

Econometrics 2: Non-parametric methods

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I have a question!

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1 Preliminaries

1.1 Probability basics

Definition 1.1 (distribution law). The distribution law of a random variable X is \mathbb{P}_X is the probability on $(\mathbb{R}^d, \mathcal{B}(\mathbb{R}^d))$ such that $\mathbb{P}_X[B] = \mathbb{P}[X \in B]$ for all $B \in \mathcal{B}(\mathbb{R}^d)$

Definition 1.2 (density). X has density f_X if $\mathbb{P}_X[B] = \int_B \underbrace{f_X(x)dx}_{d\mathbb{P}_X(x)}$ for all $B \in \mathcal{B}(\mathbb{R}^d)$

Let Y be a random variable and X be a random vector in \mathbb{R}^d defined on the same probability space $(\Omega, \mathcal{F}, \mathbb{P})$. We want to define and manipulate $\mathbb{E}[Y|X]$.

There are 2 particular cases

1. $(Y, X^\top)^\top$ has discrete support. Let $x \in \text{spt}(X)$, then $\mathbb{E}[Y|X = x] = \sum y_j \mathbb{P}(Y = y_j | X = x)$. This is well defined only when $\mathbb{P}(X = x) > 0$ in which case $\mathbb{P}(Y = y_j | X = x) = \frac{\mathbb{P}(Y = y_j, X = x)}{\mathbb{P}(X = x)}$ is well defined.
2. $(Y, X^\top)^\top$ and X have a density then $\mathbb{E}[Y|X = x] = \int y f_{Y|X=x}(y) dy$ where $f_{Y|X=x}(y) = \frac{f_{Y,X}(y, x)}{f_X(x)}$.

Proposition 1.1 (conditional expectation). The random variable $Z := \mathbb{E}[Y|X]$ is the unique random variable such that

1. $Z \in L^1(\Omega, \mathcal{F}, \mathbb{P})$, that is Z is $\sigma(X)$ -measurable.
2. $\mathbb{E}[Z \mathbb{1}_B] = \mathbb{E}[Y \mathbb{1}_B]$ for all $B \in \sigma(X)$.

unique means that if Z' is another random variable satisfying the same properties, then $Z = Z'$ a.s.

Remark. The random variable Z is $\sigma(X)$ -measurable iff $Z = \phi(X)$ for some function $\phi : (\mathbb{R}^d, \mathcal{B}(\mathbb{R}^d)) \rightarrow (\mathbb{R}, \mathcal{B}(\mathbb{R}))$. The corresponding function ϕ for $Z = \underbrace{\mathbb{E}[Y|X]}_{\text{conditional expectation}}$ is de-

noted by $\underbrace{\mathbb{E}[Y|X = x]}_{\text{conditional expectation function}}$.

Remark. The proposition 2 is equivalent to

$$\begin{aligned} \mathbb{E}[(Y - Z) \mathbb{1}_B] &= 0, \quad \forall B \in \sigma(X) \\ \Leftrightarrow \mathbb{E}[(Y - Z) \psi(X)] &= 0, \quad \forall \psi \text{ bounded and measurable.} \end{aligned}$$

Exercise. Let X and β be random vectors in \mathbb{R}^d such that X and β ARE independent, that is $\mathbb{P}_{X,\beta} = \mathbb{P}_X \times \mathbb{P}_\beta$. Let $g : \mathbb{R}^d \rightarrow \mathbb{R}$ be a bounded and measurable function. Define $Y = g(X^\top \beta)$. It can be shown that $\mathbb{E}[Y|X = x] = \mathbb{E}[g(x^\top \beta)]$ for all $x \in \text{supp}(X)$.

Proof. Let $B \in \sigma(X)$. Then

$$\mathbb{E}[Y \mathbb{1}\{X \in B\}] = \mathbb{E}[g(X^\top \beta) \mathbb{1}\{X \in B\}] = \int \int g(X^\top b) \mathbb{1}\{x \in B\} d\mathbb{P}_{\beta, X}(b, x).$$

Since X and β are independent, we have

$$\mathbb{P}_{\beta, X}(b, x) = \mathbb{P}_X(x) \mathbb{P}_\beta(b).$$

Therefore,

$$\begin{aligned} \mathbb{E}[Y \mathbb{1}\{X \in B\}] &= \int \int g(x^\top b) \mathbb{1}\{x \in B\} d\mathbb{P}_\beta(b) d\mathbb{P}_X(x) \\ &= \int \mathbb{E}[g(x^\top \beta)] \mathbb{1}\{x \in B\} d\mathbb{P}_X(x) \\ &= \mathbb{E}[\mathbb{E}[g(x^\top \beta)] \mathbb{1}\{X \in B\}] \\ &= \mathbb{E}[\phi(X) \mathbb{1}\{X \in B\}] \end{aligned}$$

where $\mathbb{E}[g(x^\top \beta)] \equiv \phi(x)$ takes expectation over β and is a function of x . By the uniqueness of conditional expectation, we have $\mathbb{E}[Y|X] = \mathbb{E}[\phi(X)]$. \square

Exercise. Let \mathcal{H} be a σ -algebra such that all elements of \mathcal{H} belong to $\sigma(X)$. We can show that

$$\mathbb{E}[\mathbb{E}[Y|X] | \mathcal{H}] = \mathbb{E}[Y | \mathcal{H}]$$

We can think of this in terms of projections. The projection of Y onto \mathcal{H} is $\mathbb{E}[Y | \mathcal{H}]$, and the projection of $\mathbb{E}[Y|X]$ onto \mathcal{H} is $\mathbb{E}[\mathbb{E}[Y|X] | \mathcal{H}]$. The equality says that the projection of the projection of Y onto \mathcal{H} is the same as the projection of Y onto \mathcal{H} .

1.2 Identification

We are given data consisting of draws from a distribution law $\mathbb{P}_{Y, X}$ where Y, X are observable vectors. An economic model consists of

1. An equation $v(Y, \gamma, X, \varepsilon; \zeta) = 0$ where v is a system of functions, γ is a vector of variables that is determined within the model but unobservable, ε is a vector of variables that is determined within the model and unobservable. ζ is a vector of functions and distributions.
2. Restrictions: $\zeta \in \mathcal{R}$ where \mathcal{R} is a set of restrictions.

For any $\zeta \in \mathcal{R}$, $\mathbb{P}_{Y, X; \zeta}$ is the distribution law of the observables generated by ζ . We denote the true structural parameter by ζ^* . **We often care about $\psi^* = \Psi(\zeta^*)$ where Ψ is a mapping from \mathcal{R} to \mathcal{P} and \mathcal{P} is the parameter space.** We define the identified set as

$$\Gamma_{Y, X}(\psi, \mathcal{R}) = \{\mathbb{P}_{Y, X; \zeta} : \zeta \in \mathcal{R} \text{ s.t. } \Psi(\zeta) = \psi\}.$$

It is the set of all distributions of the observables that are consistent with the model and the restrictions, that is, generated by ζ contained within the restriction.

Definition 1.3 (Identification). We say that ψ^* is identified if for any $\psi^* \in \mathcal{P}$ if $\Gamma_{Y,X}(\psi^*, \mathcal{R}) \cap \Gamma_{Y,X}(\psi, \mathcal{R}) \neq \emptyset$, then $\psi^* = \psi$.

Exercise. We specify a linear model $Y = f(X) + \varepsilon$ where f is continuous near $x_0 \in \text{supp}(X)$, and $\mathbb{E}[|\varepsilon| + |f(X)|] < \infty$ and $\mathbb{E}[\varepsilon|X] = 0$. We can show that $\psi^* = \mathbb{E}[f(X)]$ is identified because under these conditions $f(X) = \mathbb{E}[f(X)]$, the conditional expectation. The system of equations is $v(Y, \gamma, X, \varepsilon; \zeta) = Y - f(X) - \varepsilon = 0$. The restriction is $\zeta = (f, \mathbb{P}_{X,\varepsilon}) \in \mathcal{R}$.

Proof. Assume that there are two ζ that satisfy the restrictions and generate the same distribution of the observables.

$$(f, \mathbb{P}_{X,\varepsilon}), (f^*, \mathbb{P}_{X,\varepsilon}^*) \xrightarrow[\text{generate}]{} \mathbb{P}_{Y,X}$$

Then we have the following

- $\int \mathbb{P}_{Y,X(y,x)}(y, \cdot) dy = \int \mathbb{P}_{\varepsilon,X}(e, \cdot) de = \int \mathbb{P}_{\varepsilon,X}^*(e, \cdot) = \mathbb{P}_X(x)$
- $\mathbb{E}[Y|X] = f(X) = f^*(X)$
- Because f is identified, $\mathbb{P}_{\varepsilon,X} = \mathbb{P}_{Y-f(X),X} = \mathbb{P}_{Y-f^*(X),X} = \mathbb{P}_{\varepsilon,X}^*$

□

Now that we have introduced the basic nonparametric model, we introduce nonparametric model with instrumental variables, where $\mathbb{E}[\varepsilon|Z] = 0$. The identification requires an additional restriction – *Completeness*.

Definition 1.4 (Completeness). For any ϕ such that $\mathbb{E}[|\phi(X)|] < \infty$, $\mathbb{E}[\phi(X) | Z] = 0$ implies that $\phi(x) = 0$ on the support of X .

Discrete case When (X, Z) is discrete finite,

$$\text{supp}(X) = \{x_1, \dots, x_n\}, \text{supp}(Z) = \{z_1, \dots, z_m\}$$

We can write the completeness condition as

$$\mathbb{E}[\phi(X) | Z = z_j] = \sum \phi(x_i) \mathbb{P}_{X|Z}(X = x_i | Z = z_j) = 0 \quad \forall j$$

This is a system of m equations in n unknowns. We can show that the completeness condition is satisfied if and only if $m \geq n$. In the following section, we discuss a specific continuous case ($\text{supp}(X, Z) \in \mathbb{R}^2$) where $X = Z - \eta$ and Z is independent of ε . Both have densities and $\eta \sim \mathcal{U}([-1, 1])$.

Parametric VS Non-Parametric Models If $\zeta = (\gamma, p_\theta)$ where $\gamma \in \mathbb{R}^{d_\gamma}, \theta \in \mathbb{R}^{d_\theta}$, then the model is parametric. If $\zeta = (\gamma, p)$ where $\gamma \in \mathbb{R}^{d_\gamma}, p$ is a distribution, then the model is non-parametric. When we care about $\psi^* = \Psi(\zeta^*)$, we are in the framework of semi-parametric models.

Example 1.1. The model is $Y = \alpha + \beta X + \varepsilon$ where $\zeta = (\alpha, \beta, p_{\varepsilon,X})$ and $\psi^* = (\alpha^*, \beta^*)$. We are in semi-parametric models because α, β are finite-dimensional and $p_{\varepsilon,X}$ is infinite-dimensional.

1.3 Completeness condition

We want to understand the completeness condition when $X = Z - \eta$, where $Z \perp \eta$ and both have densities. Recall the definition of **completeness**.

Definition 1.5 (Completeness). Completeness is defined as such that

$$\forall z \in \mathbb{R}, \int_{\mathbb{R}} \varphi(x) f_{\eta}(z - x) dx = 0 \text{ implies that for all } x, \varphi(x) = 0,$$

where φ is continuous and $\int_{\mathbb{R}} |\varphi(x)| dx < \infty$.

Now, We make a detour to introduce some notations in function space.

Definition 1.6. Let f be a function defined on \mathbb{R}^d with values in \mathbb{R} or \mathbb{C} and $p \leq 1$. Then $L^p(\mathbb{R}^d)$ is defined as the space of measurable function from $(\mathbb{R}^d, \mathcal{B}(\mathbb{R}^d))$ such that $\int_{\mathbb{R}^d} |f(x)|^p dx < \infty$. If f takes value from \mathbb{C} , $|\cdot|$ is the modulus.

Definition 1.7 ($L^1(\mathbb{R})$ space). A function is in $L^1(\mathbb{R})$ if $\int_{\mathbb{R}} |f(x)| dx < \infty$.

Definition 1.8 (Fourier transform). If $f \in L^1(\mathbb{R})$, the Fourier transform of f is defined for all $w \in \mathbb{R}$ by

$$\mathcal{F}[f](w) = \int_{\mathbb{R}} e^{iwx} f(x) dx.$$

Remark. Let $t \in \mathbb{R}$, $e^{it} = \cos(t) + i \sin(t)$ and $|e^{it}|^2 = 1$

Definition 1.9 (Convolution). If f and g belong to $L^1(\mathbb{R}^d)$, the convolution of f and g is $f * g(z) = \int f(x)g(z - x)dx$.

Proposition 1.2. If f and g belong to $L^1(\mathbb{R}^d)$, then $f * g \in L^1(\mathbb{R}^d)$. Its Fourier transformation is $\mathcal{F}[f * g](w) = \mathcal{F}[f](w) \mathcal{F}[g](w)$ for all $w \in \mathbb{R}^d$

Remark. check this proposition as an exercise.

Proposition 1.3. If $f \in L^1(\mathbb{R}^d)$, then $\mathcal{F}[f]$ is continuous and $\lim_{\|w\|_2 \rightarrow \infty} \mathcal{F}[f](w) = 0$.

We introduce two properties that are useful for later cause.

Property 1.1. If $f, \mathcal{F}[f] \in L^2(\mathbb{R}) \cap L^1(\mathbb{R})$, then

1. (The Placherel equality) $\frac{1}{2\pi} \|\mathcal{F}[f]\|_2^2 = \|f\|_2^2$ (Plancherel's theorem)
2. (The Fourier inverse formula) For all $x \in \mathbb{R}$, $f(x) = \frac{1}{2\pi} \int_{\mathbb{R}} e^{-iwx} \mathcal{F}[f](w) dw$, the inversion of the Fourier transform.

Question 1. Let $Z \in L^2(\Omega, \mathcal{F}, \mathbb{P})$, then $\mathbb{E}[|z|] \leq \sqrt{\mathbb{E}[z^2]} \sqrt{\mathbb{E}[1^2]}$. Therefore, $L^2(\Omega, \mathcal{F}, \mathbb{P}) \subset L^1(\Omega, \mathcal{F}, \mathbb{P})$.

Example 1.2. Let $K(x) = \frac{1}{\sqrt{2\pi}}e^{-\frac{x^2}{2}}$, then $K \in L^1(\mathbb{R}) \cap L^2(\mathbb{R})$. Then for all $w \in \mathbb{R}$,

$$F[K](w) = e^{-\frac{w^2}{2}}.$$

Example 1.3. Let $K(x) = \frac{1}{\sqrt{2}}\mathbb{1}_{\{|x| \leq 1\}}$, then $K \in L^1(\mathbb{R}) \cap L^2(\mathbb{R})$. Then for all $w \in \mathbb{R}$,

$$\begin{aligned} F[K](w) &= 1/2 \int_{-1}^1 \cos(wx)dx + 1/2 \int_{-1}^1 \sin(wx)dx \\ &= \frac{1}{2w} [\sin(wx)] \Big|_{-1}^1 \\ &= \frac{\sin(wx)}{w} \end{aligned}$$

Here $F[K] \notin L^1(\mathbb{R})$ but $F[K] \in L^2(\mathbb{R})$. Note also that $F[K](w) = 0$ if and only if $w = \pm k\pi$ for $k \in \mathbb{N}$.

Let us check whether the functions given in Example 1.2 and 1.3 satisfy the completeness condition 1.5 for $X = Z - \eta$.

1. Since $F[f_\eta](w) > 0$ for all w . Thus, $F[\varphi](w) = 0 \Leftrightarrow \varphi(x) = 0$ for all x .
2. Similarly, $F[\varphi](w) = 0$ for all $w \in \mathbb{R} \setminus S$. Because φ is continuous, it is 0 everywhere.

2 Density function and kernel estimation

2.1 Density function

We want to estimate the density f_X of $X \in \mathbb{R}$ and will work among classes of densities. For example,

1. **continuous densities**
2. densities such that for all $x, x' \in \mathbb{R}$, $|f_X(x) - f_X(x')| \leq M|x - x'|$ for some $M > 0$
3. densities which are **monotonically increasing** on $[0, 1]$

2.2 Density function estimation

First, we introduce a **histogram estimator**. Assume $f_X : [0, 1] \rightarrow \mathbb{R}$ and divide the interval $[0, 1]$ into n equal non overlapping intervals I_1, \dots, I_n of length $h = 1/n$. Define the $n_j = \sum_{i=1}^n \mathbb{1}_{\{X_i \in I_j\}}$. The histogram estimator is defined as

$$\hat{f}_n(x) = \sum_{i=1}^n \mathbb{1}_{\{x \in I_j\}} \frac{n_j}{nh} = \frac{1}{nh} \sum_{i=1}^n \underbrace{\sum_{j=1}^n \mathbb{1}_{\{X_i \in I_j\}} \mathbb{1}_{\{x \in I_j\}}}_{K_h(X_i, x)}.$$

If X has a density f_X , then $f_X(x) = F'_X(x)$ a.e. because

$$F_X(x) = \int_{-\infty}^x f_X(t)dt = \mathbb{E} [\mathbb{1}_{\{X \leq x\}}] .$$

A natural estimator of the CDF is the **empirical CDF**, defined as

$$\hat{F}_n(x) = \frac{1}{n} \sum_{i=1}^n \mathbb{1}_{\{X_i \leq x\}} .$$

where n is the sample size. Therefore, an estimator of f_X is the derivative of the empirical CDF, which is the **empirical density function** defined as

$$\hat{f}_n(x) = \frac{\hat{F}_X(x + h/2) - \hat{F}_X(x - h/2)}{h} = \frac{1}{nh} \sum_{i=1}^n K\left(\frac{X_i - x}{h}\right)$$

where $K(x) = \mathbb{1}_{\{x \leq \frac{1}{2}\}}$ (is the **kernel** and $h > 0$ is the **bandwidth**).

Definition 2.1 (kernel function). A kernel is a function $K : \mathbb{R} \rightarrow \mathbb{R}$ such that $K \in L^1(\mathbb{R})$ and $\int K(x)dx = 1$.

Definition 2.2 (kernel density estimator with kernel K and bandwidth h). The kernel density estimator of f_X is defined as

$$\hat{f}_n(x) = \frac{1}{nh} \sum_{i=1}^n K\left(\frac{X_i - x}{h}\right)$$

2.3 Kernels estimators

Some kernels We list out some common kernels.

1. $K(x) = \frac{1}{2} \mathbb{1}_{|x| \leq \frac{1}{2}}$ the rectangular
2. $K(x) = \frac{1}{\sqrt{2\pi}} e^{-\frac{x^2}{2}}$, the Gaussian kernel
3. $K(x) = \frac{\sin(x)}{\pi x}$, the sinc kernel
4. $K(x) = \frac{3}{4} \max\{0, 1 - x^2\} = \frac{3}{4} (1 - x^2) \mathbb{1}_{|x| \leq 1}$, the Epanechnikov kernel

Remark. Note that the Gaussian kernel is both in $L^1(\mathbb{R}, \mathcal{B}(\mathbb{R}), dx)$ and in $L^2(\mathbb{R}, \mathcal{B}(\mathbb{R}), dx)$. The sinc kernel is only in $L^2(\mathbb{R}, \mathcal{B}(\mathbb{R}), dx)$ but not in $L^1(\mathbb{R}, \mathcal{B}(\mathbb{R}), dx)$, as the absolute value fails to be integrable. However, we have

$$1 = \lim_{R \rightarrow \infty} \int_{-R}^R \frac{\sin(x)}{\pi x} dx .$$

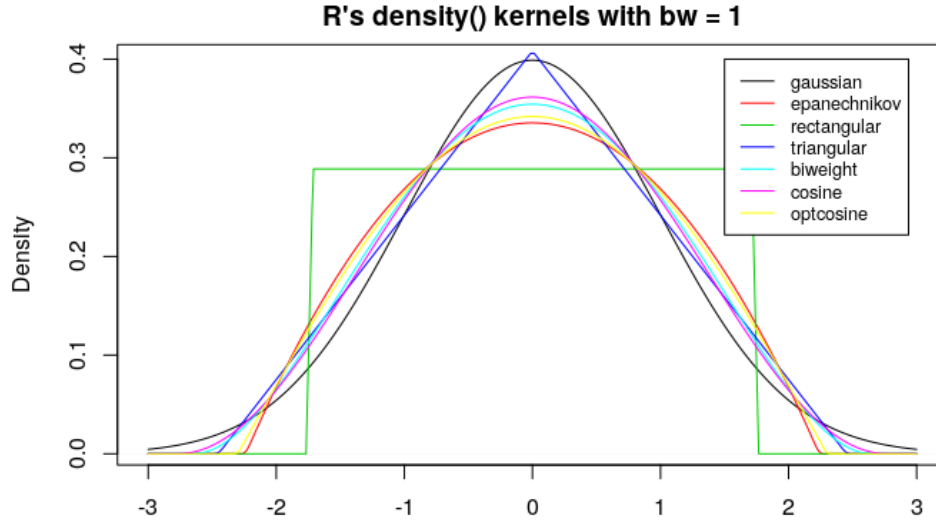


Figure 1: Kernels

Remark. The kernel estimator is unbiased when the bandwidth h goes to zero, that is,

$$\begin{aligned}
 \mathbb{E} \left[\hat{f}_X(x) \right] &= \frac{1}{hn} \sum_{i=1}^n \mathbb{E} \left[K \left(\frac{X_i - x}{h} \right) \right] \\
 &= \frac{1}{h} \mathbb{E} \left[K \left(\frac{X_1 - x}{h} \right) \right] \quad \text{by i.i.i.d.} \\
 &= \frac{1}{h} \int_{\mathbb{R}} K \left(\frac{y - x}{h} \right) f_X(y) dy \\
 &= \frac{1}{h} K \left(\frac{\cdot}{h} \right) * f_X(x) \\
 &\rightarrow f_X(x)
 \end{aligned}$$

It converges in $L^1(\mathbb{R})$ in the sense that

$$\left\| \frac{1}{h} K \left(\frac{\cdot}{h} \right) * f_X - f_X \right\|_1 \rightarrow 0$$

where $\|f\|_1 = \int_{\mathbb{R}} |f(x)| dx$

2.4 Performance analysis

For a kernel K estimator with bandwidth h , we would like to analyze its performance.

Definition 2.3. We introduce the quadratic **risk**

$$\text{MSE}(x) = \mathbb{E} \left[\left(\hat{f}_X(x) - f_X(x) \right)^2 \right],$$

where

$$\ell(x, y) = (x - y)^2$$

is the **loss** function.

Other risks include

$$\mathbb{E} \left[\sup_{x \in \mathbb{R}} \left| \hat{f}_X(x) - f_X(x) \right| \right] = \mathbb{E} \left[\left\| \hat{f}_X - f_X \right\|_{\infty} \right]$$

Note that \hat{f}_X is a function of x and the observations is $X = (X_1, \dots, X_n)$.

Definition 2.4. We define the **bias** of $\hat{f}_X(x)$ by

$$\text{Bias}(\hat{f}_X) = b(x) = \mathbb{E} \left[\hat{f}_X(x) - f_X(x) \right]$$

and we denote the **variance** of $\hat{f}_X(x)$ by $\sigma^2(x)$.

Proposition 2.1. We have

$$\text{MSE}(x) = b(x)^2 + \sigma^2(x).$$

Proof. We have

$$\begin{aligned} \text{MSE}(x) &= \mathbb{E} \left[\left(\hat{f}_X(x) - \mathbb{E}[\hat{f}_X(x)] + \mathbb{E}[\hat{f}_X(x)] - f_X(x) \right)^2 \right] \\ &= \mathbb{E} \left[\left(\hat{f}_X(x) - \mathbb{E}[\hat{f}_X(x)] \right)^2 \right] + 2 \mathbb{E} \left[\left(\hat{f}_X(x) - \mathbb{E}[\hat{f}_X(x)] \right) \underbrace{\left(\mathbb{E}[\hat{f}_X(x)] - f_X(x) \right)}_{\text{not random}} \right] \\ &\quad + \mathbb{E} \left[\left(\mathbb{E}[\hat{f}_X(x)] - f_X(x) \right)^2 \right] \\ &= \underbrace{\mathbb{E} \left[\left(\hat{f}_X(x) - \mathbb{E}[\hat{f}_X(x)] \right)^2 \right]}_{=\sigma^2(x)} + 2 \left(\mathbb{E}[\hat{f}_X(x)] - f_X(x) \right) \underbrace{\mathbb{E}[\hat{f}_X(x) - \mathbb{E}[\hat{f}_X(x)]]}_{=0} \\ &\quad + \underbrace{\left(\mathbb{E}[\hat{f}_X(x)] - f_X(x) \right)^2}_{=b(x)^2}. \end{aligned}$$

□

Proposition 2.2 (upper bound of $\sigma^2(x)$). Assume that there exists $f_{\max} \in \mathbb{R}$ such that $\forall x \in \mathbb{R}$, $f_X(x) \leq f_{\max}$ and $\int_{\mathbb{R}} K^2(u) du < \infty$. Then we have, for any bandwidth h , for any $x \in \mathbb{R}$ and for any $n \geq 1$, we have

$$\sigma^2(x) \leq \frac{C}{nh}.$$

where $C = f_{\max} \int_{\mathbb{R}} K^2(u) du$.

Proof. First observe that, by identical distribution of X_1, \dots, X_n ,

$$\mathbb{E} [\hat{f}_X(x)] = \frac{1}{n} \sum_{i=1}^n \frac{1}{h} \mathbb{E} \left[K \left(\frac{X_i - x}{h} \right) \right] = \frac{1}{h} \mathbb{E} \left[K \left(\frac{X_1 - x}{h} \right) \right]. \quad (1)$$

Now, using independence in the second line and identical distribution in the third line,

$$\begin{aligned} \sigma^2(x) &= \mathbb{E} \left[\left(\frac{1}{n} \sum_{i=1}^n \left(\frac{1}{h} K \left(\frac{X_i - x}{h} \right) \right) - \mathbb{E} [\hat{f}_X(x)] \right)^2 \right] \\ &= \mathbb{E} \left[\frac{1}{n^2} \left(\sum_{i=1}^n \left(\frac{1}{h} K \left(\frac{X_i - x}{h} \right) \right) - n \mathbb{E} [\hat{f}_X(x)] \right)^2 \right] \\ &= \frac{1}{n^2} \mathbb{E} \left[\sum_{i=1}^n \left(\frac{1}{h} K \left(\frac{X_i - x}{h} \right) - \mathbb{E} [\hat{f}_X(x)] \right)^2 \right] \text{ by independence} \\ &= \frac{1}{n^2} \sum_{i=1}^n \mathbb{E} \left[\left(\frac{1}{h} K \left(\frac{X_i - x}{h} \right) - \mathbb{E} [\hat{f}_X(x)] \right)^2 \right] \\ &= \frac{1}{n} \mathbb{E} \left[\left(\frac{1}{h} K \left(\frac{X_1 - x}{h} \right) - \mathbb{E} [\hat{f}_X(x)] \right)^2 \right] \end{aligned}$$

Inserting equation* 1,

$$\begin{aligned}
\sigma^2(x) &= \frac{1}{n} \mathbb{E} \left[\left(\frac{1}{h} K \left(\frac{X_1 - x}{h} \right) - \mathbb{E} \left[\frac{1}{h} K \left(\frac{X_1 - x}{h} \right) \right] \right)^2 \right] \\
&= \frac{1}{n} \text{Var} \left[\frac{1}{h} K \left(\frac{X_1 - x}{h} \right) \right] \\
&= \frac{1}{n} \left(\mathbb{E} \left[\frac{1}{h^2} K^2 \left(\frac{X_1 - x}{h} \right) \right] - \mathbb{E} \left[\frac{1}{h} K \left(\frac{X_1 - x}{h} \right) \right]^2 \right) \\
&\leq \frac{1}{n} \mathbb{E} \left[\frac{1}{h^2} K^2 \left(\frac{X_1 - x}{h} \right) \right] \\
&= \frac{1}{nh} \mathbb{E} \left[\frac{1}{h} K^2 \left(\frac{X_1 - x}{h} \right) \right] \\
&= \frac{1}{nh} \int_{\mathbb{R}} \frac{1}{h} K^2 \left(\frac{y - x}{h} \right) f_X(y) dy \\
&= \frac{1}{nh} \int_{\mathbb{R}} K^2(u) \underbrace{f_X(x + hu)}_{\leq f_{\max}} du \\
&\leq \frac{1}{nh} \underbrace{f_{\max} \int_{\mathbb{R}} K^2(u) du}_{=C},
\end{aligned}$$

where we used the change of variables $y = x + hu$. □

Definition 2.5 (β for a density function). Let $\beta > 0$, $L > 0$ and set $\ell = \lfloor \beta \rfloor$, by which we mean the greatest integer **strictly** less than β . The Hölder class $\Sigma(\beta, L)$ is the class of functions $f : \mathbb{R} \rightarrow \mathbb{R}$ such that $f^{(\ell)}$ exists and for all $x, x' \in \mathbb{R}$ we have

$$|f^{(\ell)}(x) - f^{(\ell)}(x')| \leq L |x - x'|^{\beta - \ell}.$$

Definition 2.6. We define

$$\mathcal{P}(\beta, L) = \left\{ f \in \Sigma(\beta, L) : f \geq 0, \int_{\mathbb{R}} f(x) dx = 1 \right\}.$$

Example 2.1. $\beta = 1$ gives the usual Hölder continuity condition: for all $x, x' \in \mathbb{R}$

$$|f(x) - f(x')| \leq L |x - x'|^\beta.$$

Remark. This Hölder condition implies continuity of f .

Definition 2.7 (β for a kernel). $K : \mathbb{R} \rightarrow \mathbb{R}$ is a kernel **of order** $\ell \in \mathbb{N}_0$ if

- $u \mapsto u^j K(u)$ is integrable for any $j \in \{0, \dots, \ell\}$,

- $\int_{\mathbb{R}} K(u) du = 1$,
- and $\int_{\mathbb{R}} u^j K(u) du = 0$ for $j \in \{1, \dots, \ell\}$.

Proposition 2.3 (upper bound of $|b(x)|$). Let $f_X \in \mathcal{P}(\beta, L)$ with $\beta, L > 0$ and K of order $\ell \geq \lfloor \beta \rfloor$ such that

$$\int_{\mathbb{R}} |u|^\beta |K(u)| du < \infty.$$

Then, for all $x \in \mathbb{R}$, $n \geq 1$ and $h > 0$, we have

$$|b(x)| \leq C_1 h^\beta,$$

where

$$C_1 = \frac{L}{\ell!} \int_{\mathbb{R}} |u|^\beta |K(u)| du.$$

Proof. Reusing equation 1 and using $1 = \int_{\mathbb{R}} K(u) du$,

$$\begin{aligned} b(x) &= \mathbb{E} \left[\hat{f}_X(x) \right] - f_X(x) \\ &= \frac{1}{h} \mathbb{E} \left[K \left(\frac{X_1 - x}{h} \right) \right] - f_X(x) \\ &= \frac{1}{h} \int_{\mathbb{R}} K \left(\frac{y - x}{h} \right) f_X(y) dy - f_X(x) \\ &= \frac{1}{h} \int_{\mathbb{R}} K \left(\frac{y - x}{h} \right) f_X(y) dy - \int_{\mathbb{R}} K(u) f_X(x) du. \end{aligned}$$

With the change of variables $y = hu + x$, we obtain

$$\begin{aligned} b(x) &= \int_{\mathbb{R}} K(u) f_X(hu + x) du - \int_{\mathbb{R}} K(u) f_X(x) du \\ &= \int_{\mathbb{R}} K(u) (f_X(hu + x) - f_X(x)) du. \end{aligned}$$

By a Taylor expansion, for some $\tau \in [0, 1]$, we obtain

$$f_X(hu + x) - f_X(x) = uh f'_X(x) + \dots + \frac{(uh)^{\ell-1}}{(\ell-1)!} f_X^{(\ell-1)}(x) + \frac{(uh)^\ell}{\ell!} f_X^{(\ell)}(x + \tau uh).$$

Thus, recalling that $\int_{\mathbb{R}} u^j K(u) du = 0$ for $j \in \{1, \dots, \ell\}$ (we use it in the second and the third step),

$$\begin{aligned} b(x) &= \int_{\mathbb{R}} K(u) \left(uh f'_X(x) + \dots + \frac{(uh)^{\ell-1}}{(\ell-1)!} f_X^{(\ell-1)}(x) + \frac{(uh)^\ell}{\ell!} f_X^{(\ell)}(x + \tau uh) \right) du \\ &= \int_{\mathbb{R}} K(u) \frac{(uh)^\ell}{\ell!} f_X^{(\ell)}(x + \tau uh) du \\ &= \int_{\mathbb{R}} K(u) \frac{(uh)^\ell}{\ell!} \left(f_X^{(\ell)}(x + \tau uh) - f_X^{(\ell)}(x) \right) du. \end{aligned}$$

Taking absolute values, using the Hölder property $f_X \in \mathcal{P}(\beta, L)$, and recalling finally $0 \leq \tau \leq 1$,

$$\begin{aligned}
|b(x)| &= \left| \int_{\mathbb{R}} K(u) \frac{(uh)^\ell}{\ell!} \left(f_X^{(\ell)}(x + \tau uh) - f_X^{(\ell)}(x) \right) du \right| \\
&\leq \int_{\mathbb{R}} |K(u)| \frac{|uh|^\ell}{\ell!} \left| f_X^{(\ell)}(x + \tau uh) - f_X^{(\ell)}(x) \right| du \\
&\leq \int_{\mathbb{R}} |K(u)| \frac{|uh|^\ell}{\ell!} L |\tau uh|^{\beta-\ell} du \\
&= \int_{\mathbb{R}} |K(u)| \frac{L |uh|^\beta}{\ell!} |\tau|^{\beta-\ell} du \\
&\leq \int_{\mathbb{R}} |K(u)| \frac{L |uh|^\beta}{\ell!} du \\
&= \frac{Lh^\beta}{\ell!} \int_{\mathbb{R}} |K(u)| |u|^\beta du.
\end{aligned}$$

This shows the claim. □

Remark. Note that the expectation

$$\mathbb{E} \left[\hat{f}_X(x) \right] = \frac{1}{h} \int_{\mathbb{R}} K \left(\frac{y-x}{h} \right) f_X(y) dy$$

is the **convolution** $\frac{1}{h} K \left(\frac{\cdot}{h} \right) * f_X$.

In general, the convolution of two integrable functions $f, g : \mathbb{R} \rightarrow \mathbb{R}$ is defined as

$$(f * g)(x) = \int_{\mathbb{R}} f(x-y) g(y) dy.$$

One interpretation of the convolution is the following: if f_X, f_Y are the densities of independent random variables X, Y , then the density of $X + Y$ is $f_X * f_Y$.

Indeed, let φ be bounded and continuous. Then, using independence and writing $u = x + y$, and using Fubini-Tonelli,

$$\begin{aligned}
\mathbb{E}[\varphi(X+Y)] &= \int_{\mathbb{R} \times \mathbb{R}} \varphi(x+y) f_{X,Y}(x,y) dx dy \\
&= \int_{\mathbb{R}} \int_{\mathbb{R}} \varphi(x+y) f_X(x) f_Y(y) dx dy \\
&= \int_{\mathbb{R}} \int_{\mathbb{R}} \varphi(u) f_X(y-u) f_Y(y) du dy \\
&= \int_{\mathbb{R}} \int_{\mathbb{R}} \varphi(u) f_X(y-u) f_Y(y) dy du \\
&= \int_{\mathbb{R}} \varphi(u) \int_{\mathbb{R}} f_X(y-u) f_Y(y) dy du \\
&= \int_{\mathbb{R}} \varphi(u) f_X * f_Y(u) du.
\end{aligned}$$

This characterises the density uniquely.

Another way to see this is to consider the characteristic function, which is the Fourier transform of the random variable, using independence:

$$\mathbb{E} [e^{it(X+Y)}] = \mathbb{E} [e^{itX} e^{itY}] = \mathbb{E} [e^{itX}] \mathbb{E} [e^{itY}].$$

The latter is the product of the characteristic functions of X and Y . The very same expression as on the right-hand side is yielded taking the characteristic function of a random variable with density $f_X * f_Y$, and the characteristic function characterises the distribution uniquely.

Result Combining proposition 2.2 and 2.3, we see

$$\text{MSE}(x) \leq C_1^2 h^{2\beta} + \frac{C}{nh}.$$

Minimizing the right-hand side in h yields $h_{\text{opt}} = \left(\frac{C}{2\beta C_1^2 n} \right)^{\frac{1}{2\beta+1}} \sim n^{-\frac{1}{2\beta+1}}$.

Plugging this back into the right-hand side, we obtain

$$\text{MSE}(x) = O\left(n^{-\frac{2\beta}{2\beta+1}}\right).$$

3 MISE and Cross validation

3.1 MISE

To define the MISE, we would like

$$\mathbb{E} \left[\left\| \hat{f}_X - f_X \right\|_2^2 \right] < \infty.$$

We assume $f_X \in L^2(\mathbb{R})$. We would like as well $\hat{f}_X \in L^2(\mathbb{R})$. This is true if $K \in L^2(\mathbb{R})$.

Indeed,

$$\begin{aligned} \left\| \hat{f}_X \right\|_2^2 &\leq \frac{2^{n-1}}{(nh)^2} \sum_{i=1}^n \int_{\mathbb{R}} K\left(\frac{X_i - x}{h}\right)^2 dx \\ &\leq \frac{2^{n-1}}{nh} \int_{\mathbb{R}} K^2(u) du < \infty. \end{aligned}$$

The idea behind this inequality is $(a+b)^2 \leq 2(a^2+b^2)$, and then by induction, $(\sum_{i=1}^n a_i)^2 \leq 2^{n-1} \sum_{i=1}^n a_i^2$.

3.2 Cross validation

Let us write

$$\begin{aligned} \text{MISE}(h) &= \mathbb{E} \left[\int_{\mathbb{R}} \left(\hat{f}_X^h(x) - f_X(x) \right)^2 dx \right] \\ &= \mathbb{E} \left[\underbrace{\int_{\mathbb{R}} \left(\hat{f}_X^h(x) \right)^2 dx - 2 \int_{\mathbb{R}} \hat{f}_X^h(x) f_X(x) dx}_{=I(h)} \right] + \int_{\mathbb{R}} f_X^2(x) dx \end{aligned}$$

introduced

$$\widehat{\text{CV}}(h) = \int_{\mathbb{R}} \hat{f}_X^2(x) dx - \underbrace{\frac{2}{n} \sum_{i=1}^n \hat{f}_{X,-i}(X_i)}_{=\hat{A}},$$

where $\hat{f}_{X,-i} = \frac{1}{(n-1)h} \sum_{j=1, j \neq i}^n K\left(\frac{X_j - x}{h}\right)$. The cross-validated bandwidth is

$$\hat{h}_{\text{CV}} = \arg \min_{h>0} \widehat{\text{CV}}(h)$$

We claim

$$\frac{1}{2} \mathbb{E} [\hat{A}] = \mathbb{E} \left[\int_{\mathbb{R}} \hat{f}_X(x) f_X(x) dx \right]$$

We have

$$\begin{aligned} \mathbb{E} [\hat{f}_{X,-1}(X_1)] &= \mathbb{E}_{\mathbb{P}_{X_2} \otimes \mathbb{P}_{X_n}} \left[\int_{\mathbb{R}} \hat{f}_{X,-1}(x) f_X(x) dx \right] \\ &= \mathbb{E} \left[\frac{1}{(n-1)h} \sum_{j=2}^n \int_{\mathbb{R}} K\left(\frac{X_j - x}{h}\right) f_X(x) dx \right] \\ &= \frac{1}{h} \int_{\mathbb{R}} \int_{\mathbb{R}} K\left(\frac{z - x}{h}\right) f_X(z) f_X(x) dz dx \end{aligned}$$

As an exercise, show that this yields the claim.

Theorem 3.1 (Oracle inequality). *Let f_{\max} be such that for all x , $f_X(x) \leq f_{\max} < \infty$. Assume the kernel K is such that $\int_{\mathbb{R}} K^2(u) du < \infty$. $\mathcal{F}[K] \geq 0$ and $\text{supp}(\mathcal{F}[K]) \subseteq [-1, 1]$. Then $\hat{f}_X^* = \hat{f}_X^{h_{\text{CV}}}$ is such that for all $0 < \delta < 1$, for all $n \geq 1$,*

$$\mathbb{E} \left[\int_{\mathbb{R}} \left(\hat{f}_X^*(x) - f_X(x) \right)^2 dx \right] \leq \left(1 + \frac{C}{n^\delta} \right) \min_{h > \frac{1}{n}} \mathbb{E} \left[\int_{\mathbb{R}} \left(\hat{f}_X^h(x) - f_X(x) \right)^2 dx \right] + \frac{C(\log n)^{\frac{\delta}{2}}}{n^{1-\delta}}$$

Remark. The cross-validation bandwidth from theorem 3.1 is random. The kind of inequality in the the theorem is called **oracle inequality**, as it is not possible to obtain the values on each side. They involve the unknown $f_X(x)$. The estimation of errors in cross-validation kernel estimation is hard, but in practice it often works well.

4 Sobolev class and symmetric kernel

4.1 Review of Fourier transform

Definition 4.1. The characteristic function of a random variable X is

$$\varphi_X(w) = \mathbb{E}[e^{iwX}] = \int_{\mathbb{R}} e^{iwX} f_X(x) dx.$$

Remark. It is possible as well to define the Fourier transform of $f \in L^2(\mathbb{R})$. Therefore, we take a sequence $f_m \in L^1(\mathbb{R}) \cap L^2(\mathbb{R})$ such that $\|f - f_m\|_2^2 \rightarrow 0$ as $m \rightarrow \infty$ and define the Fourier transform of f as the L^2 -limit of $\mathcal{F}[f_m]$. More precisely, we may take $f_m(x) = f(x) \mathbb{1}_{|x| \leq m}$. It is in L^2 as the product of an $L^2(\mathbb{R})$ function and a bounded function, and it is in $L^1(\mathbb{R})$ as a result of the Cauchy-Schwarz inequality:

$$\int_{\mathbb{R}} f(x) \mathbb{1}_{|x| \leq m} dx \leq \sqrt{\int_{\mathbb{R}} f(x)^2 dx} \sqrt{\int_{-m}^m 1 dx} = \sqrt{2m} \sqrt{\underbrace{\int_{\mathbb{R}} f(x)^2 dx}_{< \infty}}.$$

Moreover,

$$\|f_m - f\|_2^2 = \int_{-\infty}^m |f(x)|^2 dx + \int_m^{\infty} |f(x)|^2 dx \rightarrow 0 \quad (2)$$

as $m \rightarrow \infty$. By equation 2, for all m, m' , $\|f_m - f_{m'}\|_2^2 \rightarrow 0$ as $m, m' \rightarrow \infty$, i.e. (f_m) is a Cauchy sequence. By Plancherel's theorem 1.1,

$$\|\mathcal{F}[f_m] - \mathcal{F}[f_{m'}]\|_2^2 = \|\mathcal{F}[f_m - f_{m'}]\|_2^2 = 2\pi \|f_m - f_{m'}\|_2^2.$$

Thus, $\mathcal{F}[f_m]$ is a Cauchy sequence in $L^2(\mathbb{R})$, so that it admits a limit in $L^2(\mathbb{R})$, since $L^2(\mathbb{R})$ is a complete normed space. We can then define the Fourier transform of f to be this limit.

We notice that the characteristic function of a random variable is the Fourier transform of its density. Therefore, the density function f_X of a random variable X is the inverse Fourier transform of its characteristic function φ_X .

$$f_X(x) = \frac{1}{2\pi} \int_{\mathbb{R}^d} e^{-iwx} \varphi_X(w) dw.$$

We can get the derivative of the density function by differentiating the Fourier transform of the density function.

$$f_X^{(m)}(x) = \frac{1}{2\pi^d} \int_{\mathbb{R}^d} (-iw)^m e^{-iwx} \varphi_X(w) dw.$$

It means that $\mathcal{F}\left[f_X^{(m)}\right](w) = (-iw)^m \varphi_X(w)$.

4.2 Sobolev class

Building on this, we make the following definition.

Definition 4.2 (Sobolev class). Let $\beta > 0$, $L > 0$, the Sobolev class $\mathcal{P}_S(\beta, L)$ is defined as

$$\mathcal{P}_S(\beta, L) = \left\{ f : f \text{ is a density on } \mathbb{R} \text{ and } \int_{\mathbb{R}} |w|^{2\beta} |\mathcal{F}[f](w)|^2 dw \leq 2\pi L^2 \right\}.$$

The restriction is basically saying that the L^2 norm of the function $f_X^{(m)}$ is bounded by L . This is a generalization of the L^2 norm of the function $f_X(x)$, which is the L^2 norm of the density function f .

4.3 Symmetric kernel

Theorem 4.1 (Symmetric kernel). Let $f_X \in L^2(\mathbb{R})$, $K \in L^2(\mathbb{R})$ be a symmetric kernel such that

$$\sup_{w \in \mathbb{R} \setminus \{0\}} \frac{|1 - \mathcal{F}[K](w)|}{|w|^{\beta'}} \leq A < \infty$$

for some $\beta', A > 0$. Then

$$\sup_{f_X \in \mathcal{P}_S(\beta, L)} \mathbb{E} \left[\left\| \hat{f}_X - f_X \right\|_2^2 \right] \leq C n^{-\frac{2\tilde{\beta}}{2\tilde{\beta}+1}},$$

where $\tilde{\beta} = \min\{\beta, \beta'\}$, if $h = \alpha n^{-\frac{1}{2\tilde{\beta}+1}}$ for some $\alpha > 0$ and C is a constant which only depends on L, α, A, K .

Example 4.1. 1. Gaussian kernel: $K(u) = \frac{1}{\sqrt{2\pi}} e^{-\frac{u^2}{2}}$, $\mathcal{F}[K](u) = e^{-\frac{u^2}{2}}$. We have

$$\frac{|1 - e^{-w^2/2}|}{|w|^{\beta'}} \leq \begin{cases} |w|^{-\beta'}, & |w| \geq 1 \\ \frac{w^2/2}{|w|^{\beta'}}, & |w| \leq 1 \end{cases}$$

so 4.1 holds if $\beta' \leq 2$, else the sup is ∞ .

2. The sinc kernel: $K(u) = \frac{\sin(u)}{\pi u}$, $\mathcal{F}[K](w) = \mathbb{1}_{|w| \leq 1}$. We have

$$\frac{|1 - \mathcal{F}[K](w)|}{|w|^{\beta'}} \leq \begin{cases} |w|^{-\beta'}, & |w| > 1 \\ 0, & |w| \leq 1, \end{cases}$$

so 4.1 holds for all β' . Such a kernel is called an **infinite power kernel** or **superkernel**.

3. Trapeze kernel: Let

$$\mathcal{F}[K](w) = \begin{cases} 0, & |w| > a \\ 1, & |w| \leq b \\ \text{linear,} & \text{otherwise,} \end{cases}$$

a trapeze. Then 4.1 holds for all β' . Let us write K_2 for the trapeze (in Fourier space) and K_1 for the sinc Kernel (see 4.1). Then

$$K_2 = \frac{1}{2\pi} \mathcal{F}[\mathcal{F}[K_1] * F[K_1]] = \frac{1}{2\pi} \mathcal{F}[\mathcal{F}[K_1]] \mathcal{F}[\mathcal{F}[K_1]] = 2\pi K_1^2(u) = 2\pi \left(\frac{\sin u}{\pi u} \right),$$

which is in $L^1(\mathbb{R}) \cap L^2(\mathbb{R})$.

Optimal rate of convergence It can be shown that the *sinc* kernel has the optimal rate of convergence.

A Corollary of the Theorem 4.1 that we have seen for cross-validation is

Corollary 4.2. *Let K be the sinc kernel, then*

$$\sup_{f_X \in \mathcal{P}_S(\beta, L)} \mathbb{E} \left[\left\| \hat{f}_X^{\text{CV}} - f_X \right\|_2^2 \right] \leq C n^{-\frac{2\beta}{2\beta+1}}$$

for all $\beta > \frac{1}{2}, L > 0$, where C only depends on β and L .

Some people have shown:

Proposition 4.1.

$$\inf_{\hat{f}} \sup_{f_X \in \mathcal{P}_S(\beta, L)} \mathbb{E} \left[\left\| \hat{f}_X - f_X \right\|_2^2 \right] \geq C_* n^{-\frac{2\beta}{2\beta+1}}$$

for some absolute constant C_* .

This means that $n^{-\frac{2\beta}{2\beta+1}}$ is the “minimax” optimal rate of convergence and the cross-validated estimator is minimax adaptive (i.e. we can construct it with the data only).

Kernel comparison We end this section by the following table.

name	kernel	$\mathcal{F}[K]$	$\frac{ 1-\mathcal{F}[K](w) }{ w ^\beta}$
Gaussian			
Epanechnikov			
Sinc			
Trapeze			

Table 1: Summary

4.4 Extension

Remark. The condition 4.1 is satisfied for an integer β if K is a kernel of order $\beta - 1$ and $\int |u|^\beta |K(u)| < \infty$.

Remark. We can work with a smaller class of *super smooth* density functions.

1. $\mathcal{P}_{\alpha,r} = \{f \in L^2(\mathbb{R}) \text{ such that } \int \exp(\alpha |w|^2) |\phi(w)|^2 dw \leq L^2\}$ where $\phi = \mathcal{F}[f]$ is the Fourier transform of f . We can show that a MISE optimal kernel density estimation could have a risk less than $C \frac{(\log n)^{1/r}}{n}$.
2. $\mathcal{P}_{\alpha,r} = \{f \in L^2(\mathbb{R}) \text{ such that } \text{supp}(\phi) \subset [-a, a]\}$. In this case, the upper bound is $\frac{a\pi}{n}$.

5 Other types of non-parametric estimators

5.1 Orthogonal series estimators

¹ Let $f_X \in L^2([0, 1]^d)$, where $L^2([0, 1]^d)$ can be proven to be a *separable Hilbert space* when endowed with the inner product

$$\langle f, g \rangle = \int_{\mathbb{R}} f(x)g(x)dx.$$

We write

$$\|f\|_2^2 = \langle f, f \rangle.$$

Some properties are comparable to \mathbb{R}^d with $\langle x, y \rangle = x^T y$. As a separable space, $L^2([0, 1]^d)$ has a countable basis $(e_j)_{j=1}^\infty$, which is a sequence of functions in $L^2([0, 1]^d)$ such that for all

$$\langle e_j, e_k \rangle = \delta_{jk} = \begin{cases} 1, & j = k, \\ 0, & \text{else,} \end{cases}$$

and for all $f \in L^2([0, 1]^d)$,

$$f = \lim_{k \rightarrow \infty} \sum_{j=1}^k \langle f, e_j \rangle e_j.$$

Think of \mathbb{R}^d , where $(e_j)_{j=1}^d$ is a basis for $e_j = (0, \dots, 0, 1, 0, \dots, 0)$ the j -th unit vector. Then $\langle e_j, e_k \rangle = e_j^T e_k = \delta_{jk}$, and for $x \in \mathbb{R}^d$,

$$x = \sum_{j=1}^d x_j e_j = \sum_{j=1}^d x^T e_j e_j = \sum_{j=1}^d \langle x, e_j \rangle e_j.$$

¹Generalizations are called sieves (in Econometrics) or dictionaries in machine-learning.

Given $(e_j)_{j=1}^\infty$ a basis, for all $f \in L^2([0, 1]^d)$,

$$\|f\|_2^2 = \sum_{j=1}^{\infty} \langle f, e_j \rangle^2.$$

This is a version of the Pythagorean theorem. In \mathbb{R}^d ,

$$\|x\|_2^2 = \sum_{j=1}^d x_j^2 = \sum_{j=1}^d \langle x, e_j \rangle^2.$$

Back to our goal to estimate $f_X = \lim_{k \rightarrow \infty} \sum_{j=1}^{\infty} \langle f, e_j \rangle e_j$. For some $T \in \mathbb{N}$, consider $f_X^T \stackrel{\text{def}}{=} \sum_{j=1}^T \langle f, e_j \rangle e_j$. The idea is to estimate this cut-off sum instead of the limit expression for f_X . We have

$$c_j \stackrel{\text{def}}{=} \langle f_X, e_j \rangle = \int_{[0,1]^d} f_X(x) e_j(x) dx = \mathbb{E}[e_j(X)],$$

so that an unbiased estimator is

$$\hat{c}_j = \frac{1}{n} \sum_{i=1}^n e_j(X_i).$$

Thus, a candidate estimator for f_X is

$$\hat{f}_X^T = \sum_{j=1}^T \hat{c}_j e_j,$$

where

$$\mathbb{E}[\hat{f}_X^T] = \sum_{j=1}^T c_j e_j = f_X^T.$$

It is possible to write

$$\hat{f}_X^T = \frac{1}{n} \sum_{i=1}^n \underbrace{\sum_{j=1}^T e_j(X_i) e_j(x)}_{q_T(X_i, x)},$$

where $q_T(X_i, x)$ plays the role of a kernel and T plays the same role as $\frac{1}{h}$. On $L^2([0, 1]^d)$ we can use bases for which $e_j = f_{j_1} \cdots f_{j_d}$ where $(f_k)_{k=1}^\infty$ is a basis of $L^2([0, 1])$ and (j_1, \dots, j_d) plays the role of the index j^2 . For example, $f_k(x) = \sqrt{2} \sin(\pi k x)$ is a basis of $L^2([0, 1])$. This gives

$$e_{j_1, \dots, j_d}(x) = 2^{\frac{d}{2}} \prod_{k=1}^d \sin(\pi j_k x_k).$$

²Note that there exists a bijection $\mathbb{N}^d \rightarrow \mathbb{N}$.

One can check that this defines an orthogonal system (**Exercise**).

We define

$$W(\beta, L) = \left\{ f : [0, 1]^d \rightarrow \mathbb{R} \text{ with coefficients } c_{j_1, \dots, j_d} \text{ w.r.t. } (f_{j_1}, \dots, f_{j_d}) \right. \\ \left. \text{such that } \sum_{j_1=1}^{\infty} \sum_{j_2=1}^{\infty} \cdots \sum_{j_d=1}^{\infty} c_{j_1, \dots, j_d}^2 (j_1^2 + \dots + j_d^2)^\beta \leq L^2 \right\}$$

Remark. The $\|j\|^{2\beta}$ is present due to the fact that we take derivatives of our basis functions defined above till the order of β .

In $L^2(\mathbb{R}^d)$, an analogous condition (with Fourier transform instead of Fourier series) would be:

$$\int_{\mathbb{R}^d} |\mathcal{F}[f](w_1, \dots, w_d)|^2 (|w_1|^2 + \dots + |w_d|^2)^\beta dw \leq L^2.$$

Here, $\mathcal{F}[f](w_1, \dots, w_d) = \int_{\mathbb{R}} e^{iwx} f_X(x) dx$ acts like the coefficients $c_{j_1, \dots, j_d} = \int_{[0,1]^d} f_X(x) e_{j_1, \dots, j_d}(x) dx$ in the Fourier series case. Note the usual bias-variance decomposition of the mean-squared error,

$$\mathbb{E} \left[\left\| \hat{f}_X^T - f_X \right\|_2^2 \right] = \underbrace{\left\| f_X^T - f_X \right\|_2^2}_{b^2 = \text{Bias}^2} + \underbrace{\mathbb{E} \left[\left\| \hat{f}_X^T - f_X^T \right\|_2^2 \right]}_{\sigma^2}.$$

Then

$$\begin{aligned} b^2 &= \sum_{j_1=T+1}^{\infty} \cdots \sum_{j_d=T+1}^{\infty} c_{j_1, \dots, j_d}^2 \\ &\leq \sum_{j_1=T+1}^{\infty} \cdots \sum_{j_d=T+1}^{\infty} c_{j_1, \dots, j_d}^2 \left(\left(\frac{j_1}{T+1} \right)^2 + \cdots + \left(\frac{j_d}{T+1} \right)^2 \right)^\beta \\ &= \left(\frac{1}{T+1} \right)^{2\beta} \sum_{j_1=T+1}^{\infty} \cdots \sum_{j_d=T+1}^{\infty} c_{j_1, \dots, j_d}^2 (j_1^2 + \cdots + j_d^2)^\beta \\ &\leq \left(\frac{1}{T+1} \right)^{2\beta} L^2. \end{aligned}$$

Note that $\|f_{j_1} \cdots f_{j_d}\|_2^2 = \|f_{j_1}\|_2^2 \cdots \|f_{j_d}\|_2^2$, which are all = 1. Then,

$$\begin{aligned}
\sigma^2 &= \mathbb{E} \left[\left\| \hat{f}_X^T - f_X^T \right\|_2^2 \right] \\
&= \mathbb{E} \left[\sum_{j_1=1}^T \cdots \sum_{j_d=1}^T (\hat{c}_{j_1, \dots, j_d} - c_{j_1, \dots, j_d})^2 \|f_{j_1} \cdots f_{j_d}\|_2^2 \right] \\
&= \mathbb{E} \left[\sum_{j_1=1}^T \cdots \sum_{j_d=1}^T (\hat{c}_{j_1, \dots, j_d} - c_{j_1, \dots, j_d})^2 \right] \\
&= \sum_{j_1=1}^T \cdots \sum_{j_d=1}^T \text{Var}(\hat{c}_{j_1, \dots, j_d}) \\
&\leq \sum_{j_1=1}^T \cdots \sum_{j_d=1}^T \frac{2^d}{n} \\
&\leq \frac{(2(T+1))^d}{n}
\end{aligned}$$

since

$$\begin{aligned}
\text{Var}(\hat{c}_{j_1, \dots, j_d}) &= \frac{1}{n^2} \sum_{i=1}^n \text{Var}(f_{j_1} \cdots f_{j_d}(X_i)) \\
&\leq \frac{1}{n^2} \sum_{i=1}^n \mathbb{E}[f_{j_1}^2 \cdots f_{j_d}^2(X_i)] \\
&\leq \frac{2^d}{n},
\end{aligned}$$

where we used $f_{j_k}^2 \leq 2$ for our particular basis $f_{j_k}(x) = \sqrt{2} \sin(\pi j_k x)$. Remember $\sigma^2 \leq \frac{1}{nh^d}$ for kernel estimators. This gives an upper bound of the order of $n^{-\frac{2\beta}{2\beta+d}}$ if $T = \left\lfloor n^{\frac{1}{2\beta+d}} \right\rfloor$.

Remark. (Nonexaminable content)

- We can work with families of functions which may not be basis functions. We talk about series, dictionaries (machine learning).
- In the previous upper bound, the choice of T is infeasible because it depends on β which is unknown.
- It is classical to estimate many coefficients c_j , for T much larger than before (e.g. \sqrt{n}) and work with the estimators

$$\hat{f}_X^T(x) = \sum_{j_1=1}^T \cdots \sum_{j_d=1}^T \hat{\tau}(c_{j_1, \dots, j_d}) e_{j_1} \cdots e_{j_d}(x).$$

where $\tau \propto \frac{\sqrt{\log n}}{n}$. For example

- $\tau_\rho(x) = \mathbb{1}_{\{|x| \geq \rho\}}$, where ρ is a thresholding function. This is the **hard** thresholding function.
- $\tau_\rho(x) = x \max\left(1 - \frac{\rho}{|x|}, 0\right)$. This is the **soft** thresholding function.

6 Regression Function Estimation

6.1 Introduction: average effect of X on Y

The model for a nonparametric model is

$$Y = f(X) + \varepsilon,$$

where $\mathbb{E}[\varepsilon|X] = 0$ and $\mathbb{E}[|\varepsilon|] < \infty$. The goal is to estimate f . We say it has a random design if X is random, and a fixed design if X is fixed. We will focus on the random design case.

First, we define the average effect of X on Y as $\mathbb{E}[f(X)]$ if the expectation is defined.

If $f_{y|x}$, the conditional density of Y given X exists, it is given by $f_{Y|X}(y|x) = \frac{f(x,y)}{f_X(x)}$ if $f_X(x) > 0$.

Also the conditional expectation function $\mathbb{E}[Y|X=x]$ is given by

$$\mathbb{E}[Y|X=x] = \int y f_{Y|X}(y|x) dy = \frac{\int y f(x,y) dy}{f_X(x)} = \frac{\int y f(x,y) dy}{\int f(x,y) dy}.$$

A natural idea would be to use

$$\hat{f}_{Y,X}(y,x) = \frac{1}{nh^2} \sum_{i=1}^n K\left(\frac{X_i - x}{h}\right) K\left(\frac{Y_i - y}{h}\right),$$

where K is a kernel.

As an exercise, we can check that

$$\int y \hat{f}_{Y,X}(y,x) dy = \frac{1}{nh} \sum_{i=1}^n K\left(\frac{X_i - x}{h}\right) Y_i$$

and

$$\int \hat{f}_{Y,X}(y,x) dy = \frac{1}{nh} \sum_{i=1}^n K\left(\frac{X_i - x}{h}\right).$$

6.2 Nadaraya-Watson estimator

This leads to the following estimator called **Nadaraya-Watson estimator**

$$\hat{f}_X(x) = \frac{\sum_{i=1}^n K\left(\frac{X_i - x}{h}\right) Y_i}{\sum_{i=1}^n K\left(\frac{X_i - x}{h}\right)}.$$

In practice, dealing with the denominator can be tricky. We propose two ideas to deal with this issue.

1. We can work with **nonnegative** kernels because

$$\sum Y_i \underbrace{\frac{K\left(\frac{X_i-x}{h}\right)}{\sum_{i=1}^n K\left(\frac{X_i-x}{h}\right)}}_{\in [0,1]}.$$

2. We can use a trimming factor ρ and write

$$\hat{f}_X(x) = \frac{\sum_{i=1}^n K\left(\frac{X_i-x}{h}\right) Y_i}{\max\left(\sum_{i=1}^n K\left(\frac{X_i-x}{h}\right), \rho\right)}.$$

Suppose now $\text{supp}(X) = [a, b]$ and $\exists m > 0$ s.t. $f_X(x) \geq m$. Suppose I am interested in $f(b)$ and I use the rectangular kernel 1.

Then $\hat{f}(b)$ where \hat{f} is the N.W. estimator is biased. But a local polynomial estimator of order ≥ 1 is consistent and unbiased.

Remark. In TD, we will see that we can get a fast rate of convergence with nonnegative kernels (unlike in density estimation).

6.3 Local Polynomial Estimation

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7 Treatment Effects

7.1 Setup

We have a dataset $[Y_i, D_i, X_i, Z_i, W_i]_{i=1}^n$ following i.i.d. from a joint distribution.

- D is a binary treatment variable, $D \in [0, 1]$.
- Y is the outcome variable. Here Y is a random variable $Y \in \mathbb{R}$.
- X, Z, W are covariates/additional random variables.

The model equation (for each individual i) is $y_i = y(0)(1-D) + y(1)D$. The potential outcome is $y_i(1), y_i(0)$, which are not observed. We can only observe $y_i = y(D_i)$.

7.2 Parameters of Interest

- Average treatment effect (ATE): $\tau = \mathbb{E}[Y(1) - Y(0)]$
- Conditional average treatment effect (CATE): $\tau(x) = \mathbb{E}[Y(1) - Y(0) | X = x]$. It can be useful if we care about the effect of the treatment on a specific subgroup of the population.

- Average treatment effect on the treated (ATT): $\tau_{\text{ATT}} = \mathbb{E}[Y(1) - Y(0) | D = 1]$
- Average treatment effect on the untreated (ATU): $\tau_{\text{ATU}} = \mathbb{E}[Y(1) - Y(0) | D = 0]$
- Conditional average treatment effect on the treated (CATT): $\tau_{\text{ATT}}(x) = \mathbb{E}[Y(1) - Y(0) | D = 1, X = x]$

7.3 Identification

We need to impose some assumptions in order to identify the parameters.

1. $\mathbb{P}(D = 1) \in (0, 1)$
2. The covariates X, Z, W are such that if $X = X(0) + D(X(1) - X(0))$, then $X(1) = X(0)$.
3. The potential outcome $Y(1), Y(0)$ are independent of D which is $Y(1), Y(0) \perp\!\!\!\perp D$
4. The potential outcome $Y(1), Y(0)$ and X are independent of D , that is

$$Y(1), Y(0), X \perp\!\!\!\perp D$$

We introduce a new notation for the purpose of another assumption.

Definition 7.1 (Propensity score). The propensity score is defined as the conditional probability of receiving the treatment given the covariates, that is

$$\pi(x) = \mathbb{P}(D = 1 | X = x)$$

Remark. Later we will build estimators using propensity score, called **inverse propensity score weighting (IPSW) estimator**.

Common support The propensity score $\pi(x)$ continuous and bounded between 0 and 1 for all $x \in \text{supp}(X)$

Sometimes, treatment D is assigned randomly conditional on X . We introduce the following assumptions

Unconfoundedness The treatment D is unconfounded with the potential outcome Y given X if

$$Y(1), Y(0) \perp\!\!\!\perp D | X$$

Conditional mean independence The potential outcome $Y(1), Y(0)$ are independent of D given X, Z, W , that is

$$\mathbb{E}[Y(d) | D, X] = \mathbb{E}[Y(d) | X]$$

Later we will show that unconfoundedness implies conditional mean independence.

7.4 Regression discontinuity design

7.4.1 Sharp RDD

Preliminaries Previously, we consider D as a binary variable and impose certain conditions on whether $D_i = 1$ or $D_i = 0$ (for each individual). Now we specify how D is determined by a continuous variable X , that is

$$D_i = \mathbb{1} \{X_i \geq c\}$$

where c is a known threshold. The idea is that the treatment is assigned based on the value of X . We can think of X as a score, and D is assigned to those who score above a certain threshold.

For example, in the context of education, X can be the score of a student in a standardized test, and D is whether the student is admitted to a college. The threshold c is the cutoff score for admission.

Conditions Recall the definition of *unconfoundedness*

It is easy to see that since D is determined by X , the potential outcome $Y(1), Y(0)$ are independent of D given X . (Given X , D is already determined, thus a constant.) Therefore, the unconfoundedness assumption is satisfied.

Since X is a continuous variable, we make an assumption on the **average potential outcome** $\mathbb{E}[Y(j) | X = x]$ for $j = 0, 1$. We assume that the average potential outcome is continuous at the threshold c . For example, for those who have a test score slightly above and below the threshold, the average potential earning $Y(1), Y(0)$ is similar.

Remark. We assume that $\mathbb{E}[Y(j) | X = x]$ is continuous at the threshold c but not $\mathbb{E}[Y | X = x]$ at c .

Let us now check conditional ATE at the point c . Previously, we define the conditional ATE as

$$\begin{aligned} \tau(x) &= \mathbb{E}[Y(1) - Y(0) | X = x] \\ &= \mathbb{E}[Y(1) | X = x] - \mathbb{E}[Y(0) | X = x] \\ &= \underbrace{\mathbb{E}[Y | D = 1, X = x] - \mathbb{E}[Y | D = 0, X = x]}_{\text{by unconfoundedness thus mean independence}} \end{aligned}$$

Therefore,

$$\begin{aligned} \tau(c) &= \lim_{x \rightarrow c^+} \mathbb{E}[Y | D = 1, X = x] - \lim_{x \rightarrow c^-} \mathbb{E}[Y | D = 0, X = x] \\ &= \lim_{x \rightarrow c^+} \mathbb{E}[Y | X = x] - \lim_{x \rightarrow c^-} \mathbb{E}[Y | X = x] \end{aligned} \tag{3}$$

Estimation To estimate $\mathbb{E}[Y | X = x]$, we can use local polynomial of order 0 (Nadaraya-Watson estimator) or more:

$$\begin{aligned} (\hat{\alpha}_1, \hat{\beta}_1) &= \arg \min_{\alpha, \beta} \sum_{i=1, X_i \geq c}^n K\left(\frac{X_i - x}{h}\right) (Y_i - \alpha - \beta(X_i - c))^2 \\ (\hat{\alpha}_0, \hat{\beta}_0) &= \arg \min_{\alpha, \beta} \sum_{i=1, X_i < c}^n K\left(\frac{X_i - x}{h}\right) (Y_i - \alpha - \beta(X_i - c))^2 \end{aligned}$$

Then we estimate $\tau(c)$ by $\hat{\tau}(c) = \hat{\alpha}_1 - \hat{\alpha}_0$.

7.4.2 Fuzzy RDD

Preliminaries First, let's recall the definition of conditional mean independence:

Definition 7.2 (Conditional mean independence). We say U and V are conditionally mean independent given X if

$$\mathbb{E}[\phi(U)\psi(V) | X] = \mathbb{E}[\phi(U) | X] \mathbb{E}[\psi(V) | X]$$

Naturally local conditional mean independence in a neighborhood \mathcal{N} is defined as

$$\mathbb{E}[\phi(U)\psi(V) | X = x] = \mathbb{E}[\phi(U) | X = x] \mathbb{E}[\psi(V) | X = x]$$

for almost every $x \in \mathcal{N}$.

Condition Now we are ready to move from *sharp RDD* to *fuzzy RDD*. In the fuzzy RDD, the treatment D is not exactly determined by X but instead satisfies the following condition to create a discontinuity at c : The propensity score function $\pi(x)$ is continuous on $(c - \epsilon, c)$ and $(c, c + \epsilon)$ for some $\epsilon > 0$ and $\lim_{x \rightarrow c^+} \pi(x) \neq \lim_{x \rightarrow c^-} \pi(x)$. We also loosen the mean independence condition to local conditional mean independence in a neighborhood \mathcal{N} of c .

Remark. The fuzzy RDD is more general than the sharp RDD.

Since we depart from sharp RDD, the conditional ATE at c is now defined as the following:

Proposition 7.1.

$$\tau(c) = \frac{\lim_{x \rightarrow c^+} \mathbb{E}[Y | D = 1, X = x] - \lim_{x \rightarrow c^-} \mathbb{E}[Y | D = 0, X = x]}{\lim_{x \rightarrow c^+} \pi(x) - \lim_{x \rightarrow c^-} \pi(x)}$$

Proof.

□

Estimation We can make use of local polynomial estimator to estimate the denominator and the numerator separately for $\tau(c)$.

7.5 Instrumental variable

In this section, we are in the case of selection on *unobservables*. We have binary treatment D , covariates X , and a binary instrument Z , such that treatment is assigned based on Z (and maybe X). But the assigned treatment may not be taken by the individual (imperfect compliance). We have the following model:

$$Y = Y(0, 0) + Z(Y(1, 0) - Y(0, 0)) + D(Y(0, 1) - Y(0, 0)) + DZ(Y(1, 1) - Y(0, 1) - Y(1, 0) + Y(0, 0)) \quad (4)$$

and

$$D = D(0) + Z(D(1) - D(0)) \quad (5)$$

Preliminaries We define the following:

- One-sided compliance: $P(D(0) = 0) = 0$ which means there is no always taker or defiers.
- two-sided compliance: $P(D(0) = 0) \in (0, 1)$ and $P(D(1) = 1) \in (0, 1)$.

Condition We need to impose that the assignment Z is independent of the potential outcome $Y(z, d)$ and the treatment D (given X). From now on we omitted the conditioning on X for simplicity. This is the *exclusion restriction* assumption. It is similar to $\mathbb{E}[\epsilon | X, Z] = 0$ assumption in standard linear IV model.

Definition 7.3 (Local average treatment effect (LATE)). The local average treatment effect is defined as

$$\tau_{\text{LATE}} = \mathbb{E}[Y(Z, 1) - Y(Z, 0) | D(1) - D(0) = 1]$$

which is the average treatment effect for the compliers.

7.5.1 One-sided compliance

Instead of discussing ATE, we define a new set of parameters called *Intention to treat* (ITT).

Definition 7.4 (Intention to treat (ITT)). The intention to treat on treatment is defined as

$$\text{ITT}_D = \mathbb{E}[D(1) - D(0) | Z = 1]$$

The intention to treat on outcome is defined as

$$\text{ITT}_Y = \mathbb{E}[Y(1, D(1)) | Z = 1] - \mathbb{E}[Y(0, D(0)) | Z = 0]$$

The intention to treat on outcome for compliers is defined as

$$\text{ITT}_{Y|D(1)-D(0)=1} = \mathbb{E}[Y(1, 1) - Y(0, 1) | D(1) - D(0) = 1]$$

Condition Under the one-sided compliance, we need to make the following assumption:

- If $D(1) = 0$ (never taker), then $Y(1, 0) = Y(0, 0)$.
- If $D(1) = 1$ (complier), then $Y(0, d) = Y(1, d)$ for $d = 0, 1$.

Because every individual is either a never taker or a complier, we have the following:

$$Y(z, d) = Y(d)$$

This is sometimes called the an *exclusion restriction* assumption.

Then it can be shown that

$$ITT_Y = ITT_{Y,CO} ITT_D$$

Recall that LATE is the average treatment effect for the compliers. And under the condition mentioned above,

$$\begin{aligned}\tau_{LATE} &= \mathbb{E}[Y(Z, 1) - Y(Z, 0) \mid D(1) - D(0) = 1] \\ ITT_{Y|D(1)-D(0)=1} &= \mathbb{E}[Y(1, 1) - Y(0, 0) \mid D(1) - D(0) = 1]\end{aligned}$$

That is,

$$\tau_{LATE} = ITT_{Y|D(1)-D(0)=1}$$

Therefore,

$$\tau_{LATE} = \frac{ITT_Y}{ITT_D} = \frac{\mathbb{E}[Y \mid Z = 1] - \mathbb{E}[Y \mid Z = 0]}{\mathbb{E}[D \mid Z = 1] - \mathbb{E}[D \mid Z = 0]}$$

7.5.2 Two-sided compliance

Condition This is a natural assumption.

- For never takers, $Y(0, 0) = Y(1, 0)$.
- For always takers, $Y(0, 1) = Y(1, 1)$.
- For compliers (and defiers), $Y(1, d) = Y(0, d)$.

We also need to impose the *monotonicity* assumption. It states that the instrument Z has a monotonic effect on the treatment D . That is, $D(1) \geq D(0)$ a.s. or $D(1) \leq D(0)$ a.s.

Under these conditions, we also have

$$\tau_{LATE} = \frac{\mathbb{E}[Y \mid Z = 1] - \mathbb{E}[Y \mid Z = 0]}{\mathbb{E}[D \mid Z = 1] - \mathbb{E}[D \mid Z = 0]}$$

It can be shown that if there is no defiers, then $\mathbb{E}[D \mid Z = 1] - \mathbb{E}[D \mid Z = 0] = \mathbb{P}(D(1) - D(0) = 1)$.

Estimation We can estimate τ_{LATE} by

$$\frac{\mathbb{E}[Y \mid Z = 1] - \mathbb{E}[Y \mid Z = 0]}{\mathbb{E}[D \mid Z = 1] - \mathbb{E}[D \mid Z = 0]} = \frac{Cov(Y, Z)}{Cov(D, Z)}$$

7.6 Estimation methods: (Augmented) Inverse probability Weighting (AIPW)

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