorov equations	120
6.3.3	Kolmogorov's forward equations	121
6.3.4	Kolmogorov's backward equations	124
6.4	Limiting Probabilities	125
6.4.1	Limiting Probabilities	125
6.4.2	Comparison with discrete-time Markov Chain	127
6.5	Time Reversibility	128
6.6	Computing the Transition Probabilities	129
7	Brownian Motion and Stationary Process	131
7.1	Brownian motion	131
7.1.1	Joint probability density function	132
7.1.2	Markovian property of a Brownian motion	133

7.1.3	Brownian motion is a Gaussian process	134
7.2	Martingale	135
7.2.1	Filtration and filtered probability space	135
7.2.2	Introduction to Martingale	135
7.2.3	Examples of martingale in Brownian motion	136
7.3	Itô calculus	137
7.3.1	The pathwise property of Brownian motions	137
7.3.2	Stochastic Differential Equations	138
7.3.3	Itô Stochastic Process and Itô formula	141
7.3.4	History of Kiyosi Itô	142
7.4	Variation on Brownian Motion	144
7.4.1	Brownian motion with drift	144
7.4.2	Geometric Brownian motion	144
7.5	Stationary and weakly stationary processes	145
7.6	Discretization of SDEs	147
7.6.1	Euler Maruyama scheme	149
7.7	Langevin equation and the Fluctuation-Dissipation Theorem	149
7.7.1	Significance of fluctuation-dissipation theorem	151
7.8	Forward and Backward Kolmogorov equations for SDEs	151
7.8.1	Backward Kolmogorov Equation	151
7.8.2	Forward Kolmogorov Equation	153
7.9	Gaussian Processes	155
8	Time Series	156

Chapter 1

Review of Probability Theory

Randomness should be taken into account in data, model, equations (PDEs), etc. This can be realized by allowing the model to be probabilistic in nature, which is referred to as a probability model. The reference book is [10].

1.1 Probability Space

A probability theory is made up of three part, (Ω, \mathcal{F}, P) .

1.1.1 Sample space

Ω is a sample space.

Example 1.1.1 $\Omega = \{H(\text{ead}), T(\text{ail})\}$ for a coin flipping.

Example 1.1.2 $\Omega = \{(H, H), (H, T), (T, H), (T, T)\}$ for flipping two coins.

Example 1.1.3 $\Omega = \{1, 2, \dots, 6\}$ for rolling of a die.

Example 1.1.4 $\Omega = \left\{ \begin{array}{ccc} (1, 1) & \cdots & (1, 6) \\ \vdots & & \vdots \\ (6, 1) & \cdots & (6, 6) \end{array} \right\}$ for rolling of two dice.

1.1.2 Events

Subset E of Ω is known as an event.

Example 1.1.5 $E = \{H\}$

Example 1.1.6 $E = \{2, 4, 6\}$. Even number appears.

- union of events. Given $E_1 = \{1, 3, 5\}$ and $E_2 = \{1, 2, 3\}$, then $E_1 \cup E_2 = \{1, 2, 3, 5\}$.
- intersection of events. Given above, then $E_1 E_2 := E_1 \cap E_2 = \{1, 3\}$.
- complement. $E_1^c = \{2, 4, 6\}$.
- mutually exclusive. E, F, G are called mutually exclusive if $EF = \emptyset, EG = \emptyset, FG = \emptyset$. For example, $E = \{1, 2\}, F = \{3, 4\}, G = \{5, 6\}$ are mutually exclusive.

Definition 1.1.7 \mathcal{F} is a family of subsets of Ω satisfying:

- (1) $\Omega \in \mathcal{F}$.
- (2) $E \in \mathcal{F} \Rightarrow E^c \in \mathcal{F}$.
- (3) $E_j \in \mathcal{F} \Rightarrow \bigcup_{j=1}^{\infty} E_j \in \mathcal{F}$.

Then \mathcal{F} is a σ -algebra of Ω . Ω, \mathcal{F} is called a measurable space.

1.1.3 Probability space

Definition 1.1.8 P is a function defined on satisfying:

- (1) non-negative. $0 \leq P(E) \leq 1, \forall E \in \mathcal{F}$.
- (2) completeness. $P(\Omega) = 1$.
- (3) For any countable mutually exclusive sets in \mathcal{F} , $P\left(\bigcup_{j=1}^{\infty} E_j\right) = \sum_{j=1}^{\infty} P(E_j)$.

Then $P(E)$ is the prob. of E . (Ω, \mathcal{F}, P) is the triple elements of a prob. space.

Example 1.1.9 $P(\{H\}) = P(\{T\}) = \frac{1}{2}$.

Example 1.1.10 $P(\{1\}) = \dots = P(\{6\}) = \frac{1}{6}$.

$$P(\{1, 3, 5\}) = P(\{1\}) + P(\{3\}) + P(\{5\}) = \frac{1}{2}.$$

Example 1.1.11 $P(E_1 \cup E_2 \cup E_3) = P(E_1) + P(E_2) + P(E_3) - P(E_1 E_2) - P(E_1 E_3) - P(E_2 E_3) + P(E_1 E_2 E_3)$.

1.2 Conditional Prob.

Definition 1.2.1

$$P(E|F) = \frac{P(EF)}{P(F)}.$$

Example 1.2.2 Choose one number from 1-10. The number is at least five, then what is the cond. prob. that it is ten?

Sol: Let $E = \{10\}$ and $F = \{\geq 5\}$. Then

$$P(E|F) = \frac{P(EF)}{P(F)} = \frac{\frac{1}{10}}{\frac{6}{10}} = \frac{1}{6}.$$

Example 1.2.3 An urn contains 7 black balls and 5 white balls. Draw two balls without replacement. Each ball is equally drawn. What is prob. that both drawn balls are black?

Sol: $F = \{\text{first ball is black}\}$, $E = \{\text{2nd ball is black}\}$. Since $P(E|F) = \frac{6}{11}$, $P(F) = \frac{7}{12}$, then

$$P(EF) = P(F)P(E|F) = \frac{7 * 6}{12 * 11}.$$

Another solution is given directly by

$$P(\text{both black}) = \frac{C_7^2}{C_{12}^2} = \frac{7 * 6}{12 * 11}.$$

1.3 Independent events

$$\begin{aligned}
E, F \text{ are independent} &\Leftrightarrow P(EF) = P(E)P(F) \\
&\Leftrightarrow P(E|F) = P(E), \quad P(F) \neq 0 \\
&\Leftrightarrow P(F|E) = P(F), \quad P(E) \neq 0
\end{aligned}$$

Example 1.3.1 Let a ball be drawn from an urn containing 4 balls {1, 2, 3, 4}. Let $E = \{1, 2\}$, $F = \{1, 3\}$, $G = \{1, 4\}$. Then

$$\begin{aligned}
P(EF) &= P(E)P(F) = \frac{1}{4}, \\
P(EG) &= P(E)P(G) = \frac{1}{4}, \\
P(FG) &= P(F)P(G) = \frac{1}{4}.
\end{aligned}$$

However,

$$\frac{1}{4} = P(EFG) \neq P(E)P(F)P(G) = \frac{1}{8}.$$

E, F, G are not jointly independent.

1.4 Bayes' formula

- sub-additive.

$$P\left(\bigcup_{j=1}^{\infty} E_j\right) \leq \sum_{j=1}^{\infty} P(E_j).$$

- multiplicity formula.

$$P(B_1B_2 \dots B_n) = P(B_1)P(B_2|B_1) \cdots P(B_n|B_1B_2 \dots B_{n-1}).$$

- formula of total probability. Suppose that F_1, \dots, F_n are mutually exclusive events such that $\cup_{i=1}^n F_i = \Omega$ (sample space) or $E \subset \cup_{i=1}^n F_i$, then $E = \cup_{i=1}^n EF_i$. Noticing EF_i are mutually exclusive, we obtain that

$$P(E) = \sum_{j=1}^n P(EF_j) = \sum_{j=1}^n P(E|F_j)P(F_j).$$

In particular, if $n = 2$, then $E = EF \cup EF^c$ and

$$P(E) = P(EF) + P(EF^c) = P(E|F)P(F) + P(E|F^c)P(F^c).$$

Another formula is: suppose that $\cup_{i=1}^n C_i = \Omega$ (sample space) or $B \subset \cup_{i=1}^n C_i$, then if $P(A) > 0$,

$$P(B|A) = \sum_{j=1}^n P(C_j|A)P(B|AC_j).$$

Pf.

$$\begin{aligned}
\sum_{j=1}^n P(C_j|A)P(B|AC_j) &= \sum_{j=1}^n \frac{P(C_jA)}{P(A)} \frac{P(BAC_j)}{P(AC_j)} = \sum_{j=1}^n \frac{P(BAC_j)}{P(A)} \\
&= \frac{P(BA)}{P(A)} = P(B|A).
\end{aligned}$$

- Bayes' formula.

$$P(F_j|E) = \frac{P(EF_j)}{P(E)} = \frac{P(E|F_j)P(F_j)}{\sum_{j=1}^n P(E|F_j)P(F_j)}.$$

Example 1.4.1 a multiple-choice test. Let p be the prob. that the student knows the answer. Assume that a student who guesses at the answer will be correct with prob. $1/m$, where there are m choices. What is the prob. that a student knew the answer given that she answered it correctly?

Sol: Let C and K be "correct" and "know", respectively.

$$\begin{aligned} P(K|C) &= \frac{P(KC)}{P(C)} = \frac{P(C|K)P(K)}{P(C|K)P(K) + P(C|K^c)P(K^c)} \\ &= \frac{p}{p + (1/m)(1 - p)}. \end{aligned}$$

If $m = 5$ and $p = \frac{1}{2}$, $P(K|C) = 5/6$.

1.5 Discrete random variables

Definition 1.5.1 If X is discrete with prob. mass function $p(x)$, then for any real-valued function g , the expectation is defined as

$$E[g(X)] = \sum_{x:p(x)>0} g(x)p(x).$$

1.5.1 the Bernoulli random vari.

An experiment, whose outcome is either a success or a failure. $X = 1$ is a success and $X = 0$ is a failure. Then X is denoted as $X \sim B(1, p)$ and the pmf is

$$\begin{aligned} p(0) &= P(X = 0) = 1 - p, \\ p(1) &= P(X = 1) = p. \end{aligned}$$

Its expect and var is

$$\begin{aligned} EX &= 1 \cdot p + 0 \cdot q = p, \\ Var(X) &= EX^2 - (EX)^2 = 1^2 \cdot p + 0^2 \cdot q - p^2 = p(1 - p). \end{aligned}$$

1.5.2 the Binomial random vari.

Suppose there are n trials of Bernoulli experiments. That is, If X_1, \dots, X_n are samples from $B(1, p)$, then

- (1) $Y = X_1 + \dots + X_n \sim B(n, p)$.
- (2) pdf is given by $p(i) = C_n^i p^i (1 - p)^{n-i}$, $i = 0, \dots, n$.
- (3) $EY = \sum_{i=0}^n ip(i) = \sum_{i=0}^n iC_n^i p^i q^{n-i} = \sum_{i=1}^n i \frac{n!}{i!(n-i)!} p^i q^{n-i} = np \sum_{i=1}^n \frac{(n-1)!}{(i-1)!(n-i)!} p^{i-1} q^{n-i} = np$.
 $Var(Y) = npq$.

Example 1.5.2 Suppose each independent engine of an airplane will fail, when in flight, with prob. $1 - p$. Suppose that the airplane will make a successful flight if at least 50 percent of its engines remain operative. For what values of p is a four-engine plane preferable to a two-engine plane?

Sol: A four-engine plane makes a successful flight with prob.

$$\begin{aligned} & C_4^2 p^2 (1-p)^2 + C_4^3 p^3 (1-p)^1 + C_4^4 p^4 (1-p)^0 \\ &= 6p^2(1-p)^2 + 4p^3(1-p) + p^4. \end{aligned}$$

The prob. for a two-engine plane is

$$C_2^1 p^1 (1-p)^1 + C_2^2 p^2 (1-p)^0 = 2p(1-p) + p^2.$$

Hence a four-engine plane is safer if

$$\begin{aligned} 6p^2(1-p)^2 + 4p^3(1-p) + p^4 &\geq 2p(1-p) + p^2 \\ 6p^3 - 12p^2 + 6p + 4p^2 - 4p^3 + p^3 &\geq 2 - p \\ 3p^3 - 8p^2 + 7p - 2 &\geq 0 \\ (p-1)^2(3p-2) &\geq 0. \\ p &\geq \frac{2}{3}. \end{aligned}$$

1.5.3 the geometric random vari.

Suppose that independent trials, each having prob. p of being a success, are performed until a success occurs. Let X be the number of trials required until the first success. The pmf is given by

$$p(n) = P(X = n) = (1-p)^{n-1}p, \quad n = 1, 2, \dots$$

To check it is a pmf

$$\sum_{n=1}^{\infty} p(n) = p \sum_{n=1}^{\infty} (1-p)^{n-1} = \frac{p}{1-(1-p)} = 1.$$

The expect. and var is

$$\begin{aligned} EX &= \sum_{i=1}^{\infty} iq^{i-1}p = p \sum_{i=1}^{\infty} \frac{dq^i}{dq} = p \left(\frac{q}{1-q} \right)' = \frac{p}{(1-q)^2} = \frac{1}{p}. \\ Var(X) &= \frac{1-p}{p^2}. \end{aligned}$$

1.5.4 the Poisson random vari.

X is said to be a Poisson random vari. with parameter λ , denoted by $X \sim P(\lambda)$,

$$p(i) = P(X = i) = e^{-\lambda} \frac{\lambda^i}{i!}, \quad i = 0, 1,$$

Check it is pmf

$$\sum_{i=0}^{\infty} p(i) = e^{-\lambda} \sum_{i=0}^{\infty} \frac{\lambda^i}{i!} = e^{-\lambda} e^{\lambda} = 1.$$

- (1) $EX = \lambda, Var(X) = \lambda$.
(2) If X_1, \dots, X_n are indepen, $X_i \sim P(\lambda_i)$, then

$$X_1 + \dots + X_n \sim P(\lambda_1 + \dots + \lambda_n).$$

Pf. By induction.

$$\begin{aligned} P(X_1 + X_2 = i) &= \sum_{k=0}^i P(X_1 = k)P(X_2 = i - k) \\ &= \sum_{k=0}^i e^{-\lambda_1} \frac{\lambda_1^k}{k!} e^{-\lambda_2} \frac{\lambda_2^{i-k}}{(i-k)!} = e^{-(\lambda_1+\lambda_2)} \frac{1}{i!} \sum_{k=0}^i \frac{i!}{k!(i-k)!} \lambda_1^k \lambda_2^{i-k} \\ &= e^{-(\lambda_1+\lambda_2)} \frac{1}{i!} (\lambda_1 + \lambda_2)^i \sim P(\lambda_1 + \lambda_2). \end{aligned}$$

Example 1.5.3 Suppose that the number of typo errors on a single page of a book has a Poisson distr. with parameter $\lambda = 1$. Calculate the prob. that there is at least one error on this page.

Sol: $P(X \geq 1) = 1 - P(X = 0) = 1 - e^{-1} = 0.633$.

1.6 Cont. Random vari.

Definition 1.6.1 The cumulative distribution function (cdf) (or sometimes just distribution function) $F(\cdot)$ is defined by, $F(b) = P(X \leq b)$, satisfying (i) $F(b)$ is a nondecreasing function of b , (ii) $\lim_{b \rightarrow \infty} F(b) = F(\infty) = 1$, (iii) $\lim_{b \rightarrow -\infty} F(b) = F(-\infty) = 0$.

One can see obviously that

$$P(a \leq X \leq b) = F(b) - F(a), \quad \text{for all } a < b.$$

Definition 1.6.2 If there exists a nonnegative function $f(x)$, defined for all real $x \in (-\infty, \infty)$, having the property that for any set B ,

$$P(X \in B) = \int_B f(x)dx.$$

The function $f(x)$ is called the prob. density function (pdf) of X .

The relation bw the cdf F and the pdf f is

$$F(a) = P(X \leq a) = \int_{-\infty}^a f(x)dx,$$

and

$$\frac{dF(a)}{da} = f(a).$$

Definition 1.6.3 If X is cont. random vari. with pdf $f(x)$, then for any real-valued function g , its expectation is defined by

$$E[g(X)] = \int_{-\infty}^{\infty} g(x)f(x)dx.$$

1.6.1 Uniform random variable

Definition 1.6.4 A random vari X is said to be uniformly distributed over $(0, 1)$, if its pdf is given by

$$f(x) = \begin{cases} 1, & 0 < x < 1, \\ 0, & \text{otherwise.} \end{cases}$$

Denote by $X \sim \mathcal{U}(0, 1)$. Its expectation and variance are

$$EX = \int_0^1 xf(x)dx = \frac{1}{2}, \quad Var(X) = EX^2 - (EX)^2 = \frac{1}{3} - \frac{1}{4} = \frac{1}{12}.$$

1.6.2 Exponential random variable

Definition 1.6.5 A cont. random vari. whose pdf is given, for some $\lambda > 0$, by

$$f(x) = \begin{cases} \lambda e^{-\lambda x}, & \text{if } x \geq 0, \\ 0, & \text{if } x < 0. \end{cases}$$

is said to be an exponential random variable with rate parameter λ . Denote by $X \sim \mathcal{E}(\lambda)$.

The cdf can be calculated by

$$\begin{aligned} F(a) &= \int_0^a \lambda e^{-\lambda x} dx = 1 - e^{-\lambda a}, \quad a \geq 0. \\ F(\infty) &= \int_0^\infty \lambda e^{-\lambda x} dx = 1. \end{aligned}$$

- (1) $EX = \frac{1}{\lambda}$ and $Var(X) = \frac{1}{\lambda^2}$.
- (2) $P(X > t) = e^{-\lambda t}, t \geq 0$.
- (3) Y is an exponential random vari. if and only if $EY > 0$ and for $\forall s, t > 0$, such that

$$P(Y > s + t | Y > s) = P(Y > t),$$

where this condition is called **memoryless**. Denote $\bar{F}(t) = P(Y > t)$, then above Eq. is equivalent to

$$\bar{F}(t + s) = \bar{F}(t)\bar{F}(s).$$

Proof. One can easily check \Rightarrow by noticing that $\bar{F}(t) = e^{-\lambda t}$. For the opposite direction, we want to prove $\bar{F}(t)$ is an exponential function. We first prove if $f(t + s) = f(t) + f(s)$, then f is linear function. For integers t and s , we have $f(n) = nf(1)$. For rational numbers t and s , $qf(p/q) = f(p) = pf(1)$, then $f(p/q) = p/qf(1)$. Since rational numbers are dense in real numbers, one can show $f(x) = xf(1)$ for all x real. Finally, $\bar{F}(t) = e^{f(t)} = e^{tf(1)}$, which is an expon. function. ■

Example 1.6.6 Suppose a clock or a watch has a lifetime with exponential distribution with expectation 1 year. If it already works for 2 months, what's its remaining lifetime? (1 year since memoryless).

Example 1.6.7 Assume that the customer comes with interarrival time being exponential dist. If a cashier wants to go washroom, he/she goes right now or later on? (Right now since memoryless).

1.6.3 Gamma ranodm variable

Definition 1.6.8 A cont. random vari. whose pdf is given, for some $\lambda > 0, \alpha > 0$, by

$$f(x) = \begin{cases} \frac{\lambda e^{-\lambda x} (\lambda x)^{\alpha-1}}{\Gamma(\alpha)}, & \text{if } x \geq 0, \\ 0, & \text{if } x < 0. \end{cases}$$

is said to be a gamma random variable with rate parameters λ and α . Denote by $X \sim \Gamma(\alpha, \lambda)$. A Gamma function is defined by

$$\Gamma(\alpha) = \int_0^\infty e^{-x} x^{\alpha-1} dx.$$

The expectation and var of gamma vari is given by

$$EX = \frac{\alpha}{\lambda}, Var(X) = \frac{\alpha}{\lambda^2}.$$

1.6.4 Normal random variable

Definition 1.6.9 X is normal random vari. with parameters μ and σ^2 if the density of X is given by

$$f(x) = \frac{1}{\sqrt{2\pi}\sigma} e^{-(x-\mu)^2/2\sigma^2}, -\infty < x < \infty.$$

The density is bell-shaped curve that is symmetric around μ .

Definition 1.6.10 multivariate normal distribution. $\varepsilon = (\varepsilon_1, \dots, \varepsilon_m)^T$. If $X = \mu + B\varepsilon$, then

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

where $\Sigma = BB^T$ is the covariance matrix of X .

- (1) $X = (X_1, \dots, X_n)^T \sim \mathcal{N}(\mu, \Sigma)$ if and only if $\forall a_1, \dots, a_n, \sum_{j=1}^n a_j X_j$ is normally distributed.
- (2) Let $X \sim \mathcal{N}(\mu, \Sigma)$. Then X_1, \dots, X_n are independent if and only if they are uncorrelated, i.e., $Cov(X_i, X_j) = 0$ for $i \neq j$. The proof can be found following.

Many real-world quantities tend to be normally distributed—for instance, human heights and other body measurements, cumulative hydrologic measures such as annual rainfall or monthly river discharge, errors in astronomical or physical observations, and diffusion of a substance in a liquid or gas. Some things are products of many independent variables (rather than sums), and in such cases the logarithm will be approximately normal since it is a sum of many independent variables—this is often the case for economic quantities such as stock market indices, due to the effect of compound interest.

1.6.5 Inverse Gamma Random Variable

If X is Gamma distributed then the distribution of $1/X$ is called the Inverse Gamma distribution. More precisely, if $X \sim \text{Gamma}(a, b)$ and $Y = 1/X$ then $Y \sim \text{InvGamma}(a, b)$, and the p.d.f. of Y is

$$\text{InvGamma}(y|a, b) = \frac{b^a}{\Gamma(a)} y^{-a-1} \exp(-b/y).$$

So, putting a $\text{Gamma}(a, b)$ prior on the precision λ is equivalent to putting an $\text{InvGamma}(a, b)$ prior on the variance $\sigma^2 = 1/\lambda$. The Inverse Gamma can be used to define a NormalInvGamma distribution for use as a prior on (μ, σ^2) , which is sometimes more convenient than (but equivalent to) using a NormalGamma prior on (μ, λ) .

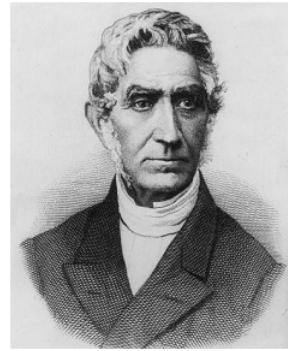
1.6.6 History of Normal distribution



Carl Friedrich Gauss



James Clerk Maxwell



Adolphe Quetelet

Figure 1.1: History of the normal distribution.

In 1809, Carl Friedrich Gauss (1777–1855) proposed the normal distribution as a model for the errors made in astronomical measurements, as a formal way of justifying the use of the sample mean, by showing it to be the most likely estimate—that is, the maximum likelihood estimate—of the true value (and more generally, to justify the method of least squares in linear regression). With astonishing speed, following Gauss' proposal, Laplace proved the central limit theorem in 1810. Laplace also calculated the normalization constant of the normal distribution, which is not a trivial task. James Clerk Maxwell (1831–1879) showed that the normal distribution arose naturally in physics, particularly in thermodynamics. Adolphe Quetelet (1796–1874) pioneered the use of the normal distribution in the social sciences. (See Fig. 1.1.)

1.7 Jointly distributed random variables

1.7.1 independent random variables

Definition 1.7.1 *The random variables X and Y are said to be independent if, for all a, b ,*

$$P(X \leq a, Y \leq b) = P(X \leq a)P(Y \leq b).$$

In terms of the joint distribution function F , we have that

$$F(a, b) = F_X(a)F_Y(b) \quad \text{for all } a, b.$$

Corollary 1.7.2 *When X and Y are discrete, the condition of indep. reduces to*

$$p(x, y) = p_X(x)p_Y(y).$$

If X and Y are jointly continuous, independence reduces to

$$f(x, y) = f_X(x)f_Y(y).$$

Pf.

$$\begin{aligned} P(X \leq a, Y \leq b) &= \sum_{y \leq b} \sum_{x \leq a} p(x, y) = \sum_{y \leq b} \sum_{x \leq a} p_X(x)p_Y(y) \\ &= \sum_{y \leq b} p_Y(y) \sum_{x \leq a} p_X(x) = P(Y \leq b)P(X \leq a). \end{aligned}$$

If X and Y are independent, then for any h and g

$$E[g(X)h(Y)] = E[g(X)]E[h(Y)].$$

Pf.

$$E[g(X)h(Y)] = \sum_y \sum_x g(x)h(y)p(x, y) = \sum_y \sum_x g(x)h(y)p_X(x)p_Y(y) = E[g(X)]E[h(Y)].$$

$$\begin{aligned} E[g(X)h(Y)] &= \int_{-\infty}^{\infty} \int_{-\infty}^{\infty} g(x)h(y)f(x, y)dxdy = \int_{-\infty}^{\infty} \int_{-\infty}^{\infty} g(x)h(y)f_X(x)f_Y(y)dxdy \\ &= \int_{-\infty}^{\infty} g(x)f_X(x)dx \int_{-\infty}^{\infty} h(y)f_Y(y)dy = E[g(X)]E[h(Y)]. \end{aligned}$$

Example 1.7.3 (Variance of a Binomial Random Variable) Compute the Variance of a Binomial Random Variable. Sol. Binomial is the sum of n independent Bernoulli.

$$\text{Var}(X) = \text{Var}(X_1) + \cdots + \text{Var}(X_n) = npq,$$

since $\text{Var}(X_i) = pq$ for each Bernoulli distribution.

1.7.2 Covariance and Variance of Sums of Random Variables

Definition 1.7.4 The covariance of any two random vari. is

$$\text{Cov}(X, Y) = E[(X - EX)(Y - EY)] = E[XY] - E[X]E[Y].$$

If X and Y are independent, then $\text{Cov}(X, Y) = 0$.

Corollary 1.7.5 Property of Covariance

- (1) $\text{Cov}(X, X) = \text{Var}(X)$,
- (2) $\text{Cov}(X, Y) = \text{Cov}(Y, X)$,
- (3) $\text{Cov}(cX + dZ, Y) = c\text{Cov}(X, Y) + d\text{Cov}(Z, Y)$.

A useful expression for the variance can be found as follows:

$$\begin{aligned} \text{Var}\left(\sum_{i=1}^n X_i\right) &= \text{Cov}\left(\sum_{i=1}^n X_i, \sum_{j=1}^n X_j\right) = \sum_{i=1}^n \sum_{j=1}^n \text{Cov}(X_i, X_j) \\ &= \sum_{i=1}^n \text{Cov}(X_i, X_i) + 2 \sum_{i < j} \text{Cov}(X_i, X_j). \end{aligned}$$

Moreover, if X_i are independent random variables, then above equation reduces to

$$\text{Var}\left(\sum_{i=1}^n X_i\right) = \sum_{i=1}^n \text{Var}(X_i).$$

Definition 1.7.6 If X_1, \dots, X_n are i.i.d., then the random variable $\bar{X} = \sum_{i=1}^n X_i/n$ is called the **sample mean**.

Proposition 1.7.7 Suppose that X_1, \dots, X_n are i.i.d. with mean μ and variance σ^2 . Then

- (a) $E[\bar{X}] = \mu$.
- (b) $Var(\bar{X}) = \sigma^2/n$.
- (c) $Cov(\bar{X}, X_i - \bar{X}) = 0, i = 1, \dots, n$.

Pf. Parts (a) and (b) are easy:

$$\begin{aligned} E[\bar{X}] &= \frac{1}{n} \sum_{i=1}^n EX_i = \mu, \\ Var[\bar{X}] &= \left(\frac{1}{n}\right)^2 Var\left(\sum_{i=1}^n X_i\right) = \frac{1}{n^2} \sum_{i=1}^n Var(X_i) = \frac{\sigma^2}{n}. \end{aligned}$$

To prove (c), we follow

$$\begin{aligned} Cov(\bar{X}, X_i - \bar{X}) &= Cov(\bar{X}, X_i) - Cov(\bar{X}, \bar{X}) = \frac{1}{n} Cov(X_i + \sum_{j \neq i} X_j, X_i) - Var[\bar{X}] \\ &= \frac{\sigma^2}{n} - \frac{\sigma^2}{n} = 0. \end{aligned}$$

Proposition 1.7.8 The **sample variance** is given by

$$S^2 = \frac{1}{n-1} \sum_{i=1}^n (X_i - \bar{X})^2.$$

Then it is unbiased, that is,

$$ES^2 = \sigma^2.$$

Pf. Notice that

$$\begin{aligned} \sum_{i=1}^n (X_i - \bar{X})^2 &= \sum_{i=1}^n (X_i - \mu + \mu - \bar{X})^2 \\ &= \sum_{i=1}^n (X_i - \mu)^2 + n(\mu - \bar{X})^2 + 2(\mu - \bar{X}) \sum_{i=1}^n (X_i - \mu) \\ &= \sum_{i=1}^n (X_i - \mu)^2 - n(\mu - \bar{X})^2. \end{aligned}$$

Then we obtain

$$\begin{aligned} E[(n-1)S^2] &= \sum_{i=1}^n E(X_i - \bar{X})^2 = \sum_{i=1}^n E(X_i - \mu)^2 - nE(\mu - \bar{X})^2 \\ &= n\sigma^2 - nVar[\bar{X}] = n\sigma^2 - n\frac{\sigma^2}{n} = (n-1)\sigma^2. \end{aligned}$$

1.7.3 Sum of two independent variables

Let us derive the formula first. Suppose that X and Y are continuous and independent, X having pdf f and Y having pdf g . Letting $F_{X+Y}(a)$ be the cdf of $X + Y$, we have

$$\begin{aligned} F_{X+Y}(a) &= P(X + Y \leq a) = \iint_{x+y \leq a} f(x)g(y)dxdy \\ &= \int_{-\infty}^{\infty} \left(\int_{-\infty}^{a-y} f(x)dx \right) g(y)dy = \int_{-\infty}^{\infty} F_X(a-y)g(y)dy. \end{aligned}$$

By differentiating above, we obtain the pdf $f_{X+Y}(a)$ of $X + Y$ given by

$$f_{X+Y}(a) = \frac{d}{da} \int_{-\infty}^{\infty} F_X(a-y)g(y)dy = \int_{-\infty}^{\infty} f(a-y)g(y)dy.$$

Thus f_{X+Y} is the **convolution** of functions f and g .

Example 1.7.9 Two uniform random vari. If X and Y are indepd. both uniformly distributed on $(0, 1)$, then calculate the pdf of $X + Y$.

Sol. The pdf's are

$$f(a) = g(a) = \begin{cases} 1, & 0 < a < 1, \\ 0, & \text{otherwise.} \end{cases}$$

we obtain

$$f_{X+Y}(a) = \int_0^1 f(a-y)g(y)dy.$$

For $0 \leq a \leq 1$, this yields

$$f_{X+Y}(a) = \int_0^a dy = a$$

since $0 \leq a - y \leq 1$ and $0 \leq y \leq 1 \Rightarrow 0 \leq y \leq a$. And for $1 \leq a \leq 2$, this yields

$$f_{X+Y}(a) = \int_{a-1}^1 dy = 2 - a$$

since $0 \leq a - y \leq 1$ and $0 \leq y \leq 1 \Rightarrow a - 1 \leq y \leq 1$. Hence,

$$f_{X+Y}(a) = \begin{cases} a, & 0 < a < 1, \\ 2 - a, & 1 < a < 2 \\ 0, & \text{otherwise.} \end{cases}$$

1.8 Moment Generating Functions

Definition 1.8.1 The moment generating function (MGF) $\phi(t)$ of random variables X is defined by

$$\phi(t) = E[e^{tX}] = \begin{cases} \sum_x e^{tx} p(x), & \text{if } X \text{ is discrete} \\ \int_{-\infty}^{\infty} e^{tx} f(x)dx, & \text{if } X \text{ is continuous} \end{cases}$$

It is called MGF because all moments of X can be obtained by successively differentiating $\phi(t)$. For example,

$$\phi'(t) = \frac{d}{dt} E[e^{tX}] = E\left[\frac{d}{dt} e^{tX}\right] = E[Xe^{tX}].$$

Hence $\phi'(0) = EX$. Similarly,

$$\phi''(t) = \frac{d}{dt} E[Xe^{tX}] = E\left[\frac{d}{dt} Xe^{tX}\right] = E[X^2 e^{tX}],$$

so that $\phi''(0) = EX^2$. One can show that $\phi^{(n)}(0) = E[X^n]$ for $n \geq 1$.

Example 1.8.2 (Poisson Distribution with mean λ).

$$\begin{aligned}\phi(t) &= E[e^{tX}] = \sum_{n=0}^{\infty} \frac{e^{tn} e^{-\lambda} \lambda^n}{n!} = e^{-\lambda} \sum_{n=0}^{\infty} \frac{(e^t \lambda)^n}{n!} = e^{-\lambda} e^{\lambda e^t} = \exp[\lambda(e^t - 1)]. \\ \phi'(t) &= \lambda e^t \exp[\lambda(e^t - 1)], \\ \phi''(t) &= (\lambda e^t)^2 \exp[\lambda(e^t - 1)] + \lambda e^t \exp[\lambda(e^t - 1)],\end{aligned}$$

Thus,

$$\begin{aligned}EX &= \phi'(0) = \lambda, EX^2 = \phi''(0) = \lambda^2 + \lambda, \\ Var(X) &= \lambda.\end{aligned}$$

Example 1.8.3 Exponential Distribution with parameter λ .

$$\begin{aligned}\phi(t) &= E[e^{tX}] = \int_0^\infty e^{tx} \lambda e^{-\lambda x} dx = \frac{\lambda}{\lambda - t} \text{ for } t < \lambda \\ \phi'(t) &= \frac{\lambda}{(\lambda - t)^2}, \phi''(t) = \frac{2\lambda}{(\lambda - t)^3}.\end{aligned}$$

Hence,

$$EX = \frac{1}{\lambda}, EX^2 = \frac{2}{\lambda^2}, Var(X) = \frac{1}{\lambda^2}.$$

Example 1.8.4 Normal distribution with mean μ and variance σ^2 . Compute the MGF of a standard normal random variable Z as follows.

$$E[e^{tZ}] = \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{\infty} e^{tx} e^{-x^2/2} dx = e^{t^2/2}.$$

If $X = \sigma Z + \mu$ is normal, then

$$\begin{aligned}\phi(t) &= E[e^{tX}] = e^{t\mu} E[e^{t\sigma Z}] = \exp\left\{\frac{\sigma^2 t^2}{2} + \mu t\right\}. \\ \phi'(t) &= (\mu + \sigma^2 t) \exp\left\{\frac{\sigma^2 t^2}{2} + \mu t\right\} \\ \phi''(t) &= (\mu + \sigma^2 t)^2 \exp\left\{\frac{\sigma^2 t^2}{2} + \mu t\right\} + \sigma^2 \exp\left\{\frac{\sigma^2 t^2}{2} + \mu t\right\}\end{aligned}$$

Hence $EX = \mu, Var(X) = \sigma^2$.

Theorem 1.8.5 A MGF uniquely determines a probability distribution.

Theorem 1.8.6 X_1, \dots, X_n are independent if and only if their MGFs satisfy

$$E e^{\vec{t} \cdot \vec{X}} := \phi(t_1, \dots, t_n) = \prod_{i=1}^n \phi_i(t_i) := \prod_{i=1}^n E e^{t_i X_i}.$$

Theorem 1.8.7 X_i are independent and have MGFs $\phi_i(t)$. Then $Y = X_1 + \dots + X_n$ has the MGF

$$\phi_Y(t) = \prod_{i=1}^n \phi_i(t).$$

Example 1.8.8 Using the results from above the example to show that $\text{Cov}(X_i, X_j) = 0$ for multivariate Gaussian distribution \Rightarrow they are independent.

Sol. Let $\vec{X} \sim \mathcal{N}(\mu, \Sigma)$, where $\vec{X} = (X_1, \dots, X_n) = \vec{\mu} + B\vec{\epsilon}$ and $\vec{\epsilon} = (\epsilon_1, \dots, \epsilon_n)$. (Think about why there is such a linear transformation? since rank of Σ is n .) One has the pdf

$$f(x) = \frac{1}{(2\pi)^{n/2} |\Sigma|^{-1/2}} \exp \left\{ -\frac{1}{2} (x - \vec{\mu})^T \Sigma^{-1} (x - \vec{\mu}) \right\}.$$

Since we have computed that $\phi_{\vec{\epsilon}}(\vec{t}) = E[e^{\vec{t} \cdot \vec{\epsilon}}] = \prod_{i=1}^n e^{t_i^2/2} = e^{\vec{t} \cdot \vec{t}/2}$, so that

$$\begin{aligned} \phi_{\vec{X}}(\vec{t}) &= E[e^{\vec{t} \cdot \vec{X}}] = E[e^{\vec{t} \cdot \vec{\mu} + \vec{t}^T B \vec{\epsilon}}] = e^{\vec{t}^T \vec{\mu} + \frac{1}{2} \vec{t}^T B B^T \vec{t}} = e^{\vec{t}^T \vec{\mu} + \frac{1}{2} \vec{t}^T \Sigma \vec{t}} \\ &= \prod_{i=1}^n e^{t_i \mu_i + \frac{1}{2} t_i^2 \sigma_i^2} = \prod_{i=1}^n \phi_i(t_i). \end{aligned}$$

Example 1.8.9 Let $X \sim N(\mu_1, \sigma_1^2)$ and $Y \sim N(\mu_2, \sigma_2^2)$ and X and Y are independent. Then

$$\phi_{X+Y}(t) = \phi_X(t)\phi_Y(t) = e^{t\mu_1 + \frac{1}{2}t^2\sigma_1^2} e^{t\mu_2 + \frac{1}{2}t^2\sigma_2^2} = e^{(\mu_1+\mu_2)t + \frac{1}{2}(\sigma_1^2+\sigma_2^2)t^2},$$

so that $X + Y \sim N(\mu_1 + \mu_2, \sigma_1^2 + \sigma_2^2)$.

1.9 Limit Theorems

Proposition 1.9.1 (Markov's Inequality) If X is a random variable that takes only nonnegative values, then for any $a > 0$

$$P\{X \geq a\} = \frac{E[X]}{a}.$$

Pf. We give a proof for the case where X is continuous with density f ,

$$\begin{aligned} E[X] &= \int_0^\infty x f(x) dx = \int_0^a x f(x) dx + \int_a^\infty x f(x) dx \\ &\geq \int_a^\infty x f(x) dx \geq a \int_a^\infty f(x) dx = a P\{X \geq a\}. \end{aligned}$$

Proposition 1.9.2 (Chebyshev's Inequality) If X is a random variable with mean μ and variance σ^2 , then, for any $k > 0$,

$$P\{|X - \mu| \geq k\} \leq \frac{\sigma^2}{k^2}.$$

Pf. We apply Markov's inequality to the nonnegative $(X - \mu)^2$,

$$P\{(X - \mu)^2 \geq k^2\} \leq \frac{E[(X - \mu)^2]}{k^2}.$$

Remark 1.9.3 The importance of Markov's and Chebyshev's inequalities is that they enable us to derive bounds on probs. when only the mean, or both the mean and the variance, are known. Of course, if the true distribution were known, then the desired probs. could be exactly computed, and we would not need to resort to bounds.

Theorem 1.9.4 (Strong Law of Large Numbers) Let X_1, X_2, \dots be a sequence of independent random variables have a common distribution, and let $E[X_t] = \mu$. Then, with probability 1, or almost surely,

$$\frac{X_1 + \dots + X_n}{n} \rightarrow \mu \text{ as } n \rightarrow \infty.$$

Definition 1.9.5 If $P(\lim_{n \rightarrow \infty} X_n = X) = 1$, then we say $X_n \rightarrow X$, a.s. (almost surely) or $X_n \rightarrow X$, w.p.1 (with probability 1).

Theorem 1.9.6 (Central Limit Theorem) Let X_1, X_2, \dots be a sequence of independent, identically distributed (i.i.d.) random variables, each with mean μ and variance σ^2 . Then the distribution of

$$\frac{X_1 + \dots + X_n - n\mu}{\sigma\sqrt{n}}$$

goes to the standard normal as $n \rightarrow \infty$. That is,

$$P\left\{\frac{X_1 + \dots + X_n - n\mu}{\sigma\sqrt{n}} \leq a\right\} \rightarrow \frac{1}{\sqrt{2\pi}} \int_{-\infty}^a e^{-x^2/2} dx = \Phi(a),$$

as $n \rightarrow \infty$

Pf. Note that the theorem holds for any distribution of the X_i s; herein lies its power.

We now present a heuristic proof the CLT. Suppose first that the X_i have mean 0 and variance 1, and then the MGF can be computed,

$$E\left[\exp\left\{t\frac{X_1 + \dots + X_n}{\sqrt{n}}\right\}\right] = E[e^{tX_1/\sqrt{n}} \dots e^{tX_n/\sqrt{n}}] = (Ee^{tX_i/\sqrt{n}})^n \text{ by independence.}$$

For large n , we obtain by Taylor expansion,

$$e^{tX_i/\sqrt{n}} = 1 + \frac{tX_i}{\sqrt{n}} + \frac{(tX_i)^2}{2n} + O(n^{-3/2}),$$

that is the reason for the **central** word. Taking expectations shows that when n is large,

$$E[e^{tX_i/\sqrt{n}}] = 1 + \frac{t^2}{2n} + O(n^{-3/2}), \text{ since } EX = 0 \text{ and } EX^2 = 1.$$

Therefore, we obtain

$$E\left[\exp\left\{t\frac{X_1 + \dots + X_n}{\sqrt{n}}\right\}\right] \approx \left(1 + \frac{t^2}{2n}\right)^n \rightarrow e^{t^2/2}.$$

Thus, the MGF of $\frac{X_1 + \dots + X_n}{\sqrt{n}}$ converges to the moment generating function of a standard normal random variable with mean 0 and variance 1. Notice that for $X \sim N(0, 1)$, its MGF $\phi(t) = e^{t^2/2}$. Hence, it can be proven that the distribution function of $\frac{X_1 + \dots + X_n}{\sqrt{n}}$ converges to the distribution function of a standard normal Φ . When X_i have mean μ and variance σ^2 , the random variables $\frac{X_i - \mu}{\sigma}$ have mean 0 and variance 1. Done.

Proposition 1.9.7 The convergence has the following relations:

$$\left. \begin{array}{l} \text{conv. in moments or } L^p \text{ converges} \\ \text{a.s. or w.p.1} \end{array} \right\} \Rightarrow \text{conv. in prob.} \Rightarrow \text{conv. in distribution.}$$

Lemma 1.9.8 (Levy-Crammer) $\{F_n\}$ is a set of distributions. If $\hat{F}_n \rightarrow \phi(t)$ conv. pointwisely, then $F_n \rightarrow F$ converges weakly, where ϕ is the character function of F and \hat{F}_n is the character function of F_n .

Theorem 1.9.9 (Linderberg-Levy CLT) check <https://zhuanlan.zhihu.com/p/69862244>. character function and Lindberg-Levy central limit theorem.

Example 1.9.10 If X is binomially distributed with parameters n and p , then X is the sum of n independent Bernoulli random variables, each with parameter p . Hence, the distribution of

$$\frac{X - E[X]}{\sqrt{Var(X)}} = \frac{\sum X_i - n\mu}{\sqrt{n}\sigma} = \frac{X - np}{\sqrt{np(1-p)}}$$

approaches the standard normal distribution as n approaches ∞ . The normal approximation will be quite good for $np(1-p) \geq 10$ or $\sqrt{Var(X)} \geq \sqrt{10}$.

Example 1.9.11 (Normal approximation to the Binomial) Let X be the number of times that a fair coin, flipped 40 times, lands heads. Find the prob. that $X = 20$.

Sol.

$$\begin{aligned} P\{X = 20\} &= P\{19.5 < X < 20.5\} \\ &= P\left\{\frac{19.5 - 20}{\sqrt{10}} < \frac{X - 20}{\sqrt{10}} < \frac{20.5 - 20}{\sqrt{10}}\right\} \\ &= P\left\{-0.16 < \frac{X - 20}{\sqrt{10}} < 0.16\right\} = \Phi(0.16) - \Phi(-0.16) \\ &= 0.1272. \end{aligned}$$

The exact result is

$$P\{X = 20\} = C_{40}^{20} \left(\frac{1}{2}\right)^{20} \left(\frac{1}{2}\right)^{20} = 0.1268.$$

Example 1.9.12 The lifetime of a battery is a random variable with mean 40 hours and standard deviation 20 hours. Assume a stockpile of 25 such batteries, approximate the probability that over 1100 hours of use can be obtained.

Sol.

$$\begin{aligned} P\{X_1 + \dots + X_{25} > 1100\} &= P\left\{\frac{X_1 + \dots + X_{25} - 25 \times 40}{20\sqrt{25}} > \frac{1100 - 25 \times 40}{20\sqrt{25}}\right\} \\ &= P\{N(0,1) > 1\} = 1 - \Phi(1) = 0.1587. \end{aligned}$$

Chapter 2

Conditional Probability

2.1 The Discrete Case

It is natural to define the conditional prob. mass function of X given that $Y = y$, by

$$p_{X|Y}(x|y) = \frac{p(x,y)}{p_Y(y)},$$

for all values of y such that $P\{Y = y\} > 0$. The conditional expectation of X given that $Y = y$ is defined by

$$E[X|Y = y] = \sum_x x p_{X|Y}(x|y).$$

Example 2.1.1 Suppose that the joint prob. mass function of X and Y is given by

$$p(1,1) = 0.5, p(1,2) = 0.1, p(2,1) = 0.1, p(2,2) = 0.3.$$

Calculate the conditional prob. mass function of X given that $Y = 1$.

Sol. We first compute

$$p_Y(1) = \sum_x p(x,1) = p(1,1) + p(2,1) = 0.6.$$

Hence

$$p_{X|Y}(1|1) = \frac{p(1,1)}{p_Y(1)} = \frac{5}{6}, p_{X|Y}(2|1) = \frac{p(2,1)}{p_Y(1)} = \frac{1}{6}.$$

Example 2.1.2 If X_1 and X_2 are independent binomial random variables with respective parameters (n_1, p) and (n_2, p) , calculate the conditional prob. mass function of X_1 given that $X_1 + X_2 = m$.

Sol. We first compute

$$\begin{aligned} P\{X_1 = k | X_1 + X_2 = m\} &= \frac{P\{X_1 = k, X_2 = m - k\}}{P\{X_1 + X_2 = m\}} \\ &= \frac{C_{n_1}^k p^k q^{n_1-k} C_{n_2}^{m-k} p^{m-k} q^{n_2-m+k}}{C_{n_1+n_2}^m p^m q^{n_1+n_2-m}} \\ &= \frac{C_{n_1}^k C_{n_2}^{m-k}}{C_{n_1+n_2}^m}. \end{aligned}$$

where we used $X_1 + X_2$ is a binomial with parameters $(n_1 + n_2, p)$.

Example 2.1.3 Three possible outcomes with prob. p_i satisfying $p_1 + p_2 + p_3 = 1$. Suppose that n independent replications are performed, and let X_i be the number of times outcome i occurs. Determine the conditional expectation of X_1 given that $X_2 = m$.

Sol. For $k \leq n - m$,

$$\begin{aligned} P\{X_1 = k | X_2 = m\} &= \frac{P\{X_1 = k, X_2 = m\}}{P\{X_2 = m\}} \\ &= \frac{P\{X_1 = k, X_2 = m, X_3 = n - m - k\}}{P\{X_2 = m, X_1 + X_3 = n - m\}} \\ &= \frac{\frac{n!}{k!m!(n-m-k)!} p_1^k p_2^m p_3^{n-m-k}}{\frac{n!}{m!(n-m)!} p_2^m (1-p_2)^{n-m}} \\ &= \frac{(n-m)!}{k!(n-m-k)!} \left(\frac{p_1}{1-p_2}\right)^k \left(\frac{p_3}{1-p_2}\right)^{n-m-k}, \end{aligned}$$

which is a Binomial with parameters $n - m$ and $\frac{p_1}{1-p_2}$. Thus, the conditional expectation is

$$E[X_1 | X_2 = m] = (n - m) \frac{p_1}{1 - p_2}$$

2.2 The Continuous Case

If X and Y have a joint density function $f(x, y)$, then the conditional prob. density function of X given that $Y = y$, is defined by

$$f_{X|Y}(x|y) = \frac{f(x, y)}{f_Y(y)},$$

for all values of y such that $f_Y(y) > 0$. The conditional expectation of X given that $Y = y$ is defined by

$$E[X | Y = y] = \int_{-\infty}^{\infty} x f_{X|Y}(x|y) dx.$$

Example 2.2.1 Suppose that the joint density of X and Y is given by

$$f(x, y) = \begin{cases} 4y(x-y)e^{-(x+y)}, & 0 < x < \infty, 0 \leq y \leq x, \\ 0, & \text{otherwise.} \end{cases}$$

Compute $E[X | Y = y]$.

Sol. The conditional density of X , given that $Y = y$, is given by

$$\begin{aligned} f_{X|Y}(x|y) &= \frac{f(x, y)}{f_Y(y)} = \frac{4y(x-y)e^{-(x+y)}}{\int_y^\infty 4w(w-y)e^{-(w+y)} dw}, x > y \\ &= \frac{(x-y)e^{-x}}{\int_y^\infty (w-y)e^{-w} dw} = \frac{(x-y)e^{-x}}{\int_0^\infty we^{-(y+w)} dw}, w = x - y \\ &= (x-y)e^{-(x-y)} \end{aligned}$$

where we used that $w \sim \mathcal{E}(1)$ and $\int_0^\infty we^{-w} dw$ is the expected value of an exponential random variable with mean 1. Thus, using $EW = 1$, $Var(W) = 1$,

$$\begin{aligned} E[X | Y = y] &= \int_y^\infty x(x-y)e^{-(x-y)} dx = \int_0^\infty (w+y)we^{-w} dw \\ &= E[W^2] + yE[W] = Var(W) + (EW)^2 + y = 2 + y. \end{aligned}$$

Example 2.2.2 The joint density of X and Y is

$$f(x, y) = \begin{cases} \frac{1}{2}ye^{-xy}, & 0 < x < \infty, 0 < y < 2, \\ 0, & \text{otherwise.} \end{cases}$$

What is $E[e^{X/2}|Y = 1]$?

Sol. The conditional density of X , given that $Y = 1$, is given by

$$f_{X|Y}(x|1) = \frac{f(x, 1)}{f_Y(1)} = \frac{\frac{1}{2}e^{-x}}{\int_0^\infty \frac{1}{2}e^{-x}dx} = e^{-x}.$$

Hence

$$E[e^{X/2}|Y = 1] = \int_0^\infty e^{x/2} f_{X|Y}(x|1) dx = \int_0^\infty e^{-x/2} dx = 2.$$

2.3 Computing Expectations by Conditioning

Double expectation formula is very important:

$$E[X] = E[E[X|Y]],$$

The conditional expectation's expectation is unconditional expectation. If Y is discrete, then

$$E[X] = \sum_y E[X|Y = y]P(Y = y).$$

If Y is continuous with density $f_Y(y)$, then

$$E[X] = \int_{-\infty}^{\infty} E[X|Y = y]f_Y(y)dy.$$

Proof for **both discrete X and Y** .

$$\begin{aligned} \sum_y E[X|Y = y]P(Y = y) &= \sum_{x,y} xP[X|Y = y]P(Y = y) \\ &= \sum_{x,y} xP[X = x|Y = y]P(Y = y) = \sum_{x,y} xP[X = x, Y = y] \\ &= \sum_x xP[X = x] = E[X]. \end{aligned}$$

Proof for **both continuous X and Y** .

$$\begin{aligned} \int_{-\infty}^{\infty} E[X|Y = y]f_Y(y)dy &= \int_{-\infty}^{\infty} dx \int_{-\infty}^{\infty} xf_{X|Y}(x|y)f_Y(y)dy \\ &= \int_{-\infty}^{\infty} dx \int_{-\infty}^{\infty} xf(x, y)dy = \int_{-\infty}^{\infty} xf_X(x)dx = E[X]. \end{aligned}$$

Example 2.3.1 Sam will choose either prob. book or history book with equal probability. If the number of misprints in prob. book is Poisson distributed with mean 2 and if the number of misprints in history book is Poisson distributed with mean 5. What is the expected number of misprints that Sam will come across?

Sol. Let X be the number of misprints and

$$Y = \begin{cases} 1, & \text{history book} \\ 2, & \text{prob. book} \end{cases}$$

then

$$\begin{aligned} EX &= E[X|Y=1]P\{Y=1\} + E[X|Y=2]P\{Y=2\} \\ &= 5\left(\frac{1}{2}\right) + 2\left(\frac{1}{2}\right) = 7/2. \end{aligned}$$

Example 2.3.2 (The Expectation of the Sum of a Random Number of Random Variables) Suppose that an industrial plant has the expected number of four accidents. Also suppose that the number of workers injured in each accident are independent random variables with a common mean 2. What is the expected number of injuries during a week?

Sol. Let N be the number of accidents and X_i be the number injured in the i th accident. Then the total number of injuries is $\sum_{i=1}^N X_i$

$$E\left[\sum_{i=1}^N X_i\right] = E\left[E\left[\sum_{i=1}^N X_i | N\right]\right].$$

The enclosed quantity can be computed by

$$E\left[\sum_{i=1}^N X_i | N = n\right] = E\left[\sum_{i=1}^n X_i\right] = nEX_i \Rightarrow E\left[\sum_{i=1}^N X_i | N\right] = NE[X_i].$$

Thus

$$E\left[\sum_{i=1}^N X_i\right] = E[NE[X_i]] = E[N]E[X_i] = 4 \times 2 = 8.$$

Definition 2.3.3 The random number N is independent of i.i.d. random variables X_i , then $\sum_{i=1}^N X_i$ is said to be a **compound random variable**.

Example 2.3.4 (The Mean of a Geometric Distribution) A coin, having prob. p of coming up heads, is to be successively flipped until the first head appears. What is the expected number of flips required?

Sol. Let N be the number of flips required, and let

$$Y = \begin{cases} 1, & \text{if the first flip is a head,} \\ 0, & \text{if the first flip is a tail.} \end{cases}$$

Now,

$$\begin{aligned} E[N] &= E[N|Y=1]P\{Y=1\} + E[N|Y=0]P\{Y=0\} \\ &= pE[N|Y=1] + (1-p)E[N|Y=0]. \end{aligned}$$

Notice that

$$E[N|Y=1] = 1, E[N|Y=0] = E[N] + 1.$$

Then

$$E[N] = p + (1-p)(E[N] + 1) \Rightarrow E[N] = 1/p,$$

which is the same with the mean of a geometric distribution.

Example 2.3.5 A miner is trapped in a mine containing three doors and only one door is out. The 1st door takes him to safety after 2 hours of travel. The 2nd door returns him to the mine after 3 hours of travel. The 3rd door returns him to the mine after 5 hours. Assume that the miner is at all times equally likely to

choose one of the doors, what is the expected length of time until the miner reaches safety?

Sol. Let X denote the time until the miner reaches safety, and let Y denote the door he initially chooses.

$$\begin{aligned} EX &= E[X|Y=1]P\{Y=1\} + E[X|Y=2]P\{Y=2\} + E[X|Y=3]P\{Y=3\} \\ &= \frac{1}{3}(E[X|Y=1] + E[X|Y=2] + E[X|Y=3]). \end{aligned}$$

Since $E[X|Y=1] = 2, E[X|Y=2] = 3 + EX, E[X|Y=3] = 5 + EX$, hence

$$EX = \frac{1}{3}(10 + 2EX) \Rightarrow E[X] = 10.$$

Example 2.3.6 Independent trials, each success with prob. p , are performed until there k consecutive successes. What is the mean number of necessary trials?

Sol. Let N_k denote the number necessary trials to obtain k consecutive successes, and let $M_k = EN_k$. Then compute by conditioning on N_{k-1} ,

$$M_k = E[N_k] = E[E[N_k|N_{k-1}]].$$

Notice that

$$E[N_k|N_{k-1}] = N_{k-1} + \{p \cdot 1 + (1-p)(1 + E[N_k])\}.$$

Taking expectations at both sides,

$$M_k = M_{k-1} + 1 + (1-p)M_k \Rightarrow M_k = \frac{1}{p} + \frac{M_{k-1}}{p}.$$

Since N_1 is geometric with parameter 1, then $M_1 = \frac{1}{p}$, and recursively

$$M_2 = \frac{1}{p} + \frac{1}{p^2},$$

In general,

$$M_k = \frac{1}{p} + \dots + \frac{1}{p^k}.$$

Another way to use conditioning is to obtain the variance of a random variable by applying the conditional variance formula. The conditional variance is defined by

$$\begin{aligned} Var(X|Y=y) &= E[(X - E[X|Y=y])^2|Y=y] \\ &= E[X^2|Y=y] - (E[X|Y=y])^2. \end{aligned}$$

Proposition 2.3.7 (The Conditional Variance Formula)

$$Var(X) = E[Var(X|Y)] + Var(E[X|Y]).$$

Pf.

$$\begin{aligned} E[Var(X|Y)] &= E\{E[X^2|Y] - (E[X|Y])^2\} = E\{E[X^2|Y]\} - E\{(E[X|Y])^2\} \\ &= E[X^2] - E\{(E[X|Y])^2\}. \end{aligned}$$

$$Var(E[X|Y]) = E\{(E[X|Y])^2\} - (E\{E[X|Y]\})^2 = E\{(E[X|Y])^2\} - (E[X])^2.$$

Thus

$$E[Var(X|Y)] + Var(E[X|Y]) = E[X^2] - (E[X])^2 = Var(X).$$

Example 2.3.8 (The Variance of a Compound Random Variable) Let X_1, X_2, \dots be i.i.d. random variables with mean μ and variance σ^2 . Assume they are independent of the nonnegative random integer N . The random variable $S = \sum_{i=1}^N X_i$ is a compound random variable. Find its variance.

Sol. Let us use the conditional variance formula. Compute

$$\text{Var}(S|N=n) = \text{Var}\left(\sum_{i=1}^N X_i | N=n\right) = \text{Var}\left(\sum_{i=1}^n X_i\right) = n\sigma^2.$$

$$E(S|N=n) = E\left(\sum_{i=1}^N X_i | N=n\right) = E\left(\sum_{i=1}^n X_i\right) = n\mu.$$

Therefore,

$$\text{Var}(S|N) = n\sigma^2, \quad E(S|N) = n\mu.$$

Using conditional variance formula,

$$\begin{aligned} \text{Var}(S) &= E[\text{Var}(S|N)] + \text{Var}[E(S|N)] = E[N]\sigma^2 + \text{Var}(N\mu) \\ &= \sigma^2 E[N] + \mu^2 \text{Var}(N). \end{aligned}$$

Another way is to compute the variance directly.

$$\begin{aligned} \text{Var}(S) &= ES^2 - (ES)^2, \\ ES &= E[E[S|N]] = \mu E[N], \\ ES^2 &= E[E[S^2|N]] = E[N^2\mu^2 + N\sigma^2] = \mu^2 E[N^2] + \sigma^2 E[N], \end{aligned}$$

where

$$\begin{aligned} E[S^2|N] &= E[(\sum_{i=1}^N X_i)^2] = E[\sum_{i=1}^N X_i^2 + 2 \sum_{i < j} X_i X_j] \\ &= NEX_i^2 + (N^2 - N)EX_i EX_j = N(\mu^2 + \sigma^2) + (N^2 - N)\mu^2 \\ &= N^2\mu^2 + N\sigma^2. \end{aligned}$$

Therefore,

$$\begin{aligned} \text{Var}(S) &= ES^2 - (ES)^2 = \mu^2 E[N^2] + \sigma^2 E[N] - (\mu E[N])^2 \\ &= \mu^2 \text{Var}(N) + \sigma^2 E[N]. \end{aligned}$$

Example 2.3.9 If N is a Poisson random variable, then $S = \sum_{i=1}^N X_i$ is called a **compound Poisson random variable**. Since

$$E[N] = \text{Var}(N) = \lambda,$$

then

$$\text{Var}(S) = \mu^2 \text{Var}(N) + \sigma^2 E[N] = \mu^2 \lambda + \sigma^2 \lambda = \lambda E[X^2].$$

We should point out that **the following derivation is wrong**:

$$\text{Var}(S) = \text{Var}\left(\sum_{i=1}^N X_i\right) = E[N]\text{Var}(X_i) = \lambda \text{Var}(X).$$

2.4 Computing Probabilities by Conditioning

We may also use conditioning approach to compute probabilities. Let E denote an arbitrary event and define the indicator random variable X by

$$X = \begin{cases} 1, & \text{if } E \text{ occurs,} \\ 0, & \text{if } E \text{ does not occur.} \end{cases}$$

It follows from the definition of X that

$$\begin{aligned} E[X] &= 1 \cdot P(E) + 0 \cdot P(E^c) = P(E), \\ E[X|Y=y] &= 1 \cdot P(E|Y=y) + 0 \cdot P(E^c|Y=y) = P(E|Y=y), \end{aligned}$$

for any random variable Y . Then from double expectation formula,

$$P(E) = \sum_y P(E|Y=y)P(Y=y), \quad \text{if } Y \text{ is discrete,}$$

$$\int_{-\infty}^{\infty} P(E|Y=y)f_Y(y)dy, \quad \text{if } Y \text{ is continuous.}$$

Example 2.4.1 Suppose X and Y are independent continuous random variables having densities f_X and f_Y . Compute $P\{X < Y\}$.

Sol. Conditioning on the value of Y yields

$$\begin{aligned} P\{X < Y\} &= \int_{-\infty}^{\infty} P(X < Y|Y=y)f_Y(y)dy \\ &= \int_{-\infty}^{\infty} P(X < y)f_Y(y)dy = \int_{-\infty}^{\infty} F_X(y)f_Y(y)dy \\ &= \int_{-\infty}^{\infty} dy \int_{-\infty}^y f_X(x)f_Y(y)dx. \end{aligned}$$

Example 2.4.2 An insurance company supposes that (1) each policyholder has number of accidents with Poisson distributed, with the random mean. (2) The mean of Poisson has a gamma distribution with density

$$g(\lambda) = \lambda e^{-\lambda}, \quad \lambda \geq 0.$$

What is the prob. that a randomly chosen policyholder has exactly n accidents next year?

Sol. Let X denote the number of accidents that a randomly chosen policyholder has next year. Let Y be the Poisson mean number of accidents for this policyholder. Then, conditioning on Y yields

$$\begin{aligned} P\{X = n\} &= \int_0^{\infty} P\{X = n|Y = \lambda\}g(\lambda)d\lambda \\ &= \int_0^{\infty} e^{-\lambda} \frac{\lambda^n}{n!} \lambda e^{-\lambda} d\lambda = \int_0^{\infty} \frac{1}{n!} \lambda^{n+1} e^{-2\lambda} d\lambda. \end{aligned}$$

By induction,

$$\begin{aligned} P\{X = 0\} &= \int_0^{\infty} \frac{1}{0!} \lambda^1 e^{-2\lambda} d\lambda = -\frac{1}{2} \int_0^{\infty} \lambda d e^{-2\lambda} \\ &= -\frac{1}{2} \left(\lambda e^{-2\lambda} \Big|_0^{\infty} - \int_0^{\infty} e^{-2\lambda} d\lambda \right) = \frac{1}{2} \int_0^{\infty} e^{-2\lambda} d\lambda \\ &= -\frac{1}{4} e^{-2\lambda} \Big|_0^{\infty} = \frac{1}{4}. \end{aligned}$$

$$\begin{aligned}
P\{X = 1\} &= \int_0^\infty \lambda^2 e^{-2\lambda} d\lambda = -\frac{1}{2} \int_0^\infty \lambda^2 de^{-2\lambda} \\
&= -\frac{1}{2} \left(\lambda^2 e^{-2\lambda} \Big|_0^\infty - 2 \int_0^\infty \lambda e^{-2\lambda} d\lambda \right) = \int_0^\infty \lambda e^{-2\lambda} d\lambda = \frac{1}{4} = \frac{2}{8}.
\end{aligned}$$

$$\begin{aligned}
P\{X = 2\} &= \frac{1}{2} \int_0^\infty \lambda^3 e^{-2\lambda} d\lambda = -\frac{1}{4} \int_0^\infty \lambda^3 de^{-2\lambda} \\
&= -\frac{1}{4} \left(\lambda^3 e^{-2\lambda} \Big|_0^\infty - 3 \int_0^\infty \lambda^2 e^{-2\lambda} d\lambda \right) = \frac{3}{4} \int_0^\infty \lambda^2 e^{-2\lambda} d\lambda = \frac{3}{16}.
\end{aligned}$$

$$\begin{aligned}
P\{X = 3\} &= \frac{1}{6} \int_0^\infty \lambda^4 e^{-2\lambda} d\lambda = -\frac{1}{12} \int_0^\infty \lambda^4 de^{-2\lambda} \\
&= -\frac{1}{12} \left(\lambda^4 e^{-2\lambda} \Big|_0^\infty - 4 \int_0^\infty \lambda^3 e^{-2\lambda} d\lambda \right) = \frac{1}{3} \int_0^\infty \lambda^3 e^{-2\lambda} d\lambda = \frac{1}{3} \frac{3}{8} = \frac{1}{8} = \frac{4}{32}.
\end{aligned}$$

By induction, we can show that $P\{X = n\} = \frac{n+1}{2^{n+2}}$. That is,

$$\begin{aligned}
P\{X = n\} &= \frac{1}{n!} \int_0^\infty \lambda^{n+1} e^{-2\lambda} d\lambda = -\frac{1}{2n!} \int_0^\infty \lambda^{n+1} de^{-2\lambda} \\
&= -\frac{1}{2n!} \left(\lambda^{n+1} e^{-2\lambda} \Big|_0^\infty - (n+1) \int_0^\infty \lambda^n e^{-2\lambda} d\lambda \right) = \frac{n+1}{2n} \int_0^\infty \frac{1}{(n-1)!} \lambda^n e^{-2\lambda} d\lambda \\
&= \frac{n+1}{2n} \frac{n}{2^{n+1}} = \frac{n+1}{2^{n+2}}.
\end{aligned}$$

Another way is to notice that

$$h(\lambda) = \frac{2e^{-2\lambda}(2\lambda)^{n+1}}{(n+1)!}$$

is the density function of a gamma $(n+2, 2)$ random variable, its integral is 1. Therefore,

$$1 = \int_0^\infty \frac{2e^{-2\lambda}(2\lambda)^{n+1}}{(n+1)!} d\lambda = \frac{2^{n+2}}{n+1} \int_0^\infty \frac{1}{n!} \lambda^{n+1} e^{-2\lambda} d\lambda$$

Thus

$$P\{X = n\} = \int_0^\infty \frac{1}{n!} \lambda^{n+1} e^{-2\lambda} d\lambda = \frac{n+1}{2^{n+2}}.$$

Example 2.4.3 Suppose that the number of people who visit a yoga studio each day is a Poisson random variable with mean λ . Suppose further that each person is independently female with prob. p or male with prob. $1-p$. Find the joint prob. that exactly n women and m men visit the academy today.

Sol. Let N_1 denote the number of women and N_2 the number of men, so that $N = N_1 + N_2$ is the total number of people. Conditioning on N gives

$$\begin{aligned}
P\{N_1 = n, N_2 = m\} &= \sum_{i=0}^{\infty} P\{N_1 = n, N_2 = m | N = i\} P\{N = i\} \\
&= P\{N_1 = n, N_2 = m | N = n+m\} P\{N = n+m\} \\
&= P\{N_1 = n, N_2 = m | N = n+m\} e^{-\lambda} \frac{\lambda^{n+m}}{(n+m)!}, \text{ Poisson} \\
&= C_{n+m}^n p^n (1-p)^m e^{-\lambda} \frac{\lambda^{n+m}}{(n+m)!}, \text{ Binomial} \\
&= e^{-\lambda p} \frac{(\lambda p)^n}{n!} e^{-\lambda(1-p)} \frac{(\lambda(1-p))^m}{m!}.
\end{aligned}$$

Since

$$\begin{aligned} P\{N_1 = n\} &= \sum_{m=0}^{\infty} P\{N_1 = n, N_2 = m\} = e^{-\lambda p} \frac{(\lambda p)^n}{n!}, \\ P\{N_2 = m\} &= \sum_{n=0}^{\infty} P\{N_1 = n, N_2 = m\} = e^{-\lambda(1-p)} \frac{(\lambda(1-p))^m}{m!}. \end{aligned}$$

so that $P\{N_1 = n, N_2 = m\} = P\{N_1 = n\}P\{N_2 = m\}$ and we can conclude that N_1 and N_2 are independent Poisson random variables with means λp and $\lambda(1 - p)$. Poisson diversion and confluence.

Example 2.4.4 (The Ballot Problem) (投票问题) In an election, candidate A receives n votes, and candidate B receives m votes where $n > m$. Assuming that all orderings are equally likely, show that the prob. that A is always ahead in the count of votes is $(n - m)/(n + m)$.

Sol. Let $P_{n,m}$ denote the desired prob. By conditioning on which candidate receives the last vote, we have

$$\begin{aligned} P_{n,m} &= P\{A \text{ always ahead} | A \text{ receives last vote}\} \frac{n}{n+m} \\ &\quad + P\{A \text{ always ahead} | B \text{ receives last vote}\} \frac{m}{n+m}. \end{aligned}$$

Notice that

$$\begin{aligned} P\{A \text{ always ahead} | A \text{ receives last vote}\} &= P_{n-1,m}, \\ P\{A \text{ always ahead} | B \text{ receives last vote}\} &= P_{n,m-1}. \end{aligned}$$

Then one obtains the recursive formula,

$$P_{n,m} = P_{n-1,m} \frac{n}{n+m} + P_{n,m-1} \frac{m}{n+m}, \quad \text{as } n-1 \geq m.$$

Notice that

$$\begin{aligned} P_{1,0} &= 1, P_{2,0} = 1, \dots, P_{n,0} = 1, \\ P_{1,1} &= 0, P_{2,2} = 0, \dots, P_{n,n} = 0. \end{aligned}$$

We now can prove the result by induction

$$\begin{aligned} P_{n,m} &= \frac{n-1-m}{n-1+m} \frac{n}{n+m} + \frac{n-m+1}{n+m-1} \frac{m}{n+m} \\ &= \frac{n^2 - n - mn + mn - m^2 + m}{(n-1+m)(n+m)} = \frac{(n-m)(n+m-1)}{(n-1+m)(n+m)} \\ &= \frac{n-m}{n+m}. \end{aligned}$$

2.5 Computing Conditional Expectation and Conditional Probability by Conditioning

The analog of

$$E[X] = \begin{cases} \sum_y E[X|Y=y]P(Y=y), & \text{if } Y \text{ is discrete,} \\ \int_{-\infty}^{\infty} E[X|Y=y]f_Y(y)dy, & \text{if } Y \text{ is continuous.} \end{cases}$$

is

$$E[X|Y = y] = \begin{cases} \sum_w E[X|Y = y, W = w]P(W = w|Y = y), & \text{if } W \text{ is discrete,} \\ \int_{-\infty}^{\infty} E[X|Y = y, W = w]f_{W|Y}(w|y)dw, & \text{if } W \text{ is continuous.} \end{cases}$$

That is, in another form,

$$E[X|Y] = E[E[X|Y, W]|Y].$$

Here is a proof for the above formula. For discrete W ,

$$\begin{aligned} \sum_w E[X|Y = y, W = w]P(W = w|Y = y) &= \sum_w \sum_x xP[X = x|Y = y, W = w]P(W = w|Y = y) \\ &= \sum_w \sum_x \frac{xP[X = x, Y = y, W = w]}{P\{Y = y, W = w\}} \frac{P(W = w, Y = y)}{P(Y = y)} \\ &= \sum_x \frac{xP[X = x, Y = y]}{P(Y = y)} = \sum_x xP[X = x|Y = y] = E[X|Y = y]. \end{aligned}$$

For continuous W ,

$$\begin{aligned} \int_w E[X|Y = y, W = w]f_{W|Y}(w|y)dw &= \int_w \int_x x f_{X|Y,W}(x|y, w) f_{W|Y}(w|y) dw dx \\ &= \int_w \int_x x \frac{f_{X,Y,W}(x, y, w)}{f_{Y,W}(y, w)} \frac{f_{Y,W}(y, w)}{f_Y(y)} dw dx \\ &= \int_w \int_x x f_{X,W|Y}(x, w|y) dw dx = \int_x x f_{X|Y}(x|y) dx = E[X|Y = y]. \end{aligned}$$

Example 2.5.1 (1) An insurance company classifies each policyholder as being one of k types. (2) Type i has the number of accidents with Poisson distribution with mean λ_i . (3) A newly policyholder is type i with prob. p_i , $\sum_{i=1}^k p_i = 1$. (4) Given that a policyholder had n accidents in her first year, the question is [1] what is the expected number that she has in her second year? [2] What is the conditional prob. that she has m accidents in her second year?

Sol. Let N_i denote the number of accidents the policyholder has in i th ($i = 1, 2$) year. Conditioning on her risk type T gives

$$\begin{aligned} E[N_2|N_1 = n] &= \sum_{j=1}^k E[N_2|T = j, N_1 = n]P\{T = j|N_1 = n\} \\ &= \sum_{j=1}^k E[N_2|T = j]P\{T = j|N_1 = n\} = \sum_{j=1}^k \lambda_j P\{T = j|N_1 = n\}, \end{aligned}$$

where the last follows the Poission mean λ_j . Using Bayes' formula

$$\begin{aligned} P\{T = j|N_1 = n\} &= \frac{P\{T = j, N_1 = n\}}{P\{N_1 = n\}} = \frac{P\{N_1 = n|T = j\}P\{T = j\}}{\sum_{j=1}^k P\{N_1 = n|T = j\}P\{T = j\}} \\ &= \frac{e^{-\lambda_j} \lambda_j^n / n! p_j}{\sum_{j=1}^k e^{-\lambda_j} \lambda_j^n / n! p_j}, \end{aligned}$$

one arrives at

$$\begin{aligned} E[N_2|N_1 = n] &= \sum_{j=1}^k \lambda_j P\{T = j|N_1 = n\} = \sum_{j=1}^k \lambda_j \frac{e^{-\lambda_j} \lambda_j^n / n! p_j}{\sum_{j=1}^k e^{-\lambda_j} \lambda_j^n / n! p_j} \\ &= \frac{\sum_{j=1}^k e^{-\lambda_j} \lambda_j^{n+1} p_j}{\sum_{j=1}^k e^{-\lambda_j} \lambda_j^n p_j}. \end{aligned}$$

The conditional prob. can also be obtained by conditioning on her type:

$$\begin{aligned} P\{N_2 = m | N_1 = n\} &= \sum_{j=1}^k P[N_2 = m | T = j, N_1 = n] P\{T = j | N_1 = n\} \\ &= \sum_{j=1}^k \frac{e^{-\lambda_j} \lambda_j^m}{m!} \frac{e^{-\lambda_j} \lambda_j^n / n! p_j}{\sum_{j=1}^k e^{-\lambda_j} \lambda_j^n / n! p_j} = \frac{\sum_{j=1}^k e^{-2\lambda_j} \lambda_j^{m+n} p_j}{m! \sum_{j=1}^k e^{-\lambda_j} \lambda_j^n p_j}. \end{aligned}$$

The second approach is to calculate the prob. $P\{N_2 = m | N_1 = n\}$ first,

$$\begin{aligned} P\{N_2 = m | N_1 = n\} &= \frac{P\{N_2 = m, N_1 = n\}}{P\{N_1 = n\}} \\ &= \frac{\sum_{j=1}^k P\{N_2 = m, N_1 = n | T = j\} p_j}{\sum_{j=1}^k P\{N_1 = n | T = j\} p_j} = \frac{\sum_{j=1}^k \frac{e^{-\lambda_j} \lambda_j^m}{m!} \frac{e^{-\lambda_j} \lambda_j^n}{n!} p_j}{\sum_{j=1}^k \frac{e^{-\lambda_j} \lambda_j^n}{n!} p_j} \\ &= \frac{1}{m!} \frac{\sum_{j=1}^k e^{-2\lambda_j} \lambda_j^{m+n} p_j}{\sum_{j=1}^k e^{-\lambda_j} \lambda_j^n p_j}. \end{aligned}$$

Then conditional expectation can be calculated by

$$\begin{aligned} E\{N_2 | N_1 = n\} &= \sum_{m=0}^{\infty} m P\{N_2 = m | N_1 = n\} = \sum_{m=1}^{\infty} \frac{m}{m!} \frac{\sum_{j=1}^k e^{-2\lambda_j} \lambda_j^{m+n} p_j}{\sum_{j=1}^k e^{-\lambda_j} \lambda_j^n p_j} \\ &= \frac{\sum_{j=1}^k \lambda_j^1 e^{-\lambda_j} \lambda_j^n p_j}{\sum_{j=1}^k e^{-\lambda_j} \lambda_j^n p_j} = \frac{\sum_{j=1}^k e^{-\lambda_j} \lambda_j^{n+1} p_j}{\sum_{j=1}^k e^{-\lambda_j} \lambda_j^n p_j}. \end{aligned}$$

where made use of $\sum_{m=1}^{\infty} m \frac{e^{-\lambda_j} \lambda_j^m}{m!} = \lambda_j$.

Chapter 3

Markov Chain

3.1 Introduction

- Stochastic Process $\{X(t), t \in T\}$ is a collection of random variables, where T is a subset of $(-\infty, \infty)$. For each $t \in T$, $X(t)$ is a random variable.
- There exists a $\omega \in \Omega$, s.t. $x_t = X_t(\omega) = X(t, \omega)$, $t \in T$. Then, $\{x_t | t \in T\}$ is an observation or a realization of $X_t | t \in T$.
- t is the time, either discrete or continuous.
- $X(t)$ is the state at time t . For example, (1) number of customers in a supermarket (2) number of accidents on a highway.
- T is the index set of the process.
 - (1) T is finite or countable \rightarrow discrete-time process or stochastic sequence.
 - (2) T is an interval of a real line \rightarrow continuous-time process.
- $\{X_n, n = 0, 1, 2, \dots\}$ is a discrete-time SP (Stochastic Process). $\{X(t), t \geq 0\}$ is a continuous-time SP.
- State space of a SP is the set of all possible values that $X(t)$ can take, either subset of \mathbb{R}^m or discrete spaces like \mathbb{Z}^m .

continuous
 $\left\{ \begin{array}{l} \text{temperature every day} \\ \text{the time spent for study} \end{array} \right.$
discrete
 $\left\{ \begin{array}{l} \text{number of customers in a supermarket} \\ \text{machine/engine works or not, 0 or 1.} \end{array} \right.$

- State space. $I = \{0, 1, 2, \dots\}$ finite or countable

$X_n = i$, in state i at time n .

- $\{X_n\}_{n=0}^{\infty}$ is a time-dependent process. X_{n+1} may depend on X_n and even earlier.

Example 3.1.1 The followings are Stochastic Processes. (1) Brownian Motion, pollen (花粉), Jean Perrin. (2) Poisson Process which is a counting process. (3) The meters one walked every day. (4) number of students in the classroom.

Definition 3.1.2 The following stochastic process is known as a **homogeneous Markov Chain**:

$$\begin{aligned} P(X_{n+1} = j | X_n = i, X_{n-1} = i_{n-1}, \dots, X_0 = i_0) &= P(X_{n+1} = j | X_n = i) \\ &= P(X_1 = j | X_0 = i) := p_{ij}, \quad i, j \in I, \end{aligned}$$

where the time-independent p_{ij} is the **one-step transition probability** (一步转移概率), and $\mathbf{P} = (p_{ij})_{i,j \in I}$ is the **one-step transition prob. matrix** (转移概率矩阵). Obviously, one has $p_{ij} \geq 0$ and by mutually exclusive property,

$$\sum_{j=0}^{\infty} p_{ij} = \sum_{j=0}^{\infty} P(X_1 = j | X_0 = i) = P\left(\bigcup_{j=0}^{\infty} \{X_1 = j\} | X_0 = i\right) = 1.$$

The row sum of a stochastic matrix (随机矩阵) is 1:

$$P = \begin{bmatrix} p_{00} & p_{01} & p_{02} & \cdots \\ p_{10} & p_{11} & p_{12} & \cdots \\ p_{20} & p_{21} & p_{22} & \cdots \\ \vdots & \vdots & \vdots & \ddots \end{bmatrix}.$$

Definition 3.1.3 The **nonhomogeneous Markov Chain** (非时齐) still has the Markovian property. In non homogeneous chains, transition probabilities can vary across time. That is to say, the Markov property is retained but the transition probabilities may depend on time.

$$P(X_{n+1} = j | X_n = i) \neq P(X_1 = j | X_0 = i),$$

or $p_{ij}(n)$ depends on time n or $p_{ij}(t)$ depends on time t .

Example 3.1.4 X_n depends on $X_{n-1}, X_{n-2}, \dots, X_{n-p}$. Markov Process, AR(1) model,

$$X_n = a_1 X_{n-1} + \varepsilon.$$

AR(p) model ($p \geq 2$). time series. non-Markov process.

$$X_n = a_1 X_{n-1} + \cdots + a_p X_{n-p} + \varepsilon.$$

Example 3.1.5 (Forecasting the weather) Suppose that if it rains today, then it will rain tomorrow with prob. α ; and if it does not rain today, then it will rain tomorrow with prob. β . Find the transition prob. matrix.

Sol. Let

$$\begin{aligned} \text{state 1,} &\quad \text{rain,} \\ \text{state 2,} &\quad \text{not rain.} \end{aligned}$$

Then

$$\mathbf{P} = \begin{bmatrix} \alpha & 1 - \alpha \\ \beta & 1 - \beta \end{bmatrix}.$$

Example 3.1.6 (Transform a non-Markov process to a Markov Chain) Suppose that whether it rains today depends on the last two days. Suppose that it has rained for the past two days, then it will rain tomorrow with prob. 0.7; if it rained today but not yesterday, then it will rain tomorrow with prob. 0.5; if it rained

yesterday but not today, then it will rain tomorrow with prob. 0.4; if it has not rained in the past two days, then it will rain tomorrow with prob. 0.2. Give the transition prob. matrix.

Sol. The condition is

	yesterday	today	tomorrow
✓	✓	✓ = 0.7, × = 0.3	
×	✓	✓ = 0.5, × = 0.5	.
✓	×	✓ = 0.4, × = 0.6	
×	×	✓ = 0.2, × = 0.8	

Let

state 0,	(rain,rain)
state 1,	(not,rain)
state 2,	(rain,not)
state 3,	(not,not)

The four-state Markov Chain has a transition prob. matrix,

$$\mathbf{P} = \begin{bmatrix} 0.7 & 0 & 0.3 & 0 \\ 0.5 & 0 & 0.5 & 0 \\ 0 & 0.4 & 0 & 0.6 \\ 0 & 0.2 & 0 & 0.8 \end{bmatrix}.$$

Example 3.1.7 (A random walk model) (随机游走模型) A Markov chain whose state space is given by the integers $i = 0, \pm 1, \pm 2, \dots$ is said to be a random walk if, for some $0 < p < 1$,

$$p_{i,i+1} = p, \quad p_{i,i-1} = 1 - p, \quad i = 0, \pm 1, \pm 2, \dots$$

Note that this is an infinite-state chain.

Example 3.1.8 (A Gambling model) Consider a gambler who, at each play, either wins 1 dollar with prob p or loses 1 dollar with prob. $1 - p$. Suppose that the gambler quits playing either when he goes broke or he attains a fortune of N dollars. Then the Markov chain has the transition prob.

$$\begin{aligned} p_{i,i+1} &= p, \quad p_{i,i-1} = 1 - p, \quad i = 1, 2, \dots, N-1, \\ p_{00} &= p_{NN} = 1. \end{aligned}$$

State 0 and N are called absorbing states (吸收态, 吸收壁) since once entered they are never left. Note that this is a finite-state chain with absorbing barriers.

Proposition 3.1.9 I is the state space of the homogeneous Markov Chain $\{X_n\}$. $A, A_j \subset I$.

- (1) When $X_n = i$ is given, the future $\{X_m : m \geq n + 1\}$ is independent of the past $\{X_j : j \leq n - 1\}$.
- (2) $P(X_{n+k} = j | X_n = i) = P(X_k = j | X_0 = i)$.
- (3) $P(X_{n+k} = j | X_n = i, X_{n-1} \in A_{n-1}, \dots, X_0 \in A_0) = P(X_k = j | X_0 = i)$.
- (4) $P(X_{n+k} \in A | X_n = i, X_{n-1} \in A_{n-1}, \dots, X_0 \in A_0) = P(X_k \in A | X_0 = i)$.

3.2 Chapman-Kolmogorov Equations

(强调我的记号)

We now define the *n-step transition prob.* $p_{ij}^{(n)}$ to be the prob. that a process in state i will be in state j after n additional transitions. That is

$$p_{ij}^{(n)} = P\{X_{n+k} = j | X_k = i\}, \quad n \geq 0, i, j \geq 0.$$

Obviously, $p_{ij}^{(1)} = p_{ij}$. In particular,

$$p_{ij}^{(0)} = P\{X_0 = j | X_0 = i\} = \begin{cases} 1, & i = j \\ 0, & i \neq j \end{cases},$$

or say $\mathbf{P}^{(1)} = \mathbf{P}$, and $\mathbf{P}^{(0)} = \mathbf{I}$ (identity matrix).

Theorem 3.2.1 (Chapman-Kolmogorov Equation) For any $m, n \geq 0$,

$$\begin{aligned} p_{ij}^{(n+m)} &= \sum_{k \in I} p_{ik}^{(n)} p_{kj}^{(m)}, \\ \mathbf{P}^{(n+m)} &= \mathbf{P}^{n+m}. \end{aligned}$$

Pf.

$$\begin{aligned} p_{ij}^{(n+m)} &= P\{X_{n+m} = j | X_0 = i\} \\ &= \sum_{k \in I} P\{X_{n+m} = j | X_n = k, X_0 = i\} P\{X_n = k | X_0 = i\} \\ &= \sum_{k \in I} p_{ik}^{(n)} p_{kj}^{(m)}. \end{aligned}$$

Corollary 3.2.2 (1) $p_{ij}^{(n+m)} \geq p_{il}^{(n)} p_{lj}^{(m)}$;

(2) $p_{ii}^{(n+k+m)} \geq p_{ij}^{(n)} p_{jl}^{(k)} p_{li}^{(m)}$;

(3) $p_{ii}^{(nk)} \geq (p_{ii}^{(n)})^k$.

Pf.

$$p_{ij}^{(n+m)} = \sum_{k \in I} p_{ik}^{(n)} p_{kj}^{(m)} \geq p_{il}^{(n)} p_{lj}^{(m)}.$$

Example 3.2.3 Consider Example 3.1.5. If $\alpha = 0.7$ and $\beta = 0.4$, then calculate the prob. that it will rain in four days given that it is raining today.

Sol. The one-step transition prob. matrix is given by

$$\mathbf{P} = \begin{bmatrix} 0.7 & 0.3 \\ 0.4 & 0.6 \end{bmatrix}.$$

Hence,

$$\begin{aligned} \mathbf{P}^2 &= \begin{bmatrix} 0.7 & 0.3 \\ 0.4 & 0.6 \end{bmatrix}^2 = \begin{bmatrix} 0.61 & 0.39 \\ 0.52 & 0.48 \end{bmatrix}, \\ \mathbf{P}^4 &= (\mathbf{P}^2)^2 = \begin{bmatrix} 0.5749 & 0.4251 \\ 0.5668 & 0.4332 \end{bmatrix}, \end{aligned}$$

so that the desired prob. $p_{00}^{(4)} = 0.5749$. (notice that row sum is always 1.)

Example 3.2.4 Consider Example 3.1.6. Given that it rained on Monday and Tuesday, what is the prob. that it will rain on Thursday?

Sol. The two-step transition matrix is given by

$$\mathbf{P}^{(2)} = \mathbf{P}^2 = \begin{bmatrix} 0.7 & 0 & 0.3 & 0 \\ 0.5 & 0 & 0.5 & 0 \\ 0 & 0.4 & 0 & 0.6 \\ 0 & 0.2 & 0 & 0.8 \end{bmatrix}^2 = \begin{bmatrix} 0.49 & 0.12 & 0.21 & 0.18 \\ 0.35 & 0.20 & 0.15 & 0.30 \\ 0.20 & 0.12 & 0.20 & 0.48 \\ 0.10 & 0.16 & 0.10 & 0.64 \end{bmatrix}.$$

The desired prob. is $p_{00}^{(2)} + p_{01}^{(2)} = 0.49 + 0.12 = 0.61$. (Notice that row sum is always 1.)

3.3 Unconditional Distribution of the State

So far, all prob. are conditional prob. If the unconditional distribution of the state at time n is desired, it is necessary to specify the prob. distribution of the initial state.

Definition 3.3.1 X_0 has the prob. distribution

$$\pi_j^{(0)} = P(X_0 = j), \quad j \in I = \{1, 2, \dots\}.$$

The initial distribution of $\{X_n\}$ is denoted by

$$\pi^{(0)} = (\pi_1^{(0)}, \pi_2^{(0)}, \dots).$$

Moreover, let the distribution at time n be

$$\begin{aligned} \pi_j^{(n)} &= P(X_n = j), \quad j \in I, \\ \pi^{(n)} &= (\pi_1^{(n)}, \pi_2^{(n)}, \dots). \end{aligned}$$

In fact, $\pi^{(n)}$ can be uniquely determined by the initial distribution $\pi^{(0)}$ and the transition matrix \mathbf{P} .

Theorem 3.3.2 Assume Markov chain $\{X_n\}$ has initial dist. $\pi^{(0)}$ and trans. prob. matrix \mathbf{P} . Then For any $0 \leq n_0 < n_1 < \dots < n_m$,

$$P(X_{n_0} = i_0, X_{n_1} = i_1, \dots, X_{n_m} = i_m) = \pi_{i_0}^{(n_0)} p_{i_0 i_1}^{(n_1 - n_0)} p_{i_1 i_2}^{(n_2 - n_1)} \dots p_{i_{m-1} i_m}^{(n_m - n_{m-1})}.$$

(2) For any $\forall n \geq 1$,

$$\begin{cases} \pi^{(n+1)} = \pi^{(n)} \mathbf{P}, & \text{where } \pi^{(n)} \text{ is a row vector,} \\ \pi^{(n)} = \pi^{(0)} \mathbf{P}^n, & \text{or say } \pi_j^{(n)} = \sum_{i \in I} \pi_i^{(0)} p_{ij}^{(n)}, \\ \pi^{(n)} = \pi^{(k)} \mathbf{P}^{n-k}, & 0 \leq k \leq n. \end{cases}$$

Proof. (1) Using the product formula,

$$P(B_1 B_2 \dots B_n) = P(B_1) P(B_2 | B_1) \dots P(B_n | B_1 B_2 \dots B_{n-1}),$$

one has

$$\begin{aligned} P(X_{n_0} = i_0, X_{n_1} = i_1, \dots, X_{n_m} = i_m) &= P(X_{n_0} = i_0) P(X_{n_1} = i_1 | X_{n_0} = i_0) \dots P(X_{n_m} = i_m | X_{n_{m-1}} = i_{m-1}) \\ &= \pi_{i_0}^{(n_0)} p_{i_0 i_1}^{(n_1 - n_0)} p_{i_1 i_2}^{(n_2 - n_1)} \dots p_{i_{m-1} i_m}^{(n_m - n_{m-1})}. \end{aligned}$$

(2)

$$\begin{aligned}\pi_j^{(n+1)} &= P(X_{n+1} = j) = \sum_{i \in I} P(X_{n+1} = j | X_n = i)P(X_n = i) \\ &= \sum_{i \in I} \pi_i^{(n)} p_{ij},\end{aligned}$$

that is $\pi^{(n+1)} = \pi^{(n)} \mathbf{P}$. ■

Example 3.3.3 An urn always contains 2 balls possibly with red and blue colors. (1) At each stage a ball is randomly chosen. (2) Then the chosen ball is replaced by the same color with prob. 0.8 and opposite color with prob. 0.2. If initially both balls are red, find the prob. that the third ball selected is red.

Sol. Define X_n to be the number of red balls in the urn after the n th section. Then the transition matrix is

$$\mathbf{P} = \begin{bmatrix} 0.8 & 0.2 & 0 \\ 0.1 & 0.8 & 0.1 \\ 0 & 0.2 & 0.8 \end{bmatrix}.$$

Then

$$\mathbf{P}^2 = \begin{bmatrix} 0.66 & 0.32 & 0.02 \\ 0.16 & 0.68 & 0.16 \\ 0.02 & 0.32 & 0.66 \end{bmatrix},$$

and the desired prob. is

$$\begin{aligned}P\{3rd \text{ section is red} | X_0 = 2\} &= \sum_i P\{3rd \text{ section is red} | X_2 = i, X_0 = 2\} P\{X_2 = i | X_0 = 2\} \\ &= (0)p_{20}^{(2)} + (0.5)p_{21}^{(2)} + (1)p_{22}^{(2)} = (0.5)(0.32) + (1)(0.66) \\ &= 0.16 + 0.66 = 0.82.\end{aligned}$$

3.4 Classification of States

See more references in [1, 12].

吸收态 i 通 j i, j 相通或互通 等价类 不可约链

Definition 3.4.1 I is the state space of Markov chain $\{X_n\}$.

- (a) If $p_{ii} = 1$, then i is called an absorbing state.
- (b) If $\exists n \geq 0$, s.t. $p_{ij}^{(n)} > 0$, then i is said to be accessible to state j . (denoted by $i \rightarrow j$)
- (c) If $i \rightarrow j$ and $j \rightarrow i$, then they are said to communicate ($i \leftrightarrow j$).
- (d) Two states that communicate are said to be in the same class.
- (e) The Markov chain is said to be irreducible if there is one class in the chain, that is, all states communicate with each other.

Remark 3.4.2 If i not accessible to j , then

$$\begin{aligned}P(\text{ever enter } j | \text{start in } i) &= P\left(\bigcup_{n=0}^{\infty} \{X_n = j\} | X_0 = i\right) \\ &\leq \sum_{n=0}^{\infty} P(X_n = j | X_0 = i) = \sum_{n=0}^{\infty} p_{ij}^{(n)} = 0.\end{aligned}$$

Proposition 3.4.3 *The relation of communication satisfies the following three properties:*

- (i) *State i communicates with state i , all $i \geq 0$.*
- (ii) *If state i communicates with state j , then state j communicates with state i .*
- (iii) *If i communicates with j , j communicates with k , then state i communicates with state k .*

Proof. (i)

$$p_{ii}^{(0)} = P(X_0 = i | X_0 = i) = 1.$$

(iii) We prove that if $i \rightarrow j, j \rightarrow k$, then $i \rightarrow k$. There exist n and m s.t. $p_{ij}^{(n)} > 0, p_{jk}^{(m)} > 0$, then

$$p_{ik}^{(n+m)} = \sum_r p_{ir}^{(n)} p_{rk}^{(m)} \geq p_{ij}^{(n)} p_{jk}^{(m)} > 0.$$

■

Example 3.4.4 $S = \{1, 2, 3\}$, where 1 is good, 2 is normal, and 3 is wrong.

$$\mathbf{P} = \begin{bmatrix} 17/20 & 2/20 & 1/20 \\ 0 & 9/10 & 1/10 \\ 0 & 0 & 1 \end{bmatrix}.$$

One can see that 3 is different from 1 and 2. Each state is a class.

Example 3.4.5 How many classes?

(1) $S = \{1, 2, 3\}$.

$$\mathbf{P} = \begin{bmatrix} 1/2 & 1/4 & 1/4 \\ 1/4 & 0 & 3/4 \\ 0 & 2/3 & 1/3 \end{bmatrix}.$$

(2) $S = \{1, 2, 3, 4, 5\}$.

$$\mathbf{P} = \begin{bmatrix} 0.6 & 0.1 & 0 & 0.3 & 0 \\ 0.2 & 0.5 & 0.1 & 0.2 & 0 \\ 0.2 & 0.2 & 0.4 & 0.1 & 0.1 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix}.$$

(3) $S = \{1, 2, 3, 4, 5\}$.

$$\mathbf{P} = \begin{bmatrix} 1/2 & 1/2 & 0 & 0 & 0 \\ 1/4 & 3/4 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 1/2 & 0 & 1/2 \\ 0 & 0 & 0 & 1 & 0 \end{bmatrix}.$$

Sol. (1) irreducible chain. One class.

(2) {1,2,3} {4} {5}

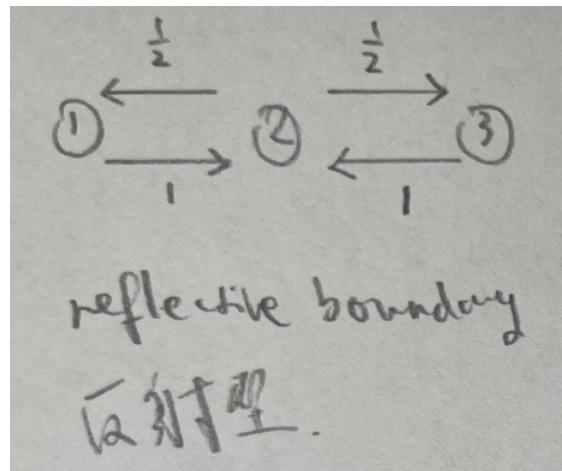
(3) {1,2} {3,4,5}.

3.4.1 recurrent and transient states and classes

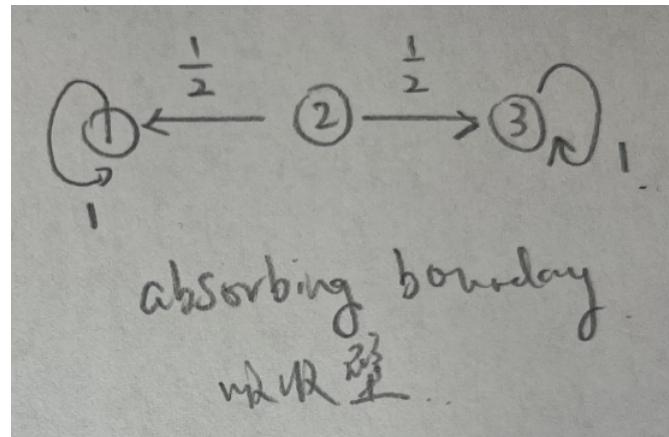
常返与非常返

- For state i , f_{ii} denotes the prob. that, starting in state i , the process will ever re-enter state i .
- Def. State i is recurrent if $f_{ii} = 1$.
State i is transient if $f_{ii} < 1$.
- Recurrent means that with prob. 1, the process will re-enter state i . By Markov chain, the process will start over again. (离开回来)
- If i is recurrent, then the process will re-enter state i again and again and again. **Infinitely many times !!!**
- Transient: With prob. f_{ii} , the process will re-enter state i , and with prob. $1 - f_{ii}$, it will never again enter state i .
 - The prob. that the process in state i for exactly n times equals $f_{ii}^{n-1}(1 - f_{ii})$, $n \geq 1$.
 - If i is transient, the number of time periods that the process will be in state i has a geometric distribution with a finite mean $\frac{1}{1-f_{ii}}$.

Example 3.4.6 All states 1,2,3 are recurrent.



Example 3.4.7 States 1 and 3 are recurrent, and state 2 is transient.



Definition 3.4.8 The first passage time probability (首达概率) is defined as

$$\begin{aligned} f_{ij}^{(1)} &= P(X_1 = j | X_0 = i), \\ f_{ij}^{(n)} &= P(X_n = j, X_k \neq j, 1 \leq k \leq n-1 | X_0 = i), n \geq 1. \end{aligned}$$

Definition 3.4.9 $f_{ij} = \sum_{n=1}^{\infty} f_{ij}^{(n)}$ represents the prob. that one starts from i and first-time arrives at j after finite-time steps. In fact, let $A_1 = \{X_1 = j\}, A_n = \{X_n = j, X_k \neq j, 1 \leq k \leq n-1\}$ be mutually exclusive events. $\bigcup_{n=1}^{\infty} A_n$ means that one ever enters j . Then

$$f_{ij} = \sum_{n=1}^{\infty} f_{ij}^{(n)} = \sum_{n=1}^{\infty} P(A_n | X_0 = i) = P\left(\bigcup_{n=1}^{\infty} A_n | X_0 = i\right) \leq 1,$$

satisfies the property of probability.

Definition 3.4.10 State i is recurrent if $f_{ii} = 1$, and state i is transient if $f_{ii} < 1$.

Remark 3.4.11 当 $f_{ii} = 1$ 时, 质点从 i 出发以概率 1 回到 i , 再出发, 再回到 i , 必然回到 i 无穷次。当 $f_{ii} < 1$ 时, 质点从 i 出发以正概率 $1 - f_{ii}$ 不再回到 i , 如果回到 i , 则再次从 i 出发以正概率 $1 - f_{ii}$ 不再回到 i , 所以只要 $f_{ii} < 1$, 质点不回到 i 总会发生, 因而最终离开 i 。

Example 3.4.12 $S = \{1, 2, 3\}$, where 1 is good, 2 is normal, and 3 is wrong.

$$\mathbf{P} = \begin{bmatrix} 17/20 & 2/20 & 1/20 \\ 0 & 9/10 & 1/10 \\ 0 & 0 & 1 \end{bmatrix}.$$

One can show that

$$\lim_{n \rightarrow \infty} \mathbf{P}^{(n)} = \begin{bmatrix} 0 & 0 & 1 \\ 0 & 0 & 1 \\ 0 & 0 & 1 \end{bmatrix},$$

by setting that $\mathbf{P} = \mathbf{V}\mathbf{D}\mathbf{V}^{-1}$, where

$$\mathbf{V} = \begin{bmatrix} 1 & 0.89 & 0.58 \\ 0 & 0.45 & 0.58 \\ 0 & 0 & 0.58 \end{bmatrix}, \mathbf{D} = \begin{bmatrix} 0.85 & 0 & 0 \\ 0 & 0.9 & 0 \\ 0 & 0 & 1 \end{bmatrix}.$$

The absorbing state i satisfying $f_{33} = f_{33}^{(1)} = 1$ is recurrent. State 1 with $f_{11} = f_{11}^{(1)} + 0 + \dots = \frac{17}{20}$ and state 2 with $f_{22} = f_{22}^{(1)} + 0 + \dots = \frac{9}{10}$ are transient.

Theorem 3.4.13 For $\forall i, j \in I, n \geq 1$,

$$p_{ij}^{(n)} = \sum_{k=1}^n f_{ij}^{(k)} p_{jj}^{(n-k)}. \quad (3.1)$$

Proof. Let $A_n = \{X_n = j, X_k \neq j, 1 \leq k \leq n-1\}$ be mutually exclusive events. Since $\{X_n = j\} \subset \bigcup_{k=1}^{\infty} A_k$, then

$$p_{ij}^{(n)} = P(X_n = j | X_0 = i) = \sum_{k=1}^n P(X_n = j | A_k, X_0 = i) P(A_k | X_0 = i) = \sum_{k=1}^n p_{jj}^{(n-k)} f_{ij}^{(k)}.$$

■

Theorem 3.4.14 *Markov chain $\{X_n\}$, then*

(1)

$$\sum_{n=0*}^{\infty} p_{ii}^{(n)} = \frac{1}{1 - f_{ii}}.$$

(2) *i is recurrent $\Leftrightarrow \sum_{n=0}^{\infty} p_{ii}^{(n)} = \infty$. i is transient $\Leftrightarrow \sum_{n=0}^{\infty} p_{ii}^{(n)} = \frac{1}{1 - f_{ii}} < \infty$.*

(3) *If i is recurrent and $i \rightarrow j$, then $i \leftrightarrow j$, and j is also recurrent. Recurrence is a class property.*

Proof. (1) Take $j = i$ in Eq. (3.1), one has

$$p_{ii}^{(n)} \rho^n = \sum_{k=1}^n f_{ii}^{(k)} \rho^k p_{ii}^{(n-k)} \rho^{n-k},$$

where $\rho \in (0, 1), 1 \leq k \leq n, 1 \leq n \leq \infty$. By summing over n , we obtain

$$\begin{aligned} G(\rho) &:= \sum_{n=0}^{\infty} p_{ii}^{(n)} \rho^n = 1 + \sum_{n=1}^{\infty} p_{ii}^{(n)} \rho^n = 1 + \sum_{n=1}^{\infty} \sum_{k=1}^n f_{ii}^{(k)} \rho^k p_{ii}^{(n-k)} \rho^{n-k} \\ &= 1 + \sum_{k=1}^{\infty} \sum_{n=k}^{\infty} f_{ii}^{(k)} \rho^k p_{ii}^{(n-k)} \rho^{n-k} = 1 + \left(\sum_{k=1}^{\infty} f_{ii}^{(k)} \rho^k \right) \left(\sum_{n=k=0}^{\infty} p_{ii}^{(n-k)} \rho^{n-k} \right) \\ &= 1 + F(\rho)G(\rho), \end{aligned}$$

where $F(\rho) = \sum_{k=1}^{\infty} f_{ii}^{(k)} \rho^k$. Then

$$G(\rho) = \frac{1}{1 - F(\rho)}.$$

Let $\rho \rightarrow 1$, done.

(2) Direct corollary.

(3) *$i \rightarrow j$ 说明从i到j的概率是正数，每次质点从i出发达到j，由于i常返，所以质点一定会回到i，因此j通i，因此 *$i \leftrightarrow j$* 。 Since i communicates with j , there exist integers k and m such that $p_{ij}^{(k)} > 0, p_{ji}^{(m)} > 0$. Notice that for any integer n ,*

$$p_{jj}^{(m+n+k)} \geq p_{ji}^{(m)} p_{ii}^{(n)} p_{ij}^{(k)}.$$

Summing over all n ,

$$\sum_{n=1}^{\infty} p_{jj}^{(m+n+k)} \geq p_{ji}^{(m)} p_{ij}^{(k)} \sum_{n=1}^{\infty} p_{ii}^{(n)} = \infty,$$

from which we can conclude that j is recurrent. ■

Remark 3.4.15 $\sum_{n=0}^{\infty} p_{ii}^{(n)}$ is the expected number of times that one starts from i and re-enter i (or say the expected number of time periods that the process is in state i). To see this, let

$$I_n = \begin{cases} 1, & X_n = i \\ 0, & X_n \neq i \end{cases}$$

$\sum_{n=0}^{\infty} I_n$ represents the number of times that the process is in state i .

$$E\left[\sum_{n=0}^{\infty} I_n | X_0 = i\right] = \sum_{n=0}^{\infty} E[I_n | X_0 = i] = \sum_{n=0}^{\infty} P[X_n = i | X_0 = i] = \sum_{n=0}^{\infty} p_{ii}^{(n)}.$$

Remark 3.4.16

- (1) recurrent $\Leftrightarrow i$ can be visited infinitely many times.
- (2) transient $\Leftrightarrow i$ can be visited finite times.

Remark 3.4.17 For finite-state MC, not all states can be transient, so that at least one of the states must be recurrent. Assume that all M states are transient, then say after T_1 , state 1 will never be visited. So for each state i . Then, after $T = \max\{T_1, \dots, T_M\}$, no state will be visited. However, the process must be in some state after T . Contradiction!

Remark 3.4.18 Transience is a class property. If state i is transient and communicates with state j , then state j must also be transient. This is a corollary of the conclusion that recurrence is a class property.

Remark 3.4.19 The result that not all states in a finite MC can be transient leads to the conclusion that all states of a finite irreducible Markov chain are recurrent.

Example 3.4.20 Let the Markov chain consisting of the states, $0, 1, 2, 3$ have the transition matrix

$$\mathbf{P} = \begin{bmatrix} 0 & 0 & \frac{1}{2} & \frac{1}{2} \\ 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 1 & 0 & 0 \end{bmatrix}.$$

Determine which states are transient and which are recurrent.

Sol. All states communicate, and hence since this is a finite chain, all states must be recurrent.

Example 3.4.21 Consider the Markov chain with states $0, 1, 2, 3, 4$ and

$$\mathbf{P} = \begin{bmatrix} \frac{1}{2} & \frac{1}{2} & & & \\ \frac{1}{2} & \frac{1}{2} & & & \\ & & \frac{1}{2} & \frac{1}{2} & \\ & & \frac{1}{2} & \frac{1}{2} & \\ \frac{1}{4} & \frac{1}{4} & & \frac{1}{2} & \end{bmatrix},$$

then the chain has three classes $\{0, 1\}$, $\{2, 3\}$, and $\{4\}$. The first two classes are recurrent and the third transient. For state 1, $f_{11} = \sum_{i \geq 1} f_{11}^{(i)} = \sum_{i \geq 1} \left(\frac{1}{2}\right)^i = 1$ since a particle once leaves state 1 and then can stay at state 2 for any times. For state 5, $f_{44} = f_{44}^{(1)} = \frac{1}{2}$.

Example 3.4.22 In the chain with $\{0, 1, \dots, n\}$ with two-sided absorbing boundaries. $\{1, 2, \dots, n-1\}$ is in the transient class, and $\{0\}$ and $\{n\}$ are in the recurrent class. For transient class, one can see that if particle starts from state 1 and then never comes back to 1 with prob. $1 - f_{11} > P\{X_1 = 0 | X_0 = 1\} = q > 0$.

Example 3.4.23 For the chain with $\{0, 1, \dots, n\}$ with two-sided reflection boundaries. Since all state communicate, they are in the same class. Moreover, due to finite number of states, thus all states are recurrent.

Example 3.4.24 (A Random Walk) Consider a Markov chain with infinite length $i = 0, \pm 1, \pm 2, \dots$ and has transition prob. by

$$p_{i,i+1} = p = 1 - p_{i,i-1}, \text{ for all } i.$$

One colorful interpretation of this process is that it represents the **wanderings of a drunken man** as he walks along a straight line. Since all states communicate, they are all either recurrent or transient. Let us consider state 0 and determine if $\sum_{n=1}^{\infty} p_{00}^{(n)}$ is finite or infinite. It is obvious for odd steps that $p_{00}^{(2n-1)} = 0$ for all n . For even steps,

$$p_{00}^{(2n)} = C_{2n}^n p^n (1-p)^n.$$

Using the Stirling formula

$$n! \sim n^{n+1/2} e^{-n} \sqrt{2\pi},$$

we obtain

$$p_{00}^{(2n)} \sim \frac{(4p(1-p))^n}{\sqrt{\pi n}}.$$

Hence, $\sum_{n=1}^{\infty} p_{00}^{(n)}$ will converge if and only if

$$\sum_{n=1}^{\infty} \frac{(4p(1-p))^n}{\sqrt{\pi n}}$$

does. Using Abel-Dirichlet determinant method, one can see that $\sum_{n=1}^{\infty} p_{00}^{(n)} = \infty$ if and only if $p = 1/2$ (recurrent). We call it a **symmetric random walk**. When $p \neq 1/2$, the chain is transient.

Example 3.4.25 For 2D symmetric random walk, the transition prob is given by

$$p_{up} = p_{down} = p_{left} = p_{right} = \frac{1}{4}.$$

The chain is irreducible. All states will be recurrent if state $(0,0)$ is recurrent. Then

$$\begin{aligned} p_{00}^{(2n)} &= \sum_{i=0}^n \frac{(2n)!}{i!i!(n-i)!(n-i)!} \left(\frac{1}{4}\right)^{2n} \\ &= \left(\frac{1}{4}\right)^{2n} \sum_{i=0}^n \frac{(2n)!}{n!n!} \frac{n!}{i!(n-i)!} \frac{n!}{i!(n-i)!} \\ &= \left(\frac{1}{4}\right)^{2n} \frac{(2n)!}{n!n!} C_{2n}^n. \end{aligned}$$

Now

$$C_{2n}^n \sim \frac{4^n}{\sqrt{\pi n}},$$

so that

$$p_{00}^{(2n)} \sim \frac{1}{\pi n}.$$

The series is divergent, and thus all states are recurrent.

Example 3.4.26 For dimension greater than or equal to 3 symmetric random walk, all states are transient.

Take $K_n = \{j, k | j, k \geq 0, j + k \leq n\}$, then

$$\begin{aligned} p_{00}^{(2n)} &= \sum_{j,k \in K_n} \frac{(2n)!}{(j!k!(n-j-k)!)^2} \left(\frac{1}{6}\right)^{2n} \\ &= \frac{1}{6^{2n}} C_{2n}^n \sum_{j,k \in K_n} \frac{n!n!}{(j!k!(n-j-k)!)^2} \\ &\leq \frac{1}{6^{2n}} C_{2n}^n \frac{n!}{\left(\frac{n}{3}\right)! \left(\frac{n}{3}\right)! \left(\frac{n}{3}\right)!} \sum_{j,k \in K_n} \frac{n!}{j!k!(n-j-k)!} \\ &= \frac{1}{6^{2n}} C_{2n}^n \frac{n!}{\left(\frac{n}{3}\right)! \left(\frac{n}{3}\right)! \left(\frac{n}{3}\right)!} (1+1+1)^n. \end{aligned}$$

Using Stirling,

$$\frac{n!}{\left(\frac{n}{3}\right)!\left(\frac{n}{3}\right)!\left(\frac{n}{3}\right)!} = O\left(\frac{n^{n+1/2}e^{-n}}{[(n/3)^{n/3+1/2}e^{-n/3}]^3}\right) = O\left(\frac{3^{n+3/2}}{n}\right),$$

then

$$p_{00}^{(2n)} = O\left(\frac{1}{6^{2n}} \frac{4^n}{\sqrt{n}} \frac{3^{n+3/2}}{n} 3^n\right) = O(n^{-3/2}),$$

eventually gives a convergent series. The symmetric random walk is transient. In general, for $d \geq 3$, $\sum_n n^{-d/2}$ is convergent, which gives transient results. However, intuitively, it is hard to think about why the behavior is quite different between $d \leq 2$ and $d \geq 3$. This result is known as Polya theorem.

3.4.2 Periodicity

Definition 3.4.27 For Markov chain $\{X_n\}$, define the periodicity:

- (a) If $\sum_{n=1}^{\infty} p_{ii}^{(n)} = 0$, then the particle can never re-enter i once leaving i and i is said to have ∞ period.
- (b) State i is said to have period d if $p_{ii}^{(n)} = 0$ whenever n is not divisible by d ($d \nmid n$), and d is the largest integer with this property. For instance, starting in i , it may be possible for the process to enter state i only at times $2, 4, 6, 8, \dots$, in which case state i has period 2.
- (c) A state with period 1 is said to be aperiodic.

非周期

Remark 3.4.28 Let d be the period of state i . If $p_{ii}^{(n)} > 0$, then $d|n$ and d must be the largest integer with this property. Moreover, when state i has period $d < \infty$, then $p_{ii}^{(nd)} > 0$ is satisfied only for some n , instead for all n .

Remark 3.4.29 Let $d < \infty$ be the period of the state i . Then d is the maximum common divisor of the set $D = \{n | p_{ii}^{(n)} > 0, n \geq 1\}$.

最大公约数

Example 3.4.30 On a line, if a particle moves forward 1-step with prob. $1/3$, backward 1-step with prob. $1/3$, moves forward 2-step with prob. $1/3$. Then all states are aperiodic. The reason is

$$\begin{aligned} p_{ii}^{(2)} &\geq p_{i,i+1}p_{i+1,i} > 0, \\ p_{ii}^{(3)} &\geq p_{i,i-2}p_{i-2,i-1}p_{i-1,i} > 0, \end{aligned}$$

so that $2 = nd, 3 = md \Rightarrow d = 1$.

Example 3.4.31 On a line, if a particle moves either forward 1-step with prob. p or backward 5-step with prob. $1-p$. Then the period is 6 for each state. If $p_{ii}^{(n)} > 0$, then assume move forward k times and backward m times resulting in $k = 5m$ so that $n = k + m = 6m$. Since $p_{ii}^{(6)} > p^5q > 0$, so $6 = md$ and thus $d \leq 6$. The period of state i is $d = 6$.

Theorem 3.4.32 Periodicity is a class property. That is, if state i has period d , and state i and j communicate, then state j also has period d .

Proof. There exist r and s such that $p_{ji}^{(r)} > 0$ and $p_{ij}^{(s)} > 0$. Let d and t be the periods of states i and j , respectively. Then $\exists n = mt$, s.t. $p_{jj}^{(n)} > 0$. Then

$$p_{ii}^{(s+n+r)} \geq p_{ij}^{(s)} p_{jj}^{(n)} p_{ji}^{(r)} > 0.$$

Thus

$$d|s + n + r.$$

Also

$$p_{ii}^{(r+s)} \geq p_{ij}^{(s)} p_{ji}^{(r)} > 0 \Rightarrow d|s + r.$$

Therefore,

$$d|n \Rightarrow d|mt.$$

- If $m = 1$, then $d|t$.
- If there exist $(m_1, m_2) = 1$ among all m_1 and m_2 , then $d|m_1 t$ $d|m_2 t \Rightarrow d|t$.
- If $(m_1, m_2) > 1$ for all m_1 and m_2 , then state j has period $(m_1, m_2)t$, contradiction.

Similarly, $t|d$. Thus, $t = d$. ■

3.4.3 positive and null recurrent states and classes

正常返和零常返

Definition 3.4.33 State i is said to be positive recurrent, if, starting in i , expected time until the process returns to state i is finite.

Definition 3.4.34 $T_i = n$ represents that one arrives at state i at n th step for the first time, that is,

$$T_i = \begin{cases} \min\{n | X_n = i, n \geq 1\}, & \text{if } \bigcup_{n=1}^{\infty} \{X_n = i\} \text{ occurs,} \\ \infty, & \text{else.} \end{cases}$$

Then,

$$f_{ii}^{(n)} = P(T_i = n | X_0 = i) = P(X_n = i, X_k \neq i, 1 \leq k \leq n-1 | X_0 = i).$$

平均回转时间

Definition 3.4.35 Let $f_{ii} = 1$ and $X_0 = i$. Denote μ_i the expected time until the process return to state i or say the first return time. Then the **mean recurrence time** or **mean first-passage time** is

$$\mu_i = E[T_i | X_0 = i] = \sum_{n=1}^{\infty} n P(T_i = n | X_0 = i) = \sum_{n=1}^{\infty} n f_{ii}^{(n)}.$$

State i is said to be positive recurrent if $\mu_i < \infty$; State i is said to be null recurrent if $\mu_i = \infty$.

Lemma 3.4.36 Let state i have period d and mean recurrence time $\mu_i = E(T_i | X_0 = i)$. Then

$$\lim_{n \rightarrow \infty} p_{ii}^{(nd)} = \frac{d}{\mu_i}.$$

Proof is hard.

Theorem 3.4.37 Let state i be recurrent.

- (1) i is null recurrent $\Leftrightarrow \lim_{n \rightarrow \infty} p_{ii}^{(n)} = 0$.
- (2) i is positive recurrent and aperiodic (ergodic) $\Leftrightarrow \lim_{n \rightarrow \infty} p_{ii}^{(n)} = \frac{1}{\mu_i} > 0$.

Proof. (1) " \Rightarrow " Since i is recurrent, then there exists some $n < \infty$ such that $p_{ii}^{(n)} > 0$. Thus, we have $d < n < \infty$.

i is null recurrent \Rightarrow by lemma $\lim_{n \rightarrow \infty} p_{ii}^{(nd)} = \frac{d}{\infty} = 0$. That is,

• when $d|m = nd \Rightarrow \lim_{m \rightarrow \infty} p_{ii}^{(m)} = 0$

• when $d \nmid m \Rightarrow p_{ii}^{(m)} = 0$ based on the definition of the period. Thus, one always has $\lim_{n \rightarrow \infty} p_{ii}^{(n)} = 0$.

" \Leftarrow " Assume that i is positive recurrent, then $\lim_{n \rightarrow \infty} p_{ii}^{(nd)} = \frac{d}{\mu_i} > 0$. Contradiction!

(2) " \Rightarrow " ergodic so that $d = 1$. By lemma, done.

" \Leftarrow " Since $\lim_{n \rightarrow \infty} p_{ii}^{(n)} = \frac{1}{\mu_i} > 0$, i is positive recurrent. Compare condition with lemma, then we see that $d = 1$. ■

Theorem 3.4.38 Let state i be recurrent.

(1) If i is null recurrent, and $i \rightarrow j$, then j is null recurrent. That is, the null recurrence is a class property.

(2) If i is positive recurrent, and $i \rightarrow j$, then j is positive recurrent. That is, the positive recurrence is a class property.

Proof. (1) First, since i recurrent and $i \rightarrow j$ so that $i \leftrightarrow j$. There exist integers m and n such that $p_{ij}^{(n)} > 0$ and $p_{ji}^{(m)} > 0$. Using the fact that $p_{ii}^{(n+k+m)} \geq p_{ij}^{(n)} p_{jj}^{(k)} p_{ji}^{(m)}$ and $\lim_{n \rightarrow \infty} p_{ii}^{(n)} = 0$, we have that

$$p_{jj}^{(k)} \leq \frac{1}{p_{ij}^{(n)} p_{ji}^{(m)}} p_{ii}^{(n+k+m)} \rightarrow 0, \text{ as } k \rightarrow \infty.$$

(2) Disproof. Assume j is null recurrent, then i is also null recurrent. However, i is positive recurrent. Contradiction! ■

3.4.4 ergodicity

Definition 3.4.39 Positive recurrent, aperiodic states are called ergodic.

Example 3.4.40 Consider the Markov chain. See Fig. 3.1. Then, states 3 and 4 are transient. States 1

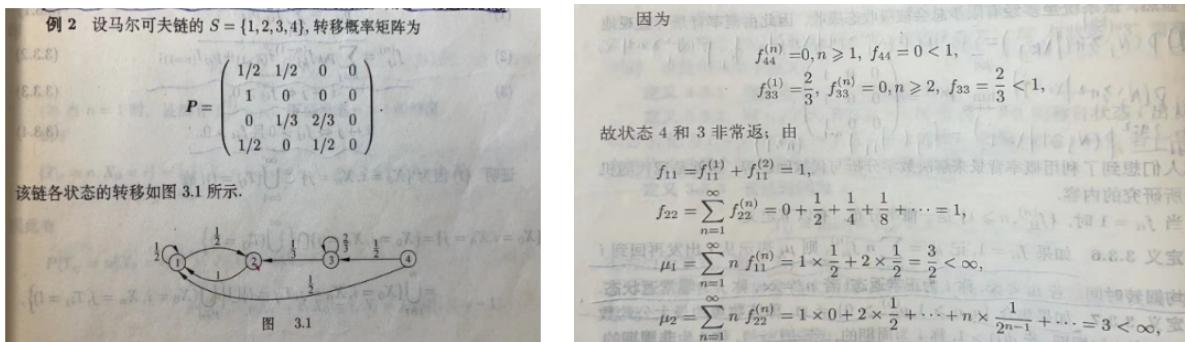


Figure 3.1: Example for classification of transient and ergodic classes.

and 2 are positive recurrent, in addition, they are aperiodic, and thus they are ergodic.

Here is the summary of this subsection (see Fig. 3.2).

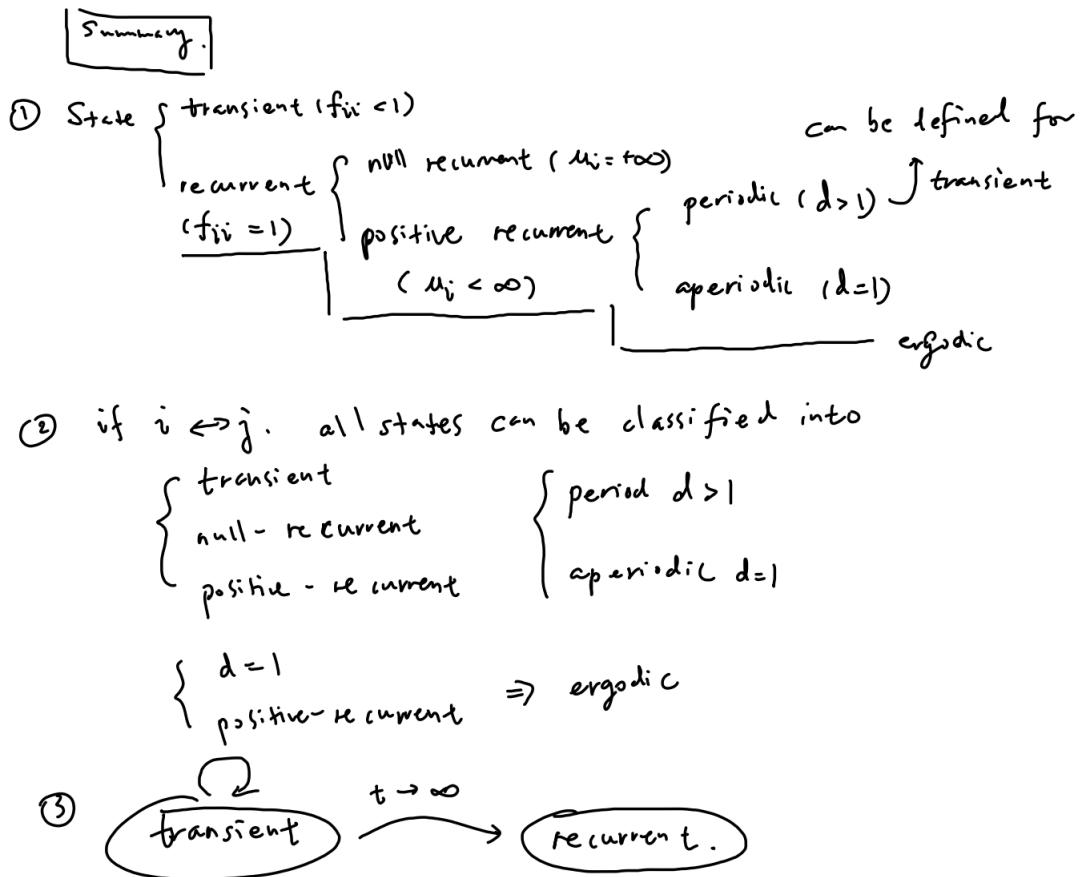


Figure 3.2: Summary of classification.

3.4.5 Classification

Definition 3.4.41 I is state space of Markov chain $\{X_n\}$ and $B \subset I$. If any state $b \in B$ is not accessible to $B^c = I - B$, then B is said to be a closed set.

Theorem 3.4.42 For Markov chain,

- (1) There is no intersection between two different classes.
- (2) Each state belongs to one of the following class, transient, null recurrent, or positive recurrent. Moreover, all states in one class have the same properties.
- (3) Recurrent class is a closed set: all particles in the recurrent class cannot leave it.
- (4) Null recurrent class has infinitely many states.
- (5) If transient class is closed, then it contains infinitely many states. (see. e.g., non-symmetric random walk with drift.)

Lemma 3.4.43 If j is transient or null recurrent, then for any $i \in I$,

$$\lim_{n \rightarrow \infty} p_{ij}^{(n)} = 0. \quad (3.2)$$

Proof. Proof for the Lemma. If j is null recurrent, then $\lim_{n \rightarrow \infty} p_{jj}^{(n)} = 0$. If j is transient, then $\sum_{n=0}^{\infty} p_{jj}^{(n)} < \infty$ gives also $\lim_{n \rightarrow \infty} p_{jj}^{(n)} = 0$. For any $i \in I$,

$$p_{ij}^{(n)} = \sum_{k=1}^m f_{ij}^{(k)} p_{jj}^{(n-k)} + \sum_{k=m+1}^n f_{ij}^{(k)} p_{jj}^{(n-k)} \leq \sum_{k=1}^m p_{jj}^{(n-k)} + \sum_{k=m+1}^n f_{ij}^{(k)}.$$

Let $n \rightarrow \infty$, then the first term on RHS $\sum_{k=1}^m p_{jj}^{(n-k)} \rightarrow 0$. Thus,

$$\lim_{n \rightarrow \infty} p_{ij}^{(n)} \leq \sum_{k=m+1}^{\infty} f_{ij}^{(k)} \leq 1, \text{ (bounded)}$$

Let $m \rightarrow \infty$, then

$$\lim_{n \rightarrow \infty} p_{ij}^{(n)} = 0.$$

■

Proof. We only prove for (4) and (5) in Theorem 3.4.42. Let C be a class of either null recurrent or closed transient states. Based on (3), null recurrent is closed. Thus, C is always closed, that is, if $i, j \in C$, then i, j can never leave the closed C . Notice that for any $i \in C$,

$$\sum_{j \in C} p_{ij}^{(n)} = \sum_{j \in I} p_{ij}^{(n)} = 1, \text{ for any } n.$$

We now use the result $\lim_{n \rightarrow \infty} p_{ij}^{(n)} = 0$ for transient or null recurrent j from above Lemma. If C contains only finite states, then

$$1 = \lim_{n \rightarrow \infty} \sum_{j \in C} p_{ij}^{(n)} = \sum_{j \in C} \lim_{n \rightarrow \infty} p_{ij}^{(n)} = \sum_{j \in C} 0 = 0.$$

Contradiction! We conclude that C contains infinitely many states. ■

Theorem 3.4.44 State space I can be decomposed as

$$I = \bigcup_{j=1}^m C_j + T, \quad m \leq \infty$$

where C_j is recurrent class and T is transient class.

Remark 3.4.45 For a transient class T ,

- (1) A particles can be in T forever, or can move from T to some C_j and then stay in C_j forever.
- (2) If T has finite states, then the particle will always leave T , enter some closed C_j and stay in C_j forever. On the other hand, for a recurrent class C_j ,
- (1) C_j has finite states \Rightarrow all states are positive recurrent.
- (2) C_j is null recurrent \Rightarrow C_j contains infinitely many states.

Remark 3.4.46 For Markov chain,

- (1) For finite-state Markov chain, either positive recurrent or transient, impossible to be null recurrent. (e.g., random walk model with absorbing boundaries)

(2) In a finite-state Markov chain, all recurrent states are positive recurrent.

(3) For irreducible finite-state Markov chain, all states are positive recurrent. (e.g., random walk model with reflection)

(4) For infinite-length irreducible Markov chain, all states can be

$$\begin{cases} \text{transient}, & \text{random walk } p \neq \frac{1}{2}, \\ \text{null recurrent}, & \text{random walk } p = \frac{1}{2}, \\ \text{positive recurrent}, & \text{e.g., see following example.} \end{cases}$$

(5) Each absorbing state is positive recurrent since $\mu_i = \sum_{n=1}^{\infty} n f_{ii}^{(n)} = 1 \cdot f_{ii}^{(1)} = 1 < \infty$. Infinite-length can have positive recurrent states. (e.g., each state is an absorbing state)

Example 3.4.47 Consider the Markov chain with infinitely many states (see Fig. 3.3). Then, all states are

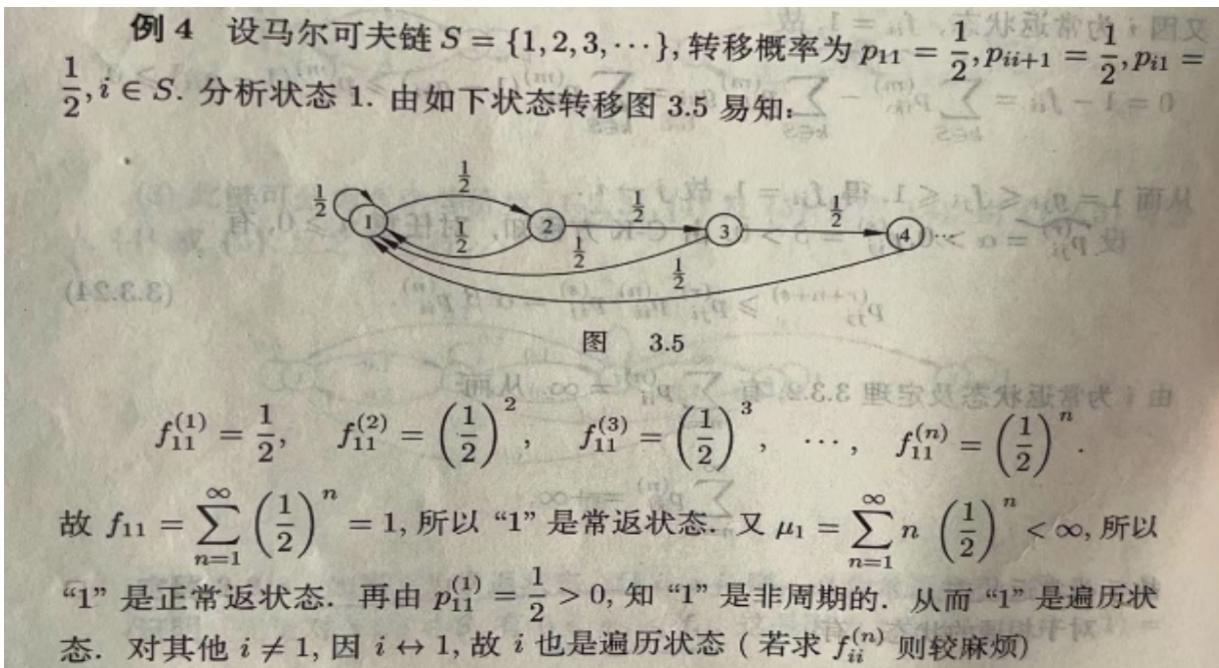


Figure 3.3: All states are ergodic (positive recurrent and aperiodic) for a Markov chain with infinitely many states.

ergodic.

3.5 Stationary Distributions, Limiting Probabilities, Limiting Distributions

For the two-state Markov chain with the transition matrix

$$\mathbf{P} = \begin{bmatrix} 0.7 & 0.3 \\ 0.4 & 0.6 \end{bmatrix}.$$

It turns out to be

$$\mathbf{P}^{(4)} = \begin{bmatrix} 0.5749 & 0.4251 \\ 0.5668 & 0.4332 \end{bmatrix}, \mathbf{P}^{(8)} = \begin{bmatrix} 0.572 & 0.428 \\ 0.570 & 0.430 \end{bmatrix}.$$

They are almost identical. In fact it seems that $p_{ij}^{(n)}$ converges to some value as $n \rightarrow \infty$. In other words, there seems to exist a limiting prob. that the process will be in state j after a large number of transitions, and this value is independent of the initial state.

The question is:

- (a) If stationary distribution exists and if it is unique? What is the value of the stationary distribution?
- (b) If limiting probability $\lim_{n \rightarrow \infty} p_{ij}^{(n)}$ exists? If exists, what is its value?
- (c) For each i , if $\lim_{n \rightarrow \infty} \pi_i^{(n)}$ exists? If $\lim_{n \rightarrow \infty} \pi^{(n)}$ exists, where $\pi^{(n)} = (\pi_1^{(n)}, \pi_2^{(n)}, \dots)$? What is its value?

3.5.1 Stationary Distributions

Definition 3.5.1 If a probability distribution $\pi = (\pi_1, \pi_2, \dots)$ satisfies

$$\sum_{j \in I} \pi_j = 1, \pi = \pi \mathbf{P},$$

or equivalently

$$\sum_{j \in I} \pi_j = 1, \pi_j = \sum_{k \in I} \pi_k p_{kj} \geq 0, j \in I,$$

then π_j is called stationary probability and π is called the stationary prob. distribution. Obviously,

$$\pi = \pi \mathbf{P} = \dots = \pi \mathbf{P}^n.$$

See reference in [2].

Lemma 3.5.2 If j is ergodic and $i \leftrightarrow j$, then

$$\lim_{n \rightarrow \infty} p_{ij}^{(n)} = \frac{1}{\mu_j}. \quad (3.3)$$

Proof. Based on the previous result, we have that

$$\lim_{n \rightarrow \infty} p_{jj}^{(n)} = \frac{1}{\mu_j},$$

where the period $d = 1$ is used. For any $i \leftrightarrow j$,

$$p_{ij}^{(n)} = \sum_{k=1}^m f_{ij}^{(k)} p_{jj}^{(n-k)} + \sum_{k=m+1}^n f_{ij}^{(k)} p_{jj}^{(n-k)}.$$

Fix m and let $n \rightarrow \infty$,

$$\lim_{n \rightarrow \infty} p_{ij}^{(n)} = \sum_{k=1}^m f_{ij}^{(k)} \frac{1}{\mu_j} + \sum_{k=m+1}^{\infty} f_{ij}^{(k)} p_{jj}^{(n-k)}.$$

Notice that

$$\sum_{k=m+1}^{\infty} f_{ij}^{(k)} p_{jj}^{(n-k)} \leq \sum_{k=m+1}^{\infty} f_{ij}^{(k)} \leq 1,$$

so that

$$\sum_{k=m+1}^{\infty} f_{ij}^{(k)} p_{jj}^{(n-k)} \rightarrow 0, \quad \text{as } m \rightarrow \infty.$$

Also notice that for recurrent j and $i \leftrightarrow j$,

$$\sum_{k=1}^{\infty} f_{ij}^{(k)} = 1.$$

Thus, let $m \rightarrow \infty$, we have

$$\lim_{n \rightarrow \infty} p_{ij}^{(n)} = \sum_{k=1}^m f_{ij}^{(k)} \frac{1}{\mu_j} + \sum_{k=m+1}^{\infty} f_{ij}^{(k)} p_{jj}^{(n-k)} \rightarrow \frac{1}{\mu_j}.$$

■

Theorem 3.5.3 Let C^+ contain all positive recurrent states of the Markov chain $\{X_n\}$ and $i \in C^+$.

Part (1) If C^+ has only one ergodic class, then the stationary prob. π_j satisfies

$$\pi_j = \lim_{n \rightarrow \infty} p_{ij}^{(n)} = \frac{1}{\mu_j}, \quad j \in I,$$

is the unique stationary distribution.

Proof. Define

$$\mu_j = \begin{cases} \infty, & j \text{ is null recurrent or transient,} \\ \text{some positive finite number,} & j \text{ is positive recurrent.} \end{cases}$$

ergodic $\Rightarrow d = 1$. Based on the result in some example (think about why for i and j , the following holds true?),

$$\lim_{n \rightarrow \infty} p_{ij}^{(n)} = \begin{cases} 0, & j \notin C^+, \forall i \in I, \\ \frac{1}{\mu_j}, & j \in C^+, i \in C^+, C^+ \text{ has only one class, that is, } i \leftrightarrow j. \end{cases}$$

Define as above for the corresponding i and j relation,

$$\pi_j := \lim_{n \rightarrow \infty} p_{ij}^{(n)},$$

we would like to show that π_j is a stationary probability, that is, π_j satisfies the equations in Definition 3.5.1.

Existence of stationary distribution. A particle starting from $i \in I$ must enter some state $j \in I$, and thus

$$\sum_{j \in I} p_{ij}^{(n)} = 1. \tag{3.4}$$

Let $n \rightarrow \infty$ in Eq. (3.4), one has for any $i \in C^+$,

$$\sum_{j \in I} \pi_j = \sum_{j \in I} \lim_{n \rightarrow \infty} p_{ij}^{(n)} = \lim_{n \rightarrow \infty} \sum_{j \in I} p_{ij}^{(n)} = 1.$$

(think about why can be interchanged? Dominated Convergence Theorem) Using Chapman-Kolmogorov equation, for a particle starting from $i \in C^+$,

$$\begin{aligned} \pi_j &= \lim_{n \rightarrow \infty} p_{ij}^{(n)} = \lim_{n \rightarrow \infty} \sum_{k \in I} p_{ik}^{(n-1)} p_{kj} = \sum_{k \in I} \lim_{n \rightarrow \infty} p_{ik}^{(n-1)} p_{kj} \\ &= \sum_{k \in C^+} \lim_{n \rightarrow \infty} p_{ik}^{(n-1)} p_{kj} = \sum_{k \in C^+} \pi_k p_{kj} = \sum_{k \in I} \pi_k p_{kj}, \end{aligned}$$

where made use of Dominated Convergence Theorem. Thus, π_j is a stationary distribution.

Uniqueness. Suppose that $\{v_j\}$ is another stationary distribution.

If $j \notin C^+$, using the property of stationary distribution $v=v\mathbf{P}^n$ and assume that $n \rightarrow \infty$ and $\sum_{k \in I}$ can be interchanged, we obtain:

$$v_j = \lim_{n \rightarrow \infty} \sum_{k \in I} v_k p_{kj}^{(n)} = \sum_{k \in I} \lim_{n \rightarrow \infty} v_k p_{kj}^{(n)} = 0 \Rightarrow v_j = 0 = \lim_{n \rightarrow \infty} p_{ij}^{(n)} := \pi_j,$$

where made use of the fact that for any $j \notin C^+$, one has $\lim_{n \rightarrow \infty} p_{ij}^{(n)} = 0$ for any $i \in I$ (see equation (3.2)). If $j \in C^+$,

$$v_j = \sum_{k \in I} v_k p_{kj}^{(n)} = \sum_{k \in C^+} v_k p_{kj}^{(n)} \rightarrow \left(\sum_{k \in C^+} v_k \right) \pi_j = \left(\sum_{k \in I} v_k \right) \pi_j = \pi_j, \text{ as } n \rightarrow \infty,$$

where the second equality holds true since $v_k = 0$ if $k \notin C^+$, and the last equality holds true since there is only one C^+ . Therefore, $v_j = \pi_j$.

Notice that when computing $\lim_{n \rightarrow \infty} p_{ij}^{(n)}$ for $j \in C^+$, the assumption $i \in C^+$ is important. Otherwise we will obtain a wrong proof:

$$v_j = \lim_{n \rightarrow \infty} \sum_{k \in I} v_k p_{kj}^{(n)} = \sum_{k \in I} \lim_{n \rightarrow \infty} v_k p_{kj}^{(n)} = \left(\sum_{k \in I} v_k \right) \pi_j = \pi_j, \text{ as } n \rightarrow \infty.$$

This proof results in that two classes of C^+ would also have a unique stationary distribution, which is a wrong statement. ■

Theorem 3.5.4 Part (2) *If C^+ is a class with period d , then*

$$\pi_j = \frac{1}{d} \lim_{n \rightarrow \infty} p_{jj}^{(nd)} = \lim_{n \rightarrow \infty} \frac{1}{d} \sum_{s=1}^d p_{ij}^{(nd+s)} = \lim_{n \rightarrow \infty} \frac{1}{n} \sum_{k=1}^n p_{ij}^{(k)}, j \in I,$$

is the unique stationary distribution. Moreover, $\pi_j = \frac{1}{\mu_j}$.

Theorem 3.5.5 Part (3) *$\{X_n\}$ has a unique stationary distribution. $\Leftrightarrow C^+$ is an equivalent class. \Leftrightarrow there exists a unique class of positive recurrent states.*

Theorem 3.5.6 Part (4) *Stationary distribution exists for $\{X_n\}$ $\Leftrightarrow C^+ \neq \emptyset$.*

Theorem 3.5.7 Part (5) *Corollary of (4): A Markov chain with finite states has at least one stationary distribution.*

Proof. $C^+ \neq \emptyset$ since at least one state is recurrent and further positive recurrent (null recurrence has infinitely many states). ■

Theorem 3.5.8 Part (6) *Corollary of (1): An irreducible and aperiodic MC with finite states has a unique stationary distribution.*

Proof. finite states + irreducible \Rightarrow positive recurrent. positive recurrent + aperiodic \Rightarrow ergodic. Thus, \exists ! stationary distribution. ■

Theorem 3.5.9 Part (7) Structure of \mathbf{P} . The transition matrix is

$$\mathbf{P} = \begin{bmatrix} C_1 & C_2 & \cdots & C_m & T \\ \mathbf{P}_1 & 0 & \cdots & \cdots & 0 \\ 0 & \mathbf{P}_2 & \ddots & \ddots & \vdots \\ \vdots & \ddots & \ddots & \ddots & \vdots \\ 0 & \cdots & 0 & \mathbf{P}_m & 0 \\ \mathbf{R}_1 & \mathbf{R}_2 & \cdots & \mathbf{R}_m & \mathbf{Q}_T \end{bmatrix} \begin{matrix} C_1 \\ C_2 \\ \vdots \\ C_m \\ T \end{matrix}$$

Since each C_j is a closed set, row sum of each \mathbf{P}_j is 1. That is to say each C_j is a irreducible MC, particle starting from C_j or move from T to C_j will always stay in C_j . One also has

$$\mathbf{P}^{(n)} = \begin{bmatrix} C_1 & C_2 & \cdots & C_m & T \\ \mathbf{P}_1^n & 0 & \cdots & \cdots & 0 \\ 0 & \mathbf{P}_2^n & \ddots & \ddots & \vdots \\ \vdots & \ddots & \ddots & \ddots & \vdots \\ 0 & \cdots & 0 & \mathbf{P}_m^n & 0 \\ \mathbf{R}_1^{(n)} & \mathbf{R}_2^{(n)} & \cdots & \mathbf{R}_m^{(n)} & \mathbf{Q}_T^n \end{bmatrix} \begin{matrix} C_1 \\ C_2 \\ \vdots \\ C_m \\ T \end{matrix}$$

after n -step transition. Row sum of each \mathbf{P}_j^n is 1. Moreover, $\lim_{n \rightarrow \infty} \mathbf{Q}_T^n = \lim_{n \rightarrow \infty} (q_{ij}^{(n)})_{i,j \in T} = 0$, any transient state can not be stayed for long time.

任何一个非常返态不可久留

Theorem 3.5.10 Part (8) Corollary of (3) and (4). If C^+ has at least two positive recurrent classes, C_1 and C_2 , then there are infinitely many stationary distribution.

Proof. Assume that C_1 and C_2 correspond to transition matrices \mathbf{P}_1 and \mathbf{P}_2 . Then there are stationary distributions, π_1 and π_2 , s.t.

$$\pi_1 = \pi_1 \mathbf{P}_1, \pi_2 = \pi_2 \mathbf{P}_2.$$

For any $r = 1 - s \in [0, 1]$, define

$$\pi = [r\pi_1, s\pi_2, 0, \dots, 0].$$

Then,

$$\begin{aligned} \pi \mathbf{P} &= [r\pi_1, s\pi_2, 0, \dots, 0] \begin{bmatrix} \mathbf{P}_1 & 0 & \cdots & \cdots & 0 \\ 0 & \mathbf{P}_2 & \ddots & \ddots & \vdots \\ \vdots & \ddots & \ddots & \ddots & \vdots \\ 0 & \cdots & 0 & \mathbf{P}_m & 0 \\ \mathbf{R}_1 & \mathbf{R}_2 & \cdots & \mathbf{R}_m & \mathbf{Q}_T \end{bmatrix} \\ &= [r\pi_1 \mathbf{P}_1, s\pi_2 \mathbf{P}_2, 0, \dots, 0] \\ &= [r\pi_1, s\pi_2, 0, \dots, 0] = \pi. \end{aligned}$$

■

3.5.2 The limiting behavior of transition probability matrix

Here we restate the results for $\lim_{n \rightarrow \infty} p_{ij}^{(n)}$.

Theorem 3.5.11 *If j is transient or null recurrent, then for any $i \in I$,*

$$\lim_{n \rightarrow \infty} p_{ij}^{(n)} = 0.$$

Corollary 3.5.12

- (1) *Finite-state Markov chain has not null recurrent states.*
- (2) *All states of a finite-state irreducible Markov chain are positive recurrent.*
- (3) *If a Markov chain has a null recurrent state, then it has infinitely many null recurrent states.*

Remark 3.5.13 *If j is positive recurrent, then $\lim_{n \rightarrow \infty} p_{ij}^{(n)}$ may not exist. Even it exists, it may be related to state i .*

Theorem 3.5.14 *If j is ergodic, then for any $i \in I$,*

$$\lim_{n \rightarrow \infty} p_{ij}^{(n)} = \frac{f_{ij}}{\mu_j}.$$

Corollary 3.5.15 *For a irreducible ergodic chain, for any $i, j \in I$,*

$$\lim_{n \rightarrow \infty} p_{ij}^{(n)} = \frac{1}{\mu_j}.$$

3.5.3 limiting probabilities and limiting distributions

Definition 3.5.16 *If $\lim_{n \rightarrow \infty} \pi_j^{(n)} = \pi_j^*$ ($j \in I$) exists, then $\pi^* = (\pi_1^*, \pi_2^*, \dots)$ is called the limiting distribution of the Markov chain.*

Theorem 3.5.17 *For an aperiodic irreducible Markov chain, all states are positive recurrent if and only if the chain has a stationary distribution. Moreover, the stationary distribution is the limiting distribution.*

Proof. \Leftarrow Let the stationary distribution be $\pi = (\pi_1, \pi_2, \dots)$, so that $\pi = \pi\mathbf{P} = \dots = \pi\mathbf{P}^n$, that is,

$$\pi_j = \sum_{i \in I} \pi_i p_{ij}^{(n)}.$$

Using Dominated Convergence Theorem, one can interchange the limitation and the summation for each $j \in I$,

$$\pi_j = \lim_{n \rightarrow \infty} \sum_{i \in I} \pi_i p_{ij}^{(n)} = \sum_{i \in I} \pi_i \lim_{n \rightarrow \infty} p_{ij}^{(n)} = \sum_{i \in I} \pi_i \cdot 0 = 0,$$

if we assume that all states are either null recurrent or transient. However, since $\sum_{j \in I} \pi_j = 1$, at least one $\pi_j > 0$ is strictly positive. Contradiction! Thus, at least one state is positive recurrent. Since the Markov chain is irreducible, all states are positive recurrent.

\Rightarrow Proved in Theorem 3.5.3.

Moreover, the conclusion that the stationary distribution is the limiting distribution is also proved in Theorem 3.5.3. ■

3.5.4 Examples for stationary distributions and limiting probabilities

Example 3.5.18 Let the state space of a Markov chain be $I = \{1, 2\}$ and the transition prob. matrix

$$\mathbf{P} = \begin{bmatrix} 3/4 & 1/4 \\ 5/8 & 3/8 \end{bmatrix}.$$

(1) Compute the stationary distribution π and the limiting $\lim_{n \rightarrow \infty} \mathbf{P}^n$;

Sol. Since $\pi = \pi \mathbf{P}$, compute the eigenvector of \mathbf{P}^T corresponding to eigenvalue 1:

$$(\mathbf{P}^T - \mathbf{I}) \rightarrow \begin{pmatrix} 1 & -5/2 \\ 0 & 0 \end{pmatrix}.$$

Thus

$$\pi \propto \begin{pmatrix} 5/2 \\ 1 \end{pmatrix} = \begin{pmatrix} 5/7 \\ 2/7 \end{pmatrix},$$

and

$$\lim_{n \rightarrow \infty} \mathbf{P}^n = \begin{pmatrix} \pi_1 & \pi_2 \\ \pi_1 & \pi_2 \end{pmatrix} = \begin{pmatrix} 5/7 & 2/7 \\ 5/7 & 2/7 \end{pmatrix}.$$

(2) Compute the mean recurrence time μ_1 and μ_2 .

Sol. Based on the formula, $\pi_j = 1/\mu_j$, so that $\mu_1 = 7/5$ and $\mu_2 = 7/2$.

Example 3.5.19 Ehrenfest model. Suppose that $2a$ molecules are distributed among two urns; and each time point one of the molecules is chosen at random, removed from its urn, and placed in the other one. Let X_n be the number of molecules in urn 1. The number of molecules in urn 1 is the Markov chain having transition prob.

$$p_{ij} = \begin{cases} \frac{2a-i}{2a}, & 0 \leq i \leq 2a-1, j = i+1. \\ \frac{i}{2a}, & 1 \leq i \leq 2a, j = i-1. \\ 0, & \text{else.} \end{cases}$$

Compute the stationary distribution π .

Sol. This is a positive recurrent Markov chain with period 2, and thus there exists a unique stationary distribution. Define $\pi_{-1} = \pi_{2a+1} = 0$. Based on $\pi = \pi \mathbf{P}$,

$$\pi_i = \pi_{i-1} p_{i-1,i} + \pi_{i+1} p_{i+1,i}, 0 \leq i \leq 2a.$$

Then

$$\pi_{i+1} = \frac{\pi_i - \pi_{i-1} p_{i-1,i}}{p_{i+1,i}}.$$

Based on computation recursively

$$\begin{aligned} \pi_1 &= \frac{\pi_0}{p_{10}} = 2a\pi_0 = C_{2a}^1 \pi_0, \\ \pi_2 &= (\pi_1 - \pi_0)2a/2 = (2a-1)a\pi_0 = C_{2a}^2 \pi_0, \\ &\dots \\ \pi_{2a} &= C_{2a}^{2a} \pi_0. \end{aligned}$$

By induction,

$$\begin{aligned}
\pi_{i+1} &= \frac{\pi_i - \pi_{i-1} p_{i-1,i}}{p_{i+1,i}} = \frac{C_{2a}^i \pi_0 - C_{2a}^{i-1} \pi_0 \frac{2a-(i-1)}{2a}}{\frac{i+1}{2a}} \\
&= \frac{\pi_0}{i+1} (2aC_{2a}^i - (2a-i+1)C_{2a}^{i-1}) = \frac{\pi_0}{i+1} \left(\frac{(2a) \cdot (2a)!}{i!(2a-i)!} - \frac{(2a) \cdot (2a+1-i)}{(i-1)!(2a+1-i)!} \right) \\
&= \frac{\pi_0}{i+1} \frac{(2a)!}{(i-1)!(2a-i)!} \left(\frac{2a}{i} - 1 \right) = \frac{\pi_0 (2a)!(2a-i)}{(i+1)(i-1)!(2a-i)!i} = C_{2a}^{i+1} \pi_0.
\end{aligned}$$

Using $\pi_0 + \dots + \pi_{2a} = 2^{2a} \pi_0 = 1$, we obtain

$$\pi_i = C_{2a}^i \left(\frac{1}{2}\right)^i \left(\frac{1}{2}\right)^{2a-i}, \quad 0 \leq i \leq 2a.$$

This is a Binomial with $B(2a, 1/2)$ since $X_1 + \dots + X_{2a} \sim B(2a, 1/2)$.

Example 3.5.20 5. Example 4.21 ignored. too simple.

Example 3.5.21 6. **The Hardy-Weinberg Law and a Markov chain in Genetics.** Consider a large population of individuals, where each individual gene is either type A or type a. Assume that the proportions of individuals whose gene pairs are AA, aa, or Aa are p_0, q_0, r_0 ($p_0 + q_0 + r_0 = 1$). When two individuals mate, each contributes one of his or her genes, chosen at random, to the offspring.

By conditioning on the gene pair of the parent, we see that for the first generation, a randomly chosen gene will be type A with prob.

$$\begin{aligned}
P\{A\} &= P\{A|AA\}P\{AA\} + P\{A|aa\}P\{aa\} + P\{A|Aa\}P\{Aa\} \\
&= p_0 + \frac{1}{2}r_0.
\end{aligned}$$

Similarly, it will be type a with prob.

$$P\{a\} = q_0 + \frac{1}{2}r_0.$$

Thus, under random mating a randomly chosen member of the next generation will be type AA with prob., where

$$p = P\{A\}P\{A\} = \left(p_0 + \frac{1}{2}r_0\right)^2.$$

The prob. for aa is

$$q = P\{a\}P\{a\} = \left(q_0 + \frac{1}{2}r_0\right)^2.$$

The prob. for Aa is

$$r = 2P\{A\}P\{a\} = 2 \left(p_0 + \frac{1}{2}r_0\right) \left(q_0 + \frac{1}{2}r_0\right),$$

where we notice that $p + q + r = 1$.

We notice that **the proportion of A will be unchanged from the previous generation:**

$$\begin{aligned}
P\{A^{new}\} &= p + \frac{1}{2}r = \left(p_0 + \frac{1}{2}r_0\right)^2 + \left(p_0 + \frac{1}{2}r_0\right) \left(q_0 + \frac{1}{2}r_0\right) \\
&= p_0 + \frac{1}{2}r_0 = P\{A^{old}\}.
\end{aligned}$$

Similarly,

$$P\{a^{new}\} = q + \frac{1}{2}r = q_0 + \frac{1}{2}r_0 = P\{a^{old}\}.$$

From this it follows that, **under random mating, in all successive generations after the initial one, the percentages of the population having gene pairs AA,aa, and Aa will remain fixed at the values p, q, and r. This is known as the Hardy-Weinberg law.**

For instance, see the following example:

	1st generation	2nd generation
$p_0 = 0.1$	$p_1 = 0.25$	$p_2 = 0.25$
$q_0 = 0.1$	$q_1 = 0.25$	$q_2 = 0.25$
$r_0 = 0.8$	$r_1 = 0.5$	$r_2 = 0.5$

Let us now see the problem from another point of view. For a given individual, let X_n denote the genetic state of her descendant in the nth generation. The transition prob. matrix of the Markov chain is

$$\begin{array}{ccc} & AA & aa \\ AA & \left[\begin{array}{ccc} p+r/2 & 0 & q+r/2 \\ 0 & q+r/2 & p+r/2 \\ p/2+r/4 & q/2+r/4 & p/2+q/2+r/2 \end{array} \right] \\ aa & & \\ Aa & & \end{array},$$

where $AA \rightarrow AA$ since AA must contribute one A and another A comes from $P\{A\} = p+r/2$, and $Aa \rightarrow AA$ since $\frac{1}{2}P\{A\} = p/2+r/4$. Let us verify that the stationary limiting prob. of this Markov chain is p, q, r . It suffices to show that

$$\begin{aligned} p &= p(p + \frac{r}{2}) + r(\frac{p}{2} + \frac{r}{4}) = (p + \frac{r}{2})^2 = (p_0 + \frac{r_0}{2})^2, \\ q &= q(q + \frac{r}{2}) + r(\frac{q}{2} + \frac{r}{4}) = (q + \frac{r}{2})^2 = (q_0 + \frac{r_0}{2})^2, \\ p + q + r &= 1. \end{aligned}$$

Example 3.5.22 7. 4.5.1 *The gambler's ruin problem ignored.* Introduced in the following.

Example 3.5.23 8. 4.5.2 *ignored since too hard.*

3.6 Mean Time Spent in Transient States

See reference in [7].

\mathbf{P} transition prob. matrix for MC $\{X_n\}$ and \mathbf{Q} is submatrix of \mathbf{P} which includes only transient states.

$$\mathbf{P} = \begin{bmatrix} \tilde{\mathbf{P}}_0 & 0 \\ \mathbf{S} & \mathbf{Q} \end{bmatrix}, \quad \mathbf{P}^n = \begin{bmatrix} \tilde{\mathbf{P}}_0^n & 0 \\ \mathbf{S}_n & \mathbf{Q}^n \end{bmatrix}.$$

Example 3.6.1 5 states with two absorbing boundaries. $S = \{0, 4, 1, 2, 3\}$.

$$\mathbf{P} = \begin{array}{cc|ccc} & & 0 & 4 & 1 & 2 & 3 \\ \hline 0 & & 1 & 0 & 0 & 0 & 0 \\ 4 & & 0 & 1 & 0 & 0 & 0 \\ \hline 1 & & \frac{1}{2} & 0 & 0 & \frac{1}{2} & 0 \\ 2 & & 0 & 0 & \frac{1}{2} & 0 & \frac{1}{2} \\ 3 & & 0 & \frac{1}{2} & 0 & \frac{1}{2} & 0 \end{array}$$

$$\mathbf{Q} = \begin{array}{cc|ccc} & & 1 & 2 & 3 \\ \hline 1 & & 0 & \frac{1}{2} & 0 \\ 2 & & \frac{1}{2} & 0 & \frac{1}{2} \\ 3 & & 0 & \frac{1}{2} & 0 \end{array}.$$

\mathbf{Q} is called a substochastic matrix, i.e., a matrix with nonnegative entries whose row sums are less than or equal to 1.

\mathbf{Q} is transient $\Rightarrow \mathbf{Q}^n = 0$ as $n \rightarrow \infty$. \Rightarrow All eigenvalues of \mathbf{Q} have absolute values strictly less than 1.
 $\Rightarrow \mathbf{I} - \mathbf{Q}$ is an invertible matrix. $\Rightarrow \mathbf{S} := (\mathbf{I} - \mathbf{Q})^{-1}$ is well-defined.

Let j be a T state and consider Y_j the total number of visits to j ,

$$Y_j = \sum_{n=0}^{\infty} I\{X_n = j\}.$$

这些事件可以相容 Since j is T , $Y_j < \infty$ with prob. 1. Suppose $X_0 = i$, where i another T . Then

$$\begin{aligned} s_{ij} &= E(Y_j | X_0 = i) = E\left(\sum_{n=0}^{\infty} I\{X_n = j\} | X_0 = i\right) \\ &= \sum_{n=0}^{\infty} P(X_n = j | X_0 = i) = \sum_{n=0}^{\infty} p_{ij}^{(n)}, \quad 0 \text{ included.} \end{aligned}$$

That is, $E(Y_j | X_0 = i)$ is the (i, j) entry of matrix

$$\mathbf{I} + \mathbf{P} + \mathbf{P}^2 + \cdots,$$

which is the same as the (i, j) entry of matrix

$$\mathbf{I} + \mathbf{Q} + \mathbf{Q}^2 + \cdots \quad (\text{because Block property}).$$

However, a simple calculation shows that

$$\mathbf{I} + \mathbf{Q} + \mathbf{Q}^2 + \cdots = (\mathbf{I} - \mathbf{Q})^{-1} = \mathbf{S}.$$

Theorem 3.6.2 Let i, j be T . Then

- (1) $s_{ij} = [\mathbf{S}]_{ij} = [(\mathbf{I} - \mathbf{Q})^{-1}]_{ij}$ is the expected number of visits to j starting at i .
- (2) The expected number of steps until the chain enters a recurrent class can be computed by summing s_{ij} over all transient j in one T class.

Example 3.6.3 *cont.*

$$\mathbf{S} = (\mathbf{I} - \mathbf{Q})^{-1} = \begin{bmatrix} 3/2 & 1 & 1/2 \\ 1 & 2 & 1 \\ 1/2 & 1 & 3/2 \end{bmatrix}.$$

The expected number of visits from 1 to 3 is 1/2. The expected number of visits from 1 until absorption is $\frac{3}{2} + 1 + \frac{1}{2} = 3$.

Another derivation of the above result. For $i, j \in T$, let s_{ij} be the expected number of visits that the MC is in j , starting from i . Let

$$\begin{aligned}\delta_{ij} &= 1, & i = j \\ \delta_{ij} &= 0, & i \neq j\end{aligned}$$

Then

$$\begin{aligned}s_{ij} &= E(Y_j | X_0 = i) = \delta_{ij} + \sum_{k \in I} E(Y_j | X_1 = k, X_0 = i) P(X_1 = k | X_0 = i) \\ &= \delta_{ij} + \sum_{k \in I} p_{ik} s_{kj} = \delta_{ij} + \sum_{k \in T} p_{ik} s_{kj} \quad (\text{since if } k \rightarrow j \in T, \text{ then } k \in T).\end{aligned}$$

In matrix form,

$$\mathbf{S} = (\mathbf{I} - \mathbf{Q})^{-1}. \quad (\text{notice that } s_{jj} \geq 1).$$

In the following, we derive another useful formula. For $i, j \in T$, the quantity f_{ij} , equal to the prob. that the MC ever makes a transition into state j given that it starts in state i . Derive an expression for s_{ij} by conditioning on whether state j is ever entered,

$$A_1 = \{\text{ever transit to } j\}, A_2 = \{\text{never transit to } j\}.$$

$$\begin{aligned}s_{ij} &= E(\text{times in } j | \text{start in } i, \text{ ever transit to } j) f_{ij} \\ &\quad + E(\text{times in } j | \text{start in } i, \text{ never transit to } j) (1 - f_{ij}) \\ &= (\delta_{ij} + s_{jj}) f_{ij} + \delta_{ij} (1 - f_{ij}) \\ &= \delta_{ij} + f_{ij} s_{jj}.\end{aligned}$$

Thus

$$f_{ij} = \frac{s_{ij} - \delta_{ij}}{s_{jj}}.$$

Example 3.6.4 (Gambler Ruin Problem) Consider a gambler who at each play has prob. p of winning one unit and prob. $q = 1 - p$ of losing one unit. Assuming that successive plays are independent. Now given $p = 0.4$ and $N = 7$. Starting with 3 units, what is the prob. that the gambler ever has a fortune of 1?

Sol. We have a 6×6 matrix

$$\mathbf{Q} = \begin{bmatrix} & & 0.4 & & & \\ & 0.6 & & 0.4 & & \\ & & 0.6 & & 0.4 & \\ & & & 0.6 & & 0.4 \\ & & & & 0.6 & \\ & & & & & 0.6 \end{bmatrix},$$

then one obtains $\mathbf{S} = (\mathbf{I} - \mathbf{Q})^{-1}$, and hence $s_{3,1} = 1.4206$ and $s_{1,1} = 1.6149$. Then

$$f_{3,1} = \frac{s_{3,1}}{s_{1,1}} = 0.8797.$$

Another way is, it is the prob. that the gambler's fortune will go down 2 before going up 4, which is the prob. that a gambler starting with 2 will go broke before reaching 6. Therefore,

$$f_{3,1} = 1 - \frac{1 - (0.6/0.4)^2}{1 - (0.6/0.4)^6} = 0.8797.$$

or

$$f_{3,1} = \frac{1 - (0.4/0.6)^4}{1 - (0.4/0.6)^6} = 0.8797.$$

Example 3.6.5 (Gambler Ruin Problem) Consider a gambler who at each play has prob. p of winning one unit and prob. $q = 1 - p$ of losing one unit. Assuming that successive plays are independent, what is the prob. that, starting with i units, the gambler's fortune will reach N before reaching 0?

Sol. Let X_n denote the fortune at time n , then $\{X_n\}$ is the MC process with transition prob.

$$\begin{aligned} p_{00} &= p_{NN} = 1, \\ p_{i,i+1} &= p = 1 - p_{i,i-1}, i = 1, 2, \dots, N-1. \end{aligned}$$

Let

$$Y = \bigcup_{n=0}^{\infty} A_n = \bigcup_{n=0}^{\infty} \{X_n = N, X_{n-1}, \dots, X_0 \neq N\}.$$

Since $\{1, 2, \dots, N-1\}$ is transient and $\{0\}, \{N\}$ are recurrent, the gambler will either attain goal N or go broke. Let P_i denote the prob. that, starting in i , the gambler's fortune will eventually reach N . By conditioning on the outcome of the initial play of the game we obtain

$$\begin{aligned} P(Y|X_0 = i) &= P(Y|X_0 = i, X_1 = i+1)P(X_1 = i+1|X_0 = i) \\ &\quad + P(Y|X_0 = i, X_1 = i-1)P(X_1 = i-1|X_0 = i). \end{aligned}$$

That is

$$\begin{aligned} P_i &= pP_{i+1} + qP_{i-1}, \quad i = 1, 2, \dots, N-1. \\ P_{i+1} - P_i &= \frac{q}{p}(P_i - P_{i-1}), \quad i = 1, 2, \dots, N-1. \end{aligned}$$

Hence, since $P_0 = 0$, we obtain

$$\begin{aligned} P_2 - P_1 &= \frac{q}{p}(P_1 - P_0) = \frac{q}{p}P_1, \\ P_N - P_{N-1} &= \frac{q}{p}(P_{N-1} - P_{N-2}) = \left(\frac{q}{p}\right)^{N-1} P_1. \end{aligned}$$

Adding together yields

$$P_i - P_1 = P_1 \left[\left(\frac{q}{p}\right) + \left(\frac{q}{p}\right)^2 + \dots + \left(\frac{q}{p}\right)^{i-1} \right].$$

or

$$P_i = \begin{cases} \frac{1-(q/p)^i}{1-(q/p)} P_1, & q/p \neq 1, \\ iP_1, & q/p = 1. \end{cases}$$

Using the fact that $P_N = 1$,

$$P_1 = \begin{cases} \frac{1-(q/p)}{1-(q/p)^N}, & q/p \neq 1, \\ 1/N, & q/p = 1. \end{cases}$$

and hence

$$P_i = \begin{cases} \frac{1-(q/p)^i}{1-(q/p)^N}, & q/p \neq 1, \\ i/N, & q/p = 1. \end{cases}$$

Note that as $N \rightarrow \infty$,

$$P_i \rightarrow \begin{cases} 1 - (q/p)^i, & p > 1/2, \\ 0, & p \leq 1/2. \end{cases}$$

Example 3.6.6 Determine the expected number of steps that an irreducible Markov chain takes to go from one state i to another state j . We change j to the first site

$$\mathbf{P} = \begin{bmatrix} p(j,j) & R \\ S & Q \end{bmatrix}.$$

We then change j to an absorbing state,

$$\tilde{\mathbf{P}} = \begin{bmatrix} 1 & 0 \\ S & Q \end{bmatrix}.$$

Let T_j be the number of steps needed to reach state j . In other words, T_j is the smallest time n such that $X_n = j$. For any other state k , let $T_{k,j}$ be the number of visits to k before reaching j (if we start at state k , we include this as one visit to k).

$$E(T_j | X_0 = i) = E\left(\sum_{k \neq j} T_{k,j} | X_0 = i\right) = \sum_{k \neq j} s_{ik}.$$

S1 gives a vector whose i th component is the number of steps starting at i until reaching j .

Example 3.6.7 A random walk with reflecting boundary, $\{0, 1, 2, 3, 4\}$.

$$\mathbf{P} = \begin{array}{c|ccccc} & 0 & 1 & 2 & 3 & 4 \\ \hline 0 & 0 & 1 & 0 & 0 & 0 \\ 1 & \frac{1}{2} & 0 & \frac{1}{2} & 0 & 0 \\ \hline 2 & 0 & \frac{1}{2} & 0 & \frac{1}{2} & 0 \\ 3 & 0 & 0 & \frac{1}{2} & 0 & \frac{1}{2} \\ 4 & 0 & 0 & 0 & 1 & 0 \end{array}.$$

If set finally at $j = 0$, then

$$\mathbf{Q} = \begin{array}{c|ccccc} & 1 & 2 & 3 & 4 \\ \hline 1 & 0 & \frac{1}{2} & 0 & 0 \\ 2 & \frac{1}{2} & 0 & \frac{1}{2} & 0 \\ 3 & 0 & \frac{1}{2} & 0 & \frac{1}{2} \\ 4 & 0 & 0 & 1 & 0 \end{array},$$

$$\mathbf{S} = (\mathbf{I} - \mathbf{Q})^{-1} = \begin{bmatrix} 2 & 2 & 2 & 1 \\ 2 & 4 & 4 & 2 \\ 2 & 4 & 6 & 3 \\ 2 & 4 & 6 & 4 \end{bmatrix}.$$

Then

$$S\mathbf{1} = \begin{bmatrix} 7 \\ 12 \\ 15 \\ 16 \end{bmatrix}.$$

Hence, the expected number of steps to get from 4 to 0 is 16.

3.7 Time Reversible Markov Chains

Consider a *stationary* ergodic Markov chain (that is, the chain has run for a long time) having transition prob. p_{ij} and stationary prob. π_i . Suppose that starting at some time n , we trace the states going backward in time. It turns out that the sequence of states is itself a Markov chain with transition prob. q_{ij} defined by

$$\begin{aligned} q_{ij} &= P\{X_m = j | X_{m+1} = i\} = \frac{P\{X_m = j, X_{m+1} = i\}}{P\{X_{m+1} = i\}} \\ &= \frac{P\{X_m = j\}P\{X_{m+1} = i | X_m = j\}}{P\{X_{m+1} = i\}} = \frac{\pi_j p_{ji}}{\pi_i}. \end{aligned}$$

To see the reversed process is Markov, we must verify that

$$P\{X_m = j | X_{m+1} = i, X_{m+2}, X_{m+3}, \dots\} = P\{X_m = j | X_{m+1} = i\}.$$

The reason is

$$P\{X_{m+2}, X_{m+3}, \dots | X_{m+1}, X_m, \dots\} = P\{X_{m+2}, X_{m+3}, \dots | X_{m+1}\} \Rightarrow X_{m+k}(X_{m+1}), k \geq 2,$$

thus

$$P\{X_m = j | X_{m+1} = i, X_{m+2}(X_{m+1}), X_{m+3}(X_{m+1}), \dots\} = P\{X_m = j | X_{m+1} = i\}.$$

Therefore, the reversed process is a Markov chain with transition prob. given by q_{ij} .

Definition 3.7.1 If $q_{ij} = p_{ij}$ for all i, j , then the Markov chain is said to be **time reversible**. The condition for time reversibility can also be expressed as a **detailed balance condition**,

$$\pi_i p_{ij} = \pi_j p_{ji}, \text{ for all } i, j. \quad (3.5)$$

The above condition can also be stated that, for all i, j , the rate at which the process goes from i to j (numely, $\pi_i p_{ij}$) is equal to the rate at which it goes from j to i (numely, $\pi_j p_{ji}$).

Proposition 3.7.2 If equation (3.5) has solution, then the solution is the limiting stationary prob. π_i . However, it is possible that (3.5) has no solution. This is so since if

$$x_i p_{ij} = x_j p_{ji}, \text{ for all } i, j, \sum_i x_i = 1,$$

then summing over i yields

$$\sum_i x_i p_{ij} = x_j \sum_i p_{ji} = x_j,$$

and, because the limiting stationary prob. π_i is the unique solution of the preceding, it follows that $x_i = \pi_i$ for all i .

Example 3.7.3 Consider a random walk with states $0, 1, \dots, M$ and transition prob.

$$\begin{aligned} p_{i,i+1} &= \alpha_i = 1 - p_{i,i-1}, \quad i = 1, \dots, M-1, \\ p_{0,1} &= \alpha_0 = 1 - p_{0,0}, \\ p_{M,M} &= \alpha_M = 1 - p_{M,M-1}. \end{aligned}$$

By observation, the MC is time reversible. This follows by noting that the number of transitions from i to $i+1$ must at all times be within 1 of the number from $i+1$ to i . This is so because between any two transitions from i to $i+1$ there must be one from $i+1$ to i . It follows that the rate of transitions from i to $i+1$ equals the rate from $i+1$ to i . Thus

$$\begin{aligned} \pi_0\alpha_0 &= \pi_1(1 - \alpha_1), \\ \pi_1\alpha_1 &= \pi_2(1 - \alpha_2), \\ &\vdots \\ \pi_i\alpha_i &= \pi_{i+1}(1 - \alpha_{i+1}), \quad i = 0, 1, \dots, M-1. \end{aligned}$$

Solve to get

$$\pi_i = \frac{\alpha_{i-1} \cdots \alpha_0}{(1 - \alpha_i) \cdots (1 - \alpha_1)} \pi_0, \quad i = 1, \dots, M.$$

Since $\sum_{i=0}^M \pi_i = 1$,

$$\pi_0 = \left[1 + \sum_{j=1}^M \frac{\alpha_{j-1} \cdots \alpha_0}{(1 - \alpha_j) \cdots (1 - \alpha_1)} \right]^{-1}$$

Example 3.7.4 One special case of above example is the Ehrenfest model. Suppose that M molecules are distributed among two urns; and each time point one of the molecules is chosen at random, removed from its urn, and placed in the other one. The number of molecules in urn 1 is a special case of the Markov chain having

$$\alpha_i = \frac{M-i}{M}, \quad i = 0, 1, \dots, M.$$

Hence

$$\begin{aligned} \pi_0 &= \left[1 + \sum_{j=1}^M \frac{(M-j+1) \cdots (M-1)M}{j(j-1) \cdots 1} \right]^{-1} \\ &= \left[\sum_{j=0}^M C_M^j \right]^{-1} = \left(\frac{1}{2} \right)^M, \end{aligned}$$

and then

$$\pi_i = C_M^i \left(\frac{1}{2} \right)^M, \quad i = 0, 1, \dots, M.$$

Example 3.7.5 Consider an undirected graph having a weight w_{ij} associated with (i, j) for each arc. If at any time the particle resides at node i , then it will next move to node j with prob. where

$$p_{ij} = \frac{w_{ij}}{\sum_j w_{ij}},$$

and where w_{ij} is 0 if (i, j) is not an arc. For instance, in Fig. 3.4, $p_{12} = 3/(3 + 1 + 2) = 1/2$. The time reversibility equation or the detailed balance condition reduces to

$$\pi_i \frac{w_{ij}}{\sum_j w_{ij}} = \pi_j \frac{w_{ji}}{\sum_i w_{ji}}.$$

Since $w_{ij} = w_{ji}$,

$$\frac{\pi_i}{\sum_j w_{ij}} = \frac{\pi_j}{\sum_i w_{ji}} = c.$$

Since $1 = \sum_i \pi_i$,

$$\pi_i = \frac{\sum_j w_{ij}}{\sum_i \sum_j w_{ij}}.$$

For the graph in Fig. 3.4, we have that

$$\pi_1 = \frac{6}{32}, \pi_2 = \frac{3}{32}, \pi_3 = \frac{6}{32}, \pi_4 = \frac{5}{32}, \pi_5 = \frac{12}{32}.$$

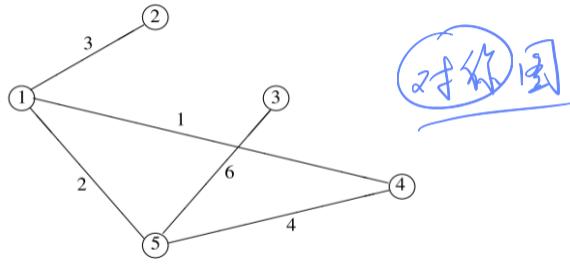


Figure 4.1 A connected graph with arc weights.

Figure 3.4: A connected graph with arc weights.

More about detailed balance condition. If we try to solve detailed balance condition for an arbitrary Markov chain with states $0, \dots, M$, it will usually turn out that no solution exists. For example,

$$\begin{aligned} x_i p_{ij} &= x_j p_{ji}, \\ x_k p_{kj} &= x_j p_{jk}, \end{aligned}$$

implies that

$$\frac{x_i}{x_k} = \frac{p_{ji} p_{kj}}{p_{ij} p_{jk}},$$

which in general not equal to p_{ki}/p_{ik} . Thus, we see that a necessary condition for time reversibility is that

$$p_{ik} p_{kj} p_{ji} = p_{ij} p_{jk} p_{ki}, \quad \text{for all } i, j, k.$$

See counterexample in Fig. 3.5.

Theorem 3.7.6 An ergodic Markov chain for which $p_{ij} = 0$ whenever $p_{ji} = 0$ is time reversible if and only if starting in state i , any path back to i has the same prob. as the reversed path. That is, if

$$p_{i,i_1} p_{i_1 i_2} \cdots p_{i_k, i} = p_{i, i_k} p_{i_k, i_{k-1}} \cdots p_{i_1, i},$$

for all states i, i_1, \dots, i_k .

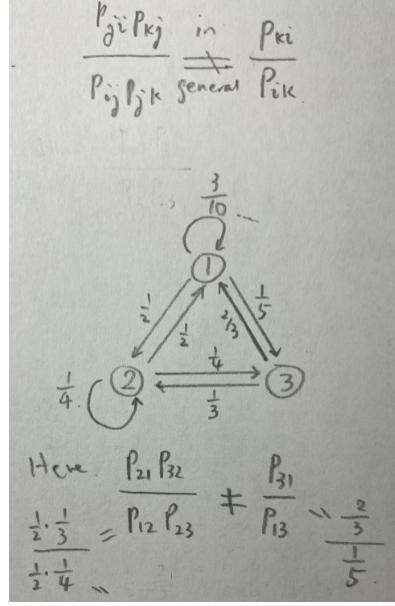


Figure 3.5: Detailed balance condition has no solution.

Proof. We have already proven necessity. To prove sufficiency, fix states i and j and rewrite above equation as

$$p_{i,i_1} p_{i_1 i_2} \cdots p_{i_k j} p_{j i} = p_{ij} p_{j,i_k} p_{i_k, i_{k-1}} \cdots p_{i_1, i}.$$

Summing the preceding over all states i_1, \dots, i_k yields

$$p_{ij}^{(k+1)} p_{ji} = p_{ij} p_{ji}^{(k+1)}.$$

Letting $k \rightarrow \infty$ yields

$$\pi_j p_{ji} = p_{ij} \pi_i,$$

which proves the theorem. ■

Example 3.7.7 (PageRank) With the development of Internet, many companies are seeking the methodology of judging the popularity of all websites. This is called PageRank, which is proposed by Page and his colleagues in 1998. They first labelled all the websites to obtain the state space $I = \{1, 2, \dots, n\}$. When i th website has link to the other $m(i)$ number of websites, then we define the transition prob. from i to j as

$$p_{ij} = \begin{cases} \frac{1}{m(i)}, & \text{when } i \text{ is accessible to } j, \\ 0, & \text{else.} \end{cases}$$

Obviously,

$$\sum_{j \in I} p_{ij} = 1.$$

Since MC has finite states, then all states are positive recurrent when all states communicate. Assume aperiodic. Thus ergodic finite-state Markov chain gives a unique stationary distribution.

For the stationary distribution $\pi = \{\pi_j\}$, π_j reflects the visiting prob. for the website j . When j th website becomes more popular, it will be visited more times, and thus π_j reflects the popularity of the website j . One

can rank the website based on the each π_j .

Not all websites communicate or some garbage websites construct many links to themselves via virus, and thus the model needs to be improved. For example, the garbage website usually is visited for a short period, so that one can design a more reasonable prob. model. Also for example, for those websites having no links to others, one can modify the transition prob. matrix to be a weighted one, $\tilde{\mathbf{P}} = \alpha \mathbf{P} + \frac{1-\alpha}{n} \mathbf{E}$.

3.8 Hidden Markov Model

3.8.1 Introduction to HMM

- Let $\{X_n, n = 1, 2, \dots\}$ be a Markov chain with transition prob. p_{ij} and initial state prob. $\pi_i^{(0)} = P(X_0 = i)$ $i \geq 0$.
- A signal from Φ is emitted each time the MC enters a state. See Fig. 3.6.

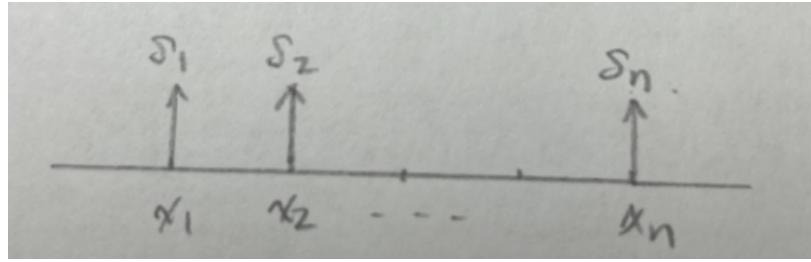


Figure 3.6: Hidden Markov Model.

- We have

$$P(S_1 = s | X_1 = j) = q(s|j), \quad \sum_{s \in \Phi} q(s|j) = 1,$$

$$P(S_n = s | X_1, S_1, \dots, X_{n-1}, S_{n-1}, X_n = j) = P(S_n = s | X_n = j) = q(s|j).$$

- S_1, S_2, \dots are observed while X_1, X_2, \dots are not.

This is called a Hidden Markov Model.

Example 3.8.1 A machine has 1 good state and 2 poor state. The transition prob. is

$$\mathbf{P} = \begin{matrix} & \begin{matrix} 1 & 2 \end{matrix} \\ \begin{matrix} 1 \\ 2 \end{matrix} & \left[\begin{matrix} 0.9 & 0.1 \\ 0 & 1 \end{matrix} \right] \end{matrix}.$$

Each item produced is of acceptable quality with prob. 0.99 when the process is in state 1.

Each item produced is of acceptable quality with prob. 0.96 when the process is in state 2.

If item accepted or not can be observed, while the state of the machine cannot be observed (HMM).

$$q(a|1) = 0.99, \quad q(u|1) = 0.01,$$

$$q(a|2) = 0.96, \quad q(u|2) = 0.04.$$

- $\{S_n, n \geq 1\}$ is not MC.
 - $\vec{S}^n = (S_1, \dots, S_n)$ is a random variable.
 - $\vec{s}_k = (s_1, \dots, s_k)$ is one observation.
- Define $F_n(j) = P(\vec{S}^n = \vec{s}_n, X_n = j)$, then

$$P(X_n = j | \vec{S}^n = \vec{s}_n) = \frac{P(X_n = j, \vec{S}^n = \vec{s}_n)}{P(\vec{S}^n = \vec{s}_n)} = \frac{F_n(j)}{\sum_i F_n(i)}.$$

Now,

$$\begin{aligned} F_n(j) &= P(\vec{S}^{n-1} = \vec{s}_{n-1}, S_n = s_n, X_n = j) \\ &= \sum_i P(\vec{S}^{n-1} = \vec{s}_{n-1}, X_{n-1} = i, S_n = s_n, X_n = j) \\ &= \sum_i P(\vec{S}^{n-1} = \vec{s}_{n-1}, X_{n-1} = i) P(S_n = s_n, X_n = j | \vec{S}^{n-1} = \vec{s}_{n-1}, X_{n-1} = i) \\ &= \sum_i F_{n-1}(i) P(S_n = s_n, X_n = j | X_{n-1} = i). \end{aligned}$$

Notice that

$$\begin{aligned} P(S_n = s_n, X_n = j | X_{n-1} = i) &= P(X_n = j | X_{n-1} = i) P(S_n = s_n | X_n = j, X_{n-1} = i) \\ &= p_{ij} q(s_n | j). \end{aligned}$$

Then

$$F_n(j) = \sum_i F_{n-1}(i) p_{ij} q(s_n | j) = q(s_n | j) \sum_i F_{n-1}(i) p_{ij}. \quad (3.6)$$

Starting with

$$\begin{aligned} F_1(i) &= P(S_1 = s_1, X_1 = i) = P(S^1 = s_1 | X_1 = i) P(X_1 = i) \\ &= \pi_i^{(0)} q(s_1 | i). \end{aligned} \quad (3.7)$$

We can use Eq. (3.6) to recursively determine $F_2(i), F_3(i), \dots$ up to $F_n(i)$.

Example 3.8.2 (Cont.) $P(X_1) = 0.8$. The first 3 items are a,u,a.

- (1) What is the prob. that the process was in good state when the 3rd item was produced?
- (2) What is the prob. that X_4 is 1?
- (3) What is the prob. that the next item produced is accepted?

Sol. $\vec{s}_3 = \{a, u, a\}$. We have

$$\begin{aligned}
F_1(i=1) &= \pi_{i=1}^{(0)} q(a|i=1) = (0.8)(0.99) = 0.792. \\
F_1(i=2) &= \pi_{i=2}^{(0)} q(a|i=2) = (0.2)(0.96) = 0.192. \\
F_2(i=1) &= q(u|1)[F_1(1)p_{11} + F_1(2)p_{21}] = (0.01)[(0.792)(0.9) + (0.192)(0)] \\
&= 0.007128. \\
F_2(i=2) &= q(u|2)[F_1(1)p_{12} + F_1(2)p_{22}] = (0.04)[(0.792)(0.1) + (0.192)(1)] \\
&= 0.010848. \\
F_3(i=1) &= q(a|1)[F_2(1)p_{11} + F_2(2)p_{21}] = (0.99)[(0.007128)(0.9) + (0.010848)(0)] \\
&= 0.006351. \\
F_3(i=2) &= q(a|2)[F_2(1)p_{12} + F_2(2)p_{22}] = (0.96)[(0.007128)(0.1) + (0.010848)(0)] \\
&= 0.011098.
\end{aligned}$$

(a) We have

$$\begin{aligned}
P(X_3=1|\vec{s}_3) &= \frac{P(X_3=1, \vec{s}_3)}{\sum_i P(X_3=i, |\vec{s}_3)} = \frac{F_3(1)}{F_3(1) + F_3(2)} = \frac{0.006351}{0.006351 + 0.011098} \\
&= 0.364.
\end{aligned}$$

(b) Conditioning on X_3 , we have

$$\begin{aligned}
P(X_4=1|\vec{s}_3) &= P(X_4=1|X_3=1, \vec{s}_3)P(X_3=1|\vec{s}_3) + P(X_4=1|X_3=2, \vec{s}_3)P(X_3=2|\vec{s}_3) \\
&= (p_{11})(0.364) + (p_{21})(1 - 0.364) = 0.3276.
\end{aligned}$$

(c) Conditioning on X_4 ,

$$\begin{aligned}
P(S_4=a|\vec{s}_3) &= P(S_4=a|X_4=1, \vec{s}_3)P(X_4=1|\vec{s}_3) + P(S_4=a|X_4=2, \vec{s}_3)P(X_4=2|\vec{s}_3) \\
&= P(S_4=a|X_4=1)(0.3276) + P(S_4=a|X_4=2)(1 - 0.3276) \\
&= (0.99)(0.3276) + (0.96)(1 - 0.3276) = 0.9698.
\end{aligned}$$

See reference in [9].

3.8.2 Key ingredients to HMM

- (1) transition prob. matrix $p_{ij} = P(X_{n+1}=j|X_n=i)$.
- (2) observation prob. distribution. $q(s|j) = P(S_n=s|X_n=j)$.
- (3) initial state distribution. $\pi_i^{(0)} = P(X_0=i)$.

Compact notation: $\lambda = (p, q, \pi)$.

3.8.3 Three Basic Problems for HMMs

Compared to the traditional Hidden Markov Model (HMM), the RNN, LSTM become more popular nowadays and receive more attentions. In order for HMM model useful in real-world application [9], we present 3 basis problems:

- (1) (probability computation problem) Given observation $\vec{s}_n = (s_1, \dots, s_n)$ and a model $\lambda = (p, q, \pi)$, how to efficiently compute $P(\vec{s}_n | \lambda)$, the prob. of the observation sequence given the model?
- (2) (decoding problem) Given observation $\vec{s}_n = (s_1, \dots, s_n)$ and a model $\lambda = (p, q, \pi)$, how to choose a state sequence $X = (X_1, \dots, X_n)$ which is optimal in some meaningful sense?
- (3) (learning problem or inverse problem) Given observation $\vec{s}_n = (s_1, \dots, s_n)$, how do we adjust or estimate the model parameter $\lambda = (p, q, \pi)$ to maximize the likelihood $P(\vec{s}_n | \lambda)$?

3.8.4 Forward and Backward Approaches for Problem 1

Forward approach

Compute $P(\vec{S}^n = \vec{s}_n)$ by $= \sum_i F_n(i)$ for $i = 1, \dots, N$ and recursive formula,

$$F_n(j) = q(s_n | j) \sum_i F_{n-1}(i)p_{ij}.$$

Notice that $F_1(i = 1), \dots, F_1(i = N)$ is $O(N)$, $F_2(i = 1), \dots, F_2(i = N)$ is $O(N)$, gives total complexity is $O(nN^2)$.

Direct computation

by conditioning on the first n states of the Markov chain.

$$\begin{aligned} P(\vec{S}^n = \vec{s}_n) &= \sum_{i_1 \dots i_n} P(\vec{S}^n = \vec{s}_n | X_1 = i_1, \dots, X_n = i_n) P(X_1 = i_1, \dots, X_n = i_n) \\ &= \sum_{i_1 \dots i_n} q(s_1 | i_1) \cdots q(s_n | i_n) \pi_{i_1}^{(0)} p_{i_1 i_2} \cdots p_{i_{n-1} i_n}, \end{aligned}$$

where made use of

$$\begin{aligned} P(X_1, \dots, X_n) &= P(X_n | X_{n-1}) P(X_{n-1} | X_{n-2}) \cdots P(X_2 | X_1) P(X_1), \\ P(S_1 \dots S_n | \underbrace{X_1 \dots X_n}_A) &= P(S_n | S_1 \dots S_{n-1} A) P(S_{n-1} | S_1 \dots S_{n-2} A) \cdots P(S_1 | A). \end{aligned}$$

The complexity is $O(N^n)$.

Backward approach

Define the quantity

$$B_k(i) = P(S_{k+1} = s_{k+1}, \dots, S_n = s_n | X_k = i).$$

A recursive formula for $B_k(i)$ can be obtained by conditioning on X_{k+1} ,

$$\begin{aligned} B_k(i) &= \sum_j P(S_{k+1} = s_{k+1}, \dots, S_n = s_n | X_k = i, X_{k+1} = j) P(X_{k+1} = j | X_k = i) \\ &= \sum_j P(S_{k+1} = s_{k+1}, \dots, S_n = s_n | X_{k+1} = j) p_{ij} \\ &= \sum_j P(S_{k+1} = s_{k+1} | X_{k+1} = j) P(S_{k+2} = s_{k+2}, \dots, S_n = s_n | S_{k+1} = s_{k+1}, X_{k+1} = j) p_{ij} \\ &= \sum_j q(s_{k+1} | j) P(S_{k+2} = s_{k+2}, \dots, S_n = s_n | X_{k+1} = j) p_{ij} \\ &= \sum_j q(s_{k+1} | j) B_{k+1}(j) p_{ij}. \end{aligned}$$

The initial condition is

$$\begin{aligned} B_{n-1}(i) &= P(S_n = s_n | X_{n-1} = i) = \sum_j P(S_n = s_n | X_{n-1} = i, X_n = j)P(X_n = j | X_{n-1} = i) \\ &= \sum_j p_{ij}q(s_n | j). \end{aligned}$$

Determine the functions $B_{n-2}(i), B_{n-3}(i), \dots$ subsequently. Finally,

$$\begin{aligned} P(\vec{S}^n = \vec{s}_n) &= \sum_i P(\vec{S}^n = \vec{s}_n | X_1 = i)\pi_i^{(0)} \\ &= \sum_i P(S_1 = s_1 | X_1 = i)P(S_2 = s_2, \dots, S_n = s_n | S_1 = s_1, X_1 = i)\pi_i^{(0)} \\ &= \sum_i q(s_1 | i)P(S_2 = s_2, \dots, S_n = s_n | X_1 = i)\pi_i^{(0)} \\ &= \sum_i q(s_1 | i)B_1(i)\pi_i^{(0)}. \end{aligned}$$

Both Forward and Backward approaches

Suppose that we have computed $F_k(j)$ and $B_k(j)$ for some k ,

$$\begin{aligned} P(\vec{S}^n = \vec{s}_n, X_k = j) &= P(\vec{S}^k = \vec{s}_k, X_k = j)P(S_{k+1}, \dots, S_n | \vec{S}^k = \vec{s}_k, X_k = j) \\ &= P(\vec{S}^k = \vec{s}_k, X_k = j)P(S_{k+1}, \dots, S_n | X_k = j) \\ &= F_k(j)B_k(j) \end{aligned}$$

Thus,

$$P(\vec{S}^n = \vec{s}_n) = \sum_j P(\vec{S}^n = \vec{s}_n, X_k = j) = \sum_j F_k(j)B_k(j), \text{ for any } k.$$

One can first parallel compute $F_k(j)$ and $B_k(j)$, recursively.

Chapter 4

Markov Chain Monte Carlo (MCMC)

Key concepts include:

- important sampling,
- Metropolis-Hastings (MH) sampling,
- Hamilton MH sampling
- 1D data sampling
- Monte Carlo sum, see [13] 数值分析书
- The algorithm and realization of MH sampling (discrete and continuous, Metropolis algorithm, q is symmetric)
- Gibbs sampling, see chapter 6 of [5],
- MCMC, advantage and disadvantage, see chapter 6 of [5],
- other notation, including burn-in. see Handbook [3],
- Hamiltonian Monte Carlo (HMC), see chapter 5 of Handbook [3],
- Vanilla MCMC, see zhihu PR II
- Choice of transition q , paper in John Harlim's book see [4].

4.1 Monte Carlo Methods

4.1.1 deterministic vs. stochastic

(1) deterministic is preferred for low dimension problems, numerical differentiation or integral for ODE and PDE problems.

(2) stochastic is preferred for high dimensional problems.

4.1.2 background

Monte Carlo方法是对任何利用随机数序列来做随机模拟的这类数值方法的统称，该类方法的使用已达数世纪之久，但是只在半个多世纪前才开始成长为应用于复杂性高的问题的数值方法之一。

Monte Carlo本是摩纳哥（Monaco）的著名赌城，第二次世界大战时期N. Metropolis在曼哈顿计划中取其博彩游戏和随机模拟算法二者的相似之处，首次借用其名来作为随机模拟算法的名称。

4.1.3 difference bw some terminology

Markov chain is a concept for the property of a time series or a continuous-time stochastic process.

Monte Carlo is terminology for all stochastic simulation, in particular, referred to the summation of an integral using stochastic approaches.

所有随机数模拟的统称

MCMC is a sampling approach.

采样方法

4.1.4 statistical mechanics

Stochastic Monte Carlo approach is different from the deterministic numerical discretization approach since the latter one is often used for ODEs and PDEs to describe the physics or mathematical systems. Monte Carlo studies the system from microscopic point of view based on the concept of pdf for many particles. In around 1950, people use pdf of many particles to calculate various macroscopic quantities, including temperature and heat capacity, which are usually high dimensional problems. Since too many particles or say the dimension of phase space is very high, the computation of expectations or numerical integrals for high dimensional space becomes extremely important.

4.1.5 applications

statistical mechanics, then quantum chemistry, material science, biological mathematics, financial mathematics, deep learning.

Example 4.1.1 (Deterministic) Consider the integral,

$$I(f) = \int_0^1 f(x)dx.$$

For equi-distance partition (see Fig. 4.1), that is, $0, 1/N, \dots, (N-1)/N, 1$, one can apply the Simpson method

$$I(f) \approx \left(\frac{1}{2}f(x_0) + \sum_{i=1}^{N-1} f(x_i) + \frac{1}{2}f(x_N) \right) h,$$

which is of order $h^2 = O(N^{-2})$. One can also apply the Darboux sum

$$I(f) \approx \sum_{i=0}^{N-1} f(x_i)h, \quad \text{or } I(f) \approx \sum_{i=1}^N f(x_i)h,$$

which is of order $h = O(N^{-1})$.

(相合估计)

Example 4.1.2 (Stochastic) $I(f) = Ef(x)$ when $X \sim U(0, 1)$. Law of large number. We have the Monte Carlo method for numerical integration:

$$I(f) \approx \frac{1}{N} \sum_{i=1}^N f(x_i) := I_N(f),$$

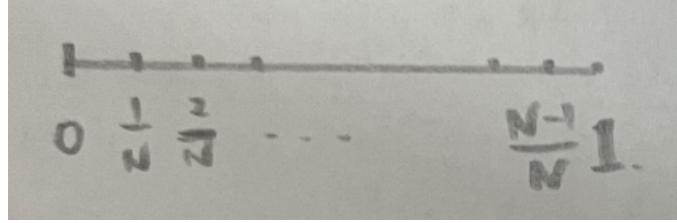


Figure 4.1: Equi-distance partition of the interval $[0, 1]$.

where X_i ($i = 1, 2, \dots, N$) are i.i.d. uniformly distributed $\sim U(0, 1)$. Obviously, $I_N(f)$ is unbiased and weakly consistent (converges in prob.)

Pf. First show that $I_N(f)$ is unbiased,

$$EI_N(f) = E \left[\frac{1}{N} \sum_{i=1}^N f(x_i) \right] = \frac{1}{N} \sum_{i=1}^N \int_0^1 f(x) dx = I(f).$$

Denote

$$e_N = I_N(f) - I(f)$$

as a random variable with $Ee_N = 0$. Then

$$\begin{aligned} E|e_N|^2 &= E(I_N(f) - I(f))^2 = E\left[\frac{1}{N} \sum_{i=1}^N \underbrace{(f(x_i) - I(f))}_{:=a_i}\right]^2 \\ &:= E\left[\frac{1}{N} \sum_{i=1}^N a_i\right]^2 = \frac{1}{N^2} E\left(\sum_{i=1}^N a_i\right)^2 \\ &= \frac{1}{N^2} \left(\sum_{i=1}^N E a_i^2 + \sum_{i < j} 2 \underbrace{E a_i a_j}_{=0} \right) = \frac{1}{N} E a_i^2 \\ &= \frac{1}{N} E(f(x_i) - I(f))^2 = \frac{1}{N} \text{Var}(f) \rightarrow 0. \end{aligned}$$

Therefore, by Chebyshev inequality, one can obtain the weak consistency:

$$P(|I_N f - If| > \varepsilon) \leq \frac{\text{Var}(e_N)}{\varepsilon^2} = \frac{E|e_N|^2}{\varepsilon^2} \rightarrow 0.$$

By Schwartz inequality,

$$E|e_N| \leq \sqrt{E|e_N|^2} = \sqrt{\frac{\text{Var}(f)}{N}} \sim N^{-1/2}.$$

If $\text{Var}(f) < \infty$, then the convergence rate for Monte Carlo is $1/2$.

Example 4.1.3 (Project 1) Consider the integral

$$\int_0^{\pi/2} \sin(x) dx = 1.$$

The Monte Carlo approximation is

$$\int_0^{\pi/2} \sin(x) dx \approx \frac{1}{N} \sum_{i=1}^N \frac{\pi}{2} \sin\left(\frac{\pi}{2} x_i\right),$$

where x_i is i.i.d. $\sim U(0, 1)$. Then fix $m = 100$ independent trials. Take many different $N = 10, 20, 40, 80, \dots, 640, 1280, 2560, 5120, 10240$. Let e_N^j be the error for the j th trial under a given N . Define

$$e_N = \frac{1}{m} \sum_{j=1}^m e_N^j.$$

Then plot $\log e_N$ vs. $\log N$ (similar to the follows in Fig. 4.2).

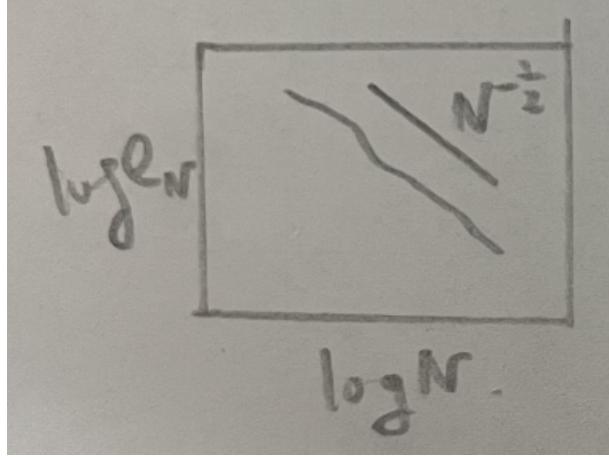


Figure 4.2: Error e_N as a function of the number of particles N .

Example 4.1.4 For general integral,

$$I(f) = \int_0^1 f(x)p(x)dx,$$

where $p(x)$ is a pdf s.t. $\int_0^1 p(x)dx = 1, p(x) \geq 0$. Then

$$I_N(f) = \frac{1}{N} \sum_{i=1}^N f(x_i), \text{ with } x_i \text{ i.i.d. with pdf } p(x).$$

Example 4.1.5 For $I(f) = \int_0^1 f(x)dx$, then

$$I_N(f) = \frac{1}{N} \sum_{i=1}^N \frac{f(x_i)}{p(x_i)}, \text{ where } x_i \text{ i.i.d. with pdf } p(x).$$

Usually, $p(x)$ can be approximated by kernel density estimation or histogram.

Example 4.1.6 Multivariate case. Consider the hyper-cubic $\Omega = [0, 1]^d$ in \mathbb{R}^d and the integral

$$I(f) = \int \cdots \int_{\Omega} f(\vec{x})p(\vec{x})d\vec{x}, \quad \vec{x} = (x_1, \dots, x_d),$$

where $p(\vec{x})$ is a pdf s.t. $\int_0^1 p(\vec{x})d\vec{x} = 1, p(\vec{x}) \geq 0$. For determinisitic approach, assume that $[0, 1]$ is equi-distance partitioned into n intervals for each dimension (see Fig. 4.3). Then the accuracy is still $O(n^{-2}) = O(N^{-\frac{2}{d}})$. However, the complexity requires $O(N = n^d)$ amounts of computation.

On the other hand, for Monte Carlo, $\{x_i\}_{i=1}^M$ i.i.d. pdf is $p(x)$. Let

$$I_M(f) := \frac{1}{M} \sum_{i=1}^M f(x_i) \rightarrow I(f),$$

with the convergence rate of $M^{-1/2}$ and the complexity of $O(M)$. If we let the same accuracy $M^{-1/2} = O(n^{-2}) = O(N^{-2/d})$, then we obtain that $M = O(N^{4/d})$. Thus when $d > 4$, we have $M < N$, which means that the complexity of Monte Carlo is smaller than the Simpson method.

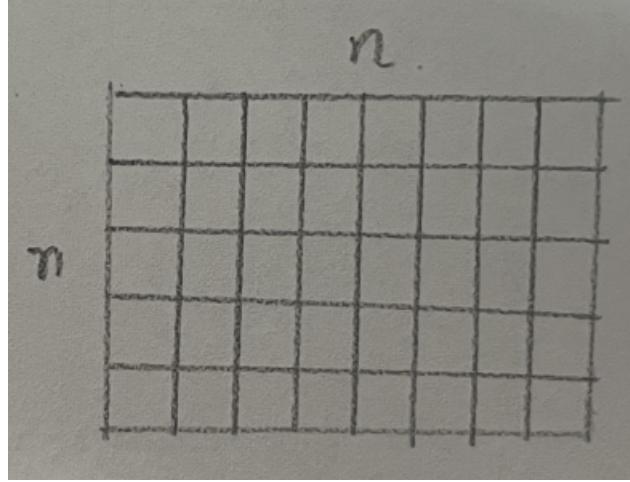


Figure 4.3: Equi-distance partition of a rectangular domain.

Example 4.1.7 In Statistical Mechanics,

$$\langle A \rangle = \frac{1}{Z} \int_{\mathbb{R}^{6N}} A(\vec{c}) e^{-\beta H(\vec{c})} d\vec{c},$$

where $\beta = (k_B T)^{-1}$ is the inverse temperature, k_B is the Boltzman constant, T temperature, $d\vec{c} = d\vec{x}_1 \cdots d\vec{x}_N d\vec{p}_1 \cdots d\vec{p}_N$, N is number of particles, and

$$Z = \int_{\mathbb{R}^{6N}} e^{-\beta H(\vec{c})} d\vec{c},$$

is the partition function. 如果模拟中取值1000个粒子，则Monte Carlo方法做100个运算量(暂且忽略产生随机数所需要的运算量)达到的数值精度基本上需要梯形公式做 10^{3000} 次运算量才可匹敌，后者的运算量对目前威力最大的计算机而言也是鞭长莫及的，可见，虽然Monte Carlo方法的半阶收敛性是非常糟糕的结果，但是在处理维数很高的问题的时候几乎是唯一的选择，因为这种方法受空间维数的制约远没有其他经典的数值离散方法大。

4.2 Generation of Random Numbers

random numbers usually are not i.i.d. in computer, but called "**pseudo random**".

4.2.1 generation of random numbers from $U(0, 1)$

1. Midsquare method

(平方取中法)

In the early development of computers, Von Neumann and others generated pseudo random numbers using midsquare method. For example, first take a 4-digit number 3333 and square it to obtain 11108889.

Take the mid 4 digits 1088 and square it to obtain 1183744, and so forth. Every number 3333, 1088, 1837, divided by 10^4 to obtain a pseudo random number between $[0, 1]$ with uniform distribution. However, the maximum cycle length of the approach is less than 10^4 , and its statistics is not very good while this algorithm was early used in the computation of nuclear reaction.

核反应计算 循环长度

2. Linear congruent algorithm

(线性同余法)

In the random number generator of $U(0, 1)$, one early popular method is called linear congruent algorithm. The algorithm is as follows:

$$X_{n+1} = aX_n + b \pmod{m},$$

where a, b, m are given integers. One important criterion for judging a random number generator is the called maximum cycle length. In the same period, the longer the maximum cycle length, the better the performance of the generator. For example,

$$m = 2^k, \quad a = 4c + 1, \quad b \text{ is odd.}$$

3. Magic "16807"

(神奇的"16807")

1969年, Lewis, Goodman和Miller提出了如下发生器:

$$X_{n+1} = aX_n \pmod{m},$$

并且取 $a = 7^5 = 16807$, $m = 2^{31} - 1 = 2147483647$. Shrage 给出了一个在计算机上高效实现上述乘法同余的算法, 这样得到的伪随机数发生器最大循环长度可达到 2.1×10^9 . 这个发生器通过了当时的所有理论测试, 被称为最小标准发生器(Minimal standard generator)(意指其他的发生器如果要被接受, 至少要能达到这一发生器的质量).

后来, 基于这一方法, L'Ecuyer 采用所谓Bays-Durham 洗牌算法(见文献[6]) 给出了一个更为强大的随机数发生器, 其最大循环长度达到约 2.3×10^{18} . 在数值计算的著名书籍[8]中, 给出了这一算法的具体实现程序ran2().该书作者声称, 如果有人能给出使用上述算法而导致系统性失败的案例, 将付款1000美元!

随机数发生器是进行随机模拟的基石, 如果没有一个可靠的随机数发生器, 一切的计算结果都不再可信, 笔者强烈建议读者使用经过大量测试的、成熟的随机数发生器程序包, 而不提倡自行编写此类程序. 另外, 对程序的选取也要慎重, 我们推荐 文献[8]中的程序以及www.netlib.org上的随机数发生器.

4.2.2 generation of random numbers with general distributions

1. Transformation method

(变换法)

Proposition 4.2.1 Let a random variable Y with the distribution function $F(y)$, that is

$$P(Y \leq y) = F(y).$$

If another random variable $X \sim U(0, 1)$, then $Y = F^{-1}(X)$ satisfies the desired distribution.

Proof. Since $X \sim U(0, 1)$ and $Y = F^{-1}(X)$, so that

$$\begin{aligned} P\{Y \leq y\} &= P\{F^{-1}(X) \leq y\} \\ &= P\{X \leq F(y)\} = F(y). \end{aligned}$$

Based on above prop., if we already have random variables $X_i (i = 1, 2, \dots)$ uniformly distributed on $(0, 1)$, then $Y_i = F^{-1}(X_i)$ is randomly distributed with distribution function $F(y)$. The larger the pdf is, the steeper the cdf is. The smaller the pdf is, the flatter the cdf is. See Fig. 4.4.

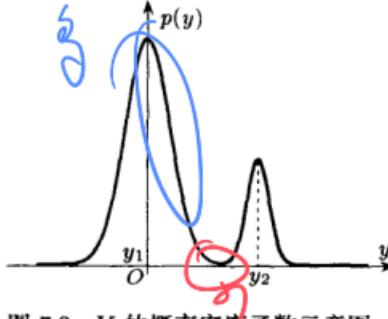


图 7.2 Y 的概率密度函数示意图

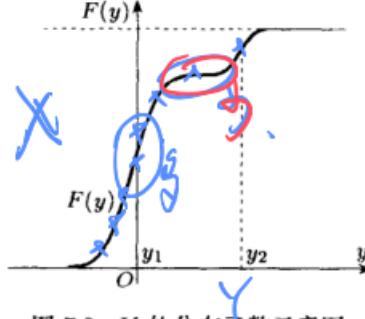


图 7.3 Y 的分布函数示意图

Figure 4.4: The probability density function (pdf) and cumulative distribution function(cdf) of a random variable Y .

Example 4.2.2 (Project 2) (Exponential Distribution) Let the exponential distribution with pdf

$$p(y) = \begin{cases} 0, & y \leq 0, \\ \lambda e^{-\lambda y}, & y > 0, \end{cases}$$

Its cdf is $F(y) = 1 - e^{-\lambda y}$. Then

$$F^{-1}(x) = -\frac{1}{\lambda} \ln(1-x), \quad x \in (0, 1).$$

Based on the transformation method, the exponentially distributed random variable can be obtained by

$$Y_i = -\frac{1}{\lambda} \ln(1-X_i), \quad i = 1, 2, \dots$$

where $X_i \sim U(0, 1)$. Check the pdf usign histogram or Kernel Density Estimation (KDE). Or check the statistics based on p-value using χ^2 -test.

Example 4.2.3 (Normal Distribution) The normally distributed variable has the pdf

$$p(x) = \frac{1}{\sqrt{2\pi}} e^{-\frac{x^2}{2}},$$

and its cdf is

$$F(x) = \int_{-\infty}^x p(y) dy = \frac{1}{2} + \frac{1}{2} \operatorname{erf}\left(\frac{x}{\sqrt{2}}\right),$$

where $\operatorname{erf}(x) = \frac{2}{\sqrt{\pi}} \int_0^x e^{-t^2} dt$ is called the error function. Thus, $F^{-1}(x) = \sqrt{2} \operatorname{erf}^{-1}(2x-1)$. However, it is directly realized during numerical implementation since erf^{-1} is difficult to compute. This means that the transformation method has limitation.

Box-Muller method

To generate normally distributed random variables, we use the following well-known Box-Muller method.

Notice that

$$\left(\int_{-\infty}^{+\infty} e^{-x^2} dx \right)^2 = \int_{-\infty}^{+\infty} \int_{-\infty}^{+\infty} e^{-(x^2+y^2)} dx dy = \int_0^{+\infty} \int_0^{2\pi} e^{-r^2} r dr d\theta = \pi,$$

on which the Box-Muller method is based. Let $(x_1, x_2) = (r \cos \theta, r \sin \theta)$, then

$$\frac{1}{2\pi} e^{-\frac{x_1^2+x_2^2}{2}} dx_1 dx_2 = \frac{1}{2\pi} e^{-\frac{r^2}{2}} r dr d\theta = \left(\frac{1}{2\pi} d\theta \right) \left(e^{-\frac{r^2}{2}} r dr \right).$$

In θ direction, $\frac{1}{2\pi}$ is the density for the uniform distribution $U(0, 2\pi)$. In r direction, $e^{-\frac{r^2}{2}} r$ is the density corresponding to the distribution function $F(r) = \int_0^r e^{-\frac{s^2}{2}} s ds = 1 - e^{-\frac{r^2}{2}}$. Thus, random variables (Y_1, Y_2) for the two-dimensional normal distribution can be generated through

$$\begin{cases} Y_1 = \sqrt{-2 \ln X_1} \cos(2\pi X_2), \\ Y_2 = \sqrt{-2 \ln X_1} \sin(2\pi X_2), \end{cases}$$

where X_1 and X_2 are independent random variables satisfying $U(0, 1)$. Notice that

$$F^{-1}(r) = \sqrt{-2 \ln(1 - r)},$$

where $1 - r \sim U(0, 1)$ and $r \sim U(0, 1)$. Then the above expression is written in this way since $Y_1 = r \cos \theta$ and $Y_2 = r \sin \theta$, where $\theta = 2\pi X_2$ and $r = \sqrt{-2 \ln X_1}$.

Acceptance-rejection method

Not introduced at this moment. The idea is kind of similar to MCMC.

4.2.3 technique for reducing the variance

The error of Monte Carlo is σ/\sqrt{N} , where $\sigma = (\text{Var}(f))^{1/2}$.

The rate $1/\sqrt{N}$ usually cannot be improved!

But the constant σ can!

Importance sampling (Project 2)

If i.i.d. $x_i \sim U(0, 1)$, then

$$If \approx I_X(f) := \frac{1}{N} \sum_{i=1}^N f(x_i).$$

On the other hand,

$$If = \int_0^1 f(y) dy = \int_0^1 \frac{f(y)}{p(y)} p(y) dy,$$

where $p(y)$ is a pdf. Then

$$I(f) \approx I_Y(f) := \frac{1}{N} \sum_{i=1}^N \frac{f(y_i)}{p(y_i)},$$

where i.i.d. $y_i \sim p(y)$.

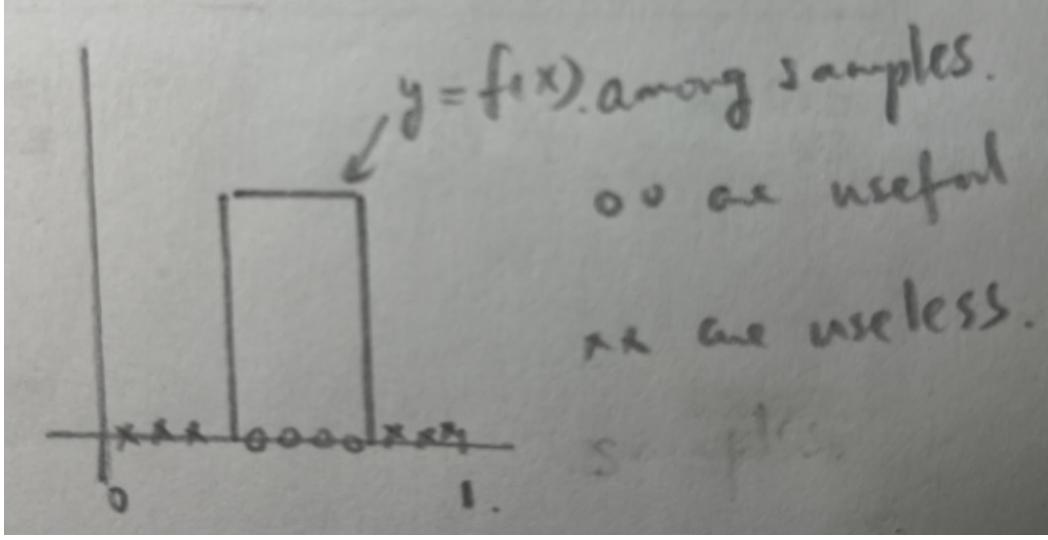


Figure 4.5: Sketch for importance sampling.

Let's see Fig. 4.5 for the sketch for importance sampling. The nonzero values of the integrand function $f(x)$ are focused in the interval (a, b) . If using $x_i \sim U(0, 1)$, then many samples of x_i lie out of the interval (a, b) and these samples are used for computing the integral $\int_0^a f(x)dx + \int_b^1 f(x)dx = 0$. This means that we spend amounts of computation to calculate the integral part that does not contribute to the final result. Thus, the efficiency is low and also the accuracy is low. On the other hand, if we take $p(y)$ proportional to $f(y)$, most samples are used to compute the dominant part of the integral. This is called the importance sampling and $p(y)$ reflects the important part of $f(x)$.

Theoretically, we can analyze the error from variance point of view:

$$\begin{aligned} Var_X(f) &= \int_0^1 [f - I(f)]^2 dx = \int_0^1 f^2 dx - I^2(f), \\ Var_Y\left(\frac{f}{p}\right) &= \int_0^1 \left(\frac{f}{p}\right)^2 pdy - I^2(f) = \int_0^1 \frac{f^2}{p} dy - I^2(f). \end{aligned} \quad (4.1)$$

Here, we notice that since

$$e_X = I_X(f) - I(f), \quad e_Y = I_Y(f) - I(f),$$

their means are

$$Ee_X = 0, \quad Ee_Y = 0,$$

and variances are

$$\begin{aligned} E|e_X|^2 &= \frac{Var_X(f)}{N}, \\ E|e_Y|^2 &= E(I_Y(f) - I(f))^2 = E\left[\frac{1}{N} \sum_{i=1}^N \underbrace{\left(\frac{f(y_i)}{p(y_i)} - I(f)\right)^2}_{:=b_i}\right] \\ &= \frac{1}{N^2} E\left(\sum_{i=1}^N b_i\right)^2 = \frac{1}{N^2} \left(\sum_{i=1}^N Eb_i^2 + \sum_{i < j} 2 \underbrace{Eb_i b_j}_{=0}\right) = \frac{1}{N} Eb_i^2 \\ &= \frac{1}{N} E\left(\frac{f(y_i)}{p(y_i)} - I(f)\right)^2 = \frac{1}{N} Var_Y\left(\frac{f}{p}\right). \end{aligned}$$

In equation (4.1), if we take an appropriate $p(y)$, s.t. $\int_0^1 \frac{f^2}{p} dy < \int_0^1 f^2 dx$, then the variance reduces $Var_Y(\frac{f}{p}) < Var_X(f)$. In particular, if $p(y) = \frac{f(y)}{I(f)} \propto f(y)$, then

$$Var_Y\left(\frac{f}{p}\right) = \int_0^1 \frac{f^2}{p} dy - I^2(f) = I(f) \int_0^1 f dy - I^2(f) = I^2(f) - I^2(f) = 0.$$

This means that when the importance of $p(x)$ is the same with that of $f(x)$, the variance becomes 0 and we obtain the exact integral value. However, in practice this is difficult to realize due to two reasons. First, for $p(y) = \frac{f(y)}{I(f)}$, it is impossible for us to prior get the value of $I(f)$ in the computation of $I_Y(f) := \frac{1}{N} \sum_{i=1}^N \frac{f(y_i)}{p(y_i)}$. Second, it is not trivial to get samples with density $p(y) \propto f(y)$ (while can be done using MCMC).

重要性抽样对实际计算有根本思想性的指导作用，它依赖于我们对一个问题先验的了解程度并以此构造适当的抽样所依据的概率密度。

Advantages of importance sampling over simple Monte Carlo (uniform sampling)

- Can significantly improve performance, by reducing the variance
- Can use samples from a different distribution, say q , to compute expectations with respect to p
- Can compute the normalization constant of p , as well as Bayes factors

Disadvantages

- Need to be able to evaluate the pdf/pmf $p(x)$ and $q(x)$, at least up to proportionality constants
- It might not be obvious how to choose a good q

Control Variate method, modification of importance sampling

控制变量法

The basic idea here is to use a random variable with given statistics (such as the mean) to control another random variable with unknown statistics. For example,

$$\int_0^1 f(x) dx = \int_0^1 [f(x) - g(x)] dx + \int_0^1 g(x) dx,$$

with $\int_0^1 g(x) dx$ is already known. Then one can use Monte Carlo to obtain

$$I_N(f) := \frac{1}{N} \sum_{i=1}^N [f(x_i) - g(x_i)] + I(g), \text{ with } x_i \sim \text{i.i.d. } U(0, 1).$$

If $Var[f - g] \leq Var[f]$, then the constant coefficient is reduced while the convergence rate is kept the same. In extreme case, when $f = g$, then $Var[f - g] = 0$.

Example 4.2.4 Consider the integral,

$$I(f) = \int_{-\infty}^{\infty} \frac{1}{\sqrt{2\pi}} (1+r)^{-1} e^{-\frac{x^2}{2}} dx,$$

where $r = e^{\sigma x}$ ($\sigma > 0$). Notice that

$$(1+r)^{-1} \approx \begin{cases} 1, & x \leq 0, \\ 0, & x > 0, \end{cases} := h(x).$$

Then $I(f)$ can be approximated by

$$I(f) = \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{\infty} [(1+r)^{-1} - h(x)] e^{-\frac{x^2}{2}} dx + \frac{1}{2}.$$

4.3 Introduction to Markov Chain Monte Carlo (MCMC)

4.3.1 Historical development of MCMC and its connection to statistical mechanics

The popular application of MCMC starts from the development of statistical mechanics. In statistical mechanics, an ensemble average is thought of as a concept that the macro state we see is the prob. average of many micro states in equilibrium (see Fig. 4.6). The prob. here usually corresponds to Gibbs measure (a more general used terminology for both finite and infinite systems) or Boltzmann distribution (a term for finite systems). For example, here is a box containing many gas particles in equilibrium and we can measure their temperature, pressure, etc. In SM, this macro system, in fact, corresponds to many micro systems, and its macro quantity is the statistical average of the quantities of these many micro systems. In mathematics, the discrete and continuous cases are

$$\langle A(\sigma) \rangle = \sum_{\sigma} \frac{\exp\{-\beta H(\sigma)\}}{Z} A(\sigma),$$

$$\langle A(\sigma) \rangle = \int_{\sigma} \frac{\exp\{-\beta H(\sigma)\}}{Z} A(\sigma) d\sigma,$$

where $Z = \sum_{\sigma} \exp\{-\beta H(\sigma)\}$ and $Z = \int_{\sigma} \exp\{-\beta H(\sigma)\} d\sigma$ are partition functions.

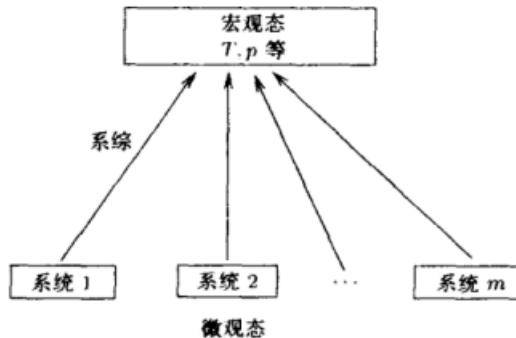


图 7.7 平衡态统计系综示意图

Figure 4.6: Ensembles in equilibrium.

The following is the statement of equilibrium statistical mechanics. Each of the m subsystems are independent. each one at a time corresponds a micro state but at any time the prob. distributions of these microstates are invariant. Metropolis has another point of view of this problem. Since macro quantities, such as temperature, are invariant in equilibrium with respect to time, this means that the **time average** should be the same as the **ensemble average**, that is

$$\langle A(\sigma) \rangle \approx \frac{1}{N} \sum_{i=1}^N A(\sigma^{(i)}),$$

where $\{\sigma^{(i)}\}_{i=1}^N$ is the time series of the states evolving under the physics law. We notice that usually a physics law gives us a Makovian process. The equality between the the ensemble average and the time average is referred to as ergodicity in SM. In mathematics, the Metropolis algorithm is based on the **ergodic theory**.

Monte Carlo simulation

Nowadays, MCMC can be used to generate samples from a given pmf or pdf, and then possibly be used for **Monte Carlo simulation**.

Let X be a discrete random vector whose set of possible values is x_j . Let the probability mass function of X be given by $P\{X = x_j\}$, and suppose that we are interested in calculating

$$\theta = E[h(X)] = \sum_j h(x_j)P\{X = x_j\},$$

for some specified function h . In situations where it is computationally difficult to evaluate the series, we often turn to simulation to approximate θ . The usual approach, called **Monte Carlo simulation**, is to use random numbers to generate a partial sequence of random vectors X_1, X_2, \dots, X_n having the mass function $P\{X = x_j\}$. Since the strong law of large numbers yields

$$\lim_{n \rightarrow \infty} \frac{1}{n} \sum_{i=1}^n h(X_i) = \theta,$$

it follows that we can estimate θ by letting n be large and using the average of the values of $h(X_i)$, $i = 1, \dots, n$ as the estimator.

4.3.2 Metropolis-Hastings Algorithm

Discrete Case

The conditions are:

(1) X is a vector of dependent random variable whose samples cannot be generated via independent sampling technique.

(2) We only know $P(X = x_j) \propto b(j)$ but do not know the normalization factor. Assume that $B = \sum_j b(j)$ is finite so that the pmf is $P(X = x_j) = b(j)/B = \pi(j)$.

The MH algorithm is given as follows:

Algorithm 4.3.1 (MH algorithm for discrete case)

Step 0. Choose a good enough initial guess X_0 .

Step 1. Specify an irreducible Markov transition prob. matrix Q with entries $q(i, j)$.

Step 2. At the n th step, given $X_n = i$, draw a proposal $Y = j$ with prob. $P(Y = j) = q(i, j)$.

Step 3. Accept the proposal $X_{n+1} = Y = j$ with prob.

$$\alpha(i, j) = \min \left(\frac{b(j)q(j, i)}{b(i)q(i, j)}, 1 \right).$$

Otherwise, reject the proposal and set $X_{n+1} = X_n = i$ with prob. $1 - \alpha(i, j)$.

Remark 4.3.2 For the MH algorithm, we do not need to know the exact the distribution π , but only need to know $b(j)$ which is different from $\pi(j)$ by a normalization constant. The MH algorithm still works when only $b(j)$ is given.

Remark 4.3.3 Numerically, draw a sample $z \sim U(0, 1)$, and set

$$X_{n+1} = \begin{cases} Y = j, & \text{if } z < \alpha(i, j), \\ X_n = i, & \text{otherwise } z \geq \alpha(i, j). \end{cases}$$

Remark 4.3.4 One can see that the sequence of states constitutes a Markov chain with transition prob. p_{ij} given by

$$\begin{aligned} p_{ij} &= q(i, j)\alpha(i, j), \quad \text{if } j \neq i, \\ p_{ii} &= q(i, i) + \sum_{k \neq i} q(i, k)(1 - \alpha(i, k)). \end{aligned}$$

Notice that p_{ij} is a pdf since

$$\begin{aligned} \sum_j p_{ij} &= \sum_{j \neq i} p_{ij} + p_{ii} = \sum_{j \neq i} q(i, j)\alpha(i, j) + q(i, i) + \sum_{j \neq i} q(i, j)[1 - \alpha(i, j)] \\ &= q(i, i) + \sum_{j \neq i} q(i, j) = 1. \end{aligned}$$

This Markov chain will be time reversible and have stationary prob. $\pi(j)$ if

$$\begin{aligned} \pi(i)p_{ij} &= \pi(j)p_{ji}, \quad \text{for } j \neq i, \\ \pi(i)q(i, j)\alpha(i, j) &= \pi(j)q(j, i)\alpha(j, i). \end{aligned} \tag{4.2}$$

We now check that the detailed balance equation (4.2) is satisfied in the MH algorithm. If

$$\alpha(i, j) = \frac{b(j)q(j, i)}{b(i)q(i, j)} < 1,$$

then $\alpha(j, i) = 1$ since $\frac{b(i)q(i, j)}{b(j)q(j, i)} > 1$ and equation (4.2) follows. If $\alpha(i, j) = 1$ then similarly again Eq. (4.2) follows.

Continuous Case

The MH method generates samples $\{x_i\}$ of a target density $\pi(x)$ (possibly different by a normalization constant) as follows:

Algorithm 4.3.5 (MH algorithm for continuous case)

- Step 0. Choose a good enough guess x_0 . (**the choice affects the convergence greatly.**)
- Step 1. Specify a Markov transition kernel density $q : E \times E \rightarrow \mathbb{R}^+$ such that $\int_E q(x, y)dy = 1$, $\forall x \in E$.
- Step 2. At the i th step, given x_{i-1} , draw a proposal $u \sim q(x_{i-1}, \cdot)$.
- Step 3. Accept the proposal, $x_i = u$ with probability $\alpha(x_{i-1}, u)$. Otherwise set $x_i = x_{i-1}$. Here,

$$\alpha(x_{i-1}, u) = \min \left(\frac{\pi(u)q(u, x_{i-1})}{\pi(x_{i-1})q(x_{i-1}, u)}, 1 \right).$$

Practically, we don't need to know the normalization constant for π since only the ratio of π is needed to determine the acceptance rate above. Numerically, this step can be realized by drawing a sample of the standard uniform distribution, $z \sim U[0, 1]$, and then setting,

$$x_i = \begin{cases} u, & \text{if } z < \alpha(x_{i-1}, u), \\ x_{i-1}, & \text{otherwise } z \geq \alpha(x_{i-1}, u). \end{cases}$$

First, let's investigate why should we believe that x_i is a realization of the chain with stationary density $\pi(x)$. First of all, let x_i be a Markov chain with transition density $p(x_{i-1}, x_i)$ and we want to show that π is

the stationary density. That is, we want to show that the chain generated by the MH method above satisfies the detailed balance condition. Notice that based on the MH method, we can write the transition kernel $p(x_{i-1}, u)$, which quantifies the probability of accepting u at the i th step given a realization of the chain x_{i-1} at the $(i-1)$ th step, as a product of the probability of proposing u and the probability of accepting u ,

$$p(x_{i-1}, u) = q(x_{i-1}, u)\alpha(x_{i-1}, u).$$

Thus,

$$\begin{aligned} \pi(x_{i-1})p(x_{i-1}, u) &= \pi(x_{i-1})q(x_{i-1}, u)\min\left(\frac{\pi(u)q(u, x_{i-1})}{\pi(x_{i-1})q(x_{i-1}, u)}, 1\right) \\ &= \min(\pi(u)q(u, x_{i-1}), \pi(x_{i-1})q(x_{i-1}, u)) \\ &= \pi(u)q(u, x_{i-1})\min(1, \frac{\pi(x_{i-1})q(x_{i-1}, u)}{\pi(u)q(u, x_{i-1})}) \\ &= \pi(u)q(u, x_{i-1})\alpha(u, x_{i-1}) \\ &= \pi(u)p(u, x_{i-1}), \end{aligned}$$

which means that x_i generated by MH method satisfies the detailed balanced condition (so that stationary) so $\pi(x)$ is the stationary density. To verify theoretically whether we obtain samples of π if we implement the MH method sufficiently large i is simply equivalent to asking whether the chain is recurrent and aperiodic (which is the hypothesis for the convergence). Detailed study of the convergence rates of some Metropolis-Hastings kernel densities were reported in [11].

Remark 4.3.6 We should remark that if the proposal transition density q is chosen to be symmetric, $q(x, y) = q(y, x)$, then the resulting method is known as the **Metropolis scheme**, with an acceptance rate,

$$\alpha(x_{i-1}, u) = \min\left(\frac{\pi(u)}{\pi(x_{i-1})}, 1\right). \quad \text{This is Metropolis scheme.}$$

Intuitively, this rate compares the probability of the proposal u to that of the previous chain value, x_{i-1} . A popular choice of symmetric proposal density is Gaussian which yields the random walk Metropolis proposal,

$$q(x_{i-1}, u) = q(u|x_{i-1}) \sim N(x_{i-1}, C),$$

for some proposal covariance matrix C . Numerically, we realize the sample as follows,

$$u = x_{i-1} + C^{1/2}\xi_i, \quad \xi_i \sim N(0, I).$$

We should choose the hyperparameter C carefully since too large C will result in low acceptance rate whereas too small C will result in long burn-in or too small range of samples.

Remark 4.3.7 There are many choice of q in history. See references in:

- (1) Tierney, 1994, Markov chain for exploring posterior distributions.
- (2) Chib and Greenberg, 1995, Understanding the Metropolis-Hastings Algorithm.
- (3) Hastings, 1970, $q(x, y) = q_2(y)$.

Remark 4.3.8 Vanilla MCMC was abandoned long time ago since its acceptance rate is too low. For Vanilla MCMC,

$$\alpha(i, j) = \pi(j)q(j, i), \quad \alpha(j, i) = \pi(i)q(i, j).$$

The **low acceptance rate** is caused by the too small values of above $\alpha(i, j)$ or $\alpha(j, i)$. Nowadays, it is still an important and popular topic to increase the **acceptance rate**, for instance, by using **Hamiltonian MCMC** and **Langevin dynamics MCMC**.

Remark 4.3.9 (*How useful MCMC is and why it works*) In many real-world applications, we have to deal with complex probability distributions on complicated **high-dimensional spaces**. On rare occasions, it is possible to sample exactly from the distribution of interest, but typically exact sampling is difficult. Further, high dimensional spaces are very large, and distributions on these spaces are hard to visualize, making it difficult to even guess where the regions of high probability are located. As a result, it may be challenging to even design a reasonable proposal distribution to use with importance sampling.

Markov chain Monte Carlo (MCMC) is a sampling technique that works remarkably well in many situations like this. Roughly speaking, my intuition for why MCMC often works well in practice is that

- (a) **the region of high probability tends to be "connected"**, that is, you can get from one point to another without going through a low-probability region, and
- (b) we tend to be interested in the expectations of functions that are relatively smooth and have lots of "symmetries", that is, one only needs to evaluate them at a small number of representative points in order to get the general picture.

MCMC constructs a sequence of correlated samples X_1, X_2, \dots that meander through the region of high probability by making a sequence of incremental movements. Even though the samples are not independent, it turns out that under very general conditions, sample averages $\frac{1}{N} \sum_{i=1}^N h(X_i)$ can be used to approximate expectations $Eh(X)$ just as in the case of simple **Monte Carlo approximation**, and by a powerful result called **the ergodic theorem**, these approximations are guaranteed to converge to the true value.

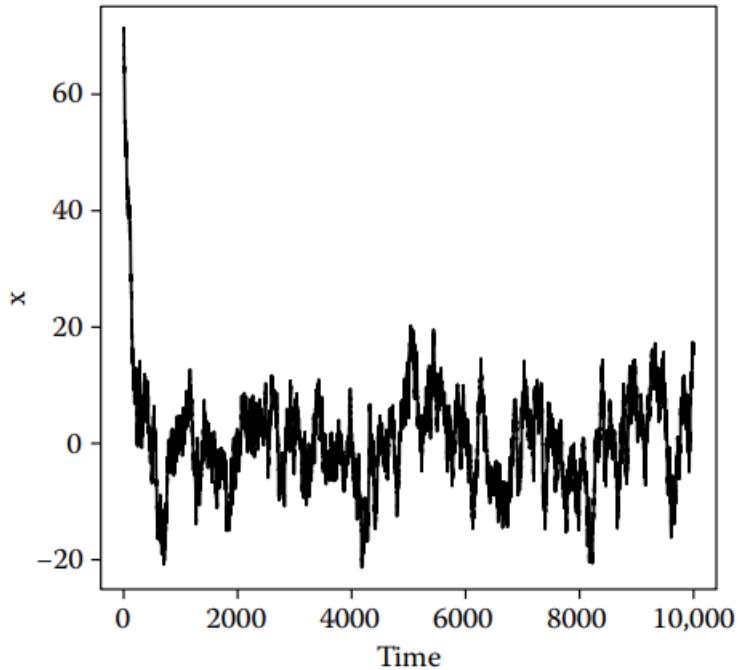


Figure 4.7: Burn-in is important for MCMC.

Remark 4.3.10 *Burn-in* is important for MCMC so that at least **one long run** is required (see Fig. 4.7). For diagnostics, many relatively-long runs are preferred for detecting the pseudo-convergence if computing resources are allowed.

Remark 4.3.11 Pseudo-convergence. A Markov chain can appear to have converged to its equilibrium distribution when it has not. This happens when parts of the state space are **poorly connected** by the Markov chain dynamics: it takes many iterations to get from one part to another. When the time it takes to transition between these parts is much longer than the length of simulated Markov chain, then the Markov chain can appear to have converged but the distribution it appears to have converged to is the equilibrium distribution conditioned on the part in which the chain was started. We call this phenomenon **pseudo-convergence**. This phenomenon has also been called "**multimodality**" since it may occur when the equilibrium distribution is multimodal. But multimodality does not cause pseudo-convergence when the troughs between modes are not severe. Nor does pseudo-convergence only happen when there is multimodality. Some of the most alarming cases of pseudo-convergence occur when the state space of the Markov chain is discrete and "modes" are not well defined (Geyer and Thompson, 1995). Hence pseudo-convergence is a better term.

Remark 4.3.12 (Advantages and disadvantages for MCMC.)

Advantages of MCMC:

- applicable even when we can't directly draw samples. When we know **nearly nothing about the model**, we can apply MCMC for initial tests to give us something while not always give correct results or accurate results.
- works for complicated distributions in **high-dimensional spaces**, even when we don't know where the **regions of high probability are** (while MCMC is not guaranteed to give correct results). Notice that in high-dimensional spaces, the density $\pi(x)$ is hard for visualization, but on the other hand, the samples $\{x_i\}_{i=1}^N$ can be used to compute expectations of interested functions $Ef(x)$ or even for visualization.
- relatively easy to implement
- fairly reliable for sufficiently long runs.

Disadvantages of MCMC:

- MCMC is slower than simple Monte Carlo or importance sampling in which i.i.d. samples are given or generated quickly by some other techniques. In addition, MCMC requires more samples for the same level of accuracy.
- computationally expensive in **high dimensional spaces**, or in **multimodality cases**, or when **model is complicated**.
- can be very difficult to assess accuracy and evaluate convergence, even empirically. There is a great deal of theory about convergence of Markov chains. Unfortunately, **none of it can be applied to get useful convergence information for most MCMC applications**. Thus most users find themselves in the following situation we call black box MCMC:

(1) **You know nothing other than that. The Markov chain is a "black box" that you cannot see inside. When run, it produces output. That is all you know.** You know nothing about the transition probabilities of the Markov chain, nor anything else about its dynamics. This Point 1 may seem extreme. You may know a lot about the particular Markov chain being used—for example, you may know that it is a Gibbs sampler—but if whatever you know is of no help in determining any convergence information about the Markov chain, then whatever knowledge you have is useless.

(2) You know nothing about the invariant distribution except what you may learn from running the Markov chain. This Point 2 may seem extreme. Many examples in the MCMC literature use small problems that can be done by independent and identically distributed (i.i.d.) Monte Carlo or even by pencil and paper and for which a lot of information about the invariant distribution is available, but in complicated applications point 2 is often simply true.

4.3.3 Gibbs sampling

The most widely used version of the Hastings-Metropolis algorithm is the **Gibbs sampler**.

Randomly chosen coordinate

Let $\vec{\mathbf{X}} = (X_1, \dots, X_n)$ be a discrete random vector with probability mass or density function $p(\vec{\mathbf{x}})$ that is only specified up to a multiplicative constant, and suppose that we want to generate a random vector whose distribution is that of $\vec{\mathbf{X}}$. That is, we want to generate a random vector having mass function

$$p(\vec{\mathbf{x}}) = Cg(\vec{\mathbf{x}})$$

where $g(\vec{\mathbf{x}})$ is known, but C is not.

Algorithm 4.3.13 (Gibbs sampling using randomly chosen coordinate)

Step 0. Choose an initial state $\vec{\mathbf{x}} = (x_1, \dots, x_n)$.

Step 1. For the current state $\vec{\mathbf{x}}$, choose a coordinate which is equally likely to be any of the coordinates $1, \dots, n$. If coordinate k is chosen, then generate a random variable Y whose probability distribution is given by

$$P\{Y = y\} = P\{X_k = y | X_j = x_j \text{ for } j = 1, \dots, n \text{ with } j \neq k\},$$

where it is assumed that the above random variable Y can be generated having the above pmf. If $Y = y$, let the candidate state $\vec{\mathbf{y}} = (x_1, \dots, x_{k-1}, y, x_{k+1}, \dots, x_n)$. Note that the above condition probability is computable since the joint pmf $g(\vec{\mathbf{x}})$ is given.

Step 2. The next state $\vec{\mathbf{x}} = (x_1, \dots, x_n)$ is set equal to $\vec{\mathbf{y}}$. Repeat step 1 with this new state $\vec{\mathbf{x}}$.

The Gibbs sampler uses the Metropolis-Hastings algorithm with the choice

$$\begin{aligned} q(\vec{\mathbf{x}}, \vec{\mathbf{y}}) &= \frac{1}{n} P\{X_k = y | X_j = x_j \text{ for } j = 1, \dots, n \text{ with } j \neq k\} \\ &= \frac{1}{n} \frac{p(\vec{\mathbf{y}})}{P\{X_j = x_j \text{ for } j = 1, \dots, n \text{ with } j \neq k\}} \end{aligned}$$

for the Markov transition matrix Q . It is not difficult to verify that for this choice the acceptance probability $\alpha(\vec{\mathbf{x}}, \vec{\mathbf{y}})$ is given by

$$\begin{aligned} \alpha(\vec{\mathbf{x}}, \vec{\mathbf{y}}) &= \min \left(\frac{p(\vec{\mathbf{y}})q(\vec{\mathbf{y}}, \vec{\mathbf{x}})}{p(\vec{\mathbf{x}})q(\vec{\mathbf{x}}, \vec{\mathbf{y}})}, 1 \right) \\ &= \min \left(\frac{p(\vec{\mathbf{y}})p(\vec{\mathbf{x}})}{p(\vec{\mathbf{x}})p(\vec{\mathbf{y}})}, 1 \right) = 1. \end{aligned}$$

Hence, when utilizing the Gibbs sampler, the candidate state is **always accepted** as the next state of the Markov chain.

Successively chosen coordinate

Suppose that $p(x, y)$ is a p.d.f. or p.m.f. that is difficult to sample from directly. Suppose, though, that we can easily sample from the **conditional distributions** $p(x|y)$ and $p(y|x)$. Roughly speaking, the Gibbs sampler proceeds as follows: set x and y to some initial starting values, then sample $x|y$, then sample $y|x$, then $x|y$, and so on. More precisely,

Algorithm 4.3.14 (Gibbs sampling using successively chosen coordinate)

Step 0. Set (x_0, y_0) to some starting value.

Step 1. Sample $x_1 \sim p(x|y_0)$, that is, from the conditional distribution $X|Y = y_0$.

Sample $y_1 \sim p(y|x_1)$, that is, from the conditional distribution $Y|X = x_1$.

Step 2. Sample $x_2 \sim p(x|y_1)$, that is, from the conditional distribution $X|Y = y_1$.

Sample $y_2 \sim p(y|x_2)$, that is, from the conditional distribution $Y|X = x_2$.

\vdots

Each iteration (1., 2., 3., ...) in the Gibbs sampling algorithm is sometimes referred to as a **sweep** or **scan**. The sampling steps within each iteration are sometimes referred to as **updates** or **Gibbs updates**. Note that when updating one variable, we always use the most recent value of the other variable (even in the middle of an iteration).

This procedure defines a sequence of pairs of random variables

$$(X_0, Y_0), (X_1, Y_1), (X_2, Y_2), (X_3, Y_3), \dots$$

which has the property of being a Markov chain—that is, the conditional distribution of (X_i, Y_i) given all of the previous pairs depends only on (X_{i-1}, Y_{i-1}) .

Gibbs sampling with more than two variables is completely straightforward—roughly speaking, we cycle through the variables, sampling each from its conditional distribution given all the rest.

Remark 4.3.15 在以上算法中，坐标轴轮换采样不是必须的，可以在坐标轴轮换中引入随机性，这时候转移矩阵 Q 中任何两个点的转移概率中就会包含坐标轴选择的概率，而在通常的 *Gibbs Sampling* 算法中，坐标轴轮换是一个确定性的过程，也就是在给定时刻 t ，在一根固定的坐标轴上转移的概率是 1。

(Project 3) Show both the Metropolis-Hastings samplers and Gibbs samplers for 2D Gaussian distribution and the distribution with the density

$$p(x, y) \propto e^{-xy} \mathbf{1}(x, y \in (0, 1)).$$

4.4 Parameter estimation problems

4.4.1 frequentist statistics vs. Bayesian statistics

Statistics have two popular methods, **frequentist statistics** and **Bayesian statistics**. The main differences are listed in Table 5.1. Once you understand the Bayesian approach, it seems so natural that it is hard to imagine any alternative. In fact, there was no satisfying alternative until the early 1900's, when Karl Pearson, Jerzy Neyman, Egon Pearson, and Ronald Fisher initiated what is now called frequentist statistics.

Both can be used for parameter estimation. For **Bayesian inference**,

$$p(\theta|y) = \frac{p(y|\theta)p(\theta)}{\int p(y|\theta)p(\theta)d\theta},$$

where y is data and θ is model (or parameters of model), MCMC is one popular approach, which can be used for uncertainty quantification. However, MCMC is not appropriate for too large amount of data or too complicated model. Another Bayesian approach is variational inference which involves Kullback-Leibler (KL) divergence.

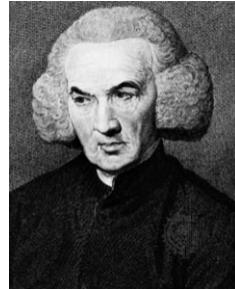
Table 4.1: Comparison between frequentists and Bayesians.

	frequentists	Bayesians
Hypothesis testing	Set null and alternative hypotheses and use statistical tests to assess evidence against the null.	Consider prior beliefs when forming hypotheses.
Probability interpretation	Frame probability in terms of objective, long-term frequencies.	Interpret probabilities subjectively and update them as new data is collected.
Sampling	Emphasize random sampling and often require fixed sample sizes.	Can adapt well to varying sample sizes since Bayesians update their beliefs as more (observed) data comes in.
Assumption	Parameters that you estimate are fixed and are a single point while samples are random variables	There is a probability distribution around both the parameters and the samples.
The regime for application	Law of large number using a large amount of data.	Probability is degree of belief. Applicable when one has limited data, priors, and computing power.

4.4.2 The Bayesian approach



$\mathbb{P}(\text{this} = \text{Bayes} | \text{data}) < 1$



Richard Price



Pierre-Simon Laplace

Figure 4.8: Founders of Bayesian statistics.

- Thomas Bayes (1701–1761) was an ordained minister who was also a talented mathematician and a Fellow of the Royal Society. Bayes came up with an ingenious solution to this problem, but died before publishing it. Fortunately, his friend Richard Price carried his work further and published it in 1764. Apparently independently, Laplace rediscovered essentially the same idea in 1774, and developed it much further. (See Figure 4.8.)

- The idea is to assume a **prior** probability distribution for θ -that is, a distribution representing the plausibility of each possible value of θ before the data is observed. Then, to make inferences about θ , one simply considers the conditional distribution of θ given the observed data. This is referred to as the **posterior** distribution, since it represents the plausibility of each possible value of θ after seeing the data.

- Mathematically, this is expressed via **Bayes' theorem**,

$$p(\theta|x) = \frac{p(x|\theta)p(\theta)}{p(x)} \propto p(x|\theta)p(\theta), \quad (4.3)$$

where x is the observed data (for example, $x = x_{1:n}$). In words, we say "the posterior is proportional to the likelihood times the prior". Bayes' theorem is essentially just the definition of conditional probability

$$P(B|A) = \frac{P(A \cap B)}{P(A)} = \frac{P(A|B)P(B)}{P(A)} \quad (4.4)$$

extended to conditional densities. (From the modern perspective, Bayes' theorem is a trivial consequence of the definition of a conditional density-however, when Bayes wrote his paper, the idea of a conditional probability density did not yet exist!)

- More generally, the Bayesian approach is to assume a prior distribution on any unknowns, and then just follow the rules of probability to answer any questions of interest. This provides a coherent framework for making inferences about unknown parameters θ as well as any future data or missing data, and for making rational decisions based on such inferences.

- The Bayesian approach can be summarized as follows: Assume a probability distribution on any unknowns (this the prior), assume the distribution of the knowns given the unknowns (this is the generating distribution or likelihood), and then just follow the rules of probability to answer any questions of interest.

What questions of interest often arise?

Here are some recurring examples:

- estimate some unknown parameter or property
- infer hidden/latent variables or missing data
- predict future data
- test a hypothesis
- choose among competing models.

What methods are employed?

In order to answer a question of interest, you usually have to get ahead of the posterior in one way or another, and compute one or more posterior expectations (integrals with respect to the posterior density). Three main categories of methods can be distinguished here: exact solution, deterministic approximation, and stochastic approximation.

1. Exact solution

In certain cases, it is computationally feasible to compute the posterior (and posterior expectations) exactly.

- Exponential families with conjugate priors often enable analytical solutions.
- Gaussians, in particular, are highly conducive to analytical solutions.

2. Deterministic approximation

Methods include:

- numerical integration, a.k.a. quadrature/cubature
- quasi-Monte Carlo (QMC), low discrepancy sequences

For low-dimensional integrals, numerical integration and QMC are superior to stochastic approximations.

QMC can sometimes perform well in high-dimensional situations as well.

3. Stochastic approximation

For high-dimensional integrals, stochastic approximations are often the only option. The basic idea is that samples from the posterior can be used to approximate posterior expectations. Methods include:

- Monte Carlo approximation, importance sampling
- Markov chain Monte Carlo (MCMC) - Gibbs sampling, Metropolis algorithm, Metropolis-Hastings algorithm, slice sampling, Hamiltonian MCMC
- sequential importance sampling, sequential Monte Carlo, population Monte Carlo
- approximate Bayesian computation (ABC)

4.4.3 Applications of Bayesian statistics

• Tracking. For vehicle guidance, navigation, and control, it is essential to know the state of the vehicle (location, orientation, velocity) of the vehicle at any given time. Usually, an array of sensors provides various kinds of information of varying quality (e.g., compass, accelerometers, gyroscope, GPS, vision, laser scanner), and this must be combined with knowledge of the vehicle's actions (e.g., wheels, propellers/turbines, rocket engines, ailerons), along with a physical model, in order to infer the state of the vehicle in real-time. In 1960, Rudolf Kalman proposed a solution using a Bayesian time-series model which became known as the Kalman filter. The Kalman filter and its successors have been extraordinarily successful-it is difficult to overstate their importance in the guidance systems of aircraft, spacecraft, and robotics.

• Phylogenetics. Understanding the evolutionary relationships among organisms—that is, the phylogenetic tree—is fundamental in nearly all biological research. Using genetic data from many organisms, along with models of how changes in the genome occur over time, researchers can infer the unknown evolutionary "family tree". Some of the dominant approaches use Bayesian inference (e.g., popular programs include MrBayes and BEAST) and these are widely used throughout biology.

• Computer science. Spam accounts for the majority of email traffic—typically between 60 to 70% of emails are spam. Yet, due to the sophisticated spam detection algorithms used by email service providers, very little spam gets through to your inbox—and only rarely is real mail classified as spam. For instance, in 2007, Gmail posted the chart, showing that the fraction of spam that gets through is very small indeed. Bayesian models are the most prominent methods for spam detection. A former Microsoft developer who moved to Google reportedly said, "Google uses Bayesian filtering the way Microsoft uses the if statement."

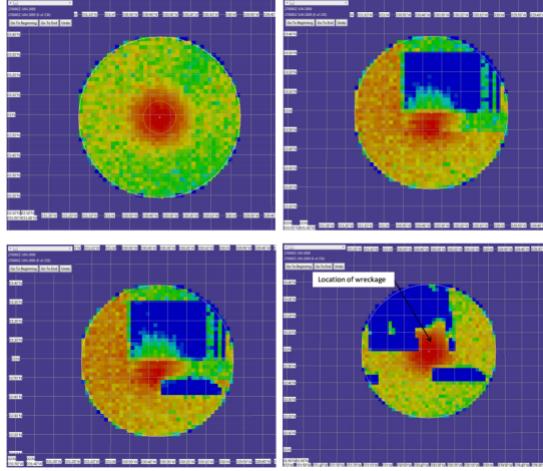


Figure 4.9: Prior and posteriors (after successive searches) for the location of the wreckage of Air France 447. (Stone et al. 2011).

- Search. On June 1, 2009, Air France Flight 447 crashed into the Atlantic Ocean, killing all aboard. Despite three intensive searches, the underwater wreckage had still not been found a year later. French authorities were eventually able to recover the wreckage with the help of a Bayesian search analysis provided by the Metron company (Figure 5). Bayesian search analysis involves formulating many hypothetical scenarios for what happened, constructing a probability distribution of the location under each scenario, and considering the posterior distribution on location given the searches conducted so far. It has also been used to find submarines and ships lost at sea.

4.4.4 Beta-Bernoulli model

Bernoulli distribution

- The Bernoulli distribution models binary outcomes, i.e., taking two possible values. The convention is to use 0 and 1.
 - It is named for Jacob Bernoulli (1655–1705), who is known for various inconsequential trivialities such as developing the foundations of probability theory (including the law of large numbers), combinatorics, differential equations, and the calculus of variations. Oh, and he discovered the constant e .
 - The Bernoulli distribution shows up everywhere, due to the ubiquity of binary outcomes (for instance, in computer vision, neuroscience, demographics and polling, public health and epidemiology, etc.).
 - The p.m.f. (probability mass function) is

$$p(x|\theta) = P(X = x|\theta) = \theta^x(1 - \theta)^{1-x} \mathbf{1}(x \in \{0, 1\}).$$

- If i.i.d. $X_1, \dots, X_n \sim \text{Bernoulli}(\theta)$ then for $x_1, \dots, x_n \in \{0, 1\}$,

$$\begin{aligned} p(x_{1:n}|\theta) &= P(X_1 = x_1, \dots, X_n = x_n|\theta) = \prod_{i=1}^n \theta^{x_i}(1 - \theta)^{1-x_i} \\ &= \theta^{\sum x_i}(1 - \theta)^{n - \sum x_i}. \end{aligned}$$

- Viewed as a function of θ , $p(x_{1:n}|\theta)$ is called the **likelihood function**. It is sometimes denoted $L(\theta; x_{1:n})$ to emphasize this. Viewed as a distribution on $x_{1:n}$, we will refer to this as the **generator** or **generating distribution** (sometimes it is referred to as the “sampling distribution”, but this becomes ambiguous when one is also sampling from the posterior).

Beta distribution

- Bayes used a uniform prior on θ , which is a special case of the beta distribution
- Given $a, b > 0$, we write $\theta \sim \text{Beta}(a, b)$ to mean θ has pdf (probability density function)

$$p(\theta) = \text{Beta}(\theta|a, b) = \frac{1}{B(a, b)} \theta^{a-1} (1-\theta)^{b-1} \mathbf{1}(0 < \theta < 1),$$

i.e., $p(\theta) \propto \theta^{a-1} (1-\theta)^{b-1}$ on the interval from 0 to 1. Here, $B(a, b)$ is Euler’s beta function defined as $\int_0^1 \theta^{a-1} (1-\theta)^{b-1} d\theta$.

- The mean is $E\theta = a/(a+b)$.

The posterior

- Using Bayes’ theorem, and plugging in the likelihood and the prior, the posterior is

$$\begin{aligned} p(\theta|x_{1:n}) &\propto p(x_{1:n}|\theta)p(\theta) \\ &= \theta^{\sum x_i} (1-\theta)^{n-\sum x_i} \frac{1}{B(a, b)} \theta^{a-1} (1-\theta)^{b-1} \mathbf{1}(0 < \theta < 1) \\ &\propto \theta^{a+\sum x_i-1} (1-\theta)^{b+n-\sum x_i-1} \mathbf{1}(0 < \theta < 1) \\ &\propto \text{Beta}(\theta|a + \sum x_i, b + n - \sum x_i). \end{aligned}$$

- So, the posterior has the same form (a Beta distribution) as the prior! When this happens, we say that the prior is **conjugate**.
- Since the posterior has such a nice form, it is easy to work with—e.g., for computing certain integrals with respect to the posterior, sampling from the posterior, and computing the posterior p.d.f. and its derivatives.

Marginal likelihood and posterior predictive

- The **marginal likelihood** is

$$p(x) = \int p(x|\theta)p(\theta)d\theta,$$

i.e., it is the marginal p.d.f./p.m.f. of the observed data, obtained by integrating θ out of the joint density $p(x, \theta) = p(x|\theta)p(\theta)$. When θ is a vector, this will be a multi-dimensional integral.

- When the data is a sequence $x = (x_1, \dots, x_n)$, the **posterior predictive distribution** is the distribution of X_{n+1} given $X_{1:n} = x_{1:n}$. When X_1, \dots, X_n, X_{n+1} are independent given $\theta = \theta$, the posterior predictive p.d.f./p.m.f. is given by

$$\begin{aligned} p(x_{n+1}|x_{1:n}) &= \int p(x_{n+1}|x_{1:n}, \theta)p(\theta|x_{1:n})d\theta \\ &= \int p(x_{n+1}|\theta)p(\theta|x_{1:n})d\theta. \end{aligned}$$

Example 4.4.1 (Beta-Bernoulli) If $\theta \sim \text{Beta}(a, b)$ and $X_1, \dots, X_n | (\theta = \theta)$ are i.i.d. $\text{Bernoulli}(\theta)$, then the marginal likelihood is

$$\begin{aligned} p(x_{1:n}) &= \int p(x_{1:n}|\theta)p(\theta)d\theta = \int_0^1 \theta^{\sum x_i} (1-\theta)^{n-\sum x_i} \frac{1}{B(a,b)} \theta^{a-1} (1-\theta)^{b-1} d\theta \\ &= \frac{B(a + \sum x_i, b + n - \sum x_i)}{B(a,b)} = \frac{B(a_n, b_n)}{B(a,b)}, \end{aligned}$$

by the integral definition of the Beta function, where we let $a_n = a + \sum x_i$ and $b_n = b + n - \sum x_i$ for brevity. Using the fact that $p(\theta|x_{1:n}) = \text{Beta}(\theta|a_n, b_n)$,

$$\begin{aligned} P(X_{n+1} = 1|x_{1:n}) &= \int P(X_{n+1} = 1|x_{1:n}, \theta)p(\theta|x_{1:n})d\theta = \int P(X_{n+1} = 1|\theta)p(\theta|x_{1:n})d\theta \\ &= \int \theta \text{Beta}(\theta|a_n, b_n) d\theta = \frac{a_n}{a_n + b_n}, \end{aligned}$$

hence, the posterior predictive pmf is

$$p(x_{n+1}|x_{1:n}) = \frac{a_n^{x_{n+1}} b_n^{1-x_{n+1}}}{a_n + b_n} \mathbf{1}(x_{n+1} \in \{0, 1\}).$$

4.4.5 Conjugate priors

Definition 4.4.2 Given a family $p(x|\theta) : \theta \in \Theta$ of generating distributions, a collection of priors $p_\alpha(\theta)$ indexed by $\alpha \in H$ is called a **conjugate prior family** if for any α and any data, the resulting posterior equals $p_{\alpha'}(\theta)$ for some $\alpha' \in H$.

Proposition 4.4.3 (Beta-Bernoulli model). The collection of $\text{Beta}(\theta|a, b)$ distributions, with $a, b > 0$, is conjugate to $\text{Bernoulli}(\theta)$, since the posterior is $p(\theta|x_{1:n}) = \text{Beta}(\theta|a + \sum x_i, b + n - \sum x_i)$.

Proposition 4.4.4 (Gamma-Exponential model). The collection of $\text{Gamma}(\theta|a, b)$ distributions, with $a, b > 0$, is conjugate to $\text{Exp}(\theta)$, since the posterior is $p(\theta|x_{1:n}) = \text{Gamma}(\theta|a + n, b + \sum x_i)$. The details are given as follows. Given the observation $x_{1:n}$, the likelihood function is

$$p(x_{1:n}|\theta) = \theta^n e^{-\theta \sum x_i}.$$

The prior $\text{Gamma}(a, b)$ is

$$p(\theta) = b e^{-b\theta} \frac{(b\theta)^{a-1}}{(a-1)!} \propto e^{-b\theta} \theta^{a-1}.$$

The posterior is also a gamma distribution,

$$p(\theta|x_{1:n}) \propto \theta^{a+n-1} e^{-(b+\sum x_i)\theta} = \text{Gamma}(\theta|a+n, b+\sum x_i).$$

Proposition 4.4.5 (Normal-Normal model) Suppose we are using an i.i.d. normal model with mean θ and precision $\lambda = \frac{1}{\sigma^2}$ (the inverse of the variance):

$$X_1, \dots, X_n \sim N(\theta, \lambda^{-1}).$$

Assume the precision $\lambda = \frac{1}{\sigma^2}$ is known and fixed, and θ is given a $N(\mu_0, \lambda_0^{-1})$ prior: $\theta \sim N(\mu_0, \lambda_0^{-1})$, i.e., $p(\theta) = N(\theta|\mu_0, \lambda_0^{-1})$. This is sometimes referred to as a Normal-Normal model. It turns out that the posterior is

$$\theta|x_{1:n} \sim N(M, L^{-1}),$$

i.e., $p(\theta|x_{1:n}) = N(\theta|M, L^{-1})$, where $L = \lambda_0 + n\lambda$ and

$$M = \frac{\lambda_0\mu_0 + \lambda \sum_{i=1}^n x_i}{\lambda_0 + n\lambda}.$$

Thus, the normal distribution is, itself, a conjugate prior for the mean of a normal distribution with known precision.

Proposition 4.4.6 (*InverseGamma-Normal model*) Now, suppose that the mean μ is given while the variance σ^2 is unknown, with i.i.d. $X_1, \dots, X_n \sim N(\mu, \sigma^2)$ as before. The likelihood function

$$p(x_{1:n}|\sigma^2) = (2\pi\sigma^2)^{-n/2} \exp\left(-\frac{\sum_{i=1}^n (x_i - \mu)^2}{2\sigma^2}\right),$$

is an Inverse-Gamma distribution on σ^2 . Recall that the density function of the inverse-Gamma distribution $Y \sim \text{InvGamma}(a, b)$ is $p(y) \propto y^{-a-1} \exp(-b/y)$, i.e.,

$$\text{InvGamma}(y|a, b) = \frac{b^a}{\Gamma(a)} y^{-a-1} \exp(-b/y).$$

In particular, the likelihood function corresponds to

$$\sigma^2 \sim \text{InvGamma}(n/2 - 1, \frac{1}{2} \sum_{i=1}^n (x_i - \mu)^2).$$

If we define a prior to be the inverse-Gamma distribution,

$$p(\sigma^2) \propto (\sigma^2)^{-(a+1)} \exp(-b/\sigma^2),$$

we can apply the Bayesian formula to obtain the posterior

$$p(\sigma^2|x_{1:n}) \propto p(x_{1:n}|\sigma^2)p(\sigma^2) \sim \text{InvGamma}\left(a + \frac{n}{2}, b + \frac{1}{2} \sum_{i=1}^n (x_i - \mu)^2\right).$$

(Project 4) Apply the Metropolis-Hastings scheme for estimating the variance of 1D Gaussian distribution if assuming that the mean of the Gaussian is given.

Proposition 4.4.7 (*NormalGamma-Normal model*) Now, suppose that both the mean μ and the precision $\lambda = \frac{1}{\sigma^2}$ are unknown, with i.i.d. $X_1, \dots, X_n \sim N(\mu, \lambda^{-1})$ as before. The NormalGamma(m, c, a, b) distribution, with $m \in \mathbb{R}$ and $c, a, b > 0$, is a joint distribution on (μ, λ) obtained by letting

$$\begin{aligned} \lambda &\sim \text{Gamma}(a, b), \\ \mu|\lambda &\sim N(m, (c\lambda)^{-1}). \end{aligned}$$

In other words, the joint pdf for the prior is

$$p(\mu, \lambda) = p(\mu|\lambda)p(\lambda) = N(\mu|m, (c\lambda)^{-1})\text{Gamma}(\lambda|a, b),$$

which we will denote by $\text{NormalGamma}(\mu, \lambda|m, c, a, b)$ following our usual convention. It turns out that this provides a conjugate prior for (μ, λ) . Indeed, the posterior is

$$\mu, \lambda|x_{1:n} \sim \text{NormalGamma}(M, C, A, B),$$

i.e., $p(\mu, \lambda | x_{1:n}) = \text{NormalGamma}(\mu, \lambda | M, C, A, B)$, where

$$\begin{aligned} M &= \frac{cm + \sum_{i=1}^n x_i}{c + n}, \\ C &= c + n, \\ A &= a + n/2, \\ B &= b + \frac{1}{2} \left(cm^2 - CM^2 + \sum_{i=1}^n x_i^2 \right). \end{aligned}$$

For interpretation, B can also be written (by rearranging terms) as

$$B = b + \frac{1}{2} \sum_{i=1}^n (x_i - \bar{x})^2 + \frac{1}{2} \frac{cn}{c+n} (\bar{x} - m)^2.$$

4.4.6 Exponential families

- Exponential families are a unifying generalization of many basic probabilistic models, and they possess many special properties. In fact, we have already encountered several exponential families—Bernoulli, Beta, Exponential, and Gamma—and there are many more. From the Bayesian perspective, a key feature of exponential families is that often the posterior has a nice form. While exponential families are useful in their own right, they can also be combined to construct more complex models in a way that is amenable to inference with Markov chain Monte Carlo, as we will see later.
- Most of the familiar distributions are exponential families, such as Bernoulli, binomial, Poisson, exponential, beta, gamma, inverse gamma, normal (Gaussian), multivariate normal, log-normal, inverse Gaussian, Dirichlet, and others. On the other hand, the Cauchy distribution and Student's t-distribution are familiar examples that are not exponential families.
- The concept of exponential families was developed by E. J. G. Pitman (1897–1993), Bernard Koopman (1900–1981), and Georges Darmois (1888–1960).
- A one-parameter exponential family is a collection of probability distributions indexed by a parameter $\theta \in \Theta$, such that the pdf/pmf are of the form

$$p(x|\theta) = \exp(\varphi(\theta)t(x) - \kappa(\theta)) h(x),$$

for some functions $\varphi(\theta)$, $t(x)$, $\kappa(\theta)$, and $h(x)$. Terminology: $t(x)$ is called the **sufficient statistic function**.

- The generalization to more than one parameter is straightforward. An **exponential family** is a collection of distributions indexed by $\theta \in \Theta$, with p.d.f.s/p.m.f.s of the form

$$p(x|\theta) = \exp(\varphi(\theta)^T t(x) - \kappa(\theta)) h(x),$$

for some vector-valued functions

$$\varphi(\theta) = \begin{pmatrix} \varphi_1(\theta) \\ \vdots \\ \varphi_k(\theta) \end{pmatrix} \quad \text{and} \quad t(x) = \begin{pmatrix} t_1(x) \\ \vdots \\ t_k(x) \end{pmatrix},$$

and some real-valued functions $\kappa(\theta)$ and $h(x)$. Here, v^T denotes the transpose of v . As before, $\kappa(\theta)$ is the log-normalization constant, and the content above regarding natural form and conjugate priors applies also to the multi-parameter case.

Chapter 5

Poisson Process

5.1 Introduction

- In making a mathematical model for a real-world phenomenon it is always necessary to make certain **simplifying assumptions** so as to render the mathematics tractable.
 - On the other hand, however, we **cannot make too many simplifying assumptions**, for then our conclusions, obtained from the mathematical model, would not be applicable to the real-world situation.
 - Thus, in short, we must make enough simplifying assumptions to enable us to handle the mathematics but not so many that the mathematical model no longer resembles the real-world phenomenon.
 - One simplifying assumption that is often made is to assume that certain random variables are exponentially distributed. The reason for this is that the exponential distribution is both relatively easy to work with and is often a good approximation to the actual distribution.

5.2 The Exponential Distribution

5.2.1 Definition

See the definition in previous subsection 1.6.2.

折扣回报

Example 5.2.1 (Exponential Random Variables and Expected Discounted Returns) Suppose that you are receiving rewards at randomly changing rates continuously throughout time. Let $r(x)$ denote the random rate at which you are receiving rewards at time x . For a value $\alpha \geq 0$, called the *discount rate*, the quantity

$$R = \int_0^\infty e^{-\alpha x} r(x) dx,$$

represents the total discounted reward. (In certain applications, α is a continuously compounded interest rate, and R is the present value of the infinite flow of rewards.) Whereas

$$E[R] = E \left[\int_0^\infty e^{-\alpha x} r(x) dx \right] = \int_0^\infty e^{-\alpha x} E[r(x)] dx,$$

is the expected total discounted reward, we will show that it is also equal to the expected total reward earned up to an exponentially distributed random time with rate α . Let T be an exponential random variable with rate α that is independent of all the random variables $r(x)$. We want to argue that

$$\int_0^\infty e^{-\alpha x} E[r(x)]dx = E \left[\int_0^T r(x)dx \right].$$

To show this define for each $x \geq 0$ a random variable $I(x)$ by

$$I(x) = \begin{cases} 1, & \text{if } x \leq T \\ 0, & \text{if } x > T \end{cases}$$

and note that

$$\int_0^T r(x)dx = \int_0^\infty r(x)I(x)dx.$$

Thus,

$$\begin{aligned} E \left[\int_0^T r(x)dx \right] &= E \left[\int_0^\infty r(x)I(x)dx \right] = \int_0^\infty E[r(x)I(x)]dx \\ &= \int_0^\infty E[r(x)]E[I(x)]dx \quad \text{by independence} \\ &= \int_0^\infty E[r(x)]P[T \geq x]dx \\ &= \int_0^\infty e^{-\alpha x} E[r(x)]dx. \end{aligned}$$

Therefore, the expected total discounted reward is equal to the expected total (undiscounted) reward earned by a random time that is exponentially distributed with a rate equal to the discount factor.

5.2.2 Properties of the Exponential Distribution

Proposition 5.2.2 A random variable X is without memory or memoryless if and only if X is exponentially distributed.

Example 5.2.3 The dollar amount of damage involved in an automobile accident is an exponential random variable with mean 1000. Of this, the insurance company only pays that amount exceeding (the deductible amount of) 400. Find the expected value and the standard deviation of the amount the insurance company pays per accident.

Sol. If X is the dollar amount of damage resulting from an accident, then the amount paid by the insurance company is $(X - 400)^+$. It is easier to condition on whether X exceeds 400. So let

$$I(x) = \begin{cases} 1, & \text{if } X > 400 \\ 0, & \text{if } x \leq 400 \end{cases}$$

Let $Y = (X - 400)^+$ be the amount paid. Then

$$\begin{aligned} E[Y|I = 1] &= 1000, \\ E[Y|I = 0] &= 0, \\ Var[Y|I = 1] &= (1000)^2, \\ Var[Y|I = 0] &= 0, \end{aligned}$$

which can be conveniently written as

$$E[Y|I] = 10^3 I, \quad \text{Var}[Y|I] = 10^6 I.$$

Because I is a Bernoullie random variable that is equal to 1 with probability $e^{-0.4}$, we obtain

$$E[Y] = E[E[Y|I]] = 10^3 E[I] = 10^3 e^{-0.4} = 670.32.$$

By conditional variance formula,

$$\begin{aligned} \text{Var}(Y) &= E[\text{Var}(Y|I)] + \text{Var}(E[Y|I]) \\ &= 10^6 e^{-0.4} + 10^6 e^{-0.4} (1 - e^{-0.4}). \end{aligned}$$

Method 2. Direct computation. $\lambda = 1/1000$. Let $Y = (X - 400)^+$. Then

$$\begin{aligned} EY &= \int_0^{400} 0 \cdot p(x)dx + \int_{400}^{\infty} (x - 400) \cdot \lambda e^{-\lambda x} dx \\ &= \int_0^{\infty} y \lambda e^{-\lambda(y+400)} dy = e^{-400\lambda} \cdot \frac{1}{\lambda} = 10^3 e^{-0.4}. \end{aligned}$$

Also

$$\begin{aligned} EY^2 &= \int_{400}^{\infty} (x - 400)^2 \cdot \lambda e^{-\lambda x} dx \\ &= e^{-400\lambda} \left(\frac{1}{\lambda^2} + \frac{1}{\lambda^2} \right). \\ \text{Var}[Y] &= e^{-400\lambda} \left(\frac{1}{\lambda^2} + \frac{1}{\lambda^2} \right) - (10^3 e^{-0.4})^2 \\ &= 10^6 \cdot 2 \cdot e^{-0.4} - 10^6 e^{-0.8}. \end{aligned}$$

Proposition 5.2.4 Let X_1, \dots, X_n be independent and identically distributed exponential random variables having mean $1/\lambda$. It is true that $X_1 + \dots + X_n$ has a gamma distribution with parameters n and λ . Let us now give a second verification of this result by using mathematical induction. Assume that $X_1 + \dots + X_{n-1}$ has density given by

$$f_{S_{n-1}}(t) = \lambda e^{-\lambda t} \frac{(\lambda t)^{n-2}}{(n-2)!}.$$

Hence

$$\begin{aligned} f_{S_n}(t) &= f_{X_n} * f_{S_{n-1}} = \int_0^{\infty} f_{X_n}(t-s) f_{S_{n-1}}(s) ds \\ &= \int_0^t \lambda e^{-\lambda(t-s)} \lambda e^{-\lambda s} \frac{(\lambda s)^{n-2}}{(n-2)!} ds = \lambda e^{-\lambda t} \frac{(\lambda t)^{n-1}}{(n-1)!}. \end{aligned}$$

Proposition 5.2.5 Another useful calculation is to determine the probability that one exponential random variable is smaller than another. That is, suppose that X_1 and X_2 are independent exponential random variables with respective means $1/\lambda_1$ and $1/\lambda_2$; what is $P\{X_1 < X_2\}$? This probability is easily calculated by conditioning on X_1 :

$$\begin{aligned} P\{X_1 < X_2\} &= \int_0^{\infty} P\{X_1 < X_2 | X_1 = x\} f_{X_1}(x) dx \\ &= \int_0^{\infty} P\{x < X_2\} \lambda_1 e^{-\lambda_1 x} dx \\ &= \int_0^{\infty} e^{-\lambda_2 x} \lambda_1 e^{-\lambda_1 x} dx = \frac{\lambda_1}{\lambda_1 + \lambda_2}. \end{aligned}$$

Proposition 5.2.6 Suppose that X_1, X_2, \dots, X_n are independent exponential random variables, with X_i having rate μ_i , $i = 1, \dots, n$. It turns out that the smallest of the X_i is exponential with a rate equal to the sum of the μ_i . This is shown as follows:

$$\begin{aligned} P\{\min(X_1, \dots, X_n) > x\} &= P(X_i > x \text{ for each } i = 1, \dots, n) \\ &= \prod_{i=1}^n P(X_i > x) \quad (\text{by independence}) \\ &= \prod_{i=1}^n e^{-\mu_i x} = \exp \left\{ - \left(\sum_{i=1}^n \mu_i \right) x \right\}. \end{aligned}$$

That is, $\min(X_1, \dots, X_n) \sim \mathcal{E}(\mu_1 + \dots + \mu_n)$.

Example 5.2.7 Suppose you arrive at a post office having two clerks at a moment when both are busy but there is no one else waiting in line. You will enter service when either clerk becomes free. If service times for clerk i are exponential with rate λ_i , $i = 1, 2$, find $E[T]$, where T is the amount of time that you spend in the post office.

Sol. Let R_i denote the remaining service time of the customer with clerk i , $i = 1, 2$, and note, by the lack of memory property of exponentials, that R_1 and R_2 are independent exponential random variables with respective rates λ_1 and λ_2 . Conditioning on which of R_1 or R_2 is the smallest yields

$$\begin{aligned} E[T] &= E[T|R_1 < R_2]P\{R_1 < R_2\} + E[T|R_2 < R_1]P\{R_2 < R_1\} \\ &= E[T|R_1 < R_2] \frac{\lambda_1}{\lambda_1 + \lambda_2} + E[T|R_2 < R_1] \frac{\lambda_2}{\lambda_1 + \lambda_2}. \end{aligned}$$

Now, with S denoting your service time

$$\begin{aligned} E[T|R_1 < R_2] &= E[R_1 + S|R_1 < R_2] \\ &= E[R_1|R_1 < R_2] + E[S|R_1 < R_2] \\ &= E[\min\{R_1, R_2\}] + \frac{1}{\lambda_1} \\ &= \frac{1}{\lambda_1 + \lambda_2} + \frac{1}{\lambda_1}. \end{aligned}$$

The final equation used that conditional on $R_1 < R_2$ the random variable R_1 is the minimum of R_1 and R_2 and is thus exponential with rate $\lambda_1 + \lambda_2$; and also that conditional on $R_1 < R_2$ you are served by server 1. As we can show in a similar fashion that

$$E[T|R_2 < R_1] = \frac{1}{\lambda_1 + \lambda_2} + \frac{1}{\lambda_2},$$

we obtain the result

$$E[T] = \frac{3}{\lambda_1 + \lambda_2}.$$

Another way to obtain $E[T]$ is to write T as a sum, take expectations, and then condition where needed. This approach yields

$$\begin{aligned} E[T] &= E[\min(R_1, R_2) + S] \\ &= E[\min(R_1, R_2)] + E[S] \\ &= \frac{1}{\lambda_1 + \lambda_2} + ES. \end{aligned}$$

To compute $E[S]$, we condition on which of R_1 and R_2 is smallest.

$$\begin{aligned} E[S] &= E[S|R_1 < R_2] \frac{\lambda_1}{\lambda_1 + \lambda_2} + E[S|R_1 \geq R_2] \frac{\lambda_2}{\lambda_1 + \lambda_2} \\ &= \frac{2}{\lambda_1 + \lambda_2}. \end{aligned}$$

5.3 The Poisson Process

5.3.1 Counting Processes

Definition 5.3.1 A stochastic process $\{N(t), t \geq 0\}$ is said to be a counting process if $N(t)$ represents the total number of “events” that occur by time t . A counting process must satisfy:

- (i) $N(t) \geq 0$.
- (ii) $N(t)$ is integer valued.
- (iii) If $s < t$, then $N(s) \leq N(t)$.
- (iv) For $s < t$, $N(t) - N(s)$ equals the number of events that occur in the interval $(s, t]$.

Some examples of counting processes are the following:

- If we let $N(t)$ equal the number of persons who enter a particular store at or prior to time t , then $\{N(t), t \geq 0\}$ is a counting process in which an event corresponds to a person entering the store. Note that if we had let $N(t)$ equal the number of persons in the store at time t , then $\{N(t), t \geq 0\}$ would not be a counting process (why not?).
- If we say that an event occurs whenever a child is born, then $\{N(t), t \geq 0\}$ is a counting process when $N(t)$ equals the total number of people who were born by time t . (Does $N(t)$ include persons who have died by time t ? Explain why it must.)
- If $N(t)$ equals the number of goals that a given soccer player scores by time t , then $\{N(t), t \geq 0\}$ is a counting process. An event of this process will occur whenever the soccer player scores a goal.

独立增量

Definition 5.3.2 A counting process is said to possess **independent increments** if the numbers of events that occur in disjoint time intervals are independent. For example, this means that the number of events that occur by time 10 (that is, $N(10)$) must be independent of the number of events that occur between times 10 and 15 (that is, $N(15) - N(10)$).

Remark 5.3.3 The assumption of independent increments might be reasonable for example (a), but it probably would be unreasonable for example (b). The reason for this is that if in example (b) $N(t)$ is very large, then it is probable that there are many people alive at time t ; this would lead us to believe that the number of new births between time t and time $t+s$ would also tend to be large (that is, it does not seem reasonable that $N(t)$ is independent of $N(t+s) - N(t)$, and so $\{N(t), t \geq 0\}$ would not have independent increments in example (b)). The assumption of independent increments in example (c) would be justified if we believed that the soccer player’s chances of scoring a goal today do not depend on “how he’s been going.” It would not be justified if we believed in “hot streaks” or “slumps.”

平稳增量

Definition 5.3.4 A counting process is said to possess **stationary increments** if the distribution of the number of events that occur in any interval of time depends only on the length of the time interval. In other words, the process has stationary increments if the number of events in the interval $(s, s+t)$ has the same distribution for all s .

Remark 5.3.5 The assumption of stationary increments would only be reasonable in example (a) if there were no times of day at which people were more likely to enter the store. Thus, for instance, if there was a rush hour (say, between 12 P.M. and 1 P.M.) each day, then the stationarity assumption would not be justified. If we believed that the earth's population is basically constant (a belief not held at present by most scientists), then the assumption of stationary increments might be reasonable in example (b). Stationary increments do not seem to be a reasonable assumption in example (c) since, for one thing, most people would agree that the soccer player would probably score more goals while in the age bracket 25–30 than he would while in the age bracket 35–40. It may, however, be reasonable over a smaller time horizon, such as one year.

5.3.2 Definition of the Poisson Process

Definition 5.3.6 The counting process $\{N(t), t \geq 0\}$ is said to be a Poisson process having rate λ , $\lambda > 0$, if

(i) $N(0) = 0$.

(ii) The process has independent increments.

(iii) The number of events in any interval of length t is Poisson distributed with mean λt . That is, for all $s, t \geq 0$

$$P\{N(t+s) - N(s) = n\} = e^{-\lambda t} \frac{(\lambda t)^n}{n!}, \quad n = 0, 1, \dots$$

Remark 5.3.7 Note that it follows from condition (iii) that a **Poisson process** has stationary increments

Remark 5.3.8 Note that $E[N(t)] = \lambda t$ and $\text{Var}[N(t)] = \lambda t$, and this is why λ is called the rate of the process since $\lambda = E[N(t)]/t$ denotes the number of events that occur in a unit time.

Example 5.3.9 For a stock market, the transactions of the stock trading are a Poisson process. There are 120,000 trading transactions every 10 minutes.

(1) Compute the rate λ for number of times per minute.

(2) The prob. for 100 trading transactions in 1 sec?

Sol. (1) Since

$$E[N(t, t+10)] = 10\lambda = 120000.$$

Then $\lambda = 12000$ per minute.

(2) The rate is $\lambda_2 = 12000/60 = 200$ per second.

$$P(N_2(1) = 100) = \frac{(\lambda_2 t)^n}{n!} e^{-\lambda_2 t} = \frac{(200)^{100}}{100!} e^{-200} = 1.88 \times 10^{-15}.$$

Remark 5.3.10 The Poisson processes are early proposed by the French scientist, Simeon-Denis Poisson, so that named after him. Poisson process is one of the widely used counting processes. Many counting processes with independent increments and stationary increments are Poisson processes. When determining if a counting process is Poisson one, one can apply the following alternative definition.

Definition 5.3.11 The function $f(\cdot)$ is said to be $o(h)$ if

$$\lim_{h \rightarrow 0} \frac{f(h)}{h} = 0.$$

Definition 5.3.12 The counting process $\{N(t), t \geq 0\}$ is said to be a Poisson process having rate λ , $\lambda > 0$, if

- (i) $N(0) = 0$.
- (ii) The process has stationary and independent increments.
- (iii) $P\{N(h) = 1\} = \lambda h + o(h)$.
- (iv) $P\{N(h) \geq 2\} = o(h)$.

Theorem 5.3.13 Definition 5.3.6 and 5.3.12 are equivalent.

Proof. We can show 5.3.6 \Rightarrow 5.3.12 by using Taylor expansion.

$$\begin{aligned} P(N(h) = 1) &= \lambda h \cdot e^{-\lambda h} = \lambda h (1 - \lambda h + o(h)) = \lambda h + o(h). \\ P\{N(h) \geq 2\} &= 1 - P(N(h) = 0) - P(N(h) = 1) \\ &= 1 - e^{-\lambda h} - \lambda h e^{-\lambda h} \\ &= 1 - [1 - \lambda h + o(h)] - [\lambda h + o(h)] = o(h). \end{aligned}$$

We then show the reverse, 5.3.12 \Rightarrow 5.3.6. To start, fix $u \geq 0$ and let

$$g(t) = E[\exp\{-uN(t)\}].$$

We derive a differential equation for $g(t)$ as follows:

$$\begin{aligned} g(t+h) &= E[\exp\{-uN(t+h)\}] \\ &= E[\exp\{-uN(t)\} \exp\{-u[N(t+h) - N(t)]\}] \\ &= E[\exp\{-uN(t)\}] E[\exp\{-u[N(t+h) - N(t)]\}] \text{ by independent increments} \\ &= g(t) E[\exp\{-uN(h)\}] \text{ by stationary increments.} \end{aligned}$$

Now, assumptions (iii) and (iv) imply that

$$P(N(h) = 0) = 1 - \lambda h + o(h).$$

Hence, conditioning on whether $N(h) = 0$ or $N(h) = 1$ or $N(h) \geq 2$ yields

$$\begin{aligned} E[\exp\{-uN(h)\}] &= 1 \cdot P\{N(h) = 0\} + e^{-u} P\{N(h) = 1\} \\ &\quad + E[\exp\{-uN(h)\}|N(h) \geq 2] P\{N(h) \geq 2\} \\ &= (1 - \lambda h) + e^{-u} \lambda h + o(h). \end{aligned}$$

Then we have

$$g(t+h) = g(t)[1 - \lambda h + e^{-u} \lambda h + o(h)],$$

implying that

$$\frac{g(t+h) - g(t)}{h} = g(t) \lambda (e^{-u} - 1) + \frac{o(h)}{h}.$$

Letting $h \rightarrow 0$ gives

$$g'(t) = g(t)\lambda(e^{-u} - 1).$$

Integrating, and using $g(0) = E(\exp(-uN(0))) = E1 = 1$, shows that

$$\begin{aligned}\log(g(t)) &= \lambda t(e^{-u} - 1), \\ g(t) &= \exp[\lambda t(e^{-u} - 1)].\end{aligned}$$

That is, the Laplace transform of $N(t)$ evaluated at u is $\exp[\lambda t(e^{-u} - 1)]$. For Poisson process, its MGF is

$$\begin{aligned}g(t) &= E[e^{-uN(t)}] = \sum_{n=0}^{\infty} \frac{(\lambda t)^n e^{-\lambda t}}{n!} e^{-un} \\ &= e^{-\lambda t} \sum_{n=0}^{\infty} \frac{(e^{-u}\lambda t)^n}{n!} = e^{-\lambda t} \exp\{e^{-u}\lambda t\} = \exp[\lambda t(e^{-u} - 1)].\end{aligned}$$

Since that is also the Laplace transform of a Poisson random variable with mean λt , the result follows from the fact that the distribution of a nonnegative random variable is uniquely determined by its Laplace transform. (Or say, that MGF has a uniqueness property, that is, if MGF exists, then there exists only one unique corresponding distribution associated with that MGF). ■

5.3.3 Interarrival and Waiting Time Distributions

Poisson 呼叫流, 到达时间间隔

Interarrival Time

Definition 5.3.14 Consider a Poisson process, and let us denote the time of the first event by T_1 . Further, for $n > 1$, let T_n denote the elapsed time between the $(n-1)$ st and the n th event. The sequence $\{T_n, n = 1, 2, \dots\}$ is called the **sequence of interarrival times**. For instance, if $T_1 = 5$ and $T_2 = 10$, then the first event of the Poisson process would have occurred at time 5 and the second at time 15.

We shall now determine the distribution of the T_n . To do so, we first note that the event $\{T_1 > t\}$ takes place if and only if no events of the Poisson process occur in the interval $[0, t]$ and thus,

$$\begin{aligned}P\{T_1 > t\} &= P\{N(t) = 0\} = e^{-\lambda t}, \\ T_1 &\sim \mathcal{E}(\lambda), \\ f_{T_1}(x) &= \lambda e^{-\lambda x}.\end{aligned}$$

Hence, T_1 has an exponential distribution with mean $1/\lambda$. Now,

$$\begin{aligned}P\{T_2 > t\} &= E[P(T_2 > t | T_1)] = \int_0^\infty P(T_2 > t | T_1 = s) f_{T_1}(s) ds \\ &= \int_0^\infty P(0 \text{ events in } (s, s+t] | T_1 = s) f_{T_1}(s) ds \\ &= \int_0^\infty P(0 \text{ events in } (s, s+t]) f_{T_1}(s) ds \\ &= \int_0^\infty P(N(t+s) - N(s) = 0) f_{T_1}(s) ds \\ &= \int_0^\infty e^{-\lambda t} f_{T_1}(s) ds = e^{-\lambda t}.\end{aligned}$$

Therefore, we conclude that T_2 is also an exponential random variable with mean $1/\lambda$,

$$\begin{aligned} T_2 &\sim \mathcal{E}(\lambda), \\ f_{T_2}(x) &= \lambda e^{-\lambda x}. \end{aligned}$$

and, furthermore, that T_2 is independent of T_1 . Repeating the same argument yields the following.

Proposition 5.3.15 $T_n, n = 1, 2, \dots$, are independent identically distributed exponential random variables having mean $1/\lambda$.

Remark 5.3.16 The proposition should not surprise us. The assumption of stationary and independent increments is basically equivalent to asserting that, at any point in time, the process probabilistically restarts itself. That is, the process from any point on is independent of all that has previously occurred (by independent increments), and also has the same distribution as the original process (by stationary increments). In other words, **the process has no memory**, and hence exponential interarrival times are to be expected.

Remark 5.3.17 The mean time for 1 event is $1/\lambda$, correspondingly, so that for 1 unit time, $E[N(t)]/t = \lambda$ events occurs.

Waiting Time

第n个事件的等待时或呼叫时或达到时

Definition 5.3.18 Another quantity of interest is S_n , the arrival time of the nth event, also called the **waiting time** until the nth event. It is easily seen that

$$S_n = \sum_{i=1}^n T_i, \quad n \geq 1.$$

and hence from Proposition 5.3.15 it follows that S_n has a gamma distribution with parameters n and λ . That is, the probability density of S_n is given by

$$f_{S_n}(t) = \lambda e^{-\lambda t} \frac{(\lambda t)^{n-1}}{(n-1)!}, \quad t \geq 0.$$

Remark 5.3.19 The above gamma distribution density may also be derived by noting that the nth event will occur prior to or at time t if and only if the number of events occurring by time t is at least n . That is,

$$\{N(t) \geq n\} \Leftrightarrow \{S_n \leq t\}$$

(Also notice that $\{N(t) = n\} \Leftrightarrow \{S_n \leq t < S_{n+1}\}$), Hence,

$$F_{S_n}(t) = P\{S_n \leq t\} = P\{N(t) \geq n\} = \sum_{j=n}^{\infty} e^{-\lambda t} \frac{(\lambda t)^j}{j!},$$

which upon differentiation, yields

$$\begin{aligned} f_{S_n}(t) &= - \sum_{j=n}^{\infty} \lambda e^{-\lambda t} \frac{(\lambda t)^j}{j!} + \sum_{j=n}^{\infty} \lambda e^{-\lambda t} \frac{(\lambda t)^{j-1}}{(j-1)!} \\ &= \lambda e^{-\lambda t} \frac{(\lambda t)^{n-1}}{(n-1)!} + \sum_{j=n+1}^{\infty} \lambda e^{-\lambda t} \frac{(\lambda t)^{j-1}}{(j-1)!} - \sum_{j=n}^{\infty} \lambda e^{-\lambda t} \frac{(\lambda t)^j}{j!} \\ &= \lambda e^{-\lambda t} \frac{(\lambda t)^{n-1}}{(n-1)!}. \end{aligned}$$

Example 5.3.20 Suppose that people immigrate into a territory at a Poisson rate $\lambda = 1$ per day.

- (a) What is the expected time until the tenth immigrant arrives?
 - (b) What is the probability that the elapsed time between the tenth and the eleventh arrival exceeds two days?
- Sol. (a) $E[S_{10}] = 10E[T_i] = 10/\lambda = 10$ days.
(b) $P(T_{11} > 2) = e^{-2\lambda} = e^{-2} \approx 0.133$.

Remark 5.3.21 Proposition 5.3.15 also gives us another way of defining a Poisson process. Suppose we start with a sequence $\{T_n, n \geq 1\}$ of independent identically distributed exponential random variables each having mean $1/\lambda$. Now let us define a counting process by saying that the n th event of this process occurs at time

$$S_n := T_1 + \cdots + T_n.$$

The resultant counting process $\{N(t), t \geq 0\}^*$ will be Poisson with rate λ .

5.3.4 Further Properties of Poisson Processes

Poisson过程的分流与合流

- Consider a Poisson process $\{N(t), t \geq 0\}$ having rate λ , and suppose that each time an event occurs it is classified as either a type I or a type II event. Suppose further that each event is classified as a type I event with probability p or a type II event with probability $1 - p$, independently of all other events.
- For example, suppose that customers arrive at a store in accordance with a Poisson process having rate λ ; and suppose that each arrival is male with probability $1/2$ and female with probability $1/2$. Then a type I event would correspond to a male arrival and a type II event to a female arrival.

Proposition 5.3.22 Let $N_1(t)$ and $N_2(t)$ denote respectively the number of type I and type II events occurring in $[0, t]$. Note that $N(t) = N_1(t) + N_2(t)$. $\{N_1(t), t \geq 0\}$ and $\{N_2(t), t \geq 0\}$ are both Poisson processes having respective rates λp and $\lambda(1 - p)$. Furthermore, the two processes are independent.

Proof. We can verify that $\{N_1(t), t \geq 0\}$ is a Poisson process with rate λp by showing that it satisfies Definition 5.3.12.

- $N_1(0) = 0$ follows from the fact that $N(0) = 0$.
- $\{N_1(t), t \geq 0\}$ inherits the stationary and independent increment properties of the process $\{N(t), t \geq 0\}$.
- $P\{N_1(h) = 1\} = P\{N_1(h) = 1 | N(h) = 1\}P\{N(h) = 1\} + P\{N_1(h) = 1 | N(h) \geq 2\}P\{N(h) \geq 2\} = p(\lambda h + o(h)) + o(h) = \lambda ph + o(h)$.
- $P\{N_1(h) \geq 2\} \leq P\{N(h) \geq 2\} = o(h)$.

Thus we see that $\{N_1(t), t \geq 0\}$ is a Poisson process with rate λp and, by a similar argument, that $\{N_2(t), t \geq 0\}$ is a Poisson process with rate $\lambda(1 - p)$. Because the probability of a type I event in the interval from t to $t + h$ is independent of all that occurs in intervals that do not overlap $(t, t + h)$, it is independent of knowledge of when type II events occur, showing that the two Poisson processes are independent. ■

Example 5.3.23 If immigrants to area A arrive at a Poisson rate of ten per week, and if each immigrant is of English descent with probability $1/12$, then what is the probability that no people of English descent will emigrate to area A during the month of February?

Sol. Since $p\lambda t = 1/12 \cdot 10 \cdot 4 = 10/3$, so that $P(N(t) = 0) = e^{-p\lambda t} = e^{-10/3}$.

可分解性

Proposition 5.3.24 *The above decomposition property of Poisson processes can be generalized to any finite number of types. That is, consider a Poisson process $\{N(t), t \geq 0\}$ having rate λ , and suppose that each time an event occurring is classified into one of m types. Suppose further that each event is independently classified as j th ($j = 1, \dots, m$) type with probability p_j satisfying $\sum_{j=1}^m p_j = 1$. Let $N_j(t)$ denote the number of type j event occurring in $[0, t]$. Note that $N(t) = \sum_{j=1}^m N_j(t)$. Then, $\{N_j(t), t \geq 0\}$ ($j = 1, \dots, m$) are all Poisson processes having rates $\lambda_j := \lambda p_j$. Furthermore, all these m number of processes are independent.*

Example 5.3.25 *At a crossing road, 6 vehicles come from the east per minute, 6.5 from the south, 9 from the west, and 8.5 from the north. At the crossing, every vehicle turns left or right with prob. 0.3, go straight forward with prob. 0.35, and turn around with prob. 0.05. Compute the rate for vehicles leaving the crossing road at each direction.*

Sol.

Table 5.1: Poisson diversion and confluence.

	toward east	toward south	toward west	toward north
from east $\lambda_1 = 6.0$	$0.05\lambda_1$	$0.30\lambda_1$	$0.35\lambda_1$	$0.30\lambda_1$
from south $\lambda_2 = 6.5$	$0.30\lambda_2$	$0.05\lambda_2$	$0.30\lambda_2$	$0.35\lambda_2$
from west $\lambda_3 = 9.0$	$0.35\lambda_3$	$0.30\lambda_3$	$0.05\lambda_3$	$0.30\lambda_3$
from north $\lambda_4 = 8.5$	$0.30\lambda_4$	$0.35\lambda_4$	$0.30\lambda_4$	$0.05\lambda_4$
leaving rate	$\lambda_E = 7.95$	$\lambda_S = 7.80$	$\lambda_W = 7.05$	$\lambda_N = 7.20$

Proposition 5.3.26 *Let $\{N_1(t)\}$ and $\{N_2(t)\}$ be independent Poisson processes with rate λ_1 and λ_2 , respectively. Then*

$$N(t) = N_1(t) + N_2(t),$$

is a Poisson process with rate $\lambda = \lambda_1 + \lambda_2$.

Proof. • $N(0) = N_1(0) + N_2(0) = 0$.

- $\{N(t), t \geq 0\}$ has the stationary and independent increment properties.
-

$$\begin{aligned} P\{N(h) = 1\} &= P\{N_1(h) = 1, N_2(h) = 0\} + P\{N_1(h) = 0, N_2(h) = 1\} \\ &= (\lambda_1 h + o(h))(1 - \lambda_2 h + o(h)) + (\lambda_1 h + o(h))(1 - \lambda_2 h + o(h)) \\ &= (\lambda_1 + \lambda_2)h + o(h) = \lambda h + o(h). \end{aligned}$$

- Since

$$\begin{aligned} P\{N(h) = 0\} &= P\{N_1(h) = 0\}P\{N_2(h) = 0\} \\ &= (1 - \lambda_1 h + o(h))(1 - \lambda_2 h + o(h)) \\ &= 1 - \lambda h + o(h), \end{aligned}$$

so that

$$\begin{aligned} P\{N(h) \geq 2\} &\geq 1 - P\{N(h) = 0\} - P\{N(h) = 1\} \\ &= 1 - [1 - \lambda h + o(h)] - [\lambda h + o(h)] = o(h). \end{aligned}$$

■ 可加性

Proposition 5.3.27 *The above confluence property of Poisson processes can be generalized to any finite number of types. That is, let $\{N_j(t)\}$ ($j = 1, \dots, m$) be independent Poisson processes with rate λ_j ($j = 1, \dots, m$). Then*

$$N(t) = \sum_{j=1}^m N_j(t),$$

is a Poisson process with rate $\lambda = \sum_{j=1}^m \lambda_j$.

Example 5.3.28 (The Coupon Collecting Problem) There are m different types of coupons. Each time a person collects a coupon it is, independently of ones previously obtained, a type j coupon with probability p_j , $\sum_{j=1}^m p_j = 1$. Let N denote the number of coupons one needs to collect in order to have a complete collection of at least one of each type. Find $E[N]$.

Sol. If we let N_j denote the number one must collect to obtain a type j coupon, then we can express N as

$$N = \max_{1 \leq j \leq m} N_j.$$

However, even though each N_j is geometric with parameter p_j , the foregoing representation of N is not that useful, because the random variables N_j are not independent.

We can, however, transform the problem into one of determining the expected value of the maximum of independent random variables. To do so, suppose that coupons are collected at times chosen according to a Poisson process with rate $\lambda = 1$. Say that an event of this Poisson process is of type j , $1 \leq j \leq m$, if the coupon obtained at that time is a type j coupon. If we now let $N_j(t)$ denote the number of type j coupons collected by time t , then it follows from Proposition 5.3.24 that $\{N_j(t), t \geq 0\}$, $j = 1, \dots, m$ are independent Poisson processes with respective rates $\lambda p_j = p_j$. Let X_j denote the time of the first event of the j th process, and let

$$X = \max_{1 \leq j \leq m} X_j,$$

denote the time at which a complete collection is amassed. Since the X_j are independent exponential random variables with respective rates p_j , it follows that

$$\begin{aligned} P\{X < t\} &= P\{\max_{1 \leq j \leq m} X_j < t\} = P\{X_j < t, \text{ for all } j = 1, \dots, m\} \\ &= \prod_{j=1}^m P\{X_j < t\} = \prod_{j=1}^m (1 - e^{-p_j t}) = \prod_{j=1}^m (1 - e^{-p_j t}). \end{aligned}$$

Therefore,

$$\begin{aligned} E[X] &= \int_0^\infty P\{X > t\} dt \\ &= \int_0^\infty \left\{ 1 - \prod_{j=1}^m (1 - e^{-p_j t}) \right\} dt, \end{aligned}$$

where the proof follows that

$$\begin{aligned} E[X] &= \int_0^\infty P\{X > t\}dt, \\ \int_0^\infty tf(t)dt &= \int_0^\infty (1 - F(t))dt, \\ \int_0^\infty [tF'(t) - 1 + F(t)]dt &= 0, \\ \int_0^\infty (tF(t) - t)'dt &= 0, \\ (tF(t) - t)|_{t=0}^\infty &= 0. \end{aligned}$$

It remains to relate $E[X]$, the expected time until one has a complete set, to $E[N]$, the expected number of coupons it takes. This can be done by letting T_i denote the i th interarrival time of the Poisson process that counts the number of coupons obtained. Then it is easy to see that X is a compound random variable,

$$X = \sum_{i=1}^N T_i.$$

Since the T_i are independent exponentials with rate 1 ($E[T_i] = 1$), and N is independent of the T_i , we see that

$$E[X|N] = E\left[\sum_{i=1}^N T_i|N\right] = N E[T_i] = N$$

Therefore, using double expectation formula,

$$E[N] = E[X] = \int_0^\infty \left\{ 1 - \prod_{j=1}^m (1 - e^{-p_j t}) \right\} dt.$$

5.3.5 Conditional Distribution of the Arrival Times

Suppose we are told that exactly one event of a Poisson process has taken place by time t , and we are asked to determine the distribution of the time at which the event occurred. Now, since a Poisson process possesses stationary and independent increments it seems reasonable that each interval in $[0, t]$ of equal length should have the same probability of containing the event. In other words, the following can be checked.

Proposition 5.3.29 *The time of the event should be uniformly distributed over $[0, t]$. This is easily checked since, for $s \leq t$,*

$$\begin{aligned} P\{T_1 < s | N(t) = 1\} &= \frac{P\{T_1 < s, N(t) = 1\}}{P\{N(t) = 1\}} \\ &= \frac{P\{1 \text{ event in } [0, s], 0 \text{ events in } [s, t]\}}{P\{N(t) = 1\}} \\ &= \frac{P\{N(s) = 1\}P\{N(t) - N(s) = 0\}}{P\{N(t) = 1\}} \\ &= \frac{(\lambda s e^{-\lambda s})(e^{-\lambda(t-s)})}{\lambda t e^{-\lambda t}} = \frac{s}{t}, \end{aligned}$$

where we recall that $P\{N(t) = n\} = e^{-\lambda t}(\lambda t)^n/n!$.

This result may be generalized, but before doing so we need to introduce the concept of order statistics.

Definition 5.3.30 Let Y_1, Y_2, \dots, Y_n be n random variables. We say that $Y_{(1)}, Y_{(2)}, \dots, Y_{(n)}$ are the **order statistics** corresponding to Y_1, Y_2, \dots, Y_n if $Y_{(k)}$ is the k th smallest value among Y_1, \dots, Y_n , $k = 1, 2, \dots, n$. For instance, if $n = 3$ and $Y_1 = 4, Y_2 = 5, Y_3 = 1$ then $Y_{(1)} = 1, Y_{(2)} = 4, Y_{(3)} = 5$. If the Y_i , $i = 1, \dots, n$, are independent identically distributed continuous random variables with probability density f , then the joint density of the order statistics $Y_{(1)}, Y_{(2)}, \dots, Y_{(n)}$ is given by

$$f(y_1, \dots, y_n) = n! \prod_{i=1}^n f(y_i), \quad y_1 < y_2 < \dots < y_n.$$

The preceding follows since

- (i) $(Y_{(1)}, Y_{(2)}, \dots, Y_{(n)})$ will equal (y_1, y_2, \dots, y_n) if (Y_1, Y_2, \dots, Y_n) is equal to any of the $n!$ permutations of (y_1, y_2, \dots, y_n) ;
- (ii) the probability density that (Y_1, Y_2, \dots, Y_n) is equal to y_{i_1}, \dots, y_{i_n} is $\prod_{j=1}^n f(y_{i_j}) = \prod_{j=1}^n f(y_j)$ when i_1, \dots, i_n is a permutation of $1, 2, \dots, n$.

Proposition 5.3.31 If the Y_i , $i = 1, \dots, n$, are uniformly distributed over $(0, t)$, then we obtain from the preceding that the joint density function of the order statistics $Y_{(1)}, Y_{(2)}, \dots, Y_{(n)}$ is

$$f(y_1, y_2, \dots, y_n) = \frac{n!}{t^n}, \quad 0 < y_1 < y_2 < \dots < y_n < t.$$

We are now ready for the following useful theorem.

Theorem 5.3.32 Given that $N(t) = n$, the n arrival times S_1, \dots, S_n have the same distribution as the order statistics corresponding to n independent random variables uniformly distributed on the interval $(0, t)$.

Proof. To obtain the conditional density of S_1, \dots, S_n given that $N(t) = n$ note that for $0 < s_1 < \dots < s_n < t$ the event that $S_1 = s_1, S_2 = s_2, \dots, S_n = s_n, N(t) = n$ is equivalent to the event that the first $n+1$ interarrival times satisfy $T_1 = s_1, T_2 = s_2 - s_1, \dots, T_n = s_n - s_{n-1}, T_{n+1} > t - s_n$. Hence, we have that the conditional joint density of S_1, \dots, S_n given that $N(t) = n$ is as follows:

$$\begin{aligned} f(s_1, \dots, s_n | n) &= \frac{f(s_1, \dots, s_n, n)}{P\{N(t) = n\}} = \frac{g(s_1)g(s_2 - s_1) \cdots g(s_n - s_{n-1})P\{N(t) - N(s_n) = 0\}}{P\{N(t) = n\}} \\ &= \frac{\lambda e^{-\lambda s_1} \lambda e^{-\lambda(s_2 - s_1)} \cdots \lambda e^{-\lambda(s_n - s_{n-1})} e^{-\lambda(t - s_n)}}{(\lambda t)^n e^{-\lambda t} / n!} \\ &= \frac{n!}{t^n}, \quad 0 < s_1 < \dots < s_n < t, \end{aligned}$$

where $g(x) = \lambda e^{-\lambda x}$ is the density of the exponential distribution for T_j . ■

5.3.6 Applications of Central Limit Theorem

Proposition 5.3.33 Let $N(t)$ be a Poisson process. Let $Y_j = N(\frac{jt}{n}) - N(\frac{(j-1)t}{n})$, $j = 1, 2, \dots, n$ be i.i.d. random variables with $Y_1 = N(\frac{t}{n})$, $EY_1 = \frac{\lambda t}{n}$, $Var(Y_1) = \frac{\lambda t}{n} < \infty$. The sum is

$$N(t) = \sum_{j=1}^n Y_j,$$

having mean λt and variance λt . By Central Limit Theorem, for large λt (either λ large or t large),

$$\xi = \frac{N(t) - n\frac{\lambda t}{n}}{\sqrt{n\frac{\lambda t}{n}}} = \frac{N(t) - \lambda t}{\sqrt{\lambda t}} = \frac{N(t) - E[N(t)]}{\sqrt{Var(N(t))}} \sim N(0, 1).$$

Example 5.3.34 For a stock market, the transactions of the stock trading are a Poisson process. There are 120,000 trading transactions every 10 minutes. What is the prob. for the number of transactions greater than $k = 1050$ in 5 seconds?

Sol. $\lambda_1 = 200$ times per second and $t = 5$. Then

$$\begin{aligned} P\{N_1(t=5) > 1050\} &= P\left\{\frac{N_1(5) - 1000}{\sqrt{1000}} > \frac{1050 - 1000}{\sqrt{1000}}\right\} \\ &\approx P\{\xi > 1.58\} = 1 - 0.9429 = 0.0571. \end{aligned}$$

Notice that the exact solution

$$\begin{aligned} P\{N_1(5) > 1050\} &= 1 - P\{N_1(5) \leq 1050\} = 1 - \sum_{j=0}^k P\{N_1(5) = j\} \\ &= 1 - \sum_{j=0}^k \frac{1000^j}{j!} e^{-1000}, \end{aligned}$$

which is not easy to compute.

5.4 Generalization of the Poisson Process

5.4.1 Nonhomogeneous Poisson Process

In this section we consider two generalizations of the Poisson process. The first of these is the nonhomogeneous, also called the nonstationary, Poisson process, which is obtained by allowing the arrival rate at time t to be a function of t .

Definition 5.4.1 The counting process $\{N(t), t \geq 0\}$ is said to be a nonhomogeneous Poisson process with intensity function $\lambda(t)$, $t \geq 0$, if

- (i) $N(0) = 0$.
- (ii) $\{N(t), t \geq 0\}$ has independent increments.
- (iii) $P\{N(t+h) - N(t) \geq 2\} = o(h)$.
- (iv) $P\{N(t+h) - N(t) = 1\} = \lambda(t)h + o(h)$.

5.4.2 Compound Poisson Process

Definition 5.4.2 A stochastic process $\{X(t), t \geq 0\}$ is said to be a compound Poisson process if it can be represented as

$$X(t) = \sum_{i=1}^{N(t)} Y_i, \quad t \geq 0,$$

where $\{N(t), t \geq 0\}$ is a Poisson process, and $\{Y_i, i \geq 1\}$ is a family of i.i.d. random variables that is also independent of $\{N(t), t \geq 0\}$. The random variable $X(t)$ is said to be a compound Poisson random variable.

- If $Y_i \equiv 1$, then $X(t) = N(t)$, and so we have the usual Poisson process.
- Suppose that buses arrive at a sporting event in accordance with a Poisson process, and suppose that the numbers of fans in each bus are assumed to be i.i.d. Then $\{X(t), t \geq 0\}$ is a compound Poisson process

where $X(t)$ denotes the number of fans who have arrived by t . Then Y_i represents the number of fans in the i th bus.

- Suppose customers leave a supermarket in accordance with a Poisson process. If the Y_i , the amount spent by the i th customer, $i = 1, 2, \dots$, are i.i.d., then $\{X(t), t \geq 0\}$ is a compound Poisson process when $X(t)$ denotes the total amount of money spent by time t .

Proposition 5.4.3 *Because $X(t)$ is a compound Poisson random variable with Poisson parameter λt , we have*

$$\begin{aligned} E[X(t)] &= E\left[E\left[\sum_{i=1}^{N(t)} Y_i | N(t)\right]\right] = E[N(t)E[Y_i]] = E[N(t)]E[Y_i] \\ &= \lambda t E[Y_i], \end{aligned}$$

and

$$\begin{aligned} Var[X(t)] &= E(Var[X(t)|N(t)]) + Var(E[X(t)|N(t)]) \\ &= E\left(Var\left[\sum_{i=1}^{N(t)} Y_i | N(t)\right]\right) + Var\left(E\left[\sum_{i=1}^{N(t)} Y_i | N(t)\right]\right) \\ &= E\{N(t)Var[Y_i]\} + Var\{N(t)E[Y_i]\} \\ &= E[N(t)]Var[Y_i] + (E[Y_i])^2 Var[N(t)] \\ &= (\lambda t)\{Var[Y_i] + (E[Y_i])^2\} \\ &= (\lambda t)(EY_i^2). \end{aligned}$$

Example 5.4.4 Suppose that families migrate to an area at a Poisson rate $\lambda = 2$ per week. If the number of people in each family is independent and takes on the values 1, 2, 3, 4 with respective probabilities $1/6, 1/3, 1/3, 1/6$, then what is the expected value and variance of the number of individuals migrating to this area during a fixed five-week period?

Sol. Letting Y_i denote the number of people in the i th family, we have

$$\begin{aligned} E[Y_i] &= 1 \cdot \frac{1}{6} + 2 \cdot \frac{1}{3} + 3 \cdot \frac{1}{3} + 4 \cdot \frac{1}{6} = \frac{5}{2}, \\ E[Y_i^2] &= 1^2 \cdot \frac{1}{6} + 2^2 \cdot \frac{1}{3} + 3^2 \cdot \frac{1}{3} + 4^2 \cdot \frac{1}{6} = \frac{43}{6}. \end{aligned}$$

Hence, letting $X(5)$ denote the number of immigrants during a five-week period, we obtain that

$$\begin{aligned} E[X(5)] &= \lambda t E[Y_i] = 2 \cdot 5 \cdot \frac{5}{2} = 25, \\ Var[X(5)] &= (\lambda t)(EY_i^2) = 2 \cdot 5 \cdot \frac{43}{6} = \frac{215}{3}. \end{aligned}$$

Proposition 5.4.5 Similarly to the result in 5.3.33. Let $X(t)$ be a compound Poisson random variable. Then

$$\frac{X(t) - E[X(t)]}{\sqrt{Var(X(t))}} \sim N(0, 1),$$

when t is large or λ is large.

Example 5.4.6 In Example 5.4.4, find the approximate probability that at least 240 people migrate to the area within the next 50 weeks.

Sol. Since $\lambda = 2$, $E[Y_i] = 5/2$, $E[Y_i^2] = 43/6$, we see that

$$E[X(50)] = 250, \quad \text{Var}[X(50)] = 4300/6$$

Now, the desired probability is

$$\begin{aligned} P\{X(50) \geq 240\} &= P\{X(50) \geq 239.5\} \\ &= P\left\{\frac{X(50) - 250}{\sqrt{4300/6}} \geq \frac{239.5 - 250}{\sqrt{4300/6}}\right\} \\ &= 1 - \phi(-0.3922) = 0.6525. \end{aligned}$$

5.5 Parameter Estimation for Poisson Processes

5.5.1 Estimate the rate λ using $N(t)$

Let $N(t)$ be the Poisson process with the rate λ . Since the distribution of $N(t)$ is

$$P\{N(t) = n\} = \frac{(\lambda t)^n}{n!} e^{-\lambda t}, \quad n = 0, 1, 2, \dots$$

the likelihood function based on $n = N(t)$ is

$$L(\lambda) = \frac{(\lambda t)^n}{n!} e^{-\lambda t}.$$

The log of likelihood is

$$l(\lambda) = \ln L(\lambda) = n \ln \lambda - \lambda t + c_0, \quad c_0 \text{ is a constant.}$$

Using $l'(\lambda) = 0$, the maximum likelihood estimate (MLE) of λ is n/t . In general, if $N(t)$ number of events occur during $[0, t]$, then the MLE of λ is

$$\hat{\lambda} = \frac{N(t)}{t}.$$

Proposition 5.5.1 The MLE $\hat{\lambda}$ is unbiased and strongly consistent.

夹逼定理 强相合估计

Proof. Since $EN(t) = \lambda t$,

$$E\hat{\lambda} = \frac{EN(t)}{t} = \frac{\lambda t}{t} = \lambda,$$

the MLE $\hat{\lambda}$ is unbiased. For each $t > 0$, there exists an integer p , such that $p - 1 < t \leq p$. Then

$$\frac{N(p-1)}{p} \leq \frac{N(t)}{t} \leq \frac{N(p)}{p-1}.$$

Due to the independent increments and stationary increments,

$$N(p) = \sum_{j=1}^p N(j-1, j],$$

where $N(j-1, j]$ are i.i.d. random variables. Using the law of large number, we have

$$\begin{aligned}\lim_{p \rightarrow \infty} \frac{N(p-1)}{p} &= \lim_{p \rightarrow \infty} \frac{N(p-1)}{p-1} \lim_{p \rightarrow \infty} \frac{p-1}{p} = EN(1) = \lambda, \text{ a.s.} \\ \lim_{p \rightarrow \infty} \frac{N(p)}{p-1} &= EN(1) = \lambda, \text{ a.s.}\end{aligned}$$

Therefore,

$$\lim_{t \rightarrow \infty} \frac{N(t)}{t} = \lambda, \text{ a.s.}$$

that is, $\hat{\lambda}$ is strongly consistent. ■

5.5.2 Estimate the rate λ using S_n

Let the time for the occurrence of the n th event, $S_n = s_n$, be observed. The likelihood function is

$$L(\lambda) = \frac{\lambda^n}{\Gamma(n)} s_n^{n-1} e^{-\lambda s_n}.$$

The log of likelihood is

$$l(\lambda) = \ln L(\lambda) = n \ln \lambda - \lambda s_n + c_0, \text{ } c_0 \text{ is a constant.}$$

Using $l'(\lambda) = 0$, the maximum likelihood estimate (MLE) of λ is n/s_n . If $S_n = s_n$ is observed, then the MLE of λ is

$$\hat{\lambda} = \frac{n}{S_n}.$$

Proposition 5.5.2 *The MLE $\hat{\lambda}$ is not unbiased but strongly consistent.*

Proof. The MLE $\hat{\lambda}$ is not unbiased. Since $S_n \sim \Gamma(\lambda, n)$, we have

$$f_{S_n}(s) = \frac{\lambda^n s^{n-1}}{\Gamma(n)} e^{-\lambda s}.$$

Then

$$\begin{aligned}E\hat{\lambda} &= E \frac{n}{S_n} = \int_0^\infty \frac{n \lambda^n s^{n-1}}{s \Gamma(n)} e^{-\lambda s} ds = \int_0^\infty \frac{n \lambda}{n-1} \frac{\lambda^{n-1} s^{n-2}}{\Gamma(n-1)} e^{-\lambda s} ds \\ &= \frac{n \lambda}{n-1} \int_0^\infty f_{S_{n-1}}(s) ds = \frac{n \lambda}{n-1} > \lambda,\end{aligned}$$

and therefore the MLE $\hat{\lambda}$ is not unbiased. Based on the law of large number, when $n \rightarrow \infty$ we have

$$\hat{\lambda}^{-1} = \frac{S_n}{n} = \frac{T_1 + \dots + T_n}{n} \rightarrow ET_i = \frac{1}{\lambda}, \text{ a.s.}$$

and thus the MLE $\hat{\lambda}$ is strongly consistent. ■

Proposition 5.5.3 *The estimator $\hat{\lambda} = \frac{n-1}{S_n}$ is unbiased and strongly consistent. The proof follows from the above.*

Chapter 6

Continuous-Time Markov Chains

- Continuous-Time Markov Chains
- Pure birth process
- Birth and death process

6.1 Continuous-Time Markov Chains

Definition 6.1.1 Suppose we have a continuous-time stochastic process $\{X(t), t \geq 0\}$ taking on values in the set of nonnegative integers. We say that the process $\{X(t), t \geq 0\}$ is a **continuous-time Markov chain** if for all $s, t \geq 0$ and nonnegative integers $i, j, x(u), 0 \leq u < s$,

$$P\{X(t+s) = j | X(s) = i, X(u) = x(u), 0 \leq u < s\} = P\{X(t+s) = j | X(s) = i\}.$$

In other words, a continuous-time Markov chain is a stochastic process having the Markovian property that the conditional distribution of the future $X(t+s)$ given the present $X(s)$ and the past $X(u)$, $0 \leq u < s$, depends only on the present and is independent of the past.

Definition 6.1.2 If

$$P\{X(t+s) = j | X(s) = i\}$$

is independent of s , then the continuous-time Markov chain is said to have **stationary or homogeneous transition probabilities**. All Markov chains considered in this text will be assumed to have stationary transition probabilities.

Remark 6.1.3 Suppose that a continuous-time Markov chain enters state i at some time, say, time 0, and suppose that the process does not leave state i (that is, a transition does not occur) during the next ten minutes. What is the probability that the process will not leave state i during the following five minutes? Since the process is in state i at time 10 it follows, by the **Markovian property**, that the probability that it remains in that state during the interval $[10, 15]$ is just the (unconditional) probability that it stays in state i for at least five minutes. That is, if we let T_i denote the amount of time that the process stays in state i before making a transition into a different state, then

$$P\{T_i > 15 | T_i > 10\} = P\{T_i > 5\},$$

or, in general, by the same reasoning,

$$P\{T_i > s + t | T_i > s\} = P\{T_i > t\},$$

for all $s, t \geq 0$. Hence the random variable T_i is **memoryless** and must thus be **exponential distributed**.

Definition 6.1.4 Here is another way of defining a continuous-time Markov chain. Namely, it is a stochastic process having the properties that each time it enters state i ,

- (i) the amount of time it spends in that state before making a transition into a different state is exponentially distributed with mean, say, $1/v_i$, and
- (ii) when the process leaves state i , it next enters state j with some probability, say, p_{ij} . Of course, the p_{ij} must satisfy

$$\begin{aligned} p_{ii} &= 0, \quad \text{all } i, \\ \sum_j p_{ij} &= 1, \quad \text{all } i, \end{aligned}$$

where **the first condition implies that it must leave state i , otherwise contradictory to the Markovian property.**

Example 6.1.5 (A Shoe Shine Shop) Consider a shoe shine establishment consisting of two chairs—chair 1 and chair 2. A customer upon arrival goes initially to chair 1 where his shoes are cleaned and polish is applied. After this is done the customer moves on to chair 2 where the polish is buffed. The service times at the two chairs are assumed to be independent random variables that are exponentially distributed with respective rates μ_1 and μ_2 . Suppose that potential customers arrive in accordance with a Poisson process having rate λ , and that a potential customer will enter the system only if both chairs are empty.

An appropriate state space consists of three states 0, 1, 2 where

State	Interpretation
0	system is empty
1	a customer is in chair 1
2	a customer is in chair 2.

Then the parameters for the Markov chain are

$$\begin{aligned} v_0 &= \lambda, \quad v_1 = \mu_1, \quad v_2 = \mu_2, \\ p_{01} &= p_{12} = p_{20} = 1. \end{aligned}$$

6.2 Birth and Death Processes

Definition 6.2.1 Consider a system whose state at any time is represented by the number of people in the system at that time. Suppose that whenever there are n people in the system, then

- (i) new arrivals enter the system at an exponential rate λ_n , and
- (ii) people leave the system at an exponential rate μ_n .

That is, whenever there are n persons in the system, then the time until the next arrival is exponentially distributed with mean $1/\lambda_n$ and is independent of the time until the next departure, which is itself exponentially distributed with mean $1/\mu_n$. Such a system is called a **birth and death process**. The parameters $\{\lambda_n\}_{n=0}^{\infty}$ and $\{\mu_n\}_{n=0}^{\infty}$ are called, respectively, the arrival (or birth) and departure (or death) rates.

Proposition 6.2.2 A birth and death process is a continuous-time Markov chain with states $\{0, 1, \dots\}$ for which transitions from state n may go only to either state $n - 1$ or state $n + 1$. The relationships between the birth and death rates and the state transition rates and probabilities are

$$\begin{aligned} v_0 &= \lambda_0, & \text{since } 0 \rightarrow 1, v_0 = \lambda_0, \\ v_i &= \lambda_i + \mu_i, \quad i > 0, & \text{since } T_{i,i+1} \sim \mathcal{E}(\lambda_i), T_{i,i-1} \sim \mathcal{E}(\mu_i) \Rightarrow \min\{T_{i,i+1}, T_{i,i-1}\} \sim \mathcal{E}(\lambda_i + \mu_i), \\ p_{01} &= 1, & 0 \rightarrow 1 \text{ for sure,} \\ p_{i,i+1} &= \frac{\lambda_i}{\lambda_i + \mu_i}, \quad i > 0, & P(T_{i,i+1} < T_{i,i-1}) = \frac{\lambda_i}{\lambda_i + \mu_i}, \\ p_{i,i-1} &= \frac{\mu_i}{\lambda_i + \mu_i}, \quad i > 0, & P(T_{i,i+1} \geq T_{i,i-1}) = \frac{\mu_i}{\lambda_i + \mu_i}. \end{aligned} \tag{6.1}$$

Example 6.2.3 (The Poisson Process) Consider a birth and death process for which

$$\begin{aligned} \mu_n &= 0, & \text{for all } n \geq 0, \\ \lambda_n &= \lambda, & \text{for all } n \geq 0. \end{aligned}$$

This is a process in which departures never occur, and the time between successive arrivals is exponential with mean $1/\lambda$. The Poisson rate is the same for n . Hence, this is just the Poisson process.

Definition 6.2.4 A birth and death process for which $\mu_n = 0$ for all n is called a **pure birth process**.

Definition 6.2.5 (Yule process, A Birth Process with Linear Birth Rate) If $\{X(t), t \geq 0\}$ is a pure birth process with $\lambda_n = n\lambda$, $n \geq 0$, this pure birth process is known as a Yule process after G. Yule, who used it in his mathematical theory of evolution.

Example 6.2.6 (A Linear Growth Model with Immigration) A model in which

$$\begin{aligned} \mu_n &= n\mu, & n \geq 1 \\ \lambda_n &= n\lambda + \theta, & n \geq 0 \end{aligned}$$

is called a **linear growth process with immigration**. Such processes occur naturally in the study of biological reproduction and population growth. Each individual in the population is assumed to give birth at an exponential rate λ ; in addition, there is an exponential rate of increase θ of the population due to an external source such as immigration. Hence, the total birth rate where there are n persons in the system is $n\lambda + \theta$. Deaths are assumed to occur at an exponential rate μ for each member of the population, so $\mu_n = n\mu$.

Let $X(t)$ denote the population size at time t . Suppose that $X(0) = i$ and let

$$M(t) = E[X(t)].$$

We start by deriving an equation for $M(t + h)$ by conditioning on $X(t)$. This yields

$$\begin{aligned} M(t + h) &= E[X(t + h)] \\ &= E[E[X(t + h)|X(t)]]. \end{aligned}$$

Now, given the size of the population at time t then, ignoring events whose probability is $o(h)$, the population at time $t + h$ will either increase in size by 1 if a birth or an immigration occurs in $(t, t + h)$, or decrease by 1 if a death occurs in this interval, or remain the same if neither of these two possibilities occurs. That is, given $X(t)$, (using definition 2 of Poisson process),

$$X(t + h) = \begin{cases} X(t) + 1, & \text{with prob. } [X(t)\lambda + \theta]h + o(h), \\ X(t) - 1, & \text{with prob. } X(t)\mu h + o(h), \\ X(t), & \text{with prob. } 1 - [\theta + X(t)\lambda + X(t)\mu]h + o(h). \end{cases}$$

Therefore,

$$\begin{aligned} E[X(t+h)|X(t)] &= (X(t)+1)([X(t)\lambda + \theta]h) + (X(t)-1)(X(t)\mu h) \\ &\quad + X(t)[1 - [\theta + X(t)\lambda + X(t)\mu]h] + o(h) \\ &= X(t) + [\theta + X(t)\lambda - X(t)\mu]h + o(h). \end{aligned}$$

Taking expectations yields

$$M(t+h) = M(t) + (\lambda - \mu)M(t)h + \theta h + o(h).$$

Taking the limit as $h \rightarrow 0$ yields the differential equation

$$M'(t) = (\lambda - \mu)M(t) + \theta.$$

If $\lambda \neq \mu$, and assume the initial condition $M(0) = m$,

$$M(t) = \frac{\theta}{\lambda - \mu}[e^{(\lambda-\mu)t} - 1] + me^{(\lambda-\mu)t}. \quad (6.2)$$

If $\lambda = \mu$, and assume the initial condition $M(0) = m$,

$$M(t) = \theta t + m.$$

Example 6.2.7 Consider now a general birth and death process with birth rates $\{\lambda_n\}$ and death rates $\{\mu_n\}$, where $\mu_0 = 0$, and let T_i denote the time, starting from state i , it takes for the process to enter state $i+1$, $i \geq 0$. We will recursively compute $E[T_i]$, $i \geq 0$, by starting with $i = 0$. Since T_0 is exponential with rate λ_0 , we have

$$E[T_0] = \frac{1}{\lambda_0}.$$

For $i > 0$, we condition whether the first transition takes the process into state $i-1$ or $i+1$. That is, let

$$I_i = \begin{cases} 1, & \text{if the first transition from } i \text{ is to } i+1, \\ 0, & \text{if the first transition from } i \text{ is to } i-1, \end{cases}$$

and note that

$$\begin{aligned} E[T_i|I_i = 1] &= \frac{1}{\lambda_i + \mu_i}, \\ E[T_i|I_i = 0] &= \frac{1}{\lambda_i + \mu_i} + E[T_{i-1}] + E[T_i]. \end{aligned}$$

This follows since, independent of whether the first transition is from a birth or death, the time until it occurs is exponential with rate $\lambda_i + \mu_i$; if this first transition is a birth, then the population size is at $i+1$, so no additional time is needed; whereas if it is death, then the population size becomes $i-1$ and the additional time needed to reach $i+1$ is equal to the time it takes to return to state i (this has mean $E[T_{i-1}]$) plus the additional time it then takes to reach $i+1$ (this has mean $E[T_i]$). Hence, since the probability that the first transition is a birth is $\lambda_i/(\lambda_i + \mu_i)$, we see that

$$\begin{aligned} E[T_i] &= E[E[T_i|I_i]] = E[T_i|I_i = 1]P(I_i = 1) + E[T_i|I_i = 0]P(I_i = 0) \\ &= \frac{1}{\lambda_i + \mu_i} \frac{\lambda_i}{\lambda_i + \mu_i} + \left(\frac{1}{\lambda_i + \mu_i} + E[T_{i-1}] + E[T_i] \right) \frac{\mu_i}{\lambda_i + \mu_i} \\ &= \frac{1}{\lambda_i + \mu_i} + \frac{\mu_i}{\lambda_i + \mu_i} (E[T_{i-1}] + E[T_i]). \end{aligned}$$

or, equivalently,

$$E[T_i] = \frac{1}{\lambda_i} + \frac{\mu_i}{\lambda_i} E[T_{i-1}], \quad i \geq 1.$$

Starting with $E[T_0] = \frac{1}{\lambda_0}$, then $E[T_i]$ can be computed recursively.

Example 6.2.8 For the birth and death process having parameters $\lambda_i \equiv \lambda$, $\mu_i \equiv \mu$,

$$E[T_i] = \frac{1}{\lambda} + \frac{\mu}{\lambda} E[T_{i-1}] = \frac{1}{\lambda} (1 + \mu E[T_{i-1}]).$$

Starting with $E[T_0] = 1/\lambda$, we see that

$$\begin{aligned} E[T_1] &= \frac{1}{\lambda} \left(1 + \frac{\mu}{\lambda}\right), \\ E[T_2] &= \frac{1}{\lambda} \left(1 + \frac{\mu}{\lambda} + \left(\frac{\mu}{\lambda}\right)^2\right), \end{aligned}$$

and, in general, for $i \geq 0$,

$$\begin{aligned} E[T_i] &= \frac{1}{\lambda} \left[1 + \frac{\mu}{\lambda} + \left(\frac{\mu}{\lambda}\right)^2 + \cdots + \left(\frac{\mu}{\lambda}\right)^i\right] \\ &= \begin{cases} \frac{1 - (\mu/\lambda)^{i+1}}{\lambda - \mu}, & \lambda \neq \mu, \\ \frac{i+1}{\lambda}, & \lambda = \mu. \end{cases} \end{aligned}$$

6.3 The Transition Probability Function $P_{ij}(t)$

6.3.1 transition probabilities

Definition 6.3.1 Let

$$P_{ij}(t) = P\{X(t+s) = j | X(s) = i\},$$

denote the probability that a process presently in state i will be in state j a time t later. These quantities are often called the **transition probabilities** of the continuous-time Markov chain

We can explicitly determine $P_{ij}(t)$ in the case of a pure birth process having **distinct** birth rates. For such a process, let T_k denote the time the process spends in state k before making a transition into state $k+1$, $k \geq 1$. Suppose that the process is presently in state i , and let $j > i$. Then, as T_i is the time it spends in state i before moving to state $i+1$, and T_{i+1} is the time it then spends in state $i+1$ before moving to state $i+2$, and so on, it follows that $\sum_{k=i}^{j-1} T_k$ is the time it takes until the process enters state j . Now, if the process has not yet entered state j by time t , then its state at time t is smaller than j , and vice versa. That is

$$X(t) < j \Leftrightarrow T_i + \cdots + T_{j-1} > t.$$

Therefore, for $i < j$, we have for a pure birth process that

$$\begin{aligned} P\{X(t) < j | X(0) = i\} &= P\left\{\sum_{k=i}^{j-1} T_k > t\right\} \\ &= \sum_{k=i}^{j-1} e^{-\lambda_k t} \prod_{\substack{r=i \\ r \neq k}}^{j-1} \frac{\lambda_r}{\lambda_r - \lambda_k}. \end{aligned}$$

where we notice that T_i, \dots, T_{j-1} are independent exponential random variables with respective different rates $\lambda_i, \dots, \lambda_{j-1}$ (see page 309 of Ross). In particular,

$$P_{ii}(t) = P(T_i > t) = e^{-\lambda_i t}.$$

We have that

$$P_{ij}(t) = P\{X(t) = j | X(0) = i\} = P\{X(t) < j+1 | X(0) = i\} - P\{X(t) < j | X(0) = i\}.$$

Therefore, for a pure birth process having $\lambda_i \neq \lambda_j$ when $i \neq j$, we have

$$\begin{aligned} P_{ij}(t) &= \sum_{k=i}^j e^{-\lambda_k t} \prod_{\substack{r=i \\ r \neq k}}^j \frac{\lambda_r}{\lambda_r - \lambda_k} - \sum_{k=i}^{j-1} e^{-\lambda_k t} \prod_{\substack{r=i \\ r \neq k}}^{j-1} \frac{\lambda_r}{\lambda_r - \lambda_k}, \\ P_{ii}(t) &= e^{-\lambda_i t}. \end{aligned}$$

Example 6.3.2 Consider the Yule process, in which $\lambda_n = n\lambda$, after tedious computation, one can arrive at

$$P_{ij}(t) = \binom{j-1}{i-1} e^{-i\lambda t} (1 - e^{-\lambda t})^{j-i}, \quad j \geq i \geq 1, \quad (6.3)$$

using above formula.

6.3.2 instantaneous transition rates and Chapman-Kolmogorov equations

We shall now derive a set of differential equations that the transition probabilities $P_{ij}(t)$ satisfy in a general continuous-time Markov chain. However, first we need a definition and a pair of lemmas.

转移转移速率

Definition 6.3.3 For any pair of states i and j , let

$$q_{ij} = v_i p_{ij}.$$

Since v_i is the rate at which the process makes a transition when in state i and p_{ij} is the probability that this transition is into state j , it follows that q_{ij} is the rate, when in state i , at which the process makes a transition into state j . The quantities q_{ij} are called the **instantaneous transition rates**. Since

$$v_i = \sum_j v_i p_{ij} = \sum_j q_{ij},$$

and

$$p_{ij} = \frac{q_{ij}}{v_i} = \frac{q_{ij}}{\sum_j q_{ij}},$$

it follows that specifying the instantaneous transition rates determines the parameters of the continuous-time Markov chain.

Lemma 6.3.4

- (a) $\lim_{h \rightarrow 0} \frac{1 - P_{ii}(h)}{h} = v_i,$
- (b) $\lim_{h \rightarrow 0} \frac{P_{ij}(h)}{h} = q_{ij}, \text{ when } i \neq j.$

p_{ij} 发生转移时的概率 仍然不考虑 $i \rightarrow i$ 停留

Remark 6.3.5

(1) $p_{ij} \neq P_{ij}(t)$, where p_{ij} is the prob. used only when the state changes and $P_{ij}(t) = P(X(t) = j | X(0) = i)$ is the transition prob.

(2) Note that $p_{ii} = 0$ as before since little p_{ij} means that the state must change from i to j .

Lemma 6.3.6 For all $s \geq 0, t \geq 0$,

$$P_{ij}(t+s) = \sum_{k=0}^{\infty} P_{ik}(t)P_{kj}(s). \quad (6.4)$$

Lemma 6.3.7 In order for the process to go from state i to state j in time $t+s$, it must be somewhere at time t and thus

$$\begin{aligned} P_{ij}(t+s) &= P\{X(t+s) = j | X(0) = i\} \\ &= \sum_{k=0}^{\infty} P\{X(t+s) = j, X(t) = k | X(0) = i\} \\ &= \sum_{k=0}^{\infty} P\{X(t+s) = j | X(t) = k, X(0) = i\} \cdot P\{X(t) = k | X(0) = i\} \\ &= \sum_{k=0}^{\infty} P\{X(t+s) = j | X(t) = k\} \cdot P\{X(t) = k | X(0) = i\} \\ &= \sum_{k=0}^{\infty} P_{kj}(s)P_{ik}(t). \end{aligned}$$

The set of Equations (6.4) is known as the **Chapman–Kolmogorov equations** for continuous-time Markov chains.

6.3.3 Kolmogorov's forward equations

A set of equations, known as **Kolmogorov's forward equations** (or **Fokker-Planck equations**), is derived as follows. From the Chapman–Kolmogorov equations (equation 6.4), we have

$$\begin{aligned} P_{ij}(t+h) - P_{ij}(t) &= \sum_{k=0}^{\infty} P_{ik}(t)P_{kj}(h) - P_{ij}(t) = \sum_{k \neq j} P_{ik}(t)P_{kj}(h) + P_{ij}(t)P_{jj}(h) - P_{ij}(t) \\ &= \sum_{k \neq j} P_{ik}(t)P_{kj}(h) - P_{ij}(t)[1 - P_{jj}(h)], \end{aligned}$$

and thus

$$\lim_{h \rightarrow 0} \frac{P_{ij}(t+h) - P_{ij}(t)}{h} = \lim_{h \rightarrow 0} \left\{ \sum_{k \neq j} P_{ik}(t) \frac{P_{kj}(h)}{h} - P_{ij}(t) \left[\frac{1 - P_{jj}(h)}{h} \right] \right\}$$

and, assuming that we can interchange limit with summation, we obtain:

Theorem 6.3.8 Under suitable regularity conditions, we have **Kolmogorov's forward equations**:

$$P'_{ij}(t) = \sum_{k \neq j} P_{ik}(t)q_{kj} - P_{ij}(t)v_j.$$

Unfortunately, we cannot always justify the interchange of limit and summation and thus the preceding is not always valid. However, they do hold in most models, including all birth and death processes and all finite state models.

Proposition 6.3.9 *We shall now solve the forward equations for the pure birth process. For this process, Equation (6.13) reduces to*

$$P'_{ij}(t) = \lambda_{j-1} P_{i,j-1}(t) - \lambda_j P_{ij}(t).$$

However, by noting that $P_{ij}(t) = 0$ whenever $j < i$ (since no deaths can occur), and further noting that

$$\begin{aligned} v_j &= \lambda_j, \\ q_{kj} &= \begin{cases} q_{j-1,j} = v_{j-1} p_{j-1,j} = v_{j-1}, & \text{if } k = j-1, \\ 0, & \text{if } k \neq j-1, \end{cases} \end{aligned}$$

we can rewrite the preceding equation to obtain

$$\begin{cases} P_{ij}(t) = 0, & \text{if } j \leq i-1, \\ P'_{ii}(t) = -\lambda_i P_{ii}(t), & \text{if } j = i, \\ P'_{ij}(t) = \lambda_{j-1} P_{i,j-1}(t) - \lambda_j P_{ij}(t), & \text{if } j \geq i+1. \end{cases}$$

Then for a pure birth process, we have

$$\begin{cases} P_{ij}(t) = 0, & \text{if } j \leq i-1, \\ P_{ii}(t) = e^{-\lambda_i t}, & \text{if } j = i, \\ P_{ij}(t) = \lambda_{j-1} e^{-\lambda_j t} \int_0^t e^{\lambda_j s} P_{i,j-1}(s) ds, & \text{if } j \geq i+1. \end{cases}$$

which can be solved using recursive formula for $j = i+1, i+2, \dots$

Proof. The fact that $P_{ii}(t) = e^{-\lambda_i t}$ follows from above equation by integrating and using the fact that $P_{ii}(0) = 1$. To prove the corresponding result for $P_{ij}(t)$, we note by above equation that

$$\begin{aligned} e^{\lambda_j t} [P'_{ij}(t) + \lambda_j P_{ij}(t)] &= e^{\lambda_j t} \lambda_{j-1} P_{i,j-1}(t), \\ \frac{d}{dt} [e^{\lambda_j t} P_{ij}(t)] &= \lambda_{j-1} e^{\lambda_j t} P_{i,j-1}(t). \end{aligned}$$

Hence, since $P_{ij}(0) = 0$, we obtain the desired results. ■

Example 6.3.10 Consider the Yule process with $\lambda_n = n\lambda$. Then, for $j = i+1$,

$$\begin{aligned} P_{i,i+1}(t) &= \lambda_i e^{-\lambda_{i+1} t} \int_0^t e^{\lambda_{i+1}s} e^{-\lambda_i s} ds = \frac{\lambda_i}{\lambda} e^{-\lambda_{i+1} t} \int_0^t e^{\lambda s} d\lambda s \\ &= ie^{-\lambda_{i+1} t} (e^{\lambda t} - 1) = ie^{-i\lambda t} (1 - e^{-\lambda t}). \end{aligned}$$

For $j = i+2$,

$$\begin{aligned} P_{i,i+2}(t) &= \lambda_{i+1} e^{-\lambda_{i+2} t} \int_0^t e^{\lambda_{i+2}s} ie^{-i\lambda s} (1 - e^{-\lambda s}) ds \\ &= (i+1)i\lambda e^{-(i+2)\lambda t} \int_0^t e^{2\lambda s} (1 - e^{-\lambda s}) ds \\ &= \frac{(i+1)i}{2} e^{-(i+2)\lambda t} (e^{\lambda t} - 1)^2. \end{aligned}$$

One can use induction to show that

$$P_{ij}(t) = \binom{j-1}{i-1} e^{-i\lambda t} (1 - e^{-\lambda t})^{j-i}, \quad j \geq i \geq 1.$$

Assume that the above hold true for $j-1$, then for j ,

$$\begin{aligned} P_{ij}(t) &= \lambda_{j-1} e^{-\lambda_j t} \int_0^t e^{\lambda_j s} P_{i,j-1}(s) ds = (j-1)\lambda e^{-j\lambda t} \int_0^t e^{j\lambda s} C_{j-2}^{i-1} e^{-i\lambda s} (1 - e^{-\lambda s})^{j-i-1} ds \\ &= \frac{(j-1)(j-2)!}{(i-1)!(j-i-1)!} \lambda e^{-j\lambda t} \int_0^t (e^{\lambda s})^{j-i-1} e^{\lambda s} (1 - e^{-\lambda s})^{j-i-1} ds \\ &= \frac{(j-1)!(j-i)}{(i-1)!(j-i)!} \lambda e^{-j\lambda t} \int_0^t (e^{\lambda s} - 1)^{j-i-1} e^{\lambda s} ds \\ &= C_{j-1}^{i-1} (j-i) e^{-j\lambda t} \frac{1}{j-i} (e^{\lambda s} - 1)^{j-i} \Big|_{s=0}^t \\ &= C_{j-1}^{i-1} e^{-j\lambda t} (e^{\lambda t} - 1)^{j-i} = C_{j-1}^{i-1} e^{-i\lambda t} (1 - e^{-\lambda t})^{j-i}. \end{aligned}$$

This results is consistent with (6.3).

Example 6.3.11 For Yule process, first, notice that

$$\sum_{j=i}^{\infty} P_{ij}(t) = \sum_{j=i}^{\infty} P_{i+1,j+1}(t) = 1,$$

where $P_{ij}(t) = P(X(t) = j | X(0) = i) = C_{j-1}^{i-1} e^{-i\lambda t} (1 - e^{-\lambda t})^{j-i}$ for $j \geq i \geq 1$. Then

$$\begin{aligned} M(t) &= EX(t) = \sum_{j \geq i} j P_{ij}(t) = \sum_{j \geq i} j C_{j-1}^{i-1} e^{-i\lambda t} (1 - e^{-\lambda t})^{j-i} \\ &= \sum_{j \geq i} i C_j^i e^{-(i+1)\lambda t} e^{\lambda t} (1 - e^{-\lambda t})^{j-i} = ie^{\lambda t} \sum_{j \geq i} C_j^i e^{-(i+1)\lambda t} (1 - e^{-\lambda t})^{j-i} \\ &= ie^{\lambda t}. \end{aligned}$$

This result is consistent with (6.2).

Example 6.3.12 (Forward Equations for Birth and Death Process) The forward equations (Equation 6.13) for the general birth and death process become

$$\begin{aligned} P'_{i0}(t) &= \sum_{k \neq 0} q_{k0} P_{ik}(t) - \lambda_0 P_{i0}(t) = q_{10} P_{i1}(t) - \lambda_0 P_{i0}(t) \\ &= \mu_1 P_{i1}(t) - \lambda_0 P_{i0}(t). \end{aligned}$$

$$\begin{aligned} P'_{ij}(t) &= \sum_{k \neq j} q_{kj} P_{ik}(t) - (\lambda_j + \mu_j) P_{ij}(t) \\ &= q_{j-1,j} P_{i,j-1}(t) + q_{j+1,j} P_{i,j+1}(t) - P_{ij}(t)(\lambda_j + \mu_j) \\ &= \lambda_{j-1} P_{i,j-1}(t) + \mu_{j+1} P_{i,j+1}(t) - (\lambda_j + \mu_j) P_{ij}(t). \end{aligned}$$

where we notice that

$$\begin{aligned} q_{10} &= v_1 p_{10} = (\lambda_1 + \mu_1) \frac{\mu_1}{\lambda_1 + \mu_1} = \mu_1, \\ q_{j-1,j} &= v_{j-1} p_{j-1,j} = (\lambda_{j-1} + \mu_{j-1}) \frac{\lambda_{j-1}}{(\lambda_{j-1} + \mu_{j-1})}, \\ q_{j+1,j} &= v_{j+1} p_{j+1,j} = (\lambda_{j+1} + \mu_{j+1}) \frac{\mu_{j+1}}{(\lambda_{j+1} + \mu_{j+1})}. \end{aligned}$$

Remark 6.3.13 Fokker-Planck equation is to see the time forward evolution of prob. $P_{ij}(t) = ?$ as a function of time given the initial condition at $t = 0$.

6.3.4 Kolmogorov's backward equations

From CK equation, we obtain

$$\begin{aligned} P_{ij}(h+t) - P_{ij}(t) &= \sum_{k=0}^{\infty} P_{ik}(h)P_{kj}(t) - P_{ij}(t) \\ &= \sum_{k \neq i} P_{ik}(h)P_{kj}(t) - [1 - P_{ii}(h)]P_{ij}(t) \end{aligned}$$

and thus

$$\lim_{h \rightarrow 0} \frac{P_{ij}(h+t) - P_{ij}(t)}{h} = \lim_{h \rightarrow 0} \left\{ \sum_{k \neq i} \frac{P_{ik}(h)}{h} P_{kj}(t) - \left[\frac{1 - P_{ii}(h)}{h} \right] P_{ij}(t) \right\}$$

Now, assuming that we can interchange the limit and the summation in the preceding, we obtain

$$P'_{ij}(t) = \sum_{k \neq i} q_{ik} P_{kj}(t) - v_i P_{ij}(t).$$

It turns out that this interchange can indeed be justified and, hence, we have the following theorem.

Theorem 6.3.14 (Kolmogorov's Backward Equations) For all states i, j , and times $t \geq 0$,

$$P'_{ij}(t) = \sum_{k \neq i} q_{ik} P_{kj}(t) - v_i P_{ij}(t).$$

Example 6.3.15 The backward equations for the pure birth process become

$$P'_{ij}(t) = \lambda_i P_{i+1,j}(t) - \lambda_i P_{ij}(t),$$

where for pure birth process, $v_i = \lambda_i$ and

$$q_{ik} = \begin{cases} q_{i,i+1} = v_i p_{i,i+1} = \lambda_i, & \text{if } k = i + 1, \\ 0, & \text{else } k. \end{cases}$$

Example 6.3.16 The backward equations for the birth and death process become

$$\begin{aligned} P'_{0j}(t) &= \lambda_0 P_{1j}(t) - \lambda_0 P_{0j}(t), \\ P'_{ij}(t) &= (\lambda_i + \mu_i) \left[\frac{\lambda_i}{\lambda_i + \mu_i} P_{i+1,j}(t) + \frac{\mu_i}{\lambda_i + \mu_i} P_{i-1,j}(t) \right] - (\lambda_i + \mu_i) P_{ij}(t), \end{aligned}$$

or equivalently,

$$\begin{aligned} P'_{0j}(t) &= \lambda_0 P_{1j}(t) - \lambda_0 P_{0j}(t), \\ P'_{ij}(t) &= \lambda_i P_{i+1,j}(t) + \mu_i P_{i-1,j}(t) - (\lambda_i + \mu_i) P_{ij}(t), \quad i > 0, \end{aligned}$$

where we notice that

$$\begin{aligned} q_{0k} &= \begin{cases} q_{01} = v_0 p_{01} = \lambda_0, & \text{if } k = 1, \\ 0, & \text{else } k. \end{cases} \\ v_i &= \lambda_i + \mu_i, \\ q_{ik} &= \begin{cases} q_{i,i+1} = v_i p_{i,i+1} = (\lambda_i + \mu_i) \frac{\lambda_i}{\lambda_i + \mu_i}, & \text{if } k = i + 1, \\ q_{i,i-1} = v_i p_{i,i-1} = (\lambda_i + \mu_i) \frac{\mu_i}{\lambda_i + \mu_i}, & \text{if } k = i - 1, \\ 0, & \text{else } k. \end{cases} \end{aligned}$$

Remark 6.3.17 The backward equation solves the problem backward in time, that is from the target time $t = t_0$ to the initial $t = 0$.

6.4 Limiting Probabilities

6.4.1 Limiting Probabilities

Definition 6.4.1 In analogy with a basic result in discrete-time Markov chains, the probability that a continuous-time Markov chain will be in state j at time t often converges to a limiting value that is independent of the initial state. That is, if we call this value P_j , then

$$P_j = \lim_{t \rightarrow \infty} P_{ij}(t),$$

where we are assuming that the limit exists and is independent of the initial state i .

Proposition 6.4.2 To derive a set of equations for the P_j , we let $t \rightarrow \infty$ or $P'_{ij}(t) \rightarrow 0$ in the forward equations. Then

$$P_j v_j = \sum_{k \neq j} P_k q_{kj}, \quad \text{all state } j, \tag{6.5}$$

along with the equation

$$\sum_j P_j = 1, \tag{6.6}$$

can be used to solve for the limiting probabilities.

Remark 6.4.3 We have assumed that the limiting probabilities P_j exist. A sufficient condition for this is that

- (a) all states of the Markov chain communicate in the sense that starting in state i , there is a positive probability of ever being in state j , for all i, j and
- (b) the Markov chain is positive recurrent in the sense that, starting in any state, the mean time to return to that state is finite.

If conditions (a) and (b) hold, then the limiting probabilities will exist and satisfy Equations for limiting prob. In addition, P_j also will have the interpretation of being the long-run proportion of time that the process is in state j .

Remark 6.4.4 Equations (6.5) and (6.6) have a nice interpretation: In any interval $(0, t)$ the number of transitions into state j must equal to within 1 the number of transitions out of state j (why?). Hence, in the long run, the rate at which transitions into state j occur must equal the rate at which transitions out of state j occur. When the process is in state j , it leaves at rate v_j , and, as P_j is the proportion of time it is in state j , it thus follows that

$$v_j P_j = \text{rate at which the process leaves state } j.$$

Similarly, when the process is in state k , it enters j at a rate q_{kj} . Hence, as P_k is the proportion of time in state k , we see that the rate at which transitions from k to j occur is just $q_{kj} P_k$; thus

$$\sum_{k \neq j} q_{kj} P_k = \text{rate at which the process enters state } j$$

So, Equation (6.5) is just a statement of the equality of the rates at which the process enters and leaves state j . Because it balances (that is, equates) these rates, Equation (6.5) is sometimes referred to as a set of “balance equations.”

Remark 6.4.5 When the limiting probabilities P_j exist, we say that the chain is ergodic. The P_j are sometimes called stationary probabilities since it can be shown that (as in the discrete-time case) if the initial state is chosen according to the distribution $\{P_j\}$, then the probability of being in state j at time t is P_j , for all t .

Example 6.4.6 Let us now determine the limiting probabilities for a birth and death process. From Equation (6.5) or equivalently, by equating the rate at which the process leaves a state with the rate at which it enters that state, we obtain from (6.1)

$$\begin{array}{ll} \text{State} & \text{Rate at which leave} = \text{rate at which enter} \\ 0 & \lambda_0 P_0 = \mu_1 P_1, \\ 1 & (\lambda_1 + \mu_1) P_1 = \mu_2 P_2 + \lambda_0 P_0, \\ 2 & (\lambda_2 + \mu_2) P_2 = \mu_3 P_3 + \lambda_1 P_1, \\ n, n \geq 1 & (\lambda_n + \mu_n) P_n = \mu_{n+1} P_{n+1} + \lambda_{n-1} P_{n-1}. \end{array}$$

By adding to each equation the equation preceding it, we obtain

$$\begin{aligned} \lambda_0 P_0 &= \mu_1 P_1, \\ \lambda_1 P_1 &= \mu_2 P_2, \\ \lambda_2 P_2 &= \mu_3 P_3, \\ &\vdots \\ \lambda_n P_n &= \mu_{n+1} P_{n+1}, \quad n \geq 0. \end{aligned}$$

Solving in terms of P_0 yields

$$\begin{aligned} P_1 &= \frac{\lambda_0}{\mu_1} P_0, \\ P_2 &= \frac{\lambda_1}{\mu_2} P_1 = \frac{\lambda_1 \lambda_0}{\mu_2 \mu_1} P_0, \\ &\vdots \\ P_n &= \frac{\lambda_{n-1}}{\mu_n} P_{n-1} = \frac{\lambda_{n-1} \lambda_{n-2} \cdots \lambda_0}{\mu_n \cdots \mu_2 \mu_1} P_0. \end{aligned}$$

And by using the fact that $\sum_{n=0}^{\infty} P_n = 1$, we obtain

$$\begin{aligned} 1 &= P_0 + P_0 \sum_{n=1}^{\infty} \frac{\lambda_{n-1} \lambda_{n-2} \cdots \lambda_0}{\mu_n \cdots \mu_2 \mu_1}, \\ P_0 &= \frac{1}{1 + \sum_{n=1}^{\infty} \frac{\lambda_{n-1} \lambda_{n-2} \cdots \lambda_0}{\mu_n \cdots \mu_2 \mu_1}}, \\ P_n &= \frac{\lambda_{n-1} \lambda_{n-2} \cdots \lambda_0}{\mu_n \cdots \mu_2 \mu_1 \left(1 + \sum_{n=1}^{\infty} \frac{\lambda_{n-1} \lambda_{n-2} \cdots \lambda_0}{\mu_n \cdots \mu_2 \mu_1}\right)}, \quad n \geq 1. \end{aligned}$$

The foregoing equations also show us what condition is necessary for these limiting probabilities to exist. Namely, it is necessary that

$$\sum_{n=1}^{\infty} \frac{\lambda_{n-1} \lambda_{n-2} \cdots \lambda_0}{\mu_n \cdots \mu_2 \mu_1} < \infty.$$

This condition also may be shown to be sufficient.

Example 6.4.7 Let us reconsider the shoe shine shop of Example 6.1.5, and determine the proportion of time the process is in each of the states 0, 1, 2. Because this is not a birth and death process (since the process can go directly from state 2 to state 0), we start with the balance equations for the limiting probabilities.

State	Rate at which leave = rate at which enter
0	$\lambda P_0 = \mu_2 P_2$,
1	$\mu_1 P_1 = \lambda P_0$,
2	$\mu_2 P_2 = \mu_1 P_1$.

Solving in terms of P_0 yields

$$P_2 = \frac{\lambda}{\mu_2} P_0, \quad P_1 = \frac{\lambda}{\mu_1} P_0,$$

which implies, since $P_0 + P_1 + P_2 = 1$, that

$$P_0 \left[1 + \frac{\lambda}{\mu_2} + \frac{\lambda}{\mu_1} \right] = 1.$$

Thus

$$\begin{aligned} P_0 &= \frac{\mu_1 \mu_2}{\mu_1 \mu_2 + \lambda(\mu_1 + \mu_2)}, \\ P_1 &= \frac{\lambda \mu_2}{\mu_1 \mu_2 + \lambda(\mu_1 + \mu_2)}, \\ P_2 &= \frac{\lambda \mu_1}{\mu_1 \mu_2 + \lambda(\mu_1 + \mu_2)}. \end{aligned}$$

6.4.2 Comparison with discrete-time Markov Chain

Consider a continuous-time Markov chain that is ergodic and let us consider the limiting probabilities P_i from a different point of view than previously. If we consider the sequence of states visited, ignoring the amount of time spent in each state during a visit, then this sequence constitutes a discrete-time Markov chain with transition probabilities p_{ij} . Let us assume that this discrete-time Markov chain, called the embedded chain, is ergodic and denote by π_i its limiting probabilities. That is, the π_i are the unique solution of

$$\begin{aligned} \pi_i &= \sum_j \pi_j p_{ji}, \quad \text{all } i, \\ \sum_i \pi_i &= 1. \end{aligned}$$

π_i is i 的次数, P_i 是*i*的时间

Now, since π_i represents the proportion of transitions that take the process into state i , and because $1/v_i$ is the mean time spent in state i during a visit, it seems intuitive that P_i , the proportion of time in state i , should be a weighted average of the π_i where π_i is weighted proportionately to $1/v_i$. That is, it is intuitive that

$$P_i = \frac{\pi_i / v_i}{\sum_j \pi_j / v_j}. \tag{6.7}$$

To check the preceding, recall that the limiting probabilities P_i must satisfy

$$P_i v_i = \sum_{j \neq i} P_j q_{ji}, \quad \text{all } i,$$

or equivalently, since $p_{ii} = 0$

$$v_i P_i = \sum_j P_j v_j p_{ji}, \quad \text{all } i. \quad (6.8)$$

Hence, for the P_i s to be given by Equation (6.7) and (6.8), the following would be necessary:

$$\pi_i = \sum_j \pi_j p_{ji}, \quad \text{all } i.$$

But this, of course, follows since it is in fact the defining equation for the π_i s.

6.5 Time Reversibility

The sequence of states visited by the reversed process constitutes a discrete-time Markov chain with transition probabilities \hat{q}_{ij} given by

$$\hat{q}_{ij} = \frac{\pi_j p_{ji}}{\pi_i}.$$

Hence, we see from the preceding that the reversed process is a continuous time Markov chain with the same transition rates as the forward-time process and with one-stage transition probabilities \hat{q}_{ij} . Therefore, the continuous-time Markov chain will be **time reversible**, in the sense that the process reversed in time has the same probabilistic structure as the original process, if the embedded chain is time reversible. That is, if
→离散

$$\pi_i p_{ij} = \pi_j p_{ji}, \quad \text{for all } i, j.$$

Now, using the fact that $P_i = (\pi_i/v_i)/(\sum_j \pi_j/v_j)$, we see that the preceding condition is equivalent to →连续

$$\begin{aligned} P_i v_i (\sum_k \pi_k/v_k) p_{ij} &= P_j v_j (\sum_k \pi_k/v_k) p_{ji}, \\ P_i q_{ij} &= P_j q_{ji}, \quad \text{for all } i, j. \end{aligned}$$

Since P_i is the proportion of time in state i and q_{ij} is the rate when in state i that the process goes to j , the condition of time reversibility is that **the rate at which the process goes directly from state i to state j is equal to the rate at which it goes directly from j to i** . It should be noted that this is exactly the same condition needed for an ergodic discrete-time Markov chain to be time reversible.

Proposition 6.5.1 *An ergodic birth and death process is time reversible.*

Proof. We must show that the rate at which a birth and death process goes from state i to state $i + 1$ is equal to the rate at which it goes from $i + 1$ to i . In any length of time t the number of transitions from i to $i + 1$ must equal to within 1 the number from $i + 1$ to i (since between each transition from i to $i + 1$ the process must return to i , and this can only occur through $i + 1$, and vice versa). Hence, as the number of such transitions goes to infinity as $t \rightarrow \infty$, it follows that the rate of transitions from i to $i + 1$ equals the rate from $i + 1$ to i .

Another way for the proof is to check directly by

$$P_i q_{ij} = P_j q_{ji}, \quad \text{for all } i, j.$$

Since

$$\begin{aligned} P_n &= \frac{\lambda_{n-1}\lambda_{n-2}\cdots\lambda_0}{\mu_n\cdots\mu_2\mu_1(1+M)}, \\ q_{i,i+1} &= v_i \frac{\lambda_i}{\lambda_i + \mu_i} = \lambda_i, \\ q_{i,i-1} &= v_i \frac{\mu_i}{\lambda_i + \mu_i} = \mu_i, \end{aligned}$$

with $M = \sum_{n=1}^{\infty} \frac{\lambda_{n-1}\lambda_{n-2}\cdots\lambda_0}{\mu_n\cdots\mu_2\mu_1}$. Then we can check

$$P_i q_{i,i+1} = \frac{\lambda_0 \cdots \lambda_{i-1}}{\mu_1 \cdots \mu_i(1+M)} \lambda_i = \frac{\lambda_0 \cdots \lambda_i}{\mu_1 \cdots \mu_{i+1}(1+M)} \mu_{i+1} = P_{i+1} q_{i+1,i}.$$

Similarly, one can check $P_i q_{i,i-1} = P_{i-1} q_{i-1,i}$. ■

Proposition 6.5.2 If for some set $\{P_i\}$

$$\sum_i P_i = 1, \quad P_i \geq 0,$$

and

$$P_i q_{ij} = P_j q_{ji} \quad \text{for all } i \neq j, \tag{6.9}$$

then the continuous-time Markov chain is time reversible and P_i represents the limiting probability of being in state i .

Proof. For fixed i we obtain upon summing Equation (6.9) over all $j : j \neq i$

$$\sum_{j \neq i} P_i q_{ij} = \sum_{j \neq i} P_j q_{ji}$$

or, since $\sum_{j \neq i} q_{ij} = v_i$,

$$v_i P_i = \sum_{j \neq i} P_j q_{ji}.$$

Hence, the P_i s satisfy the balance equations and thus represent the limiting probabilities. Because Equation (6.9) holds, the chain is time reversible. ■

6.6 Computing the Transition Probabilities

For any pair of states i and j , let

$$r_{ij} = \begin{cases} q_{ij}, & \text{if } i \neq j, \\ -v_i, & \text{if } i = j. \end{cases}$$

Using this notation, we can rewrite the Kolmogorov backward equations

$$P'_{ij}(t) = \sum_{k \neq i} q_{ik} P_{kj}(t) - v_i P_{ij}(t),$$

and the forward equations

$$P'_{ij}(t) = \sum_{k \neq j} P_{ik}(t) q_{kj} - P_{ij}(t) v_j,$$

as follows:

$$\begin{aligned} P'_{ij}(t) &= \sum_k r_{ik} P_{kj}(t), \quad (\text{backward}), \\ P'_{ij}(t) &= \sum_k P_{ik}(t) r_{kj}, \quad (\text{forward}). \end{aligned}$$

This representation is especially revealing when we introduce matrix notation. Define the matrices \mathbf{R} and $\mathbf{P}(t)$, $\mathbf{P}'(t)$ by letting the element in row i , column j of these matrices be, respectively, r_{ij} , $P_{ij}(t)$, and $P'_{ij}(t)$. Since the backward equations say that the element in row i , column j of the matrix $\mathbf{P}'(t)$ can be obtained by multiplying the i th row of the matrix \mathbf{R} by the j th column of the matrix $\mathbf{P}(t)$, it is equivalent to the matrix equation

$$\mathbf{P}'(t) = \mathbf{R}\mathbf{P}(t).$$

Similarly, the forward equation can be written as

$$\mathbf{P}'(t) = \mathbf{P}(t)\mathbf{R}.$$

It can be shown that the solution of both of the matrix differential Equations for backward and forward is given by

$$\mathbf{P}(t) = \mathbf{P}(0)e^{\mathbf{R}t}.$$

Since $\mathbf{P}(0) = \mathbf{I}$ (the identity matrix), this yields that

$$\mathbf{P}(t) = e^{\mathbf{R}t},$$

where the matrix $e^{\mathbf{R}t}$ is defined by

$$e^{\mathbf{R}t} = \sum_{n=0}^{\infty} \mathbf{R}^n \frac{t^n}{n!},$$

where \mathbf{R}^n being the matrix multiplication of \mathbf{R} by itself n times.

Chapter 7

Brownian Motion and Stationary Process

7.1 Brownian motion

- Let us start by considering the symmetric random walk. Now suppose that we speed up this process by taking smaller and smaller steps in smaller and smaller time intervals. If we now go to the limit in the right manner what we obtain is Brownian motion.
- More precisely, suppose that each Δt time unit we take a step of size Δx either to the left or the right with equal probabilities. If we let $X(t)$ denote the position at time t then

$$X(t) = \Delta x (X_1 + \dots + X_{[t/\Delta t]}),$$

where

$$X_i = \begin{cases} +1, & \text{if the } i\text{th step is to the right,} \\ -1, & \text{if the } i\text{th step is to the left.} \end{cases}$$

As $EX_i = 0, Var(X_i) = 1$, so that

$$\begin{aligned} EX(t) &= 0, \\ Var(X(t)) &= \Delta x^2 \left[\frac{t}{\Delta t} \right]. \end{aligned}$$

- If we let $\Delta x = O(\Delta t)$ (advection) and $\Delta t \rightarrow 0$, then both mean and variance would converge to 0 (trivial).

- If we let $\Delta x = \sigma\sqrt{\Delta t}$ (diffusion) and $\Delta t \rightarrow 0$, then

$$\begin{aligned} EX(t) &= 0, \\ Var(X(t)) &\rightarrow \sigma^2 t. \end{aligned}$$

独立增量，严平稳

Definition 7.1.1 $\{X(t), t \geq 0\}$ has independent increments, that is, for all $t_1 < t_2 < \dots < t_n$,

$$X(t_n) - X(t_{n-1}), X(t_{n-1}) - X(t_{n-2}), \dots, X(t_2) - X(t_1), X(t_1),$$

are independent.

Definition 7.1.2 $\{X(t), t \geq 0\}$ has stationary increments, that is, the distribution of $X(t+s) - X(t)$ does not depend on t .

Definition 7.1.3 A stochastic process $\{B(t), t \geq 0\}$ is said to be a Brownian motion if

- (i) $B(0) = 0$;
- (ii) $\{B(t), t \geq 0\}$ has stationary and independent increments;
- (iii) for every $t > 0$, $B(t)$ is normally distributed with mean 0 and variance $\sigma^2 t$. Also denoted by a Wiener process $W(t)$.

• In 1827, the English botanist Robert Brown discovered the random-walk motion exhibited by a small particle that is totally immersed in a liquid or gas. In 1905, Einstein developed a theory for the interpretation for collision of water molecules. Later on, Norbert Wiener gave the motion a rigorous mathematical formulation in his 1918 dissertation and in later papers. Since then, the process has been used beneficially in such areas as statistical testing of goodness of fit, analyzing the price levels on the stock market, and quantum mechanics.

标准布朗运动

Remark 7.1.4 When $\sigma = 1$, the process is called standard Brownian motion. Because any Brownian motion can be converted to the standard process by letting $B(t)/\sigma$ we shall, unless otherwise stated, suppose throughout this chapter that $\sigma = 1$.

处处连续, 但处处不可导

Proposition 7.1.5 The interpretation of Brownian motion as the limit of the random walks suggests that $B(t)$ should be a continuous function of t . This turns out to be the case, and it may be proven that, with probability 1, $B(t)$ is indeed a **continuous** function of t but is **not a differentiable function of t almost everywhere**. This fact is quite deep, and no proof shall be attempted.

7.1.1 Joint probability density function

Theorem 7.1.6 The joint density of $B(t_1), \dots, B(t_n)$ is given by

$$\begin{aligned} f(x_1, \dots, x_n) &= f_{t_1}(x_1)f_{t_2-t_1}(x_2 - x_1) \cdots f_{t_n-t_{n-1}}(x_n - x_{n-1}) \\ &= \frac{\exp\left\{-\frac{1}{2}\left[\frac{x_1^2}{t_1} + \frac{(x_2-x_1)^2}{t_2-t_1} + \cdots + \frac{(x_n-x_{n-1})^2}{t_n-t_{n-1}}\right]\right\}}{(2\pi)^{n/2}[t_1(t_2-t_1) \cdots (t_n-t_{n-1})]^{1/2}}. \end{aligned}$$

Proof. As $B(t)$ is normal with mean 0 and variance t , its density function is given by

$$f_t(x) = \frac{1}{\sqrt{2\pi t}} \exp\left\{-\frac{x^2}{2t}\right\}.$$

First note that the set of equalities

$$\begin{aligned} B(t_1) &= x_1, \\ B(t_2) &= x_2, \\ &\vdots \\ B(t_n) &= x_n, \end{aligned}$$

is equivalent to

$$\begin{aligned} B(t_1) - B(t_0) &= 0 = x_1 - (x_0 = 0), \\ B(t_2) - B(t_1) &= x_2 - x_1, \\ &\vdots \\ B(t_n) - B(t_{n-1}) &= x_n - x_{n-1}, \end{aligned}$$

Then, by inde by the independent increment assumption it follows that $B(t_1) - B(t_0)$, $B(t_2) - B(t_1)$, ..., $B(t_n) - B(t_{n-1})$, are independent and, by the stationary increment assumption, that $B(t_k) - X(t_{k-1})$ is normal with mean 0 and variance $t_k - t_{k-1}$. Hence, the joint density is given in the theorem. ■

Corollary 7.1.7 Given $B(t_0) = x_0$, the conditional pdf of $B(t)$ ($t > t_0$) is

$$p(x, t|x_0, t_0) = \frac{1}{\sqrt{2\pi(t-t_0)}} \exp\left\{-\frac{(x-x_0)^2}{2(t-t_0)}\right\} = \frac{1}{\sqrt{2\pi(t-t_0)}} \exp\left\{-\frac{(x-x_0)^2}{2(t-t_0)}\right\}.$$

Thus, we obtain the symmetry of a Brownian motion,

$$P(B(t) > x_0 | B(t_0) = x_0) = P(B(t) \leq x_0 | B(t_0) = x_0) = \frac{1}{2}.$$

One can further check that $p(x, t|x_0, t_0)$ satisfies the diffusion equation

$$\frac{\partial p}{\partial t} = \frac{1}{2} \frac{\partial^2 p}{\partial x^2},$$

with the initial condition $p(x, t=t_0|x_0, t_0) = \delta(x_0)$. This is the Fokker-Planck equation.

7.1.2 Markovian property of a Brownian motion

正向马尔可夫性

- forward Markovian process

$\forall t_1 < t_2 < \dots < t_n$, given $B(t_1), \dots, B(t_{n-1})$, conditional pdf of $B(t_n)$ is the same with the condition pdf of $B(t_n)$ given only $B(t_{n-1})$.

逆向马尔可夫性

- backward Markovian process

$\forall t_1 < t_2 < \dots < t_n$, given $B(t_2), \dots, B(t_n)$, conditional pdf of $B(t_1)$ is the same with the condition pdf of $B(t_1)$ given only $B(t_2)$.

中间关于两边的马尔可夫性

- intermediate Markovian process

$\forall t_1 < t_2 < \dots < t_n$, given $B(t_1), \dots, B(t_{i-1}), B(t_{i+1}), \dots, B(t_n)$, conditional pdf of $B(t_i)$ only depends on $B(t_{i-1})$ and $B(t_{i+1})$.

Theorem 7.1.8 $\forall 0 \leq t_1 < t < t_2$. Let $B(t_1) = a, B(t_2) = b, B(0) = 0$. Then, conditional pdf of $B(t)$ is normal with mean $a + (b-a)(t-t_1)(t_2-t_1)^{-1}$ and variance $(t_2-t)(t-t_1)(t_2-t_1)^{-1}$.

Proof. The joint pdf of $B(t_1), B(t), B(t_2)$ is

$$f(x_1, x, x_2) = \frac{\exp\left\{-\frac{1}{2}\left[\frac{x_1^2}{t_1} + \frac{(x-x_1)^2}{t-t_1} + \frac{(x_2-x)^2}{t_2-t}\right]\right\}}{(2\pi)^{3/2}[t_1(t-t_1)(t_2-t)]^{1/2}},$$

and the joint pdf of $B(t_1), B(t_2)$ is

$$f(x_1, x_2) = \frac{\exp\left\{-\frac{1}{2}\left[\frac{x_1^2}{t_1} + \frac{(x_2-x_1)^2}{t_2-t_1}\right]\right\}}{(2\pi)^{2/2}[t_1(t_2-t_1)]^{1/2}}.$$

Thus,

$$\begin{aligned} f_{B(t)|B(t_1)=a,B(t_2)=b}(x|a, b) &\propto \exp\left\{-\frac{1}{2}\left[\frac{x^2}{t_1} + \frac{(x-x_1)^2}{t-t_1} + \frac{(x_2-x)^2}{t_2-t}\right]\right\} \quad \text{only } x \text{ remains} \\ &\propto \exp\left\{-\frac{1}{2}\left[x^2\left(\frac{1}{t-t_1} + \frac{1}{t_2-t}\right) - x\left(\frac{2x_1}{t-t_1} + \frac{2x_2}{t_2-t}\right)\right]\right\} \\ &= \exp\left\{-\frac{1}{2}\left[\frac{t_2-t_1}{(t-t_1)(t_2-t)}x^2 - \left(\frac{x_1t_2-x_1t+x_2t-x_2t_1}{(t-t_1)(t_2-t)}\right)2x\right]\right\} \\ &\propto \exp\left\{-\frac{1}{2}\left[\frac{t_2-t_1}{(t-t_1)(t_2-t)}\left(x - \frac{x_1t_2-x_1t+x_2t-x_2t_1}{(t_2-t_1)}\right)^2\right]\right\}. \end{aligned}$$

The mean is

$$\begin{aligned} \text{mean} &= \frac{x_1t_2-x_1t+x_2t-x_2t_1-x_1t_1+x_1t_1}{(t_2-t_1)} \\ &= \frac{1}{t_2-t_1}(x_1(t_2-t_1)+(x_2-x_1)t-(x_2-x_1)t_1) \\ &= x_1 + \frac{(x_2-x_1)(t-t_1)}{t_2-t_1} = a + \frac{(b-a)(t-t_1)}{t_2-t_1}, \end{aligned}$$

and the variance is

$$\text{var} = \sigma^2 = \frac{(t-t_1)(t_2-t)}{t_2-t_1}.$$

■

Corollary 7.1.9 If $t_1 = 0, a = 0$, then mean = $\frac{bt}{t_2}$, var = $\frac{t(t_2-t)}{t_2}$.

7.1.3 Brownian motion is a Gaussian process

Definition 7.1.10 A stochastic process $X(t), t \geq 0$ is called a Gaussian, or a normal process if $X(t_1), \dots, X(t_n)$ has a multivariate normal distribution for all t_1, \dots, t_n .

Remark 7.1.11 If $\{B(t), t \geq 0\}$ is a Brownian motion process, then because each of $B(t_1), B(t_2), \dots, B(t_n)$ can be expressed as a linear combination of the independent normal random variables $B(t_1), B(t_2) - B(t_1), B(t_3) - B(t_2), \dots, B(t_n) - B(t_{n-1})$ it follows that Brownian motion is a Gaussian process.

Theorem 7.1.12 $\{B(t), t \geq 0\}$ is a Brownian motion if and only if $\{B(t), t \geq 0\}$ is a Gaussian process with $B(0) = 0, \forall s, t > 0, E(B(t)) = 0, E(B(s)B(t)) = \min\{t, s\}$, and $B(t)$ has a continuous path.

Proof. Only " \Rightarrow " is proved. $E(B(t)) = 0, \forall s \leq t$,

$$\begin{aligned} E(B(s)B(t)) &= E[B(s)(B(t) - B(s) + B(s))] = E[B(s)(B(t) - B(s))] + E[B^2(s)] \\ &= E[B(s)]E[(B(t) - B(s))] + s = s. \end{aligned}$$

The remaining proofs are ignored here. ■

7.2 Martingale

7.2.1 Filtration and filtered probability space

See references in arxiv.org/pdf/0712.0622

<https://www.math.nyu.edu/~varadhan/course/PROB.ch5.pdf>

https://www.math.arizona.edu/~tgk/464_07/cond_exp.pdf

<https://www.cs.yale.edu/homes/aspnes/pinewiki/Martingales.html>

Filtrations have been introduced by Doob and have been a fundamental feature of the theory of stochastic processes. Most basic objects, such as martingales, semimartingales, stopping times or Markov processes involve the notion of filtration.

Definition 7.2.1 Let (Ω, \mathcal{F}, P) be a probability space. A filtration on (Ω, \mathcal{F}, P) is an increasing family $(\mathcal{F}_t)_{t \geq 0}$ of sub- σ -algebras of \mathcal{F} . In other words, for each t , \mathcal{F}_t is a σ -algebra included in \mathcal{F} and if $s \leq t$, $\mathcal{F}_s \subset \mathcal{F}_t$. A probability space (Ω, \mathcal{F}, P) endowed with a filtration $(\mathcal{F}_t)_{t \geq 0}$ is called a filtered probability space, denoted by $(\Omega, \mathcal{F}, (\mathcal{F}_t)_{t \geq 0}, P)$.

7.2.2 Introduction to Martingale

A martingale is a model of a fair game. While discrete-time and continuous-time martingale are defined similarly, we here still define them separately.

discrete-time case

Definition 7.2.2 A stochastic process $X = (X_n, \mathcal{F}_n)$, $n \subseteq \{0, 1, 2, \dots\}$, defined on a filtered probability space $(\Omega, \mathcal{F}, (\mathcal{F}_n)_{n \geq 0}, P)$, that is, a probability space (Ω, \mathcal{F}, P) with a non-decreasing family of σ -algebras $(\mathcal{F}_n)_{n \in \{0, 1, 2, \dots\}}$, $\mathcal{F}_s \subseteq \mathcal{F}_t \subseteq \mathcal{F}$, $s \leq t$, such that $E|X_n| < \infty$, X_n is \mathcal{F}_n -measurable and

$$E(X_t | \mathcal{F}_s) = X_s.$$

Remark 7.2.3 The information that we have at time n is the value of all the variables in \mathcal{F}_n . $\mathcal{F}_s \subseteq \mathcal{F}_t$ for $s \leq t$ means that we do not lose information. A random variable X is said to be \mathcal{F}_n -measurable means that we can determine the value of X if we know the value of all the random variables in \mathcal{F}_n .

Proposition 7.2.4 By induction on k , we can show that for $k \geq 0$, $E[X_{t+k} | \mathcal{F}_t] = X_t$. This can be shown by observing that

$$E[X_{t+k} | \mathcal{F}_t] = E\{E[X_{t+k} | \mathcal{F}_{t+k-1}] \mathcal{F}_t\} = E\{X_{t+k-1} | \mathcal{F}_t\} = \dots = E\{X_{t+1} | \mathcal{F}_t\} = X_t.$$

Proposition 7.2.5 We notice that

$$E(X_t) = E\{E[X_t | \mathcal{F}_0]\} = E(X_0), \quad \forall t = 0, 1, 2, \dots$$

In other words, martingales never go anywhere, at least in expectation.

continuous-time case

Definition 7.2.6 A stochastic process $X = (X_t, \mathcal{F}_t)$, $t \in T \subseteq [0, \infty)$, defined on a filtered probability space $(\Omega, \mathcal{F}, (\mathcal{F}_t)_{t \geq 0}, P)$, that is, a probability space (Ω, \mathcal{F}, P) with a non-decreasing family of σ -algebras $(\mathcal{F}_t)_{t \in T}$, $\mathcal{F}_s \subseteq \mathcal{F}_t \subseteq \mathcal{F}$, $s \leq t$, such that $E|X_t| < \infty$, X_t is \mathcal{F}_t -measurable and

$$E(X_t | \mathcal{F}_s) = X_s.$$

7.2.3 Examples of martingale in Brownian motion

Theorem 7.2.7 $\{B(t), t \geq 0\}$ is a Brownian motion. Notice that $\forall t \geq t_0$

$$E(B(t)|B(t_0) = x_0) = E(B(t) - B(t_0)|B(t_0) = x_0) + B(t_0) = x_0 = B(t_0),$$

so that $B(t)$ is a martingale.

Theorem 7.2.8 $\{B(t), t \geq 0\}$ is a Brownian motion. Then

- (1) $\{B(t)\}$,
- (2) $\{B^2(t) - t\}$,
- (3) $\{e^{\lambda B(t) - \frac{1}{2}\lambda^2 t}\}$, $\forall \lambda \in R$,
- (4) $\{e^{i\lambda B(t) + \frac{1}{2}\lambda^2 t}\}$, $\forall \lambda \in R$

are all martingales.

Proof. We only prove (2) and (3). For (2),

$$\begin{aligned} E(B^2(t) - t | B(t_0) = x_0) &= E(B^2(t) - B^2(t_0) + B^2(t_0) - t | B(t_0) = x_0) \\ &= E[B^2(t) - B^2(t_0) | B(t_0) = x_0] + B^2(t_0) - t \\ &= E[(B(t) - B(t_0))^2 + 2B(t)B(t_0) - 2B^2(t_0) | B(t_0)] + B^2(t_0) - t \\ &= E[(B(t) - B(t_0))^2 | B(t_0)] + 2B(t_0)E[B(t) - B(t_0) | B(t_0)] + B^2(t_0) - t \\ &= t - t_0 + 0 + B^2(t_0) - t = B^2(t_0) - t_0. \end{aligned}$$

For (3), first notice that

$$\begin{aligned} Ee^{\lambda(B(t) - B(t_0))} &= \int e^{\lambda(x - x_0)} \frac{1}{\sqrt{2\pi(t - t_0)}} e^{-\frac{(x - x_0)^2}{2(t - t_0)}} dx \\ &= \frac{1}{\sqrt{2\pi(t - t_0)}} \int e^{-\frac{1}{2(t - t_0)}(x - x_0 - \lambda(t - t_0))^2} dx \cdot e^{\frac{1}{2}\lambda^2(t - t_0)} \\ &= e^{\frac{1}{2}\lambda^2(t - t_0)}. \end{aligned}$$

Then

$$\begin{aligned} E[e^{\lambda B(t) - \frac{1}{2}\lambda^2 t} | B(t_0) = x_0] &= E[e^{\lambda B(t) - \lambda B(t_0) + \lambda B(t_0) - \frac{1}{2}\lambda^2 t} | B(t_0) = x_0] \\ &= e^{\lambda B(t_0) - \frac{1}{2}\lambda^2 t} E[e^{\lambda B(t) - \lambda B(t_0)} | B(t_0) = x_0] \\ &= e^{\lambda B(t_0) - \frac{1}{2}\lambda^2 t} e^{\frac{1}{2}\lambda^2(t - t_0)} = e^{\lambda B(t_0) - \frac{1}{2}\lambda^2 t_0}. \end{aligned}$$

■

7.3 Itó calculus

7.3.1 The pathwise property of Brownian motions

均方收敛

Definition 7.3.1 We say that X_n is mean-square convergent to X if

$$\lim_{n \rightarrow \infty} \|X_n - X\|_{L^2} = \lim_{n \rightarrow \infty} (E(X_n - X)^2)^{1/2} = 0,$$

denoted by

$$\lim_{n \rightarrow \infty} X_n \xrightarrow{\text{m.s.}} X.$$

Lemma 7.3.2 Assume that $B(t)$ is a standard Brownian motion and $B(0) = 0$. Fix t . Let $0 = t_0 < t_1 < \dots < t_n = t$ and $\lambda = \max_{1 \leq k \leq n} (t_k - t_{k-1})$. Then, the following mean-square convergence holds:

$$\lim_{\lambda \rightarrow 0} \sum_{k=1}^n (B(t_k) - B(t_{k-1}))^2 \xrightarrow{\text{m.s.}} t.$$

Proof. We need to prove that

$$\lim_{\lambda \rightarrow 0} E \left[\sum_{k=1}^n (B(t_k) - B(t_{k-1}))^2 - t \right]^2 = 0.$$

Notice that $\Delta B_k := B(t_k) - B(t_{k-1}) \sim N(0, t_k - t_{k-1})$, so that

$$\begin{aligned} E[(\Delta B_k)^2] &= t_k - t_{k-1}, \\ E[(\Delta B_k)^4] &= 3(t_k - t_{k-1})^2, \end{aligned}$$

and for $k \neq l$,

$$E[(\Delta B_k)^2 (\Delta B_l)^2] = E[(\Delta B_k)^2] E[(\Delta B_l)^2] = (t_k - t_{k-1})(t_l - t_{l-1}).$$

Therefore,

$$\begin{aligned} E \left[\sum_{k=1}^n (\Delta B_k)^2 - t \right]^2 &= E \left[\sum_{k=1}^n \Delta B_k^2 \right]^2 - 2t E \sum_{k=1}^n \Delta B_k^2 + t^2 \\ &= \sum_{k=1}^n 3(t_k - t_{k-1})^2 + 2 \sum_{i < j} (t_i - t_{i-1})(t_j - t_{j-1}) - 2t \left[\sum_{k=1}^n (t_k - t_{k-1}) \right] + t^2 \\ &= \sum_{k=1}^n 2(t_k - t_{k-1})^2 + \left[\sum_{k=1}^n (t_k - t_{k-1}) \right]^2 - 2t^2 + t^2 \\ &= \sum_{k=1}^n 2(t_k - t_{k-1})^2. \end{aligned}$$

Since

$$\sum_{k=1}^n (t_k - t_{k-1})^2 \leq \lambda \sum_{k=1}^n (t_k - t_{k-1}) = \lambda t \rightarrow 0, \quad \text{as } \lambda \rightarrow 0 \text{ and } t \text{ fixed,}$$

the proof is finished. ■

Remark 7.3.3 Roughly, we have

$$[dB(t)]^2 \approx dt.$$

Theorem 7.3.4 We have that

$$\lim_{\lambda \rightarrow 0} \sum_{k=1}^n B(t_{k-1})(B(t_k) - B(t_{k-1})) \stackrel{\text{m.s.}}{=} \frac{1}{2} B^2(t) - \frac{1}{2} t, \quad (7.1)$$

$$\lim_{\lambda \rightarrow 0} \sum_{k=1}^n B(t_k)(B(t_k) - B(t_{k-1})) \stackrel{\text{m.s.}}{=} \frac{1}{2} B^2(t) + \frac{1}{2} t. \quad (7.2)$$

Here, the first quantity later on is defined to be the Itô integral $\int_0^t B(s) dB(s)$.

Proof. Eq. (7.1) + Eq. (7.2): we need to prove that

$$\sum_{k=1}^n (B(t_k) + B(t_{k-1})) (B(t_k) - B(t_{k-1})) = \sum_{k=1}^n B^2(t_k) - B^2(t_{k-1}) = B^2(t),$$

which is obviously true. Eq. (7.1) – Eq. (7.2): we need to prove that

$$\lim_{\lambda \rightarrow 0} \sum_{k=1}^n (B(t_k) - B(t_{k-1}))^2 \stackrel{\text{m.s.}}{=} t,$$

which has been proved in above lemma. Done. ■

Remark 7.3.5 Itô integral, named after Kiyosi Itô, does not obey the Fundamental Theorem of Calculus, since for a deterministic function $B(s)$, we have

$$\int_0^t B(s) dB(s) = \frac{1}{2} B^2(t).$$

Theorem 7.3.6 Brownian motion $\{B(t)\}$ is continuous a.e. but not differentiable a.e. for all paths ω . (not proved here.)

Remark 7.3.7 Let us compare the two famous stochastic calculus. For Itô integral, one has

$$\int_0^t Y(s) dX(s) = \lim_{\lambda \rightarrow 0} \sum_{k=1}^n Y(t_{k-1})(X(t_k) - X(t_{k-1})),$$

which is constructed via the forward Euler discretizations. For Stratonovich integral, one has

$$\int_0^t Y(s) \circ dX(s) = \lim_{\lambda \rightarrow 0} \sum_{k=1}^n \frac{Y(t_k) + Y(t_{k-1})}{2} (X(t_k) - X(t_{k-1})),$$

which is constructed via the trapezoidal or the midpoint method.

7.3.2 Stochastic Differential Equations

Definition 7.3.8 $H := \{X : E|X|^2 < \infty\}$, that is, H contains random variables whose second-order moments exist.

均方连续

Definition 7.3.9 Let $\{X(t) \in H : \forall t \in [0, T]\}$ be a stochastic process. If for $t_0 \in [0, T]$, $\lim_{t \rightarrow t_0} X(t) \stackrel{\text{m.s.}}{=} X(t_0)$, i.e.,

$$\lim_{t \rightarrow t_0} \|X(t) - X(t_0)\|_{L^2} = \lim_{t \rightarrow t_0} (E(X(t) - X(t_0))^2)^{1/2} = 0,$$

then $X(t)$ is said mean-square continuous at t_0 .

均方导数

Definition 7.3.10 Let $\{X(t) \in H : \forall t \in [0, T]\}$ be a stochastic process. $X'(t_0)$ is said to be the mean-square derivative of $X(t)$ at $t_0 \in [0, T]$ if

$$\lim_{h \rightarrow 0} \frac{X(t_0 + h) - X(t_0)}{h} \stackrel{\text{m.s.}}{=} X'(t_0).$$

$X(t)$ is said to be mean-square differentiable at $t_0 \in [0, T]$.

Theorem 7.3.11 Let $X(t)$ be mean-square continuous on $[a, b]$.

- (1) $X(t)$ is integrable on $[a, b]$, i.e., $\int_a^b X(t) dt$ exists.
- (2) $\left\| \int_a^b X(t) dt \right\|_{L^2} \leq \int_a^b \|X(t)\|_{L^2} dt$, where $\|X(t)\|_{L^2} = (E(X(t))^2)^{1/2}$ for any $t \in [a, b]$.
- (3) Let $Y(t) := \int_a^t X(u) du$ ($a \leq t \leq b$). Then $\{Y(t) \in H : t \in [a, b]\}$ is mean-square continuous and mean-square differentiable on $[a, b]$. Moreover, $Y'(t) = X(t)$.
- (4) If $X'(t)$ is mean-square continuous on $[a, b]$, then

$$X(t) - X(a) = \int_a^t X'(u) du,$$

for $a \leq t \leq b$.

Definition 7.3.12 Let $\{\mathcal{F}_t, t \geq 0\}$ is a set of monotonically increasing sub- σ field of \mathcal{F} , i.e., $\mathcal{F}_{t_1} \subset \mathcal{F}_{t_2} \subset \mathcal{F}$ for $\forall t_1 < t_2$. Then $\{\mathcal{F}_t, t \geq 0\}$ is called a filtration.

Proposition 7.3.13 $\{B(t), t \geq 0\}$ is a Brownian motion. Then, for $\forall 0 \leq s \leq t$,

- (1) $B(s)$ is measurable w.r.t. \mathcal{F}_t .
- (2) $E(B(t)|\mathcal{F}_s) = B(s)$.
- (3) $E(B(t) - B(s)|\mathcal{F}_s) = 0$.

Definition 7.3.14 The set of stochastic process $\{g(t, \omega), t \geq 0\}$ is denoted by L_T^2 if

- (1) $g(t, \omega)$ is measurable w.r.t. $[0, T] \times \Omega$.
- (2) $\forall t \in [0, T]$, $g(t, \cdot)$ is measurable w.r.t. the filtration \mathcal{F}_t .
- (3) $\int_0^T E[g(t, \omega)]^2 dt < \infty$ and $E[g(t, \omega)]^2 < \infty$ for $\forall t \in [0, T]$.

Definition 7.3.15 $\{B(t), t \geq 0\}$ is a standard Brownian motion. Let $\{g(t, \omega) \in L_T^2, t \geq 0\}$ satisfy the above definition. Let $0 = t_0 < t_1 < \dots < t_n = t$ and $\lambda = \max_{1 \leq k \leq n} (t_k - t_{k-1})$. If

$$\lim_{\lambda \rightarrow 0} \sum_{k=1}^n g(t_{k-1})(B(t_k) - B(t_{k-1})) \stackrel{\text{m.s.}}{=} I_g(t),$$

exists, then

$$I_g(t) := \int_0^t g(s, \omega) dB(s),$$

is said to be the Itô integral of $\{g(t, \omega) \in L_T^2, t \geq 0\}$ w.r.t. $\{B(t), t \geq 0\}$ on $[0, t]$.

Theorem 7.3.16 For $\forall g \in L_T^2$,

- (1) $EI_g(t) = 0$.
- (2) $EI_g^2(t) = \int_0^t Eg^2(s)ds$, and $Cov(I_g(s), I_g(t)) = \int_0^s Eg^2(u)du$ for any $0 \leq s \leq t$.
- (3) $\{I_g(t) = \int_0^t g(s)dB(s), t \geq 0\}$ is martingale w.r.t. \mathcal{F}_t .

Proof. (1) Ignore the limitation of $\lambda \rightarrow 0$, then

$$I_g(t) = \sum_{k=1}^n g(t_{k-1})(B(t_k) - B(t_{k-1})).$$

Thus, denote $g(t_{k-1})$ by g_{k-1} ,

$$\begin{aligned} EI_g(t) &= \sum_{k=1}^n E[g_{k-1}(B(t_k) - B(t_{k-1}))] = \sum_{k=1}^n E\left\{E[g_{k-1}(B(t_k) - B(t_{k-1}))|B(t_{k-1})]\right\} \\ &= \sum_{k=1}^n E\left\{g_{k-1}E[(B(t_k) - B(t_{k-1}))|B(t_{k-1})]\right\} \\ &= \sum_{k=1}^n E\left\{g_{k-1} \cdot 0\right\} = 0. \end{aligned}$$

(2) We only prove the first part.

$$\begin{aligned} EI_g^2(t) &= E\left\{\sum_{k=1}^n [g_{k-1}(B(t_k) - B(t_{k-1}))]\right\}^2 \\ &= E\left\{\sum_{k=1}^n [g_{k-1}^2(B(t_k) - B(t_{k-1}))^2] + 2 \sum_{i < j} [g_{i-1}g_{j-1}(B(t_i) - B(t_{i-1}))(B(t_j) - B(t_{j-1}))]\right\} \\ &= \sum_{k=1}^n E\left\{E[g_{k-1}^2(B(t_k) - B(t_{k-1}))^2]|B(t_{k-1})\right\} \\ &\quad + 2 \sum_{i < j} E\left\{E[g_{i-1}g_{j-1}(B(t_i) - B(t_{i-1}))(B(t_j) - B(t_{j-1}))]|B(t_1) \cdots B(t_{j-1})\right\} \\ &= \sum_{k=1}^n E\left\{g_{k-1}^2 E[(B(t_k) - B(t_{k-1}))^2]|B(t_{k-1})\right\} \\ &\quad + 2 \sum_{i < j} E\left\{g_{i-1}g_{j-1}(B(t_i) - B(t_{i-1})) \underbrace{E[(B(t_j) - B(t_{j-1}))]}_{=0} |B(t_1) \cdots B(t_{j-1})\right\} \\ &= \sum_{k=1}^n E\left\{g_{k-1}^2(t_k - t_{k-1})\right\} \\ &= \int_0^t E(g^2(s))ds. \end{aligned}$$

(3) Let $\{\phi_n(t, \omega)\}$ be a sequence of simple measurable functions mean-square converging to $g(t, \omega)$ in the definition of the Itô integral. If we define

$$M_t^n(\omega) = \int_0^t \phi_n(s, \omega) dB(s, \omega),$$

it follows easily from the Itô isometry that $\|M_t^n - I_g(t)\|_{L^2} \rightarrow 0$ for all $t \in [0, T]$. Therefore, $\{I_g(t), \mathcal{F}_t\}$ is a martingale w.r.t. $\{\mathcal{F}_t\}$ if $\{M_t^n, \mathcal{F}_t\}$ is a martingale. From the definition of ϕ_n , M_t^n is clearly \mathcal{F}_t -measurable

and in $L^2(\Omega, \mathcal{F}, P)$. Assume that $0 \leq s < t \leq T$. Then, by the linearity of the conditional expectation and the Itô integral,

$$\begin{aligned} E(M_t^n | \mathcal{F}_s) &= E(M_s^n | \mathcal{F}_s) + E\left(\int_s^t \phi_n(r) dB(r) | \mathcal{F}_s\right) \\ &= M_s^n + E\left(\int_s^t \phi_n(r) dB(r) | \mathcal{F}_s\right). \end{aligned}$$

It remains to prove that the last term is 0. The integral only consists of terms of the form

$$\xi_k(B_{t_{k+1}} - B_{t_k}),$$

where $s \leq t_k < t_{k+1} \leq t$ and ξ_k is \mathcal{F}_{t_k} -measurable. Since $\mathcal{F}_s \subset \mathcal{F}_{t_k}$, we have

$$\begin{aligned} E\left(\int_s^t \phi_n(r) dB(r) | \mathcal{F}_s\right) &= E\left(\sum_{k=1}^n \xi_k(B_{t_{k+1}} - B_{t_k}) | \mathcal{F}_s\right) \\ &= \sum_{k=1}^n E(\xi_k(B_{t_{k+1}} - B_{t_k}) | \mathcal{F}_s) = \sum_{k=1}^n E[E(\xi_k(B_{t_{k+1}} - B_{t_k}) | \mathcal{F}_{t_k}) | \mathcal{F}_s] \\ &= \sum_{k=1}^n E[\xi_k E((B_{t_{k+1}} - B_{t_k}) | \mathcal{F}_{t_k}) | \mathcal{F}_s] = \sum_{k=1}^n E[\xi_k \cdot 0 | \mathcal{F}_s] = 0. \end{aligned}$$

The proof is complete. ■

7.3.3 Itô Stochastic Process and Itô formula

Definition 7.3.17 $\{X(t), t \geq 0\}$ is said to be a Itô Stochastic Process if $X(t)$ satisfies

$$X(t) - X(t_0) = \int_{t_0}^t b(s, X(s)) ds + \int_{t_0}^t \sigma(s, X(s)) dB(s),$$

or

$$dX(t) = b(t, X(t)) dt + \sigma(t, X(t)) dB(t).$$

Itô's lemma or Itô formula is the **chain rule** for stochastic calculus.

Theorem 7.3.18 (Itô's lemma or Itô formula) Let $\{X(t), t \geq 0\}$ be a Itô stochastic process as above. Let $f(t, x) \in C^{1,2}$ and let $Y(t) := f(t, X(t))$. Then the stochastic process $\{Y(t), t \geq 0\}$ satisfies

$$Y(t) - Y(t_0) = \int_{t_0}^t \left(\frac{\partial f}{\partial t} + b \frac{\partial f}{\partial x} + \frac{\sigma^2}{2} \frac{\partial^2 f}{\partial x^2} \right) (s, X(s)) ds + \int_{t_0}^t \sigma \frac{\partial f}{\partial x} (s, X(s)) dB(s),$$

or

$$dY(t) = \left(\frac{\partial f}{\partial t} + b \frac{\partial f}{\partial x} + \frac{\sigma^2}{2} \frac{\partial^2 f}{\partial x^2} \right) (t, X(t)) dt + \sigma \frac{\partial f}{\partial x} (t, X(t)) dB(t). \quad (\text{a.e.})$$

Proof. The expansion of $f(t, x)$ in a Taylor series is

$$df = \frac{\partial f}{\partial t} dt + \frac{\partial f}{\partial x} dx + \frac{1}{2} \frac{\partial^2 f}{\partial x^2} dx^2 + \dots$$

Substituting $X(t)$ for x and therefore the governing equation gives

$$df = \frac{\partial f}{\partial t} dt + \frac{\partial f}{\partial x} [bdt + \sigma dB(t)] + \frac{1}{2} \frac{\partial^2 f}{\partial x^2} [b^2 dt^2 + 2b\sigma dt dB(t) + \sigma^2 dB^2(t)] + \dots$$

In the limit $dt \rightarrow 0$, the terms dt^2 and $dtdB(t)$ tend to zero faster than $dB^2(t)$ and notice that $dB^2(t) = O(dt)$. Setting the dt^2 and $dtdB(t)$ terms to zero, and substituting dt for $dB^2(t)$ (due to the quadratic variation of a Wiener process), and collecting the dt and $dB(t)$ terms, we obtain

$$df = \left(\frac{\partial f}{\partial t} + b \frac{\partial f}{\partial x} + \frac{\sigma^2}{2} \frac{\partial^2 f}{\partial x^2} \right) dt + \sigma \frac{\partial f}{\partial x} dB(t).$$

7.3.4 History of Kiyosi Itô

Kiyosi Ito



Figure 7.1: Kiyosi Itô.

- See reference on the website <https://mathshistory.st-andrews.ac.uk/Biographies/Ito>. See Fig. 7.1 for Kiyosi Itô.

• Kiyosi Itô was born on 7 September 1915, in Hokusei-cho (now Inabe, Mie Prefecture), Japan. He died on 10 November 2008, in Kyoto, Japan. Kiyosi Itô was a Japanese mathematician who pioneered the theory of stochastic integration and stochastic differential equations. He won the Gauss prize in 2006.

• Kiyosi Itô studied mathematics in the Faculty of Science of the Imperial University of Tokyo. It was during his student years that he became attracted to probability theory. He explains how this came about: *Ever since I was a student, I have been attracted to the fact that statistical laws reside in seemingly random phenomena. Although I knew that probability theory was a means of describing such phenomena, I was not satisfied with contemporary papers or works on probability theory, since they did not clearly define the random variable, the basic element of probability theory. At that time, few mathematicians regarded probability theory as an authentic mathematical field, in the same strict sense that they regarded differential and integral calculus. With clear definition of real numbers formulated at the end of the 19th century, differential and integral calculus had developed into an authentic mathematical system. When I was a student, there were few researchers in probability; among the few were Kolmogorov of Russia, and Paul Levy of France.*

• In 1938 Itô graduated from the University of Tokyo and in the following year he was appointed to the Cabinet Statistics Bureau. He worked there until 1943 and it was during this period that he made his most outstanding contributions: *During those five years I had much free time, thanks to the special consideration given me by the then Director Kawashima ... Accordingly, I was able to continue studying probability theory, by reading Kolmogorov's Basic Concept of Probability Theory and Levy's Theory of Sum of Independent Random Variables. At that time, it was commonly believed that Levy's works were extremely difficult, since Levy,*

a pioneer in the new mathematical field, explained probability theory based on his intuition. I attempted to describe Levy's ideas, using precise logic that Kolmogorov might use. Introducing the concept of regularisation, developed by Doob of the United States, I finally devised stochastic differential equations, after painstaking solitary endeavours. My first paper was thus developed; today, it is common practice for mathematicians to use my method to describe Levy's theory.

- In 1940 he published On the probability distribution on a compact group on which he collaborated with Yukiyosi Kawada. The background to Itô's famous 1942 paper On stochastic processes (Infinitely divisible laws of probability) which he published in the Japanese Journal of Mathematics is given: *Brown, a botanist, discovered the motion of pollen particles in water. At the beginning of the twentieth century, Brownian motion was studied by Einstein, Perrin and other physicists. In 1923, against this scientific background, Wiener defined probability measures in path spaces, and used the concept of Lebesgue integrals to lay the mathematical foundations of stochastic analysis. In 1942, Dr. Itô began to reconstruct from scratch the concept of stochastic integrals, and its associated theory of analysis. He created the theory of stochastic differential equations, which describe motion due to random events.*

- Although today we see this paper as a fundamental one, it was not seen as such by mathematicians at the time it was published. Itô, who still did not have a doctorate at this time, would have to wait several years before the importance of his ideas would be fully appreciated and mathematicians would begin to contribute to developing the theory. In 1943 Itô was appointed as Assistant Professor in the Faculty of Science of Nagoya Imperial University. This was a period of high activity for Itô, and when one considers that this occurred during the years of extreme difficulty in Japan caused by World War II, one has to find this all the more remarkable.

- In 1945 Itô was awarded his doctorate. He continued to develop his ideas on stochastic analysis with many important papers on the topic. Among them were On a stochastic integral equation (1946), On the stochastic integral (1948), Stochastic differential equations in a differentiable manifold (1950), Brownian motions in a Lie group (1950), and On stochastic differential equations (1951).

- In 1952 Itô was appointed to a Professorship at Kyoto University. In the following year he published his famous text Probability theory. In this book, Itô develops the theory on a probability space using terms and tools from measure theory. The years 1954-56 Itô spent at the Institute for Advanced Study at Princeton University. An important publication by Itô in 1957 was Stochastic processes. This book contained five chapters, the first providing an introduction, then the remaining ones studying processes with independent increments, stationary processes, Markov processes, and the theory of diffusion processes. In 1960 Itô visited the Tata Institute in Bombay, India, where he gave a series of lectures surveying his own work and that of other on Markov processes, Levy processes, Brownian motion and linear diffusion.

- Nowadays, Dr. Itô's theory is used in various fields, in addition to mathematics, for analysing phenomena due to random events. Calculation using the "Itô calculus" is common not only to scientists in physics, population genetics, stochastic control theory, and other natural sciences, but also to mathematical finance in economics. In fact, experts in financial affairs refer to Itô calculus as "Itô's formula." Dr. Itô is the father of the modern stochastic analysis that has been systematically developing during the twentieth century. This ceaseless development has been led by many, including Dr. Itô, whose work in this regard is remarkable for its mathematical depth and strong interaction with a wide range of areas. His work deserves special mention as involving one of the basic theories prominent in mathematical sciences during this century.

- A recent monograph entitled Itô's Stochastic Calculus and Probability Theory (1996), dedicated to

Itô on the occasion of his eightieth birthday, contains papers which deal with recent developments of Itô's ideas: *Professor Kiyosi Itô is well known as the creator of the modern theory of stochastic analysis. Although Itô first proposed his theory, now known as Itô's stochastic analysis or Itô's stochastic calculus, about fifty years ago, its value in both pure and applied mathematics is becoming greater and greater. For almost all modern theories at the forefront of probability and related fields, Itô's analysis is indispensable as an essential instrument, and it will remain so in the future. For example, a basic formula, called the Itô formula, is well known and widely used in fields as diverse as physics and economics.*

7.4 Variation on Brownian Motion

7.4.1 Brownian motion with drift

Definition 7.4.1 We say that $\{X(t), t \geq 0\}$ is a Brownian motion process with drift coefficient μ and variance parameter σ^2 if

- (i) $X(0) = 0$;
- (ii) $\{X(t), t \geq 0\}$ has stationary and independent increments;
- (iii) $X(t)$ is normally distributed with mean μt and variance $t\sigma^2$.

Definition 7.4.2 An equivalent definition is to let $\{B(t), t \geq 0\}$ be standard Brownian motion and then define

$$X(t) = \sigma B(t) + \mu t.$$

7.4.2 Geometric Brownian motion

Definition 7.4.3 If $\{Y(t), t \geq 0\}$ is a Brownian motion process with drift coefficient μ and variance parameter σ^2 , i.e., $Y(t) = \sigma B(t) + \mu t$, then the process $\{X(t), t \geq 0\}$ defined by

$$X(t) = e^{Y(t)},$$

is called **geometric Brownian motion**.

- For the geometric Brownian motion $\{X(t)\}$, the mean, second-order moment, variance at time t given the history of the process up to time s can be computed (see details in [10]),

$$\begin{aligned} E[X(t)|X(u), 0 \leq u \leq s] &= X(s)e^{(t-s)(\mu+\sigma^2/2)}, \\ E[X^2(t)|X(u), 0 \leq u \leq s] &= X^2(s)e^{(t-s)(2\mu+2\sigma^2)}, \\ Var[X(t)|X(u), 0 \leq u \leq s] &= X^2(s)e^{(t-s)(2\mu+\sigma^2)}(e^{\sigma^2(t-s)} - 1), \end{aligned}$$

since $Y(t) - Y(s)$ is normal with mean $\mu(t-s)$ and variance $\sigma^2(t-s)$.

- Geometric Brownian motion is a martingale only when $\mu + \sigma^2/2 = 0$.
- In particular, geometric Brownian motion is not a Gaussian process.
- Geometric Brownian motion $\{X(t) = e^{\mu t + \sigma B(t)}, t \geq 0\}$ satisfies the stochastic differential equation

$$dX(t) = (\mu + \frac{\sigma^2}{2})X(t)dt + \sigma X(t)dB(t).$$

- Application in financial mathematics and Black-Scholes PDE model. For instance, Geometric Brownian motion is useful in the modeling of stock prices over time when you feel that the percentage changes are independent and identically distributed. For instance, suppose that X_n is the price of some stock at time n . Then, it might be reasonable to suppose that X_n/X_{n-1} , $n \geq 1$, are independent and identically distributed. Let

$$Z_n = X_n/X_{n-1} \Rightarrow X_n = Z_n X_{n-1}.$$

Iterating this equality gives

$$X_n = Z_n Z_{n-1} \cdots Z_1 X_0.$$

Thus,

$$\log(X_n) = \sum_{i=1}^n \log(Z_i) + \log(X_0),$$

Since $\log(Z_i)$, $i \geq 1$ are independent and identically distributed, $\{\log(X_n)\}$ will, when suitably normalized by central limit theorem, approximately be Brownian motion with a drift, and so $\{X_n\}$ will be approximately geometric Brownian motion.

7.5 Stationary and weakly stationary processes

严平稳过程

Definition 7.5.1 A stochastic process $\{X(t), t \geq 0\}$ is said to be a stationary process if for all s, n, t_1, \dots, t_n the random vectors $X(t_1), \dots, X(t_n)$ and $X(t_1 + s), \dots, X(t_n + s)$ have the same joint distribution. In other words, a process is stationary if, in choosing any fixed point s as the origin, the ensuing process has the same probability law.

Two examples of stationary processes are:

- An ergodic continuous-time Markov chain $\{X(t), t \geq 0\}$ when

$$P\{X(0) = j\} = \pi_j, \quad j \geq 0,$$

where $\{\pi_j, j \geq 0\}$ are the limiting probabilities or the stationary distribution.

- $\{X(t), t \geq 0\}$ when $X(t) = N(t + L) - N(t)$, $t \geq 0$, where $L > 0$ is a fixed constant and $\{N(t), t \geq 0\}$ is a Poisson process having rate λ .

宽平稳过程

Definition 7.5.2 The condition for a process to be stationary is rather stringent and so we define the process $\{X(t), t \geq 0\}$ to be a **second-order stationary** or a **weakly stationary** process if $E[X(t)] = c$ and $Cov[X(t), X(t + s)]$ does not depend on t . That is, a process is second-order stationary if the first two moments of $X(t)$ are the same for all t and the covariance between $X(s)$ and $X(t)$ depends only on $|t - s|$. For a second-order stationary process, define $R(s) = Cov[X(t), X(t + s)]$.

Remark 7.5.3 As the finite dimensional distributions of a Gaussian process (being multivariate normal) are determined by their means and covariance, it follows that a second-order stationary Gaussian process is stationary. See the following example.

Example 7.5.4 (The Ornstein-Uhlenbeck Process) Let $\{X(t), t \geq 0\}$ be a standard Brownian motion process, and define, for $\alpha > 0$,

$$V(t) = e^{-\alpha t/2} B(e^{\alpha t})$$

The process $\{V(t), t \geq 0\}$ is called the Ornstein-Uhlenbeck (OU) process. It has been proposed as a model for describing the velocity of a particle immersed in a liquid or gas, and as such is useful in statistical mechanics. Let us compute its mean and covariance function.

$$\begin{aligned} E(V(t)) &= e^{-\alpha t/2} E[B(e^{\alpha t})] \\ &= e^{-\alpha t/2} \int_{-\infty}^{\infty} x f_{e^{\alpha t}}(x) dx \\ &= 0, \\ Cov[V(t), V(t+s)] &= e^{-\alpha t/2} e^{-\alpha(t+s)/2} Cov[B(e^{\alpha t}), B(e^{\alpha(t+s)})] \\ &= e^{-\alpha t - \alpha s/2} \min\{e^{\alpha t}, e^{\alpha(t+s)}\} \\ &= e^{-\alpha t - \alpha s/2} e^{\alpha t} = e^{-\alpha s/2}, \end{aligned}$$

where $f_t(x) = \frac{1}{\sqrt{2\pi t}} e^{-x^2/2t}$. Hence, $\{V(t), t \geq 0\}$ is weakly stationary and as it is clearly a Gaussian process (since Brownian motion is Gaussian) we can conclude that it is stationary.

For the OU process, another more usually used form is the corresponding SDE,

$$dV(t) = -\frac{\alpha}{2} V(t) dt + \sqrt{\alpha} dB(t).$$

For this SDE in equilibrium or in long time limit, one can show that

$$\begin{aligned} E(V(t)) &= 0, \\ Cov[V(t), V(t+s)] &= \frac{(\sqrt{\alpha})^2}{2(\frac{\alpha}{2})} e^{-\frac{\alpha}{2}s} = e^{-\alpha s/2}. \end{aligned}$$

See equation (1.30) in reference [14].

We now see the equivalence between above two forms of the OU process from another view point. For the analytic solution $V(t) = e^{-\alpha t/2} B(e^{\alpha t})$, we see that

$$d\left(e^{\alpha t/2} V(t)\right) = dB(e^{\alpha t}),$$

so that for the RHS we have

$$EB(e^{\alpha t}) = 0, E[B^2(e^{\alpha t})] = e^{\alpha t}, Cov(B(e^{\alpha t}), B(e^{\alpha(s+t)})) = e^{\alpha t}.$$

For the PDE form, it can be rewritten as

$$d\left(e^{\alpha t/2} V(t)\right) = e^{\alpha t/2} \sqrt{\alpha} dB(t) := g(t) dB(t).$$

We now use the results in Theorem 7.3.16. Denote $I_g(t) = \int_{-\infty}^t g(s) dB(s)$, (the reason for choosing starting from $-\infty$ is that we hope to ignore the effect of the initial condition, that is, we hope to consider the system for sufficient long time to let the system in equilibrium. See also reference [14] for the relation between long time limit and the equilibrium state.) then

$$\begin{aligned} EI_g(t) &= 0, \\ EI_g^2(t) &= \int_{-\infty}^t Eg^2(s) ds = \int_{-\infty}^t \left(e^{\alpha s/2} \sqrt{\alpha}\right)^2 ds = e^{\alpha t}, \\ Cov(I_g(t), I_g(t+s)) &= \int_{-\infty}^t Eg^2(s) ds = e^{\alpha t}, \end{aligned}$$

which is exactly the same as above results. Since both systems are weakly stationary, Gaussian processes and they have the same moments up to second order, they are identically stationary Gaussian processes.

As the following examples show, there are many types of second-order stationary processes that are not stationary.

Example 7.5.5 (An Autoregressive Process) Let Z_0, Z_1, Z_2, \dots be uncorrelated random variables with $E[Z_n] = 0$, $n \geq 0$ and

$$Var(Z_n) = \begin{cases} \sigma^2/(1 - \lambda^2), & n = 0, \\ \sigma^2, & n \geq 1. \end{cases}$$

where $\lambda^2 < 1$. Define

$$\begin{aligned} X_0 &= Z_0, \\ X_n &= \lambda X_{n-1} + Z_n, \quad n \geq 1. \end{aligned}$$

The process $\{X_n, n \geq 0\}$ is called a first-order autoregressive process. It says that the state at time n (that is, X_n) is a constant multiple of the state at time $n - 1$ plus a random error term Z_n . Iterating above equation yields

$$\begin{aligned} X_n &= \lambda(\lambda X_{n-1} + Z_{n-1}) + Z_n \\ &= \sum_{i=0}^n \lambda^{n-i} Z_i, \end{aligned}$$

and so

$$\begin{aligned} Cov(X_n, X_{n+m}) &= Cov\left(\sum_{i=0}^n \lambda^{n-i} Z_i, \sum_{i=0}^{n+m} \lambda^{n+m-i} Z_i\right) \\ &= \sum_{i=0}^n \lambda^{n-i} \lambda^{n+m-i} Cov(Z_i, Z_i) \\ &= \sigma^2 \lambda^{2n+m} \left(\frac{1}{1 - \lambda^2} + \sum_{i=1}^n \lambda^{-2i} \right) \\ &= \sigma^2 \lambda^{2n+m} \left(\frac{1}{1 - \lambda^2} + \lambda^{-2n} \frac{1 - \lambda^{2n}}{1 - \lambda^2} \right) \\ &= \frac{\sigma^2 \lambda^m}{1 - \lambda^2}, \end{aligned}$$

where the preceding uses the fact that Z_i and Z_j are uncorrelated when $i \neq j$. As $E[X_n] = 0$, we see that $\{X_n, n \geq 0\}$ is weakly stationary (the definition for a discrete time process is the obvious analog of that given for continuous time processes). It is easy to see that $\{X_n, n \geq 0\}$ is not necessarily to be stationary by taking Z_0 and other Z_1, Z_2, \dots to be Bernoulli distribution. In detail, if taking $X_0 = Z_0 = \pm \sqrt{\sigma^2/(1 - \lambda^2)}$, $Z_1 = \pm \sigma$ with $1/2$ probability, then the distribution of $X_1 = \lambda Z_0 + Z_1$ is different from that of X_0 .

7.6 Discretization of SDEs

See references, http://www.spec.gmu.edu/~pparis/classes/notes_630/class4_2018.pdf and https://www.seas.upenn.edu/~ese3030/homework/week_11/week_11_white_gaussian_noise.pdf.

Definition 7.6.1 $\xi(t)$ is a standard Gaussian white noise (GWN) if $\xi(t)$ is a Gaussian process such that

$$\langle \xi(t) \rangle = E\xi(t) = 0, \quad \langle \xi(t_1), \xi(t_2) \rangle = \text{Cov}(\xi(t_1), \xi(t_2)) = \delta(t_1 - t_2).$$

Lemma 7.6.2 (Wiener-Khinchin theorem) Let $f(t)$ be a stationary process. Define its inverse Fourier transform by

$$\hat{f}(\omega) := F^{-1}[f(t)](\omega) = \int f(t)e^{i\omega t} dt.$$

Then, the Fourier transform takes the form

$$f(t) = F[\hat{f}(\omega)](t) = \frac{1}{2\pi} \int \hat{f}(\omega)e^{-i\omega t} d\omega.$$

Define the auto-correlation function

$$C(t) = \int f(\tau)^* f(t + \tau) d\tau.$$

Then the Wiener-Khinchin theorem states that

$$C(t) = F[S(\omega)](t) := F[|\hat{f}(\omega)|^2](t),$$

where $S(\omega) := |\hat{f}(\omega)|^2$ is called the power spectral density or power spectrum.

Proof. We have

$$\begin{aligned} C(t) &= \int f(\tau)^* f(t + \tau) d\tau = \frac{1}{(2\pi)^2} \int \hat{f}(\omega_1) e^{-i\omega_1(t+\tau)} \hat{f}^*(\omega_2) e^{i\omega_2\tau} d\omega_1 d\omega_2 d\tau \\ &= \frac{1}{(2\pi)^2} \int \hat{f}(\omega_1) e^{-i\omega_1 t} \hat{f}^*(\omega_2) (2\pi) \delta(\omega_1 - \omega_2) d\omega_1 d\omega_2 \\ &= \frac{1}{2\pi} \int |\hat{f}(\omega)|^2 e^{-i\omega t} d\omega = F\left[|\hat{f}(\omega)|^2\right](t). \end{aligned}$$

■

Remark 7.6.3 The Wiener-Khinchin theorem is a special case of the cross-correlation theorem with $f = g$.

Remark 7.6.4 The spectral contents of a non-stationary process will be time-varying.

Theorem 7.6.5 The standard Gaussian white noise has constant power spectral density, i.e., $S(\omega) \equiv 1$.

Proof. Since

$$\begin{aligned} C(t) &= \langle \xi(t), \xi(0) \rangle = \delta(t) = F[S(\omega)](t) = \frac{1}{2\pi} \int S(\omega) e^{-i\omega t} d\omega, \\ \delta(t) &= \frac{1}{2\pi} \int e^{-i\omega t} d\omega. \end{aligned}$$

Thus $S(\omega) \equiv 1$ is proved. ■

Definition 7.6.6 Let $\xi(t)$ be a standard Gaussian white noise, then the random process

$$B(t) = \int_0^t \xi(s) ds$$

is the standard Brownian motion or standard Wiener process.

Remark 7.6.7 Roughly speaking, one has

$$\frac{dB(t)}{dt} = \xi(t).$$

Definition 7.6.8 Define the discrete time process ΔB_n as the integral of $\xi(t)$ over the interval $[t_n, t_{n+1}] = [nh, (n+1)h]$, i.e.,

$$\Delta B_n = \int_{nh}^{(n+1)h} \xi(s) ds.$$

7.6.1 Euler Maruyama scheme

Consider the SDE

$$dX_t = a(X_t, t)dt + b(X_t, t)dB_t,$$

where $X_t = X(t)$ is the stochastic process and $B_t = B(t)$ is the standard Brownian motion, with initial condition $X_0 = x_0$. Suppose that we wish to solve this SDE on some interval of time $[0, T]$. Then the Euler-Maruyama approximation to the true solution X is the Markov chain Y defined as follows:

- partition the interval $[0, T]$ into N equal subintervals of width $\Delta t > 0$:

$$0 = t_0 < t_1 < \dots < t_N = T \text{ and } \Delta t = T/N;$$

- set $Y_0 = x_0$;
- recursively define Y_n for $0 \leq n \leq N - 1$ by

$$Y_{n+1} = Y_n + a(Y_n, t_n)\Delta t + b(Y_n, t_n)\Delta B_n,$$

where

$$\Delta B_n = B_{t_{n+1}} - B_{t_n} = \varepsilon\sqrt{\Delta t}, \quad \text{with } \varepsilon \sim N(0, 1),$$

that is, the random variables ΔB_n are i.i.d. normal random variables with expected value 0 and variance Δt .

7.7 Langevin equation and the Fluctuation-Dissipation Theorem

- The random motion of a small particle immersed in a fluid is a Brownian motion.
- Observations were made on pollen grains, dust particles, and other objects of colloidal size.
- "Brownian particle" is not a real particle, but instead some collective property of a macroscopic system.
- Newton's equation or Hamiltonian dynamics.
- One-dimensional motion of a spherical particle (Radius a , mass m , position x , velocity v) in a fluid medium (with viscosity η).

$$m \frac{dv}{dt} = F_{total}(t).$$

- The force is dominated by a frictional force $-\gamma v$, where $\gamma = 6\pi\eta a$ is given by Stoke's law:

$$\begin{aligned} m \frac{dv}{dt} &\cong -\gamma v, \\ v(t) &= e^{-\frac{\gamma t}{m}} v(0). \end{aligned}$$

However, $\langle \frac{1}{2}mv^2 \rangle_{eq} = \frac{1}{2}k_B T \Rightarrow F_{total}$ should be modified!

- random Langevin equation for a Brownian particle:

$$m \frac{dv}{dt} = -\gamma v + \sigma \xi(t), \quad (7.3)$$

where $\xi(t)$ is i.i.d. Gaussian white noise with

$$\langle \xi(t) \rangle = 0, \quad \langle \xi(t) \xi(t') \rangle = \delta(t - t').$$

The solution to equation (7.3) is

$$\begin{aligned} \frac{dv}{dt} + \frac{\gamma}{m} v &= \frac{\sigma}{m} \xi(t), \\ e^{\frac{\gamma t}{m}} \left(\frac{dv}{dt} + \frac{\gamma}{m} v \right) &= e^{\frac{\gamma t}{m}} \frac{\sigma}{m} \xi(t), \\ \frac{d}{dt} \left(e^{\frac{\gamma t}{m}} v \right) &= e^{\frac{\gamma t}{m}} \frac{\sigma}{m} \xi(t), \\ e^{\frac{\gamma t}{m}} v(t) - v(0) &= \int_0^t e^{\frac{\gamma t'}{m}} \frac{\sigma}{m} \xi(t') dt', \\ v(t) &= e^{-\frac{\gamma t}{m}} v(0) + \int_0^t e^{-\frac{\gamma(t-t')}{m}} \frac{\sigma}{m} \xi(t') dt'. \end{aligned}$$

- Only 3 terms possibly contribute to the $\langle \frac{1}{2}mv^2 \rangle_{eq}$.

- For the multiplication of two noise terms,

$$\begin{aligned} &\left\langle \int_0^t e^{-\frac{\gamma(t-t')}{m}} \frac{\sigma}{m} \xi(t') dt' \int_0^t e^{-\frac{\gamma(t-t'')}{m}} \frac{\sigma}{m} \xi(t'') dt'' \right\rangle \\ &= \int_0^t dt' \int_0^t dt'' e^{-\frac{\gamma(t-t')}{m}} \frac{\sigma}{m} e^{-\frac{\gamma(t-t'')}{m}} \frac{\sigma}{m} \langle \xi(t') \xi(t'') \rangle \\ &= \int_0^t dt' \int_0^t dt'' e^{-\frac{\gamma(t-t')}{m}} e^{-\frac{\gamma(t-t'')}{m}} \frac{\sigma^2}{m^2} \delta(t'' - t') \\ &= \int_0^t dt' e^{-\frac{2\gamma(t-t')}{m}} \frac{\sigma^2}{m^2} = \int_0^t \frac{\sigma^2}{m^2} e^{-\frac{2\gamma(t-t')}{m}} d \frac{2\gamma t'}{m} \frac{m}{2\gamma} \\ &= \frac{\sigma^2}{2m\gamma} e^{\frac{2\gamma(t-t')}{m}}|_{t'=0} = \frac{\sigma^2}{2m\gamma} \left(1 - e^{-\frac{2\gamma t}{m}} \right). \end{aligned}$$

In the limit of $t \rightarrow \infty$ (that is, for sufficient long time, the system will be in equilibrium),

$$\langle v(t)^2 \rangle = \frac{\sigma^2}{2m\gamma}.$$

- <http://vallance.chem.ox.ac.uk/pdfs/Equipartition.pdf>

The equipartition theorem states that energy is shared equally amongst all energetically accessible degrees of freedom of a system. The equipartition theorem can go further than simply predicting that the available energy will be shared evenly amongst the accessible modes of motion, and can make quantitative predictions about how much energy will appear in each degree of freedom. Specifically, it states that each quadratic degree of freedom will, on average, possess an energy $\frac{1}{2}k_B T$. We can use the Maxwell-Boltzmann distribution of molecular speeds to determine the average kinetic energy of a particle in a gas, and show that it agrees with the equipartition result. The average translational kinetic energy of a particle in a gas is $\frac{3}{2}k_B T$, or $\frac{1}{2}k_B T$ per translational degree of freedom, in agreement with the equipartition theorem. Then, we have $\langle \frac{1}{2}mv^2 \rangle = \frac{1}{2}k_B T$, and thus

$$\langle v(t)^2 \rangle = \frac{\sigma^2}{2m\gamma} = \frac{k_B T}{m}.$$

Therefore, we obtain the **fluctuation-dissipation theorem** or **Einstein relation**:

$$\sigma^2 = 2\gamma k_B T.$$

(普适) 理想气体常数, 阿伏伽德罗常数, 克拉伯龙方程

- <https://www.physik.uni-augsburg.de/theo1/hanggi/History/Renn.pdf>

In Einstein 1905's paper about Brownian motion, σ^2, γ, T are macroscopic quantities, which can be measured. Thus, the Boltzmann constant can be approximated. Furthermore, it was already well-known that

$$k_B = \frac{R}{N_A},$$

where R is the gas constant and N_A is the Avogadro's number. Recall that R appears in the Clausius-Clapeyron equation, $pV = nRT$.

花粉

• In 1926, Perrin studied the Brownian motion using pollen grains and eventually won the Nobel prize. In his paper, he claimed that his experiments are based on Einstein's theoretical prediction ($x \sim \sqrt{t}$).

7.7.1 Significance of fluctuation-dissipation theorem

- (1) The concept of atoms and molecules become universally accepted.
- (2) Deduction of the value of the Avogadro's number.
- (3) The theory applies to diffusion phenomena, including mixing of gases and liquids, atom motion in solids, spread of the black plague, etc.

7.8 Forward and Backward Kolmogorov equations for SDEs

see Kolmogorov_handout_Lecture10_2019%20[Good].pdf

E, W., Li, T., and Vanden-Eijnden, E. (2014). Applied Stochastic Analysis.

Consider the SDE

$$dX_t = b(X_t, t)dt + \sigma(X_t, t)dB_t, \quad X_t, B_t \in \mathbb{R}^d. \quad (7.4)$$

We assume throughout this lecture that b, σ satisfy a global Lipschitz condition and linear growth condition. We work with a vector-valued process here, since it will be no more complicated than a scalar one.

7.8.1 Backward Kolmogorov Equation

Definition 7.8.1 Define the linear operator \mathcal{L} by

$$(\mathcal{L}f)(x, t) := b(x, t) \cdot \nabla f(x, t) + a(x, t) : \nabla^2 f(x, t),$$

with $a(x, t) = \frac{1}{2}\sigma(x, t)\sigma^T(x, t)$ is a tensor.

Theorem 7.8.2 Backward Kolmogorov Equation (time-inhomogeneous).

(i) Let X_t solve SDE (7.4). Let $u(y, s) = E^{y,s}f(X_t) = E[f(X_t)|X_s = y]$, (notice that the conditioning $X_s = y$ here is the independent variable of u), where $f \in C_c^2(\mathbb{R}^d)$. Then

$$\begin{aligned} \partial_s u(y, s) + \mathcal{L}_y u(y, s) &= 0, \quad s < t, \\ u(y, t) &= f(y). \end{aligned}$$

(ii) Let $p(x, t|y, s)$ be the transition density of X_t solving (7.4). Then

$$\begin{aligned}\partial_s p + \mathcal{L}_y p &= 0, \quad s < t, \\ p(x, t|y, t) &= \delta(x - y).\end{aligned}$$

The operator \mathcal{L}_y acts on the y -variables of p .

Remark 7.8.3 Notice that Equations (7.3) and (7.3) are solved **backward** in time, with the given terminal conditions. This is the origin of the name "backward equation."

Proof. (from E et al. (2014)) We first prove (i). Using Itô formula, we can show for any function $u(x, r)$ that

$$\begin{aligned}du(X_r, r) &= (\partial_r u + b \cdot \nabla_{X_r} u + a : \nabla_{X_r}^2 f)(X_r, r)dt + \nabla_{X_r} u \cdot \sigma \cdot dB_r, \\ du(X_r, r) &= (\partial_r u + \mathcal{L}u)(X_r, r)dr + \nabla u \cdot \sigma \cdot dB_r.\end{aligned}$$

Integrate from s to p and take the expectation conditional on $X_s = y$ to obtain

$$E^{y,s} \int_s^p du(X_r, r) = E^{y,s} \int_s^p \{(\partial_r u + \mathcal{L}u)(X_r, r)dr + \nabla u \cdot \sigma \cdot dB_r\}.$$

Notice that by martingale property in Theorem 7.3.16 (proof in (3)), the last term vanishes, that is,

$$E^{y,s} \int_s^p \nabla u \cdot \sigma \cdot dB_r = 0.$$

Then, divide by $p - s$ and take the limit $p \rightarrow s$,

$$\begin{aligned}\lim_{p \rightarrow s} \frac{1}{p-s} E^{y,s} [u(X_p, p) - u(X_s, s)] &= \lim_{p \rightarrow s} \frac{1}{p-s} E^{y,s} \int_s^p (\partial_r u + \mathcal{L}u)(X_r, r)dr, \\ \lim_{p \rightarrow s} \frac{1}{p-s} [E^{y,s} u(X_p, p) - E[u(X_s, s)|X_s = y]] &= E^{y,s} \lim_{p \rightarrow s} \frac{1}{p-s} \int_s^p (\partial_r u + \mathcal{L}u)(X_r, r)dr, \\ \lim_{p \rightarrow s} \frac{1}{p-s} [E^{y,s} u(X_p, p) - u(y, s)] &= E^{y,s} (\partial_s u + \mathcal{L}_{X_s} u) = E[(\partial_s u + \mathcal{L}_{X_s} u)(X_s, s)|X_s = y], \\ &= (\partial_s u + \mathcal{L}_y u)(y, s).\end{aligned}$$

Next we would like to prove that the LHS vanishes. In the definition $u(y, s) = E^{y,s} f(X_t) = E[f(X_t)|X_s = y]$, we take $s = t$ and $y = X_t$ to obtain

$$u(X_t, t) = E[f(X_t)|X_t = X_t] = f(X_t).$$

Take $E^{y,s}$ at both sides to obtain

$$E^{y,s} u(X_t, t) = E^{y,s} f(X_t) = u(y, s),$$

where the last equation follows from the definition of $u(y, s)$. Therefore, the LHS vanishes.

For the initial condition, take $s = t$ in the definition of $u(y, s)$ to obtain

$$u(y, t) = E^{y,t} f(X_t) = E[f(X_t)|X_t = y] = f(y).$$

To show (ii), use the definition $u(y, s) = E[f(X_t)|X_s = y]$ to write $u(y, s) = \int f(x)p(x, t|y, s)dx$ (notice that t is a constant), and compute $\partial_s u$ and $\mathcal{L}_y u$. Substitute into $\partial_s u(y, s) + \mathcal{L}_y u(y, s) = 0$ to obtain

$$\partial_s \int f(x)p(x, t|y, s)dx + \mathcal{L}_y \int f(x)p(x, t|y, s)dx = 0.$$

If $p(\cdot|y, s)$ has continuous bounded derivatives in s and y , then we can move the derivatives inside the integral to obtain

$$\int f(x) [\partial_s p(x, t|y, s) + \mathcal{L}_y p(x, t|y, s)] dx = 0.$$

This holds for all test function f , so $\partial_s p(x, t|y, s) + \mathcal{L}_y p(x, t|y, s) = 0$ as claimed. The terminal condition follows from the terminal condition on u ,

$$u(y, t) = \int f(x)p(x, t|y, t)dx = f(y). \Rightarrow p(x, t|y, t) = \delta(x - y).$$

■ For a time-homogeneous equation, we have an equation that is solved forward in time for the Backward Kolmogorov Equation.

Theorem 7.8.4 Backward Kolmogorov Equation (time-homogeneous). *Let X_t solve a time homogeneous SDE (7.4). Let $u(x, t) = E^x f(X_t) = E[f(X_t)|X_0 = x]$, where $f \in C_c^2(\mathbb{R}^d)$ is bounded with two bounded derivatives. Then*

$$\partial_t u(x, t) = \mathcal{L}_x u(x, t), \quad u(x, 0) = f(x), \quad t > 0.$$

Proof. One can derive this from the equation for a time-inhomogeneous process by a change of variables. Consider the function $v(x, s) = u(x, t - s)$. Then $\partial_s v(x, s) = -\partial_s u(x, t - s)$, and $\mathcal{L}v(x, s) = \mathcal{L}u(x, t - s)$, so v solves

$$\partial_s v(x, s) + \mathcal{L}v(x, s) = 0, \quad v(x, 0) = f(x).$$

which is the desired equation. Furthermore, we have by definition and the time-homogeneity of the process that $v(x, s) = E^{x,t-s} f(X_t) = E^{x,0} f(X_s)$. ■

There is a different proof, which shows the result directly (Varadhan [2007], section 6.3 p.95-96.)

Example 7.8.5 (Brownian motion). *Consider a Brownian motion, which solves the SDE $dX_t = dB_t$. The generator is $\mathcal{L} = \frac{1}{2}\Delta$ so the Backward equation is*

$$\partial_t u = \frac{1}{2}\Delta u, \quad u(x, 0) = f(x).$$

This is a heat equation for u , with initial condition $f(x)$.

7.8.2 Forward Kolmogorov Equation

Now we consider the forward equation.

Definition 7.8.6 *The forward Kolmogorov equation will be formulated in terms of \mathcal{L}^* , the formal adjoint of \mathcal{L} . This is the operator that satisfies*

$$\langle \mathcal{L}f, g \rangle = \langle f, \mathcal{L}^*g \rangle,$$

for all f, g in a suitable function space, and where $\langle f, g \rangle = \int fg dx$ is the L^2 -inner product. Let's work out

the formal expression for \mathcal{L}^* . We have

$$\begin{aligned}
\langle \mathcal{L}f, g \rangle &= \int g(b \cdot \nabla f + a : \nabla^2 f) dx = \int (gb \cdot \nabla f + ga : \nabla^2 f) dx \\
&= \int (\nabla \cdot (fgb) - f \nabla \cdot (gb)) dx + \int (\nabla \cdot (ga \cdot \nabla f) - (\nabla \cdot (ga)) \cdot \nabla f) dx \\
&= \int (\nabla \cdot (fgb) - f \nabla \cdot (gb)) dx + \int (\nabla \cdot (ga \cdot \nabla f) - [\nabla \cdot (f \nabla \cdot (ga)) - f \nabla \cdot \nabla \cdot (ga)]) dx \\
&= \int [-f \nabla \cdot (gb) + f \nabla \cdot \nabla \cdot (ga)] dx = \int f[-\nabla \cdot (bg) + \nabla^2 : (ag)] dx = \int f \mathcal{L}^* g dx = \langle f, \mathcal{L}^* g \rangle.
\end{aligned}$$

In the third line, we used the Divergence theorem and assumed that all the terms vanish at infinity. We are left with

$$\mathcal{L}_x^* g(x, t) = -\nabla \cdot (b(x, t)g(x, t)) + \nabla^2 : (a(x, t)g(x, t)). \quad (7.5)$$

Theorem 7.8.7 Forward Kolmogorov Equation (Fokker-Planck equation).

(i) Let $\rho(x, t)$ the probability density of X_t at time t , and let $\rho_0(x)$ be its initial probability density. Then ρ solves

$$\partial_t \rho = \mathcal{L}_x^* \rho, \quad \rho(x, 0) = \rho_0(x).$$

The operator \mathcal{L}_x^* is defined in (7.5).

(ii) Suppose $\partial_t p(x, t|y, s)$ is continuous in t . Then the transition probability density $p(x, t|y, s)$ is a weak solution to

$$\partial_t p = \mathcal{L}_x^* p, \quad (x, t = s|y, s) = \delta(x - y). \quad (7.6)$$

Remark 7.8.8 The forward equation (sometimes called the Fokker-Planck equation) is the same whether the process is time-homogeneous or not.

Proof. We show (ii) first. The first step of the proof is similar to that in Kolmogorov Backward Equation. For any $f(x) \in C^2$ not depending on time t , we apply the Itô formula:

$$df(X_r) = (\mathcal{L}_{X_r} f)(X_r) dr + \nabla f \cdot \sigma \cdot dB_r.$$

Take the integral and then take the expectation:

$$\begin{aligned}
E^{y,s}[f(X_t) - f(X_s)] &= E^{y,s} \int_s^t (\mathcal{L}_{X_r} f)(X_r) dr. \\
\int f(x)p(x, t|y, s) dx - f(y) &= \int_s^t dr \int \mathcal{L}f(x)p(x, r|y, s) dx.
\end{aligned}$$

Take ∂_t of the equation, and assume that $\partial_t p$ is continuous so we can interchange ∂_t and \int , to find

$$\int f(x)\partial_t p(x, t|y, s) dx = \int \mathcal{L}f(x)p(x, t|y, s) dx.$$

This holds for all suitable test functions f and defines a weak solution to (7.6). Indeed, integrate by parts on the RHS of above equation to get

$$\int f(x)(\partial_t p(x, t|y, s) - \mathcal{L}^* p(x, t|y, s)) dx.$$

This holds for all test functions f , which implies the result in (ii). Notice that we assumed that all the boundary terms vanished when we integrated by parts.

To show (i), notice that $p(x, t|y, s)$ is the fundamental solution, so that write

$$\rho(x, t) = \int p(x, t|y, 0)\rho_0(y)dy.$$

Integrate both sides of the result in (ii) over $\rho_0(y)$ to get the result in (i). (From PDE view point, this is just the convolution formulation for the solution and one can now understand the solution of the PDE problem from the transition probability matrix view point.) ■

Example 7.8.9 (*Brownian motion*). For a Brownian motion solving $dX_t = dB_t$, we have $\mathcal{L} = \frac{1}{2}\Delta$, so $\mathcal{L}^* = \frac{1}{2}\Delta$. The probability density therefore solves a heat equation

$$\partial_t \rho = \frac{1}{2}\Delta \rho, \quad \rho|_{t=0} = \rho_0.$$

In this example, we had $\mathcal{L} = \mathcal{L}^*$ but that almost always won't be the case.

Remark 7.8.10

- Boundary condition is complicated.
- Microscopic and macroscopic views are quite different! For micro view, we have Newton's law and Hamiltonian system, which is time reversible. For macro view we have heat equation and energy dissipation, which is time irreversible.

7.9 Gaussian Processes

Read more references.

Chapter 8

Time Series

Kalman filter

Bibliography

- [1] 林元烈. 应用随机过程. 应用随机过程, 2002.
- [2] 何书元. 随机过程. 北京大学出版社, 2008.
- [3] S. Brooks, A. Gelman, G. Jones, and X.-L. Meng. *Handbook of Markov Chain Monte Carlo*. CRC Press, 2011.
- [4] J. Harlim. *Data-Driven Computational Methods: Parameter and Operator Estimations*. Cambridge University Press, UK, 2018.
- [5] W. M. Jeffrey. Lecture notes on bayesian statistics. *Duke University, Durham, NC*, 2015.
- [6] D. E. Knuth. *The Art of Computer Programming: Sorting and Searching*. Massachusetts: Addison-Wesley, 1973.
- [7] G. F. Lawler. *Introduction to Stochastic Processes, Second Edition*. Introduction to Stochastic Processes, Second Edition, 2006.
- [8] W. H. Press, B. P. Flannery, S. A. Teukolsky, and et al. *Numerical Recipes: The Art of Scientific Computing*. Cambridge: Cambridge University Press, 1986.
- [9] R. Rabiner, L. A tutorial on hidden markov models and selected applications in speech recognition. *Proceedings of the IEEE*, 1989.
- [10] S. M. Ross. Introduction to probability models, ninth edition. In *Academic Press, Inc.*, 2006.
- [11] L. Tierney. Markov chains for exploring posterior distributions. *Ann. Statist.*, 22(4):1701–1728, 1994.
- [12] C. Tijms, Henk. *A First Course in Stochastic Models*. Wiley, 2003.
- [13] P. Zhang and T. Li. 数值分析. 北京大学数学教学系列丛书, 2007.
- [14] R. Zwanzig. *Nonequilibrium Statistical Mechanics*. Oxford Univ Press, Oxford, 2001.