

# Matching

EC 425/525, Set 7

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# Prologue

# Schedule

## Last time

- The conditional independence assumption:  $(Y_{0i}, Y_{1i}) \perp\!\!\!\perp D_i | X_i$
- Omitted variable bias
- Good vs. bad controls

## Today

- Return first round of project proposals.
- Matching estimators (*MHE* 3.2 and Cameron and Trivedi 25.4).

## Upcoming

- Admin: Assignment and midterm
- Next round of the project proposal

# Follow up

## OLS weighting

At the beginning of the lecture, we discussed OLS weights—especially for heterogeneous treatment effects.

We should keep our questions clear.

1. Which weights on  $\beta_1$  and  $\beta_2$  recover  $\beta_{12}$ , where  $\beta_i$  comes from a regression using observations in group  $i$ ?
2. What does  $\beta$  represent when the treatment effect is heterogeneous?

More soon.

# Matching

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## The gist

Remember the **conditional independence assumption<sup>†</sup>** in a setting—i.e., treatment is as-good-as random conditional on a known set of covariates?

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$$\tau(x) = E[Y_{1i} - Y_{0i} | X_i = x]$$

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*The idea:* Estimate a treatment effect only using observations with (nearly?) identical values of  $X_i$ . The CIA buys us causality within these groups.

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## Goals

Let's return to **the fundamental problem of causal inference** for a moment.

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Matching is no different.

We match untreated observations to treated observations using  $X_i$ , i.e., calculate a  $\widehat{Y}_{0i}$  for each  $Y_{1i}$ , based upon "matched" untreated individuals.

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- with  $N_C$  weights in each set:  $w_i(j)$  ( $i = 1, \dots, N_T; j = 1, \dots, N_C$ )

Assume  $\sum_j w_i(j) = 1$ . Our estimate for the counterfactual of treated  $i$  is

$$\widehat{\mathbf{Y}}_{0i} = \sum_{j \in (D=0)} w_i(j) \mathbf{Y}_j$$

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If our estimated counterfactual for treated individual  $i$  is

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then our estimated treatment effect (for individual  $i$ ) is

$$\hat{\tau}_i = Y_{1i} - \widehat{Y}_{0i} = Y_{1i} - \sum_j w_i(j) Y_j$$

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$\therefore$  a generic matching estimator for the treatment effect on the treated is

$$\hat{\tau}_M = \frac{1}{N_T} \sum_{i \in (D=1)} \left( Y_{1i} - \widehat{Y}_{0i} \right) = \frac{1}{N_T} \sum_{i \in (D=1)} \left( Y_{1i} - \sum_{j \in (D=0)} w_i(j) Y_j \right)$$

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So all we need is those weights and we're done.<sup>††</sup>

 Plus an interesting, policy-relevant setting with valid conditional independence. And data.

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**Q** Where does one find these handy weights?

**A** You've got options, but you need to choose carefully/responsibly.

E.g., if  $w_i(j) = \frac{1}{N_c}$  for all  $(i, j)$ , then we're back to a difference in means.

This weighting doesn't abide by our conditional independence assumption.

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This weighting doesn't abide by our conditional independence assumption.

*The plan* Choose weights  $w_i(j)$  that indicate **how close**  $\mathbf{X}_j$  is to  $\mathbf{X}_i$ .

 Plus an interesting, policy-relevant setting with valid conditional independence. And data.

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If  $\mathbf{X}$  is **discrete**, then we can consider equality, *i.e.*,  $w_i(j) = \mathbb{I}(\mathbf{X}_i = \mathbf{X}_j)$ , scaling as necessary to get  $\sum_j w_i(j) = 1$ .

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*Nearest-neighbor matching* chooses the single closest control observation using the Euclidean distance between  $\mathbf{X}_i$  and  $\mathbf{X}_j$ , i.e.,

$$d_{i,j} = (\mathbf{X}_i - \mathbf{X}_j)' (\mathbf{X}_i - \mathbf{X}_j)$$

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- $\hat{\tau}_i = \mathbf{Y}_{1i} - \mathbf{Y}_{0j}^i$ , where  $\mathbf{Y}_{0j}^i$  is  $i$ 's nearest neighbor in the control group.
- **Estimator:**  $\hat{\tau}_M = \frac{1}{N_T} \sum_i \hat{\tau}_i$
- Produces causal estimates if CIA is valid and we have sufficient overlap.
- Suffers from arbitrary choices of units.

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*Nearest-neighbor matching with Mahalanobis distance* chooses the single closest control using **Mahalanobis** distance between  $\mathbf{X}_i$  and  $\mathbf{X}_j$ , i.e.,

$$d_{i,j} = (\mathbf{X}_i - \mathbf{X}_j)' \Sigma_X^{-1} (\mathbf{X}_i - \mathbf{X}_j)$$

where  $\Sigma_X^{-1}$  is the covariance matrix of  $\mathbf{X}$ .

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- Produces causal estimates if CIA is valid *and* we have sufficient overlap.
- Does not suffer from arbitrary choices of units.

# Matching

## More neighbors?

Why limit ourselves to a **single** "best" match?

If we're going to let a function/algorithm choose the *nearest* match, can't we also let the function/algorithm choose *how many* matches?

Furthermore, if  $N_C \gg N_T$ , it we're throwing away *a lot* of information.

We could instead use this information and be more efficient.

# Matching

## More neighbors!

Kernel matching gives positive weight to all control observations within some **bandwidth**  $h$ , with higher weight for closer matches determined by some **kernel function**  $K(\cdot)$ ,

$$w_i(j) = \frac{K\left(\frac{\mathbf{X}_j - \mathbf{X}_i}{h}\right)}{\sum_{j \in (D=0)} K\left(\frac{\mathbf{X}_j - \mathbf{X}_i}{h}\right)}$$

# Matching

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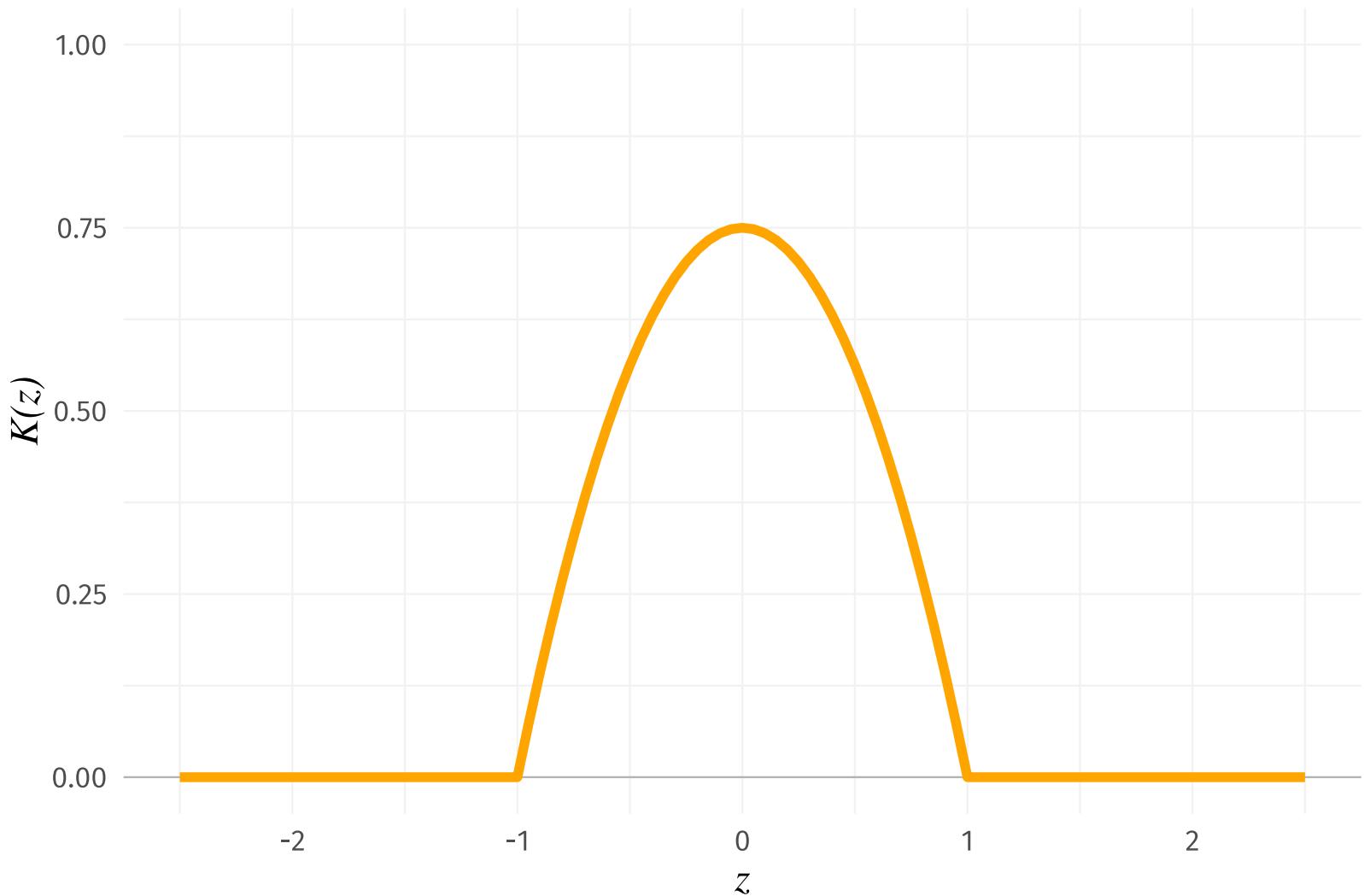
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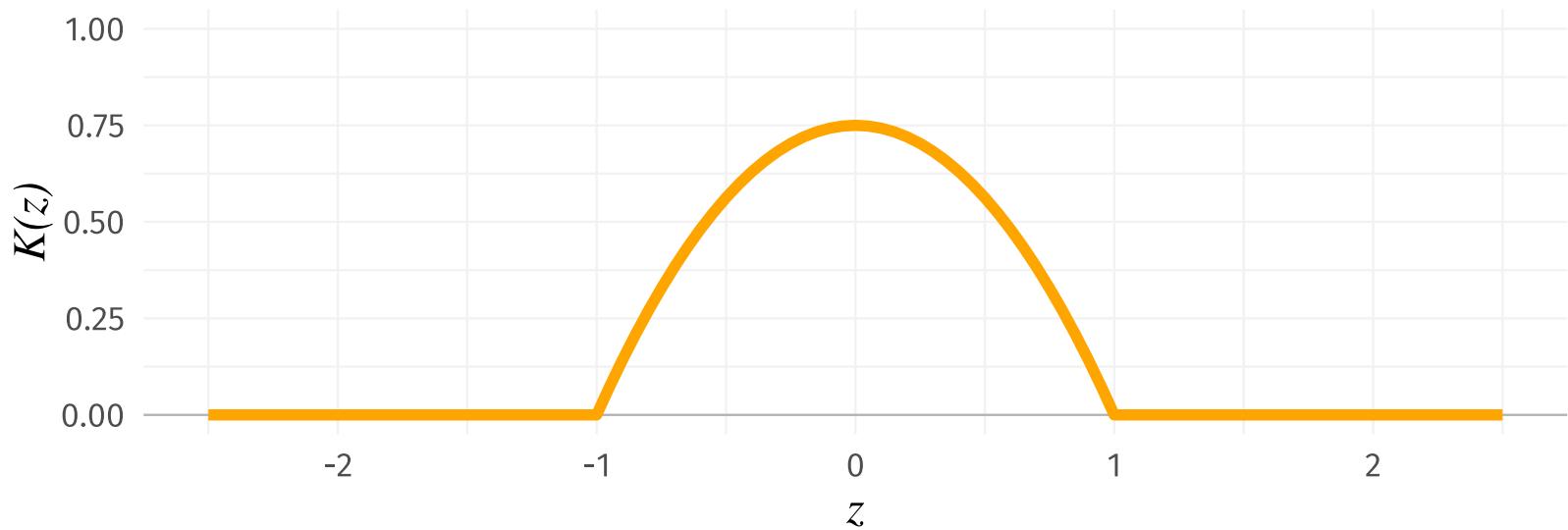
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Example The *Epanechnikov kernel* is defined as

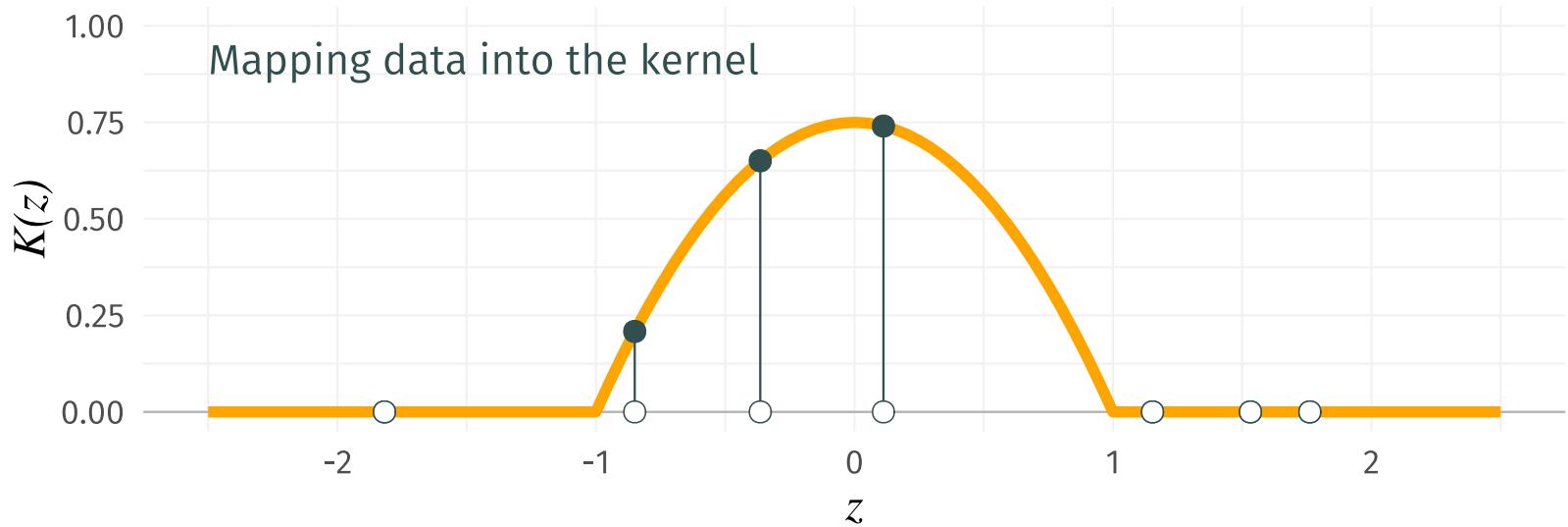
$$K(z) = \frac{3}{4}(1 - z^2) \times \mathbb{I}(|z| < 1)$$

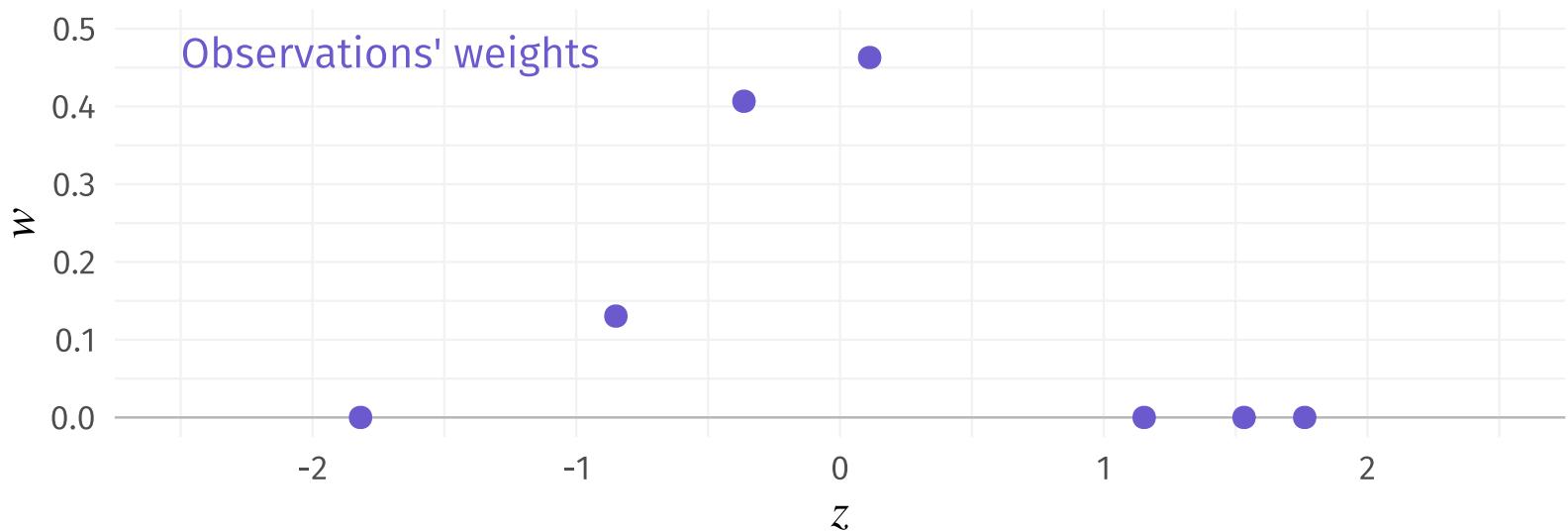
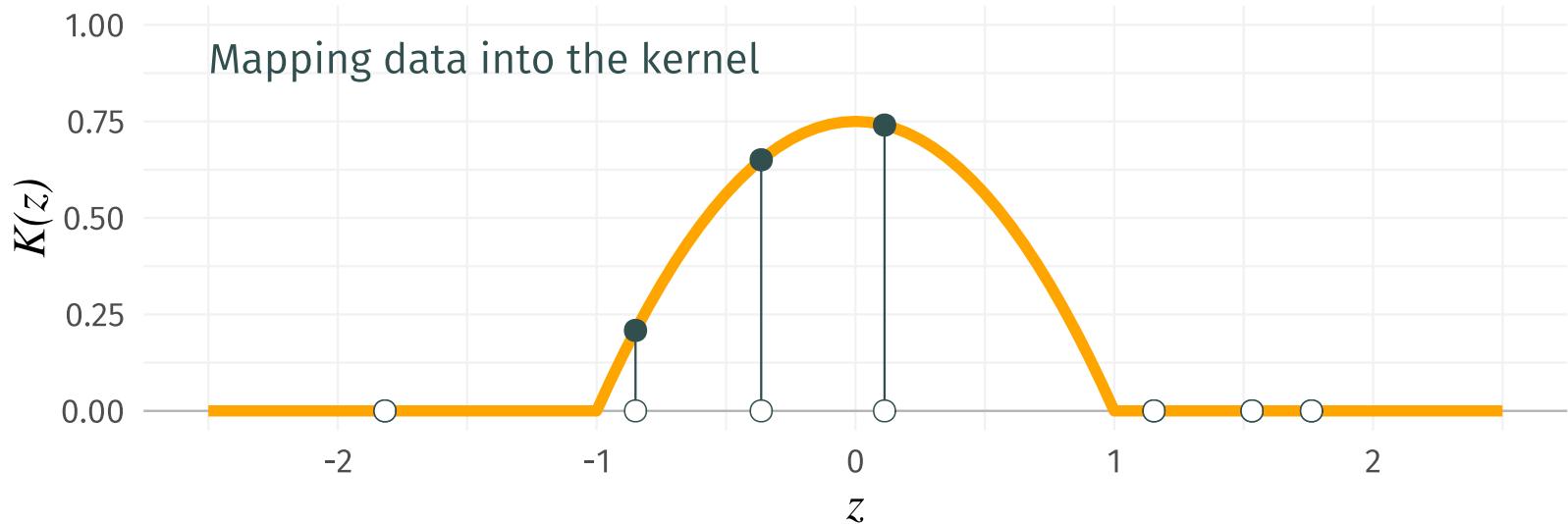
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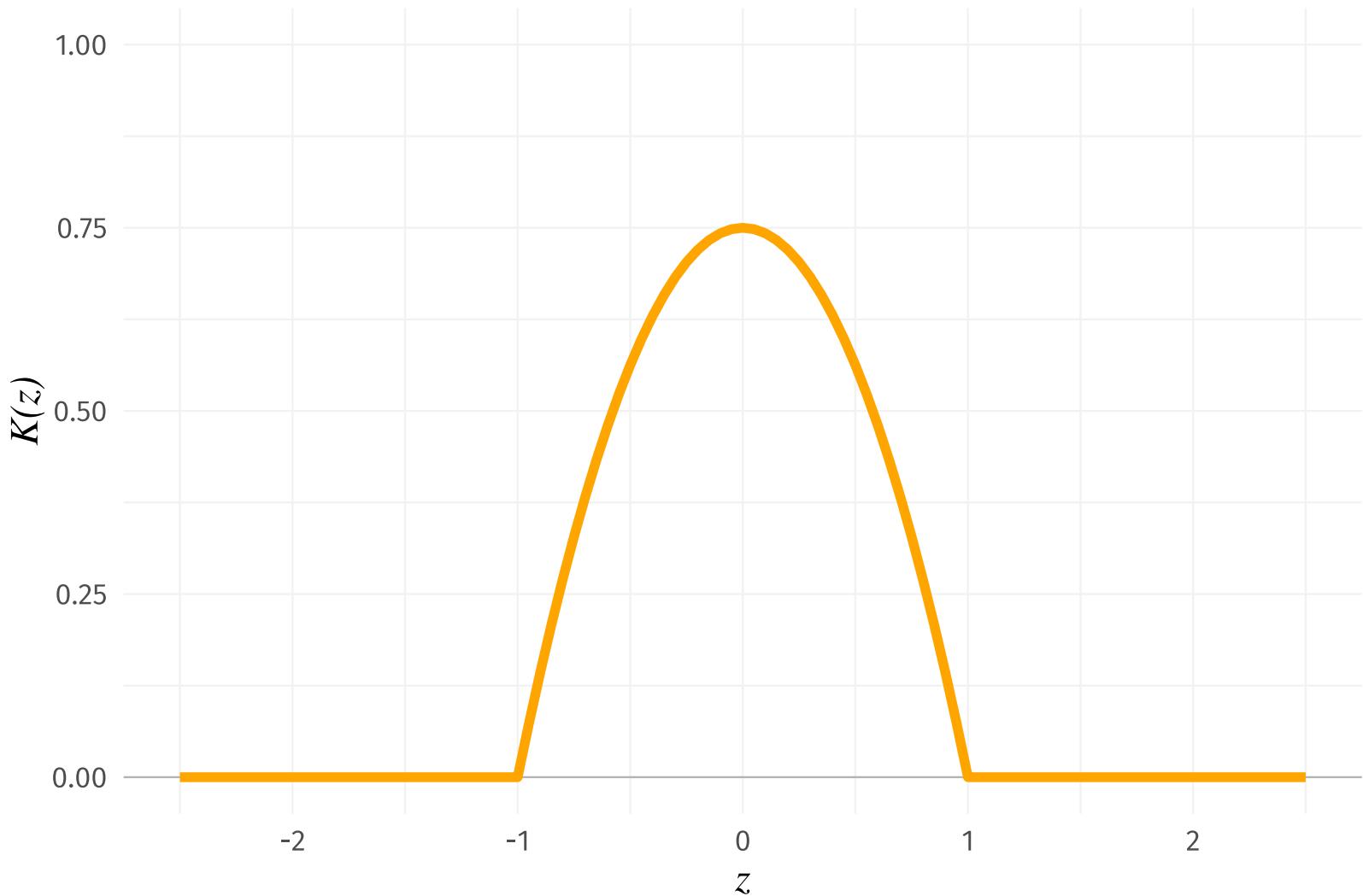


## Mapping data into the kernel

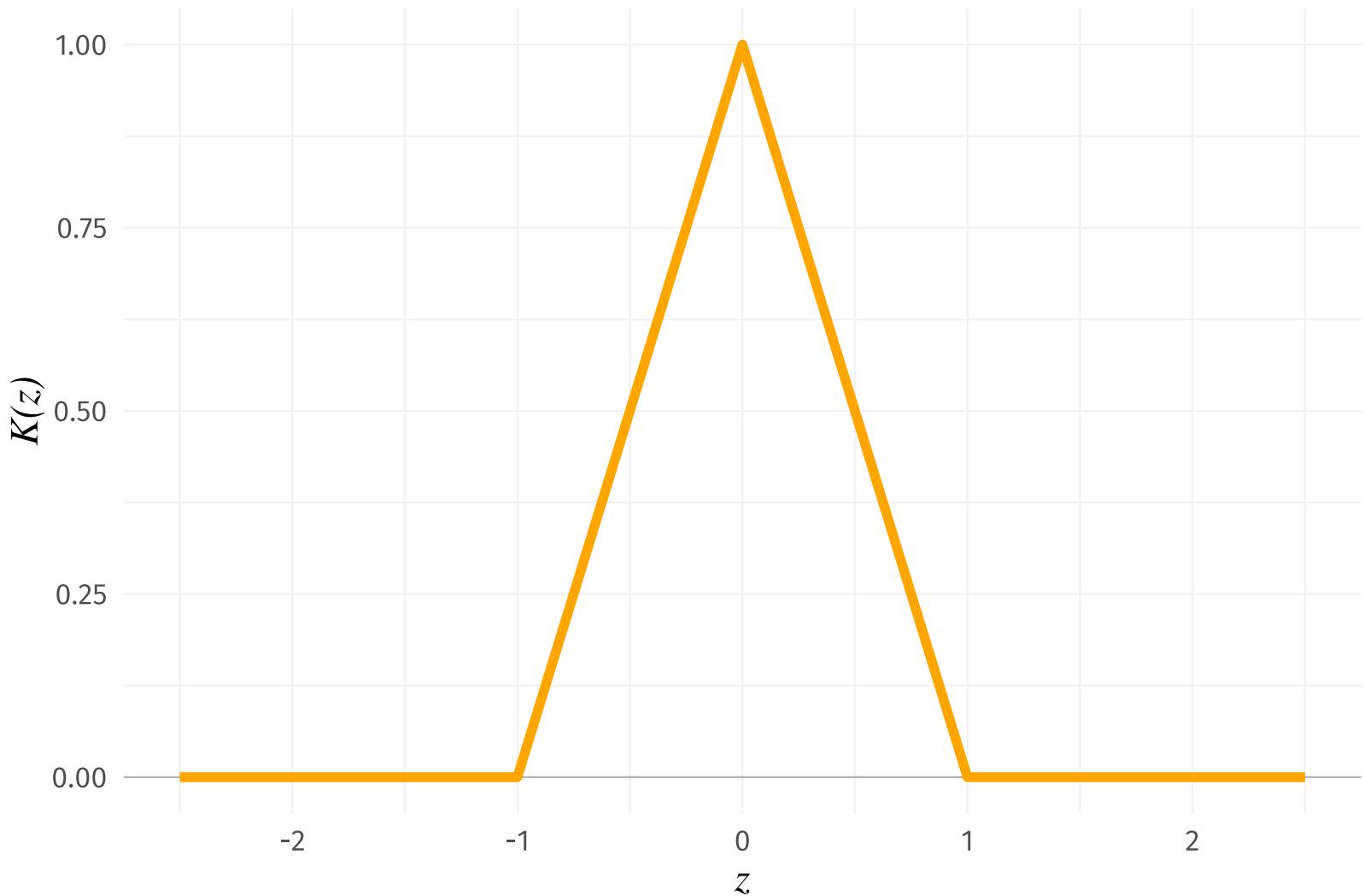




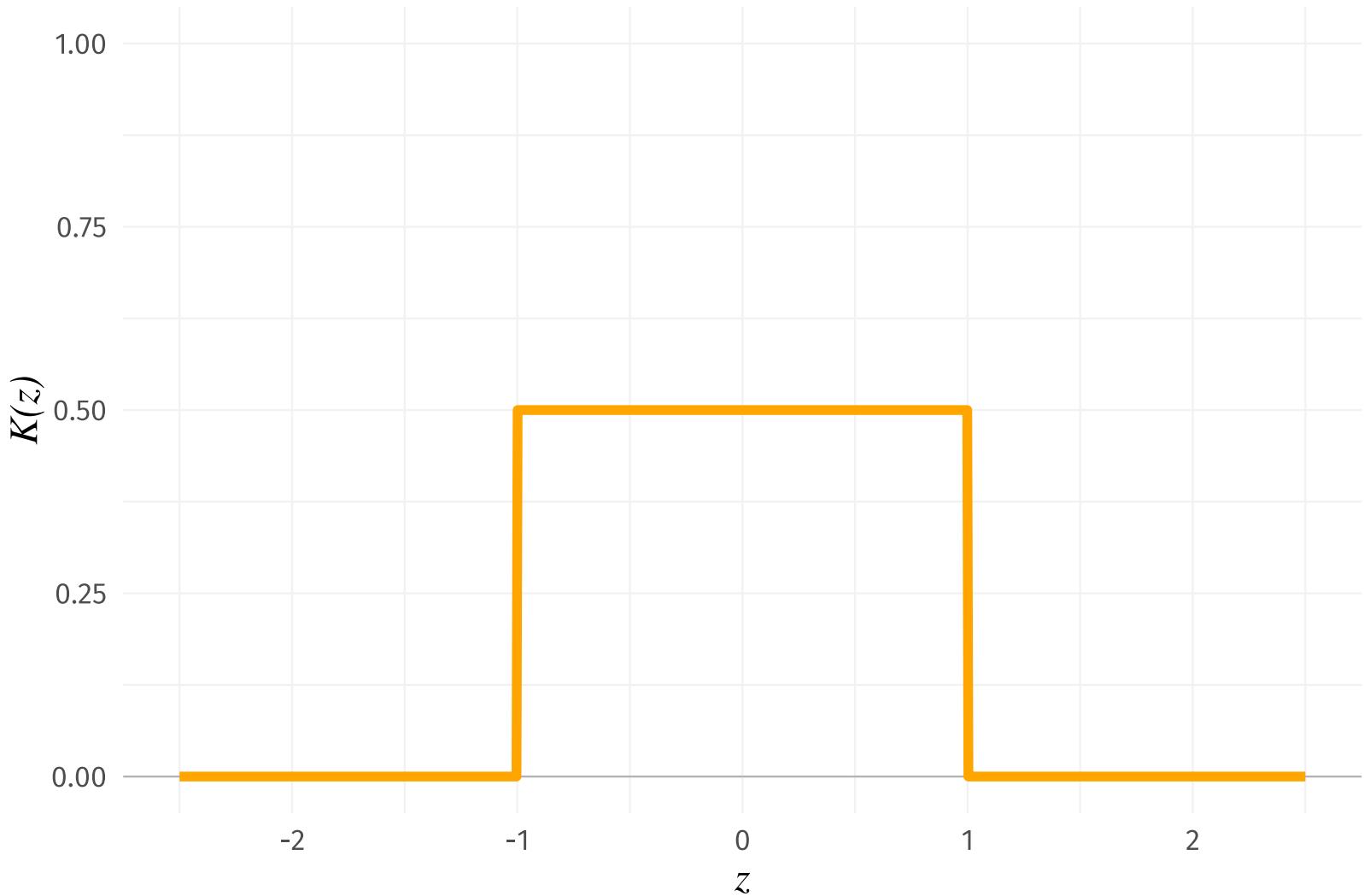
**The Epanechnikov kernel**  $K(z) = \frac{3}{4}(1 - z^2) \times \mathbb{I}(|z| < 1)$



**The Triangle kernel**  $K(z) = (1 - |z|) \times \mathbb{I}(|z| < 1)$



**The Uniform kernel**  $K(z) = \frac{1}{2} \times \mathbb{I}(|z| < 1)$



**The Gaussian kernel**  $K(z) = (2\pi)^{-1/2} \exp(-z^2/2)$



# Kernels

## Aside

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You will most commonly see/use them smoothing out densities—providing a smooth, moving-window average.

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geom\_density() defaults to kernel = "gaussian", but you can specify many other kernel functions (including "epanechnikov").

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E.g., R's (`ggplot2`'s) smooth, density-plotting function `geom_density()`.

`geom_density()` defaults to `kernel = "gaussian"`, but you can specify many other kernel functions (including `"epanechnikov"`).

You can also change the `bandwidth` argument. The default is a bandwidth-choosing function called `bw.nrd0()`.

# Matching

## Adding neighbors

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CIA requires that we're actually conditioning on the observables—it does not allow us to take a simple average across all control observations.

# Matching

## The curse of dimensionality<sup>†</sup>

It turns out kernel- and bandwidth-selection are not our biggest enemies.

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As the dimension of  $\mathbf{X}$  expands (matching on more variables), it becomes **harder and harder to find a nice, close control** for each treated unit.

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We need a way to shrink the dimensionality of  $\mathbf{X}$ .

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# Propensity-score methods

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## Setup

Let's begin with two assumptions—one old and one new.

1. **Conditional independence:**  $(Y_{0i}, Y_{1i}) \perp\!\!\!\perp D_i | X_i$
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**Propensity scores provide a solution** to this mess.

# Propensity-score methods

## The magic

It turns out that if  $(Y_{0i}, Y_{1i}) \perp\!\!\!\perp D_i | X_i$ , then we actually only need to match/condition on  $p(X_i) = E[D_i | X_i]$ .

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**Propensity-score theorem** If  $(Y_{0i}, Y_{1i}) \perp\!\!\!\perp D_i | X_i$ , then  $(Y_{0i}, Y_{1i}) \perp\!\!\!\perp D_i | p(X_i)$ .

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This theorem extends our CIA to a one-dimensional score, avoiding the curse of dimensionality.

# Propensity-score methods

*Theorem* If  $(Y_{0i}, Y_{1i}) \perp\!\!\!\perp D_i | X_i$ , then  $(Y_{0i}, Y_{1i}) \perp\!\!\!\perp D_i | p(X_i)$ .

## Proof

# Propensity-score methods

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## Proof

To prove this theorem, we will show  $\Pr(D_i = 1 | Y_{0i}, Y_{1i}, p(X_i)) = p(X_i)$ ,  
*i.e.*,  $D_i$  is independent of  $(Y_{0i}, Y_{1i})$  after conditioning on  $p(X_i)$ .

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$$\Pr[D_i = 1 | Y_{0i}, Y_{1i}, p(X_i)]$$

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$$\begin{aligned} & \Pr[D_i = 1 | Y_{0i}, Y_{1i}, p(X_i)] \\ &= E[D_i | Y_{0i}, Y_{1i}, p(X_i)] \\ &= E\left[E\left(D_i | Y_{0i}, Y_{1i}, p(X_i), X_i\right) | Y_{0i}, Y_{1i}, p(X_i)\right] \end{aligned}$$

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$$\Pr[D_i = 1 | Y_{0i}, Y_{1i}, p(X_i)] = \dots = E \left[ E(D_i | Y_{0i}, Y_{1i}, X_i) | Y_{0i}, Y_{1i}, p(X_i) \right]$$

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# Propensity-score methods

## Intuition

Q What's going on here?

$\mathbf{X}_i$  carries way more information than  $p(\mathbf{X}_i)$ , so how can we still get conditional independence of treatment by only conditioning on  $p(\mathbf{X}_i)$ ?

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**A<sub>1</sub>** Conditional independence of treatment isn't about extracting all of the information possible from  $\mathbf{X}_i$ . We actually only care about creating a situation in which  $\mathbf{D}_i | \text{something}$  is independent of  $(\mathbf{Y}_{0i}, \mathbf{Y}_{1i})$ .

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**A<sub>2</sub>** Back to our main concern: **selection bias**. People select into treatment. If  $\mathbf{X}$  says two people were equally likely to be treated, and if  $\mathbf{X}_i$  explains all of selection (CIA), then there cannot be selection between these two people.

# Propensity-score methods

## Estimation

So where do propensity scores come from?

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We estimate them—and there are a lot of ways to do that.

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**Q** Can we just use plain OLS (linear probability model)?

**A** Sort of. Think about FWL. This route is going to be the same as a regression conditioning on  $\mathbf{X}_i$ .

# Propensity-score methods

## Estimation

From *MHE* (p. 83)

### Question

A big question here is how to best model and estimate  $p(\mathbf{X}_i)$ ...

### Answer

The answer to this is inherently application-specific. A growing empirical literature suggests that a logit model for the propensity score with a few polynomial terms in continuous covariates works well in practice...

# Propensity-score methods

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*Option 1* Conditioning via regression

*Option 1a* Use a **regression to condition** on  $p(\mathbf{X}_i)$ , i.e.,

$$\mathbf{Y}_i = \alpha + \delta \mathbf{D}_i + \beta p(\mathbf{X}_i) + u_i \quad (1a)$$

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*Option 1b* If we think treatment effects are heterogeneous and may covary with  $\mathbf{X}$ , then we might want to also **interact** treatment with  $p(\mathbf{X}_i)$ , i.e.,

$$\mathbf{Y}_i = \alpha + \delta_1 \mathbf{D}_i + \delta_2 \mathbf{D}_i p(\mathbf{X}_i) + \beta p(\mathbf{X}_i) + u_i \quad (1b)$$

# Propensity-score methods

## Heterogeneity with regression

Let's think a bit more about heterogeneous treatment effects in this setting.

$$\begin{aligned} Y_{0i} &= \alpha + \beta X_i + u_i \\ Y_{1i} &= Y_{0i} + \delta_1 + \delta_2 X_i \end{aligned}$$

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# Propensity-score methods

## Heterogeneity

This final equation

$$Y_i = \alpha + \delta_1 D_i + \delta_2 D_i X_i + \beta X_i + u_i$$

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$$Y_i = \alpha + \delta_1 D_i + \delta_2 D_i p(X_i) + \beta p(X_i) + u_i \quad (1b)$$

which yields

1. a **group-specific treatment effect**  $\delta_1 + \delta_2 X_i$  for each  $X_i$
2. an **average treatment effect**  $\delta_1 + \delta_2 \bar{p}(X_i)$

# Propensity-score methods

## More flexibility

We motivated propensity scores with a desire to reduce dimensionality and estimate/choose/assume fewer parameters.

Adding  $p(\mathbf{X}_i)$  and  $\mathbf{D}_i p(\mathbf{X}_i)$  as covariates in a linear regression doesn't quite exhaust our potential for flexible/nonparametric estimation.

# Propensity-score methods

## Blocking

*Option 2* Block (stratify) on propensity scores.

# Propensity-score methods

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1. Divide the range of  $\hat{p}(\mathbf{X}_i)$  into  $K$  blocks (e.g., 0.05-wide blocks).
2. Place each observation into a block via its  $\hat{p}(\mathbf{X}_i)$ .
3. Calculate  $\hat{\tau}_k$  for each block via difference in means.
4. Average the  $\hat{\tau}_k$  using their shares of the sample, *i.e.*,

$$\hat{\tau}_{\text{Block}} = \sum_{k=1}^K \hat{\tau}_k \frac{N_{1k} + N_{0k}}{N}$$

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*Note* Blocking is similar to NN/kernel matching using  $p(\mathbf{X}_i)$  as distance.

# Propensity-score methods

## Choosing blocks

Blocking on propensity scores requires defining defining blocks.

One common route involves some iteration.

1. **Choose blocks.**
2. Check the **balance of the covariates** within each block.<sup>†</sup>
  - If covariates are **not balanced**, then split your blocks and repeat.
  - If covariates are **balanced**, then stop.

<sup>†</sup> Keep multiple-hypothesis testing in mind. With many covariates and many blocks, you are bound to find statistically significant relationships—even if you are balanced in truth.

# Propensity-score methods

## Overlap

Blocking emphasizes our overlap assumption, *i.e.*,  $0 < \Pr(D_i | X_i) < 1$ .

If a block contains zero treated/control units, we cannot calculate  $\hat{\tau}_k$ .

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*Common practice* Empirically enforce overlap:

- Drop control units with  $\hat{p}(X_i)$  below the minimum propensity score in the treatment group.
- Drop treated units with  $\hat{p}(X_i)$  above the maximum propensity score in the control group.

# Propensity-score methods

## Weighting

*Option 3* Weight observations by the inverse propensity score.

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**Q** How does weighting by  $1/\hat{p}(X_i)$  make sense?

**A** Consider our old (likely biased) friend the difference in means, i.e.,

$$\hat{\tau}_{\text{Diff}} = \bar{Y}_T - \bar{Y}_C = \frac{\sum_i D_i Y_i}{\sum_i D_i} - \frac{\sum_i (1 - D_i) Y_i}{\sum_i (1 - D_i)}$$

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which we've discussed is biased due to selection into treatment, *i.e.*,

$$E[Y_{0i}|D_i = 1] \neq E[Y_{0i}]$$

# Propensity-score methods

## Weighting, justified

Suppose we know  $p(\mathbf{X}_i)$  and we weight each **treated** individual by  $1/p(\mathbf{X}_i)$

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Similarly, weighting **control** individuals by  $1/(1 - p(\mathbf{X}_i))$  yields

$$E\left[\frac{(1 - D_i) Y_i}{1 - p(\mathbf{X}_i)}\right] = \textcolor{blue}{E[Y_{0i}]}$$

# Propensity-score methods

## Weighting: The estimator

Thus, we can estimate an unbiased treatment effect via

$$\hat{\tau}_{p\text{Weight}} = \frac{1}{N} \sum_{i=1}^N \left[ \frac{D_i Y_i}{p(X_i)} - \frac{(1 - D_i Y_i)}{1 - p(X_i)} \right]$$

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*Intuition* We're trying to overcome selection bias, *i.e.*, treated individuals were more likely to be treated as a function of  $\mathbf{X}_i$ —producing higher  $p(\mathbf{X}_i)$ .

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*Intuition* We're trying to overcome selection bias, *i.e.*, treated individuals were more likely to be treated as a function of  $\mathbf{X}_i$ —producing higher  $p(\mathbf{X}_i)$ .

We want to get back to *as-good-as random* variation in treatment.

So we upweight (1) **treated** individuals with low  $p(\mathbf{X}_i)$  and (2) **control** observations with high  $p(\mathbf{X}_i)$ .

# Propensity-score methods

## Weighting: The example

Suppose for some individual  $i$ ,  $p(\mathbf{X}_i) = 0.80$ .

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- If  $i$  is **treated**, then her weight is  $1/p(\mathbf{X}_i) = 1/0.80 = 1.25$

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This propensity score says someone with this set of  $\mathbf{X}_i$  was four-times more likely to be **treated** than **control**.

Our weights fix this imbalance for each  $\mathbf{X}_i$ .

- If  $i$  is **treated**, then her weight is  $1/p(\mathbf{X}_i) = 1/0.80 = 1.25$
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Applying the normalized (and estimated) propensity scores

$$\hat{\tau}_{p\text{Weight}} = \sum_{i=1}^N \frac{\frac{D_i Y_i}{\hat{p}(\mathbf{X}_i)}}{\sum_i \frac{D_i}{\hat{p}(\mathbf{X}_i)}} - \sum_{i=1}^N \frac{\frac{(1 - D_i) Y_i}{1 - \hat{p}(\mathbf{X}_i)}}{\sum_i \frac{(1 - D_i)}{1 - \hat{p}(\mathbf{X}_i)}}$$

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Hirano, Imbens, and Ridder (2003) suggests this estimator is efficient.

# Propensity-score methods

## Why choose one?

There's nothing special about weighted averages—regression can weight.

Thus, a **regression-based estimate**

$$Y_i = \alpha + X_i\beta + \tau D_i + u_i$$

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offers a *doubly robust* property—you have two chances to be right:  $p(X_i)$  or the regression specification.

# Propensity-score methods

## Why choose one? Part two

An alternative, doubly robust method combines propensity-score blocking with regression.

# Propensity-score methods

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An alternative, doubly robust method combines propensity-score blocking with regression.

*Step 1* For each block  $k$ , we run the regression

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*Step 1* For each block  $k$ , we run the regression

$$Y_i = \alpha_k + X_i\beta_k + \tau_k D_i + u_i$$

*Step 2* Aggregate block-level treatment-effect estimates

$$\hat{\tau} = \sum_{k=1}^K \hat{\tau}_k \frac{N_{1k} + N_{0k}}{N}$$

# Propensity-score methods

## Major requirements

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We still have two **major** requirements for any of these methods to work?

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## Major requirements

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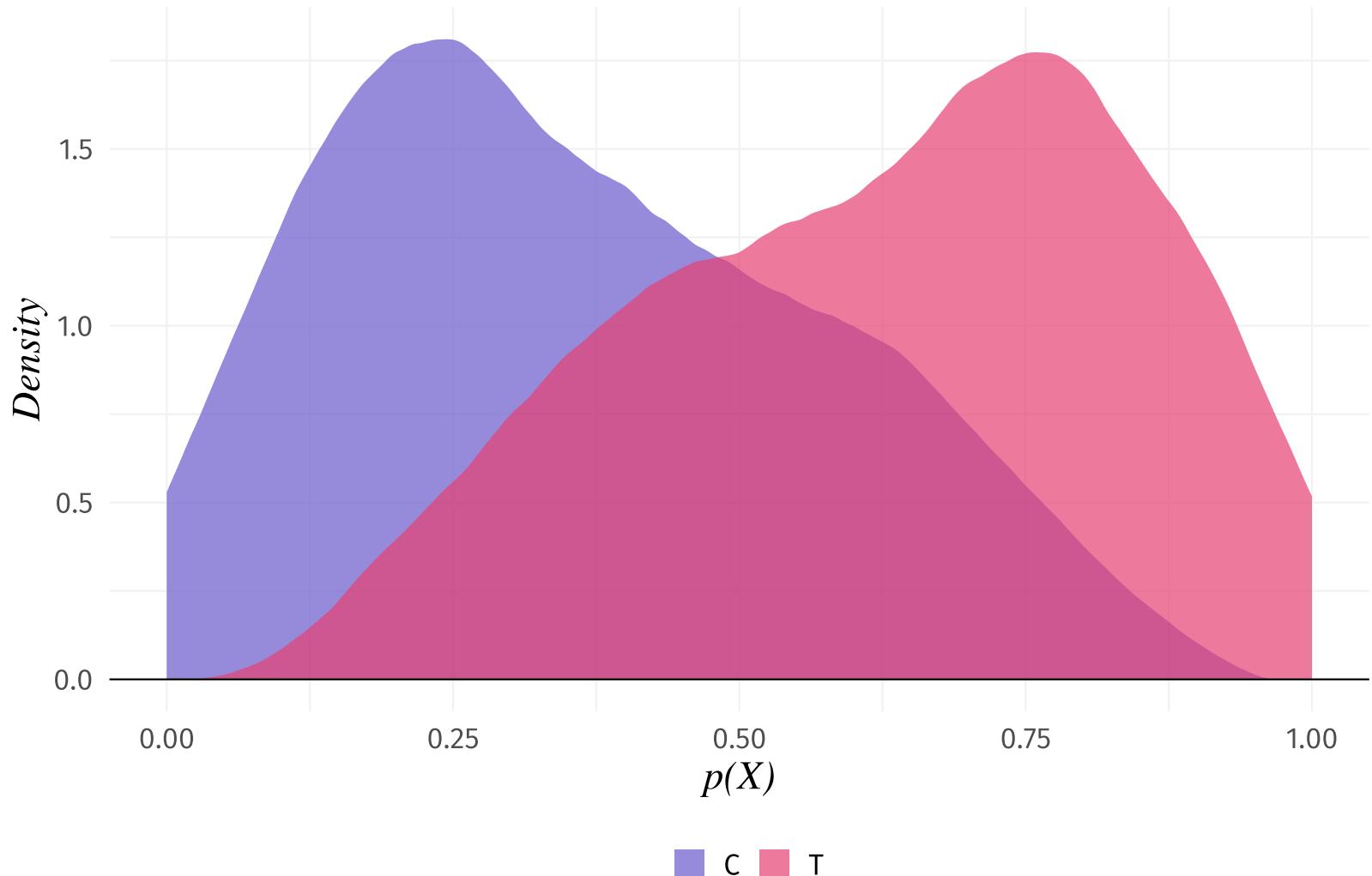
1. Is the **conditional-independence assumption** true?
2. Do we have **overlap** between treatment and control units.

We can look for evidence of (2) in the data—particularly if we're using propensity-score methods.<sup>†</sup>

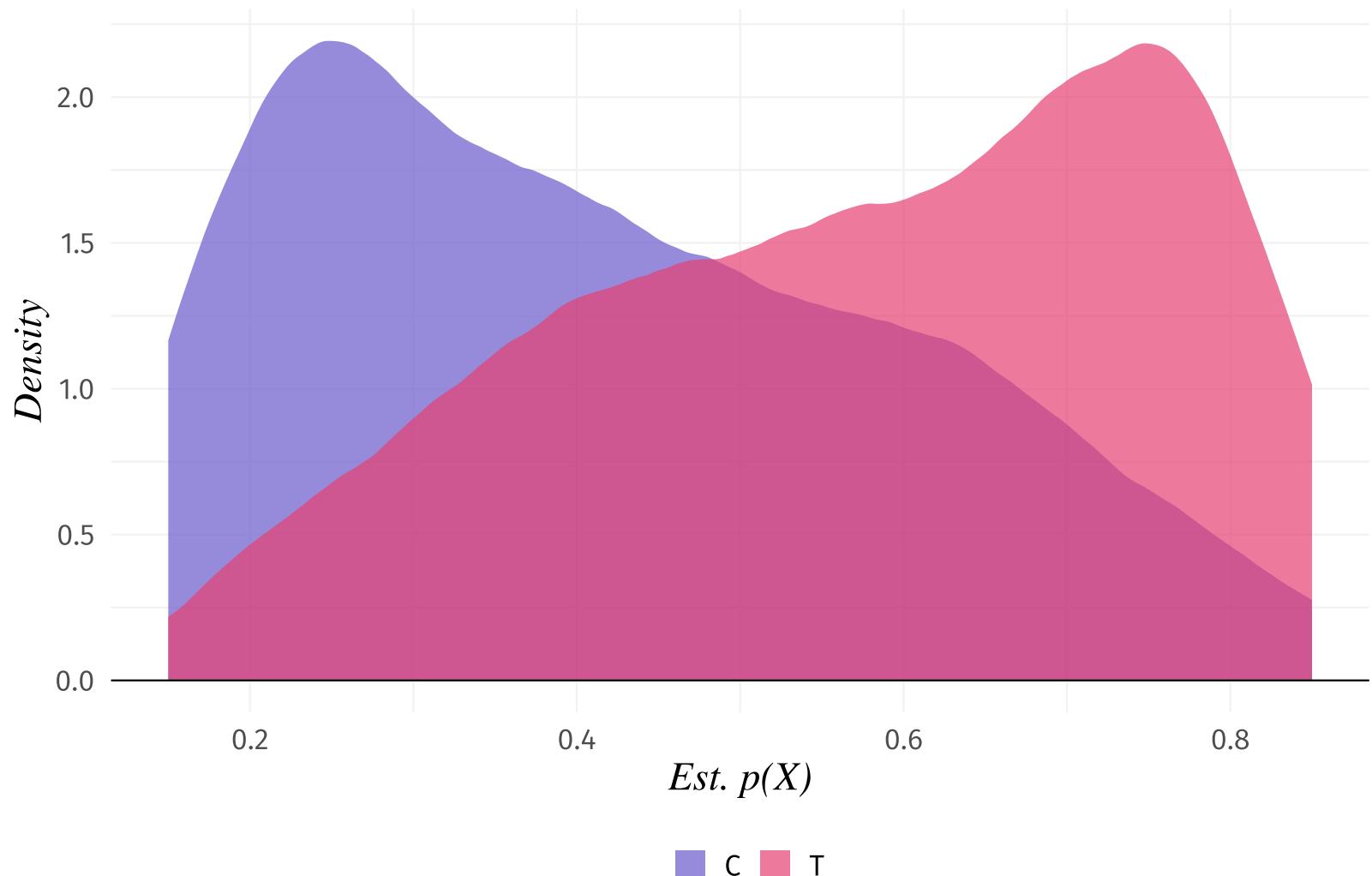
How? Plot the distributions of  $p(\mathbf{X}_i)$  for **T** and **C**.

<sup>†</sup> Checking for overlap in  $\mathbf{X}$ -space, can be tough as the dimensions of  $\mathbf{X}$  expand.

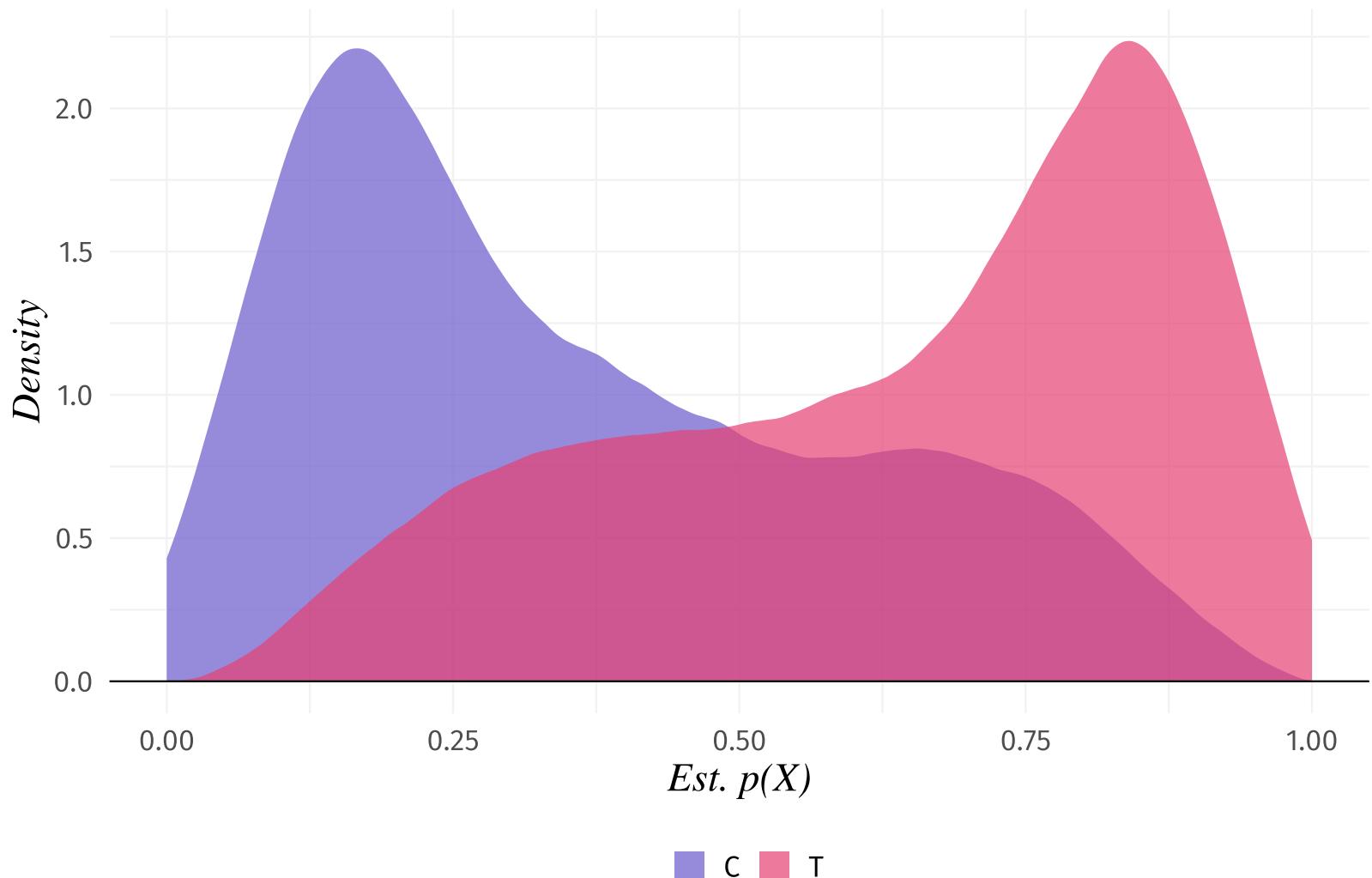
Missing overlap in  $p(\mathbf{X}_i)$



Authentic (enforced) overlap in  $p(\mathbf{X}_i)$



Logit-based  $\hat{p}(\mathbf{X}_i)$  hiding some of the missing overlap in  $p(\mathbf{X}_i)$



Overlap in one dimension does not guarantee in two dimensions.

Note Shading denotes **share of treatment**: white =0% and pink=100%.



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