

Lecture 1: Properties of A Random Sample

Econ 205A: Econometric Methods I

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1 Basic Concepts of a Random Sample

Random Sample and Statistic

- The random variables X_1, \dots, X_n are called a **random sample** of size n from the population $f(x)$ if X_1, \dots, X_n are:
 - mutually independent random variables, and
 - the marginal pdf or pmf of each X_i is the same function $f(x)$.
- Sometimes a random sample is denoted as $\{X_i\}_{i=1}^n$.
- They are also called **independent and identically distributed** random variables with pdf or pmf $f(x)$, denoted as **i.i.d.** random variables.

- Let X_1, \dots, X_n be a random sample of size n from a population, and let $T(x_1, \dots, x_n)$ be a real-valued (or vector-valued) function whose domain includes the sample space of (X_1, \dots, X_n) . Then the random variable (or random vector) $Y \equiv T(X_1, \dots, X_n)$ is called a **statistic**.
 - **Example 1** Sample mean $\bar{X} \equiv n^{-1} \sum_{i=1}^n X_i$, sample variance $S^2 \equiv (n-1)^{-1} \sum_{i=1}^n (X_i - \bar{X})^2$ and order statistic $X_{(1)} \equiv \min_{1 \leq i \leq n} X_i$ are all statistics. We often use \bar{X}_n and S_n^2 to emphasize the sample size.
 - Population mean μ , population variance σ^2 and population median are not statistics. They are parameters.
- The probability distribution of a statistic Y is called the **sampling distribution** of Y .
- Statistics are functions of and only of the sample, not of the parameters. However, their sampling distributions generally depend on parameters.
- Why are we interested in sampling distribution? Because they form the basis of statistical inference and hypothesis testing.

2 Sample Mean and Sample Variance

Sample Mean and Sample Variance

- Now let's prove some results that we knew from undergraduate econometrics class (but probably didn't know why).
- **Theorem 1** Let X_1, \dots, X_n be a random sample from a population with mean μ and variance $\sigma^2 < \infty$. Then:
 - (a) $\mathbb{E}(\bar{X}) = \mu$;
 - (b) $\text{var}(\bar{X}) = \sigma^2/n$;
 - (c) $\mathbb{E}(S^2) = \sigma^2$.
- *Proof.* To prove (a),

$$\mathbb{E}(\bar{X}) = \mathbb{E}\left(n^{-1} \sum_{i=1}^n X_i\right) = n^{-1} \sum_{i=1}^n \mathbb{E}(X_i) = \frac{n}{n} \mathbb{E}(X_1) = \mu.$$

To prove (b),

$$\begin{aligned} \text{var}(\bar{X}) &= \text{var} \left(n^{-1} \sum_{i=1}^n X_i \right) = n^{-2} \text{var} \left(\sum_{i=1}^n X_i \right) \\ &= n^{-2} \sum_{i=1}^n \text{var}(X_i) = \frac{n}{n^2} \text{var}(X_1) = \frac{\sigma^2}{n}. \end{aligned}$$

To prove (c),

$$\begin{aligned} \mathbb{E}(S^2) &= \mathbb{E} \left(\frac{1}{n-1} \sum_{i=1}^n (X_i - \bar{X})^2 \right) = \mathbb{E} \left(\frac{1}{n-1} \left[\sum_{i=1}^n X_i^2 - n\bar{X}^2 \right] \right) \\ &= \frac{1}{n-1} (n\mathbb{E}(X_1^2) - n\mathbb{E}(\bar{X}^2)) = \frac{n}{n-1} \left((\mu^2 + \sigma^2) - (\mu^2 + \frac{\sigma^2}{n}) \right) = \sigma^2. \end{aligned}$$

- **Lemma 1 (*Jensen's*) Inequality** If X is a random variable and $\varphi(\cdot)$ is a convex function, then $\varphi(\mathbb{E}(X)) \leq \mathbb{E}(\varphi(X))$. The direction of the inequality is reversed for concave function.
- Using the Jensen's inequality, it is easy to show that $\mathbb{E}(S) \leq \sigma$.

3 Sampling from Normal Distribution

Sampling Distribution of Sample Mean

- Throughout this section, we assume that the random sample $\{X_i\}_{i=1}^n$ is drawn from a $\mathcal{N}(\mu, \sigma^2)$ population, and $\sigma^2 < \infty$.
- **Lemma 2** Let X_1, \dots, X_n be a random sample from a population with mgf $M_X(t)$, then the mgf of the sample mean is

$$M_{\bar{X}}(t) = [M_X(t/n)]^n.$$

- **Theorem 2** \bar{X}_n has a $\mathcal{N}(\mu, \sigma^2/n)$ distribution.

- *Proof.* Recall that the mgf of a $\mathcal{N}(\mu, \sigma^2)$ random variable is $M_X(t) = \exp\left(\mu t + \frac{1}{2}\sigma^2 t^2\right)$, then by Lemma 2, the mgf of the sample mean is

$$\begin{aligned} M_{\bar{X}}(t) &= \left[\exp\left(\mu \frac{t}{n} + \frac{1}{2}\sigma^2 \left(\frac{t}{n}\right)^2\right) \right]^n = \exp\left[n\left(\mu \frac{t}{n} + \frac{1}{2}\sigma^2 \left(\frac{t}{n}\right)^2\right)\right] \\ &= \exp\left(\mu t + \left(\frac{1}{2}\frac{\sigma^2}{n}\right)t^2\right). \end{aligned}$$

Sampling Distribution of Sample Variance

- **Theorem 3** \bar{X} and S^2 are independent random variables.
- *Proof.* Recall that a linear combination of normal random variables is still a normal random variable. Hence both \bar{X} and $X_i - \bar{X}$ ($i = 1, \dots, n$) are normal. Their covariance is

$$\begin{aligned} \text{cov}(\bar{X}, X_i - \bar{X}) &= \text{cov}(\bar{X}, X_i) - \text{cov}(\bar{X}, \bar{X}) \\ &= \text{cov}\left(\frac{X_1 + \dots + X_n}{n}, X_i\right) - \text{var}(\bar{X}) \\ &= 0 + \dots + \frac{\sigma^2}{n} + \dots + 0 - \frac{\sigma^2}{n} \\ &= 0. \end{aligned}$$

Moreover, recall that uncorrelated normal random variables are independent, so we have proven that \bar{X} and $X_i - \bar{X}$ ($i = 1, \dots, n$) are independent. Finally, since S^2 is a function of $X_i - \bar{X}$ ($i = 1, \dots, n$), then \bar{X} and S^2 are independent.

- **Theorem 4** $(n-1)S^2/\sigma^2$ has chi-squared distribution with $n-1$ degrees of freedom.
- *Proof.* Recall that $\bar{X} \sim \mathcal{N}(\mu, \sigma^2/n)$, then $\frac{\bar{X}-\mu}{\sigma/\sqrt{n}} \sim \mathcal{N}(0, 1)$. As a result, $\frac{(\bar{X}-\mu)^2}{\sigma^2/n} \sim \chi_1^2$. Similarly, since $X_i \sim \mathcal{N}(\mu, \sigma^2)$, then $\frac{X_i-\mu}{\sigma} \sim \mathcal{N}(0, 1)$ and

$\frac{\sum_{i=1}^n (X_i - \mu)^2}{\sigma^2} \sim \chi_n^2$. Note that

$$\begin{aligned} \underbrace{\frac{\sum_{i=1}^n (X_i - \mu)^2}{\sigma^2}}_A &= \frac{\sum_{i=1}^n (X_i - \bar{X} + \bar{X} - \mu)^2}{\sigma^2} \\ &= \frac{\sum_{i=1}^n (X_i - \bar{X})^2 + \sum_{i=1}^n (\bar{X} - \mu)^2 + 2(\bar{X} - \mu) \sum_{i=1}^n (X_i - \bar{X})}{\sigma^2} \\ &= \frac{\sum_{i=1}^n (X_i - \bar{X})^2 + \sum_{i=1}^n (\bar{X} - \mu)^2}{\sigma^2} \\ &= \underbrace{\frac{\sum_{i=1}^n (X_i - \bar{X})^2}{\sigma^2}}_B + \underbrace{\frac{(\bar{X} - \mu)^2}{\sigma^2/n}}_C. \end{aligned}$$

We have shown above that $A \sim \chi_n^2$ and $C \sim \chi_1^2$, so their mgf are $M_A(t) = (1 - 2t)^{-n/2}$ and $M_C(t) = (1 - 2t)^{-1/2}$. Moreover, note that B is a function of $X_i - \bar{X}$ ($i = 1, \dots, n$) and C is a function of \bar{X} . In the proof of Theorem 3 we showed that \bar{X} and $X_i - \bar{X}$ are independent, so B and C are independent. This implies that

$$M_B(t) = M_A(t)/M_C(t) = (1 - 2t)^{-(n-1)/2},$$

which is the mgf of a chi-squared distribution with df $n - 1$. Note that $B = (n - 1)S^2/\sigma^2$ and this completes the proof.

Sampling Distribution of t and F Statistics

- The sampling distributions of sample mean \bar{X} and the sample variance S^2 enable us to prove some familiar results from undergraduate econometrics class.
- **Definition 1** If $W \sim \mathcal{N}(0, 1)$, $U \sim \chi_r^2$, and W and U are independent, then

$$T \equiv \frac{W}{\sqrt{U/r}}$$

has a **(Student) t distribution** with r degrees of freedom.

- **Definition 2** If $U \sim \chi_{r_1}^2$ and $V \sim \chi_{r_2}^2$ and they are independent, then

$$F \equiv \frac{U/r_1}{V/r_2}$$

has an **F distribution** with (r_1, r_2) degrees of freedom .

- Proof of the following theorems are left as exercises.
- **Theorem 5 (*t statistic*)** *The random variable*

$$T \equiv \frac{\bar{X} - \mu}{S/\sqrt{n}}$$

has a Student t distribution with $n - 1$ degrees of freedom.

- **Theorem 6 (*F statistic*)** *We have two random samples X_1, \dots, X_n and Y_1, \dots, Y_m . Suppose their population variances are σ_X^2 and σ_Y^2 , respectively. And let S_n^2 and S_m^2 be their respective sample variances. Then the random variable*

$$F \equiv \frac{\sigma_Y^2 S_n^2}{\sigma_X^2 S_m^2}$$

has an F distribution with $(n - 1, m - 1)$ degrees of freedom.

4 Sampling Distributions Without Normality

Why Do We Need Asymptotic Distributions

- In the previous section, we learned the sampling distributions of certain statistics under the assumption that the random sample is drawn from normally distributed population.
- Those results hold regardless of the sample size n . They are called **finite sample sampling distributions**.
- If the population does not have normal distribution, then in general the statistics computed from finite samples do not have simple known distributions.
- In such general cases, we need to invoke **asymptotic (sampling) distributions**, which are the *limits* of (unknown) finite sample sampling distributions of the statistics as the sample size n approaches infinity.
- Asymptotic distributions are approximations of finite sample distributions.

- In order to study the asymptotic distributions, we need to take a not so short digression on different convergence concepts.
- Asymptotic properties (mean, variance, sampling distribution, etc.) are often referred to as **large sample properties** as well.

Notation and Definitions

- Let $\{X_n\}$ for $n = 1, 2, 3, \dots$ denote a sequence of random variables on a common probability space $\{S, \mathcal{F}, P\}$ and let X denote another random variable defined on the same probability space.
- Let F_n denote the distribution function of X_n and let F denote the distribution function X .
- **Definition 3** Let $\{a_n\}$ be a sequence of non-stochastic real numbers. If there exists a real number a such that for any $\epsilon > 0$, there exists a finite integer $N(\epsilon)$ such that for all $n \geq N(\epsilon)$, we have $|a_n - a| < \epsilon$, then a is called the **limit** of the sequence $\{a_n\}$. We write $a_n \rightarrow a$ as $n \rightarrow \infty$, or $\lim_{n \rightarrow \infty} a_n = a$.
- **Definition 4** A function $g(\cdot)$ is **continuous at** a if and only if for any $\epsilon > 0$, there exists a positive real number $\delta(\epsilon)$ such that $|a_n - a| < \delta(\epsilon)$ implies $|g(a_n) - g(a)| < \epsilon$.
- Equivalently, the function $g(\cdot)$ is continuous at a if and only if for any sequence $\{a_n\}$, $a_n \rightarrow a$ as $n \rightarrow \infty$ implies $g(a_n) \rightarrow g(a)$ as $n \rightarrow \infty$.
- **Definition 5** A function $g(\cdot)$ is called a **continuous function** if it is continuous at any value in its domain.

Convergence in Probability

- **Definition 6** A sequence of random variable X_1, X_2, \dots **converges in probability** to a random variable X if, for any $\epsilon > 0$,

$$\lim_{n \rightarrow \infty} P(|X_n - X| \geq \epsilon) = 0, \text{ or equivalently, } \lim_{n \rightarrow \infty} P(|X_n - X| < \epsilon) = 1.$$

This is denoted as $X_n \xrightarrow{P} X$, or $\text{plim}_{n \rightarrow \infty} X_n = X$.

- If X is a degenerated random variable, i.e. it is a constant (we denote it by μ), then we say that the sequence of random variables X_1, X_2, \dots converges in probability to a constant μ . This is denoted as $X_n \xrightarrow{p} \mu$ or $\text{plim}_{n \rightarrow \infty} X_n = \mu$.
- **Lemma 3 (Chebyshev's inequality)** Let X be a random variable with finite mean μ and finite variance σ^2 . Then for any $\epsilon > 0$, we have

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

- **Theorem 7 (WLLN)** Let X_1, X_2, \dots be i.i.d. random variables with mean μ and variance $\sigma^2 < \infty$. Then for any $\epsilon > 0$,

$$\lim_{n \rightarrow \infty} P(|\bar{X}_n - \mu| \geq \epsilon) = 0;$$

that is, \bar{X}_n converges in probability to μ .

- *Proof.* The proof is a straightforward application of Chebyshev's inequality. Recall that $\mathbb{E}(\bar{X}_n) = \mu$ and $\text{var}(\bar{X}_n) = \sigma^2/n$. We have, for any $\epsilon > 0$,

$$P(|\bar{X}_n - \mu| \geq \epsilon) \leq \frac{\sigma^2}{n\epsilon^2} \rightarrow 0 \text{ as } n \rightarrow \infty.$$

Since probability is always non-negative, we have $\lim_{n \rightarrow \infty} P(|\bar{X}_n - \mu| \geq \epsilon) = 0$.

- Weak Law of Large Numbers (WLLN) states that under quite mild conditions, the sample mean approaches the population mean as $n \rightarrow \infty$.
- There is a more general version of WLLN, where we only need to assume that μ is finite.
- **Definition 7** Let X_1, \dots, X_n be a random sample from a population $f(x; \theta)$, where θ is the parameter of the pdf (or pmf). Let T_n be a statistic computed using the sample of size n . T_n is called a **consistent estimator** of θ if $T_n \xrightarrow{p} \theta$.
- WLLN states that under certain conditions, sample mean is a consistent estimator of population mean.

- In the above, we used Chebyshev's inequality to prove WLLN. This argument suggests a general approach of showing that a statistic T_n is a consistent estimator of parameter θ . We only need three steps:

- (i) Show that $\mathbb{E}(T_n) = \theta$;
- (ii) Show that $\text{var}(T_n) \rightarrow 0$ as $n \rightarrow \infty$;
- (iii) Invoke Chebyshev's inequality to show that for any $\epsilon > 0$,

$$P(|T_n - \theta| \geq \epsilon) \leq \frac{\text{var}(T_n)}{\epsilon^2} \rightarrow 0 \text{ as } n \rightarrow \infty.$$

- **Example 2 (Consistency of S_n^2)** Now we follow this approach to show the consistency of sample variance. Suppose X_1, X_2, \dots is a sequence of i.i.d. random variable with mean μ and variance $\sigma^2 < \infty$. By Theorem 1 (c), we have $\mathbb{E}(S_n^2) = \sigma^2$. Then as long as $\text{var}(S_n^2) \rightarrow 0$ as $n \rightarrow \infty$, by Chebyshev's inequality we have

$$P(|S_n^2 - \sigma^2| \geq \epsilon) \leq \frac{\text{var}(S_n^2)}{\epsilon^2} \rightarrow 0 \text{ as } n \rightarrow \infty.$$

- **Theorem 8** Suppose that X_1, X_2, \dots converges in probability to a random variable X and $g(\cdot)$ is a continuous function. Then $g(X_1), g(X_2), \dots$ converges in probability to $g(X)$.
- This theorem is a special case of the **continuous mapping theorem**.
- **Example 3** Suppose $X_n \xrightarrow{p} X$ and a is a constant. Then $aX_n \xrightarrow{p} aX$.
- **Example 4** Suppose $X_n \xrightarrow{p} c$ (c is a constant) and the function $g(\cdot)$ is continuous at c . Then $g(X_n) \xrightarrow{p} g(c)$.
- **Theorem 9** Suppose $X_n \xrightarrow{p} X$ and $Y_n \xrightarrow{p} Y$. Then $X_n + Y_n \xrightarrow{p} X + Y$.
- **Theorem 10** Suppose $X_n \xrightarrow{p} X$ and $Y_n \xrightarrow{p} Y$. Then $X_n Y_n \xrightarrow{p} XY$.

- **Example 5** If X_1, X_2, \dots are i.i.d. uniform $(0, \theta)$ where θ is unknown, and let $X_{(n)} \equiv \max_{1 \leq i \leq n} X_i$. Then $X_{(n)}$ is consistent for θ . To see this,

$$\begin{aligned}
P(|X_{(n)} - \theta| \geq \epsilon) &= P(X_{(n)} \geq \theta + \epsilon) + P(X_{(n)} \leq \theta - \epsilon) \\
&= 0 + P(X_{(n)} \leq \theta - \epsilon) \\
&= P\left(\max_{1 \leq i \leq n} X_i \leq \theta - \epsilon\right) \\
&= P(X_1 \leq \theta - \epsilon, \dots, X_n \leq \theta - \epsilon) \\
&= \prod_{i=1}^n P(X_i \leq \theta - \epsilon) \\
&= \left(1 - \frac{\epsilon}{\theta}\right)^n \\
&\rightarrow 0 \text{ as } n \rightarrow \infty \text{ for any } 0 < \epsilon \leq \theta.
\end{aligned}$$

And it equals zero for $\epsilon > \theta$.

Almost Sure Convergence

- **Definition 8** A sequence of random variables X_1, X_2, \dots **converges almost surely** to a random variable X if, for any $\epsilon > 0$,

$$P\left(\lim_{n \rightarrow \infty} |X_n - X| < \epsilon\right) = 1, \text{ or equivalently, } P\left(\lim_{n \rightarrow \infty} X_n = X\right) = 1.$$

This is denoted as $X_n \xrightarrow{a.s.} X$.

- Almost sure convergence is also known as **convergence with probability 1**.
- Recall that a random variable is a function on the sample space, i.e., $X_i: S \rightarrow \mathbb{R}$. For random variables X_1, X_2, \dots , every point $s \in S$ generate a particular sequence of real numbers x_1, x_2, \dots . Almost sure convergence is similar to pointwise convergence of a sequence of functions, with one exception: the convergence is allowed to fail on a zero probability subset of S .
- Almost sure convergence is stronger than convergence in probability, in the sense that $X_n \xrightarrow{a.s.} X$ implies $X_n \xrightarrow{p.} X$, but not vice versa. However, if a sequence converges in probability, it is possible to find a subsequence that converges almost surely.

- Examples of sequence that \xrightarrow{p} but not $\xrightarrow{a.s.}$ will be given in TA sessions.
- **Theorem 11 (SLLN)** Let X_1, X_2, \dots be i.i.d. random variables with mean μ and variance $\sigma^2 < \infty$. Then for any $\epsilon > 0$,

$$P\left(\lim_{n \rightarrow \infty} |\bar{X} - \mu| < \epsilon\right) = 1;$$

that is, \bar{X} converges almost surely to μ .

- Continuous mapping theorem also applies to almost sure convergence.

Convergence in Distribution

- **Definition 9** A sequence of random variables X_1, X_2, \dots **converges in distribution** to a random variable X if

$$\lim_{n \rightarrow \infty} F_{X_n}(x) = F_X(x)$$

at all points x where $F_X(x)$ is continuous. This is denoted as $X_n \xrightarrow{d} X$.

- **Example 6** (Continuation of Example 5) Previously we have shown that for any $0 < \epsilon \leq \theta$,

$$P(X_{(n)} \leq \theta - \epsilon) = \left(1 - \frac{\epsilon}{\theta}\right)^n.$$

If we take $\epsilon = t\theta/n$ (note that for fixed t , $t\theta/n < \theta$ for large enough n), then we have

$$P\left(X_{(n)} \leq \theta \left(1 - \frac{t}{n}\right)\right) = \left(1 - \frac{t}{n}\right)^n \rightarrow e^{-t},$$

which, upon rearranging, yields

$$P\left(n \left(1 - \frac{X_{(n)}}{\theta}\right) \leq t\right) \rightarrow 1 - e^{-t};$$

that is, the random variable $n(1 - X_{(n)}/\theta)$ converges in distribution to an exponential(1) random variable.

- Although we talk of a sequence of random variables converging in distribution, it is really the *cdfs* that converge, not the random variables. In this fundamental way convergence in distribution is quite different from convergence in probability or almost sure convergence.
- **Theorem 12** *If the sequence of random variables X_1, X_2, \dots converges in probability to a random variable X , the sequence also converges in distribution to X .*
- **Theorem 13** *The sequence of random variables X_1, X_2, \dots converges in probability to a constant c if and only if the sequence also converges in distribution to c ; that is, the statement*

$$P(|X_n - c| > \epsilon) \rightarrow 0 \text{ for any } \epsilon > 0$$

is equivalent to

$$P(X_n \leq x) \rightarrow \begin{cases} 0 & \text{if } x < c \\ 1 & \text{if } x \geq c \end{cases}$$

- The asymptotic distribution of sample mean is summarized in the most important theorems in statistics, the Central Limit Theorems (CLT). The following is the simplest version of CLT (there are many others).
- **Theorem 14** *Let X_1, X_2, \dots , be a sequence of i.i.d. random variables whose mgfs exist in a neighborhood of 0 (that is, $M_{X_i}(t)$ exists for $|t| < h$ for some positive h). Let $\mathbb{E}(X_i) = \mu$ and $\text{var}(X_i) = \sigma^2 > 0$. (Both μ and σ^2 are finite since the mgf exists.) Define $\bar{X}_n \equiv n^{-1} \sum_{i=1}^n X_i$. Let $G_n(x)$ denote the cdf of $\sqrt{n}(\bar{X}_n - \mu)/\sigma$. Then for any x such that $-\infty < x < \infty$,*

$$\lim_{n \rightarrow \infty} G_n(x) = \int_{-\infty}^x \frac{1}{\sqrt{2\pi}} e^{-y^2/2} dy;$$

that is, the asymptotic distribution of $\sqrt{n}(\bar{X}_n - \mu)/\sigma$ is standard normal.

- An equivalent way of stating the result of this CLT is: define

$$Z_n \equiv \frac{\bar{X}_n - \mathbb{E}(\bar{X}_n)}{\sqrt{\text{var}(\bar{X}_n)}} = \frac{\bar{X}_n - \mu}{\sigma/\sqrt{n}} = \frac{\sqrt{n}(\bar{X}_n - \mu)}{\sigma} \xrightarrow{d} Z \sim \mathcal{N}(0, 1).$$

- To prove this CLT, we need one important property of mgf: if two distributions have the same mgf, then they are identical at almost all points. Equivalently, if for all values of t ,

$$M_X(t) = M_Y(t),$$

then

$$F_X(x) = F_Y(x)$$

for all values of x ; that is, X and Y have the same distribution.

- *Proof.* We will show that for $|t| < h$, the mgf of $\sqrt{n}(\bar{X}_n - \mu)/\sigma$ converges to $e^{t^2/2}$, the mgf of a $\mathcal{N}(0, 1)$ random variable. Define $Y_i \equiv (X_i - \mu)/\sigma$, and let $M_Y(t)$ the common mgf of Y_i s, which exists for $|t| < \sigma h$. Note that Y_i has mean 0 and variance 1, then by the property of mgf, we have

$$\begin{aligned} M_Y^{(1)}(0) &\equiv \frac{d}{dt} M_Y(t)|_{t=0} = 0, \\ M_Y^{(2)}(0) &\equiv \frac{d^2}{dt^2} M_Y(t)|_{t=0} = 1. \end{aligned}$$

Since

$$\frac{\sqrt{n}(\bar{X}_n - \mu)}{\sigma} = \frac{1}{\sqrt{n}} \sum_{i=1}^n Y_i,$$

we have, by Lemmas 2 and 4,

$$\begin{aligned} M_{\sqrt{n}(\bar{X}_n - \mu)/\sigma}(t) &= M_{\sum_{i=1}^n Y_i/\sqrt{n}}(t) \\ &= M_{\sum_{i=1}^n Y_i}(t/\sqrt{n}) \\ &= [M_Y(t/\sqrt{n})]^n. \end{aligned}$$

Expanding $M_Y(t/\sqrt{n})$ in a Taylor series around 0 gives

$$\begin{aligned} [M_Y(t/\sqrt{n})]^n &= 1 + M_Y^{(1)}(0) \frac{t/\sqrt{n}}{1!} + M_Y^{(2)}(0) \frac{(t/\sqrt{n})^2}{2!} + R_Y(t/\sqrt{n}) \\ &= 1 + \frac{(t/\sqrt{n})^2}{2!} + R_Y(t/\sqrt{n}) \end{aligned}$$

By the definition of the Taylor remainder term, it is easy to see that for fixed t , $\lim_{n \rightarrow \infty} R_Y(t/\sqrt{n})/(\sqrt{n})^2 = 0$. As a result,

$$\lim_{n \rightarrow \infty} [M_Y(t/\sqrt{n})]^n = \lim_{n \rightarrow \infty} \left[1 + \frac{(t/\sqrt{n})^2}{2!} \right]^n = \lim_{n \rightarrow \infty} \left[1 + \frac{t^2}{2n} \right]^n = e^{t^2/2}.$$

This completes the proof.

- **Lemma 4** For any constants a and b , the mgf of the random variable $aX + b$ is given by

$$M_{aX+b}(t) = e^{bt}M_X(at).$$

- Even though we prove the previous CLT using mgf, the CLT is valid in much more generality. In particular, all of the assumptions about mgfs are not needed, similar proof goes through with characteristic functions.
- **Theorem 15 (Lindeberg-Levy CLT)** Let X_1, X_2, \dots be a sequence of i.i.d. random variables with mean μ and variance $0 < \sigma^2 < \infty$. Define $\bar{X}_n \equiv n^{-1} \sum_{i=1}^n X_i$. Let $G_n(x)$ denote the cdf of $\sqrt{n}(\bar{X}_n - \mu)/\sigma$. Then, for any x , $-\infty < x < \infty$,

$$\lim_{n \rightarrow \infty} G_n(x) = \int_{-\infty}^x \frac{1}{\sqrt{2\pi}} e^{-y^2/2} dy;$$

that is, the asymptotic distribution of $\sqrt{n}(\bar{X}_n - \mu)/\sigma$ is standard normal.

- Continuous mapping theorem also applies to convergence in distribution.
- **Example 7** If $Z_n \xrightarrow{d} Z \sim \mathcal{N}(0, 1)$, then the continuous mapping theorem implies that $Z_n^2 \xrightarrow{d} Z^2 \sim \chi_1^2$.
- However, $X_n \xrightarrow{d} X$ and $Y_n \xrightarrow{d} Y$ do not necessarily imply that $X_n + Y_n \xrightarrow{d} X + Y$ or $X_n Y_n \xrightarrow{d} XY$. (Can you come up with a counterexample?)

Slutsky's Theorem

- Slutsky's theorem is an important tool when deriving the asymptotic distributions for statistics.
- **Theorem 16** If $X_n \xrightarrow{d} X$ and $Y_n \xrightarrow{p} c$, where c is a constant. Then:

$$(a) Y_n X_n \xrightarrow{d} cX;$$

$$(b) X_n + Y_n \xrightarrow{d} X + c.$$

- **Example 8 (*t statistic*)** Suppose X_1, X_2, \dots is a random sequence of i.i.d random variables with mean μ and variance σ^2 such that $0 < \sigma^2 < \infty$, and suppose that $\text{var}(S_n^2) \rightarrow 0$ as $n \rightarrow \infty$. Then by CLT we have

$$\frac{\sqrt{n}(\bar{X}_n - \mu)}{\sigma} \xrightarrow{d.} \mathcal{N}(0, 1);$$

and by the result in Example 2 and the continuous mapping theorem we have

$$\frac{\sigma}{S_n} \xrightarrow{p.} 1.$$

Then by Slutsky's theorem, we can conclude that

$$\frac{\sqrt{n}(\bar{X}_n - \mu)}{S_n} \xrightarrow{d.} \mathcal{N}(0, 1).$$

Cramér-Wold Theorem

- We have learned the tools that allow us to show convergence in distribution of a sequence of random variables. Let $\{\mathbf{b}_n\}$ be a sequence of $k \times 1$ random vectors. In general, convergence in distribution of each element in the vectors does not imply the *joint* convergence in distribution of the vectors. But, we have the following result that gives the condition for the joint convergence to hold.
- **Theorem 17 (*Cramér-Wold Theorem*)** Let $\{\mathbf{b}_n\}$ be a sequence of $k \times$ random vectors and suppose that for any real $k \times 1$ vector λ such that $\lambda' \lambda = 1$ and $\lambda' \mathbf{b}_n \xrightarrow{d.} \lambda' \mathbf{Z}$, where \mathbf{Z} is a $k \times 1$ vector with joint distribution function F . Then the limit distribution function of \mathbf{b}_n exists and equals F .
- Cramér-Wold theorem implies that \mathbf{b}_n is asymptotically jointly normal if any linear combination of its elements is asymptotically normal.
- Consider a special case where $\lambda = (1, 0, \dots, 0)'$, then $\lambda' \mathbf{b}_n \xrightarrow{d.} \lambda' \mathbf{Z}$ is just the convergence of the first element of \mathbf{b}_n to that of \mathbf{Z} . Similar results hold for $\lambda = (0, 1, 0, \dots, 0)'$ and so on. Therefore, convergence of the elements is a necessary but not sufficient condition of the convergence of the vectors.

Delta Method

- CLT tells us that if a statistic takes the form of a sample mean, then under mild conditions its asymptotic distribution is normal. However, sometimes we are not interested in the sample means themselves, but some functions of the sample means.

- **Theorem 18 (*Delta Method*)** Let Y_n be a sequence of random variables that satisfies $\sqrt{n}(\bar{Y}_n - \theta) \xrightarrow{d} \mathcal{N}(0, \sigma^2)$. For a given function g and a specific value of θ , suppose that $g'(\theta)$ exists and is not 0. Then

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

- *Proof.* The first order Taylor expansion of $g(Y_n)$ around $Y_n = \theta$ implies

$$\sqrt{n}[g(Y_n) - g(\theta)] = g'(\theta)\sqrt{n}(Y_n - \theta) + \text{remainder},$$

where the remainder term takes the form of $c_n\sqrt{n}(Y_n - \theta)$, where $c_n \xrightarrow{p} 0$. By Slutsky's theorem, the remainder term $\xrightarrow{p} 0$. Then again by Slutsky's theorem, we have

$$\sqrt{n}[g(Y_n) - g(\theta)] \xrightarrow{d} g'(\theta) \cdot \mathcal{N}(0, \sigma^2) \stackrel{d}{=} \mathcal{N}(0, g'(\theta)^2 \sigma^2).$$

($\stackrel{d}{=}$ means that the left hand side and the right hand side have the same distribution.)

- **Example 9** Let X_1, X_2, \dots be a sequence of i.i.d. random variables with mean $\mu \neq 0$ and variance $0 < \sigma^2 < \infty$. Consider

$$\sqrt{n} \left(\frac{1}{\bar{X}_n} - \frac{1}{\mu} \right) = \sqrt{n}[g(\bar{X}_n) - g(\mu)],$$

where the function $g(x) = 1/x$. Note that $g'(x) = -1/x^2$, then using the delta method,

$$\sqrt{n} \left(\frac{1}{\bar{X}_n} - \frac{1}{\mu} \right) \xrightarrow{d} \mathcal{N}(0, \sigma^2/\mu^4).$$

Standardizing the statistic, we get

$$Z_n \equiv \frac{\sqrt{n} \left(\frac{1}{\bar{X}_n} - \frac{1}{\mu} \right)}{\sigma/\mu^2} \xrightarrow{d} \mathcal{N}(0, 1);$$

and by Slutsky's theorem,

$$t_n \equiv \frac{\sqrt{n} \left(\frac{1}{\bar{X}_n} - \frac{1}{\mu} \right)}{S_n / (\bar{X}_n)^2} \xrightarrow{d} \mathcal{N}(0, 1).$$

- If $g'(\theta) = 0$, then we often need the second-order delta method.
- **Theorem 19 (Second-Order Delta Method)** Let Y_n be a sequence of random variables that satisfies $\sqrt{n}(Y_n - \theta) \xrightarrow{d} \mathcal{N}(0, \sigma^2)$. For a given function g and a specific value of θ , suppose that $g'(\theta) = 0$ and $G''(\theta)$ exists and is not 0. Then

$$n[g(Y_n) - g(\theta)] \xrightarrow{d} \frac{\sigma^2 g''(\theta)}{2} \chi_1^2.$$

- *Proof.* Here we only provide a sketch of proof. When $g'(\theta) = 0$, we take one more term in the Taylor expansion to get

$$\begin{aligned} g(Y_n) - g(\theta) &= g'(\theta)(Y_n - \theta) + \frac{g''(\theta)}{2}(Y_n - \theta)^2 + \text{remainder} \\ &= \frac{g''(\theta)}{2}(Y_n - \theta)^2 + \text{remainder}. \end{aligned}$$

Recall that the square of a $\mathcal{N}(0, 1)$ variable has χ_1^2 distribution, which implies that

$$\frac{n(Y_n - \theta)^2}{\sigma^2} \xrightarrow{d} \chi_1^2.$$

Then the result of the theorem can be established by applying Slutsky's theorem.

- **Theorem 20 (Multivariate Delta Method)** Let $\{\mathbf{X}_n\}$ be a sequence of $k \times 1$ random vectors such that $\sqrt{n}(\mathbf{X}_n - \theta) \xrightarrow{d} \mathcal{N}(0, \Sigma)$, where Σ is a $k \times k$ positive semi-definite matrix. Suppose that the function $g: \mathbb{R}^k \rightarrow \mathbb{R}$ is differentiable at θ and $\nabla g(\theta)$ denotes the corresponding gradient vector. Then

$$\sqrt{n}[g(\mathbf{X}_n) - g(\theta)] \xrightarrow{d} \mathcal{N}(0, \nabla g(\theta)^T \Sigma \nabla g(\theta)).$$

Bounded in Probability

- **Definition 10** A sequence of random variables $\{X_n\}$ is called to be **bounded in probability** if for any $\epsilon > 0$ there exists a constant $B_\epsilon > 0$ and an integer N_ϵ such that for all $n \geq N_\epsilon$, we have

$$P(|X_n| \leq B_\epsilon) \geq 1 - \epsilon$$

- **Theorem 21** Let $\{X_n\}$ be a sequence of random variables and let X be a random variable. If $X_n \xrightarrow{d} X$, then $\{X_n\}$ is bounded in probability.
- **Theorem 22** Let $\{X_n\}$ be a sequence of random variables bounded in probability and let $\{Y_n\}$ be a sequence of random variables which converge to 0 in probability. Then $X_n Y_n \xrightarrow{p} 0$.

Big O and Small o Notation

- In this subsection, we introduce the big O small o notation. It is very useful in econometrics.
- **Definition 11 (Big O Notation)** For a sequence of random variables and a sequence a_n (random variables or positive constants), the notation

$$X_n = O_p(a_n)$$

means that the sequence X_n/a_n is bounded in probability. We say that X_n is at most of order in magnitude of a_n .

- **Definition 12 (Small o Notation)** The notation

$$X_n = o_p(a_n)$$

means that the sequence X_n/a_n converges to zero in probability. We say that X_n is of smaller order in magnitude than a_n .

- **Example 10** $X_n = o_p(1)$ is equivalent to $X_n \xrightarrow{p} 0$ (when $a_n = 1$).
- **Theorem 23** If $X_n = o_p(1)$, then $X_n = O_p(1)$.
- **Example 11** If $X_n = O_p(a_n)$ and $a_n \rightarrow 0$ (or $a_n = o_p(1)$), then $X_n = o_p(1)$.

- CLT implies that $X_n - \mu = O_p(n^{-1/2})$; the delta method implies that $g(Y_n) - g(\theta) = O_p(n^{-1/2})$ if $g'(\theta) \neq 0$.
- The second-order delta method implies that $g(Y_n) - g(\theta) = O_p(n^{-1})$ if $g'(\theta) = 0$ and $g''(\theta) \neq 0$.
- **Example 12** *Using this notation, the result of Theorems 21 and 22 can be written succinctly as:*

(a) If $X_n \xrightarrow{d} X$, then $X_n = O_p(1)$.

(b) $O_p(1) \cdot o_p(1) = o_p(1)$.

- The following results hold:
 - (a) $O_p(1) + O_p(1) = O_p(1)$, $O_p(1) \cdot O_p(1) = O_p(1)$;
 - (b) $o_p(1) + o_p(1) = o_p(1)$, $o_p(1) \cdot o_p(1) = o_p(1)$;
 - (c) $O_p(1) + o_p(1) = O_p(1)$, $O_p(1) \cdot o_p(1) = o_p(1)$;
 - (d) $O_p(a_n) + O_p(b_n) = O_p(\max\{a_n, b_n\})$, $O_p(a_n) \cdot O_p(b_n) = O_p(a_n \cdot b_n)$;
 - (e) $o_p(a_n) + o_p(b_n) = o_p(\max\{a_n, b_n\})$, $o_p(a_n) \cdot o_p(b_n) = o_p(a_n \cdot b_n)$;
 - (f) $O_p(a_n) + o_p(b_n) = ?$ (it depends on the exact order in magnitude of each), $O_p(a_n) \cdot o_p(b_n) = o_p(a_n \cdot b_n)$;

5 Exercises

1. Prove Lemma 2.
2. Find $\text{var}(S^2)$ for a normal random sample.
3. Prove Theorems 5 and 6.
4. Prove the result in Example 4 using the definition of convergence in probability, that is, without invoking the continuous mapping theorem.
5. Prove Theorem 9. Hint: what is the relationship between $P(|(X_n + Y_n) - (X + Y)| \geq \epsilon)$, $P(|X_n - X| \geq \epsilon/2)$ and $P(|Y_n - Y| \geq \epsilon/2)$?
6. Prove Theorem 12 for the case where X_n and X are continuous random variables. Follow these steps:

- (a) Given t and ϵ , show that $P(X \leq t - \epsilon) \leq P(X_n \leq t) + P(|X_n - X| \geq \epsilon)$. This gives a lower bound on $P(X_n \leq t)$;
 - (b) Use a similar strategy to get an upper bound on $P(X_n \leq t)$;
 - (c) By pinching, deduce that $P(X_n \leq t) \rightarrow P(X \leq t)$.
7. Generate 1000 random samples $\{X_i\}_{i=1}^n$ of size n from each of the following distributions. Try different sample sizes $n = 1, 5, 10, 25, 50, 100$.
- (a) Bernoulli distribution with $p = 0.78$;
 - (b) Uniform distribution on $[0, 2]$;
 - (c) Exponential distribution with mean 0.5;
 - (d) Distribution with the pdf

$$f_X(x) = \frac{1}{\sqrt{2\pi}} \frac{1}{x} \exp[-(\ln x)^2/2];$$

- (e) t distribution with df 2;
- (f) Cauchy distribution.

Find μ and σ^2 for each distribution. Plot the histogram of $\sqrt{n}(\bar{X}_n - \mu)/\sigma$ for each n , and compare the histograms with pdf of the standard normal distribution.

8. Let X_1, X_2, \dots be a random sample from Bernoulli(p) distribution, and let $Y_n \equiv n^{-1} \sum_{i=1}^n X_i$.
- (a) Show that $\sqrt{n}(Y_n - p) \xrightarrow{d} \mathcal{N}(0, p(1 - p))$;
 - (b) Show that for $p \neq 1/2$, the estimate of variance $Y_n(1 - Y_n)$ satisfies $\sqrt{n}[Y_n(1 - Y_n) - p(1 - p)] \xrightarrow{d} \mathcal{N}(0, (1 - 2p)^2 p(1 - p))$;
 - (c) Show that for $p = 1/2$, $n[Y_n(1 - Y_n) - 1/4] \xrightarrow{d} -\chi_1^2/4$.
9. Prove Theorem 23.