Algorithm is Experiment:
Machine Learning, Market Design, and Policy Eligibility Rules

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Abstract

Algorithms produce a growing portion of decisions and recommendations both in policy and business. Such algorithmic decisions are natural experiments (conditionally quasi-randomly assigned instruments) since the algorithms make decisions based only on observable input variables. We use this observation to develop a treatment-effect estimator for a class of stochastic and deterministic algorithms. Our estimator is shown to be consistent and asymptotically normal for well-defined causal effects. A key special case of our estimator is a high-dimensional regression discontinuity design. The proofs use tools from differential geometry and geometric measure theory, which may be of independent interest.

The practical performance of our method is first demonstrated in a high-dimensional simulation resembling decision-making by machine learning algorithms. Our estimator has smaller mean squared errors compared to alternative estimators. We finally apply our estimator to evaluate the effect of the Coronavirus Aid, Relief, and Economic Security (CARES) Act, where more than $10 billion worth of relief funding is allocated to hospitals via an algorithmic rule. The estimates suggest that the relief funding has little effect on COVID-19-related hospital activity levels. Naive OLS and IV estimates exhibit substantial selection bias.

Keywords: Algorithm, Causal Inference, Regression Discontinuity Design, Propensity Score, Instrumental Variable, Artificial Intelligence, Fiscal Stimulus, Health Care

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1 Introduction

Today’s society increasingly resorts to algorithms for decision-making and resource allocation. For example, judges in the US make legal decisions aided by predictions from supervised machine learning algorithms. Supervised learning is also used by governments to detect potential criminals and terrorists, and by banks and insurance companies to screen potential customers. Tech companies like Facebook, Microsoft, and Netflix allocate digital content by reinforcement learning and bandit algorithms. Retailers and e-commerce platforms engage in algorithmic pricing. Similar algorithms are encroaching on high-stakes settings, such as in education, healthcare, and the military.

Other types of algorithms also loom large. School districts, college admissions systems, and labor markets use matching algorithms for position and seat allocations. Objects worth astronomical sums of money change hands every day in algorithmically run auctions. Many public policy domains like Medicaid often use algorithmic rules to decide who are eligible.

All of the above, diverse examples share a common trait: a decision-making algorithm makes decisions based only on its observable input variables. Thus conditional on the observable variables, algorithmic treatment decisions are (quasi-)randomly assigned. That is, they are independent of any potential outcome or unobserved heterogeneity. This property turns algorithm-based treatment decisions into instrumental variables (IVs) that can be used for measuring the causal effect of the final treatment assignment. The algorithm-based instrument may produce stratified randomization (e.g., certain bandit and reinforcement learning algorithms), regression-discontinuity-style local variation (e.g., machine judges), or some combination of the two.

This paper shows how to use data obtained from algorithmic decision-making to identify and estimate causal effects. In our framework, the analyst observes a random sample \{\(Y_i, X_i, D_i, Z_i\)\}_{i=1}^{n}, where \(Y_i\) is the outcome of interest, \(X_i \in \mathbb{R}^p\) is a vector of pre-treatment covariates (algorithm’s input variables), \(D_i\) is the binary treatment assignment, possibly made by humans, and \(Z_i\) is the binary treatment recommendation made by a known algorithm. The algorithm takes \(X_i\) as input and computes the probability of the treatment recommendation \(A(X_i) = \Pr(Z_i = 1|X_i)\). \(Z_i\) is then randomly determined based on the known probability \(A(X_i)\) independently of everything else conditional on \(X_i\). The algorithm’s recommendation \(Z_i\) may influence the final treatment assignment \(D_i\), determined as \(D_i = Z_iD_i(1) + (1 - Z_i)D_i(0)\), where \(D_i(z)\) is the potential treatment assignment that would be realized if \(Z_i = z\). Finally, the observed outcome \(Y_i\) is determined as \(Y_i = D_iY_i(1) + (1 - D_i)Y_i(0)\), where \(Y_i(1)\) and \(Y_i(0)\) are potential outcomes that would be realized if the individual were treated and not treated, respectively. This setup is an IV model where the IV satisfies the conditional independence condition but may not satisfy the overlap (full-support) condition. To our knowledge, there is no standard estimator for this setup.

Within this framework, we first characterize the sources of causal-effect identification for a class of data-generating algorithms. This class includes both stochastic and deterministic algorithms.\(^1\) The sources of causal-effect identification turn out to be summarized by a suitable

\(^1\)This class includes all of the aforementioned examples, thus nesting existing insights on quasi-experimental variation in particular algorithms, such as surge pricing (Cohen, Hahn, Hall, Levitt and Metcalfe, 2016), bandit (Li, Chu, Langford and Schapire, 2010), reinforcement learning (Precup, 2000), supervised learning (Cowgill, 2018; Bundorf, Polyakova and Tai-Seale, 2019), and market-design algorithms (Abdulkadiroğlu, Angrist, Narita, Narita, ...
modification of the Propensity Score (Rosenbaum and Rubin, 1983). We call it the Approximate Propensity Score (APS). For each covariate value \( x \), the Approximate Propensity Score is the average probability of a treatment recommendation in a shrinking neighborhood around \( x \), defined as

\[
p^A(x) = \lim_{\delta \to 0} \frac{\int_{B(x,\delta)} A(x^*) dx^*}{\int_{B(x,\delta)} dx^*},
\]

where \( B(x,\delta) \) is a \( p \)-dimensional ball with radius \( \delta \) centered at \( x \). The Approximate Propensity Score provides an easy-to-check condition for what causal effects the data from an algorithm allow us to identify. In particular, we show that the conditional local average treatment effect (LATE; Imbens and Angrist, 1994) at covariate value \( x \) is identified if and only if the Approximate Propensity Score is nondegenerate, i.e., \( p^A(x) \in (0, 1) \).

The identification analysis suggests a way of estimating treatment effects using the algorithm-produced data. The treatment effects can be estimated by two-stage least squares (2SLS) where we regress the outcome on the treatment with the algorithm’s recommendation as an IV. To make the algorithmic recommendation a conditionally independent IV, we propose to control for the Approximate Propensity Score. Our proposed estimator is as follows.

1. For small bandwidth \( \delta > 0 \) and a large number of simulation draws \( S \), compute

\[
p^s(X_i; \delta) = \frac{1}{S} \sum_{s=1}^{S} A(X_{i,s}^*),
\]

where \( X_{i,1}^*, ..., X_{i,S}^* \) are \( S \) independent simulation draws from the uniform distribution on \( B(X_i,\delta) \). This \( p^s(X_i; \delta) \) is a simulation-based approximation to the Approximate Propensity Score \( p^A(x) \).

2. Using the observations with \( p^s(X_i; \delta) \in (0, 1) \), run the following 2SLS IV regression:

\[
D_i = \gamma_0 + \gamma_1 Z_i + \gamma_2 p^s(X_i; \delta) + \nu_i \quad \text{(First Stage)}
\]
\[
Y_i = \beta_0 + \beta_1 D_i + \beta_2 p^s(X_i; \delta) + \epsilon_i \quad \text{(Second Stage)}.
\]

Let \( \hat{\beta}_1 \) be the estimated coefficient on \( D_i \).

2 To make common \( \delta \) for all dimensions reasonable, we standardize each characteristic \( X_{ij} \) to have mean zero and variance one for each \( j = 1, ..., p \), where \( p \) is the number of input characteristics. For the bandwidth \( \delta \), we suggest that the analyst considers several different values and check if the 2SLS estimates are robust to bandwidth changes, as we often do in regression discontinuity design (RDD) applications. It is hard to pick \( \delta \) in a data-driven way. Common methods for bandwidth selection in univariate RDDs include Imbens and Kalyanaraman (2012) and Calonico, Cattaneo and Titunik (2014), who estimate the bandwidth that minimizes the asymptotic mean squared error (AMSE). It is not straightforward to estimate the AMSE-optimal bandwidth in our setting with many running variables and complex IV assignment, since it requires nonparametric estimation of functions on the multidimensional covariate space such as conditional mean functions, their derivatives, the curvature of the RDD boundary, etc.
As the main theoretical result, we prove the 2SLS estimator \( \hat{\beta}_1 \) is a consistent and asymptotically normal estimator of a well-defined causal effect (weighted average of conditional local average treatment effects). We also show that inference based on the conventional 2SLS heteroskedasticity-robust standard errors is asymptotically valid as long as the bandwidth \( \delta \) goes to zero at an appropriate rate. We prove the asymptotic properties by exploiting results from differential geometry and geometric measure theory, which may be of independent interest. There appears to be no existing estimator with these asymptotic properties even for the multidimensional RDD, a special case of our framework where the decision-making algorithm is deterministic and uses multiple input (running) variables for assigning treatment recommendations. Moreover, our result applies to much more general settings with stochastic algorithms, deterministic algorithms, and combinations of the two.

The practical performance of our estimator is demonstrated through simulation and an original application. We first conduct a Monte Carlo simulation mimicking real-world decision-making based on machine learning algorithms. We consider a data-generating process combining stochastic and deterministic algorithms. Treatment recommendations are randomly assigned for a small experimental segment of the population and are determined by a high-dimensional, deterministic machine learning algorithm for the rest of the population. Our estimator is shown to have smaller mean squared errors relative to alternative estimators.

Our empirical application is an analysis of COVID-19 hospital relief funding. The Coronavirus Aid, Relief, and Economic Security (CARES) Act and Paycheck Protection Program designated $175 billion for COVID-19 response efforts and reimbursement to health care entities for expenses or lost revenues (Kakani, Chandra, Mullainathan and Obermeyer, 2020). This policy intended to help hospitals hit hard by the pandemic, as “financially insecure hospitals may be less capable of investing in COVID-19 response efforts” (Khullar, Bond and Schpero, 2020). We ask whether this problem is alleviated by the relief funding to hospitals.

We identify the causal effects of the relief funding by exploiting the funding eligibility rule. The government employs an algorithmic rule to decide which hospitals are eligible for funding. This fact allows us to apply our method to estimate the effect of relief funding. Specifically, our 2SLS estimators use eligibility status as an instrumental variable for funding amounts, while controlling for the Approximate Propensity Score induced by the eligibility-determining algorithm.

The resulting estimates suggest that COVID-19 relief funding has little to no effect on outcomes, such as the number of hospitalized COVID-19 patients at each hospital. The estimated causal effects of relief funding are much smaller and less significant than the naive ordinary least squares (OLS) (with and without controls) or 2SLS estimates with no controls. Our finding provides causal evidence for the concern that funding in the CARES Act was not well targeted to the clinics and hospitals with the greatest needs.\(^3\)

Related Literature

Theoretically, our framework integrates the classic propensity-score (selection-on-observables) scenario with a multidimensional extension of the RDD. We analyze this integrated setup in the IV world with noncompliance. This general setting appears to have no prior established estimator. Armstrong and Kolesár (2020) provide an estimator for a related setting with perfect compliance. The Approximate Propensity Score developed in this paper shares its spirit with the local random assignment interpretation of the RDD, discussed by Frölich (2007), Cattaneo, Frandsen and Titiunik (2015), Cattaneo, Titiunik and Vazquez-Bare (2017), Frandsen (2017), Sekhon and Titiunik (2017), Frölich and Huber (2019), Abduladiroğlu et al. (Forthcoming) and Eckles, Ignatiadis, Wager and Wu (2020). These papers consider settings that fit into this paper’s framework.

When we adapt our estimator to the multidimensional RDD case, our estimator has three features. First, it is a consistent and asymptotically normal estimator of a well-interpreted causal effect (average of conditional treatment effects along the RDD boundary) even if treatment effects are heterogeneous. Second, it uses observations near all the boundary points as opposed to using only observations near one specific boundary point, thus avoiding variance explosion even when \( X_i \) has many elements. Third, it can be easily implemented even in cases with many covariates and complex algorithms (RDD boundaries). Our method circumvents the difficulty of identifying the decision boundary from a complicated decision-making algorithm. No prior estimator appears to have all of these properties (Papay, Willett and Murnane, 2011; Zajonc, 2012; Keele and Titiunik, 2015; Cattaneo, Titiunik, Vazquez-Bare and Keele, 2016; Imbens and Wager, 2019). Appendix A.1 provides a detailed review of the most closely related papers on the multidimensional RDD.

Our empirical application uses the proposed method to study hospitals receiving CARES Act relief funding. Our empirical finding contributes to emerging work on how health care providers respond to financial shocks (Duggan, 2000; Adelino, Lewellen and Sundaram, 2015; Dranove, Garthwaite and Ody, 2017; Adelino, Lewellen and McCartney, 2021).

A focal group of decision-making algorithms are machine learning algorithms, as illustrated in our machine-learning simulation. While we are interested in machine learning as a data-production tool, the existing literature (except the above mentioned strand) focuses on machine learning as a data-analysis tool. For example, a set of predictive studies applies machine learning to make predictions important for social policy questions (Kleinberg, Lakkaraju, Leskovec, Ludwig and Mullainathan, 2017; Einav, Finkelstein, Mullainathan and Obermeyer, 2018). Another set of causal and structural work repurposes machine learning to aid with causal inference and structural econometrics (Athey and Imbens, 2017; Belloni, Chernozhukov, Fernández-Val and Hansen, 2017; Bonhomme, Lamadon and Manresa, 2019; Mullainathan and Spiess, 2017). We supplement these studies by highlighting the role of machine learning as a data-production tool.\(^4\)

\(^4\)Building on their prior work (Armstrong and Kolesár, 2018), Armstrong and Kolesár (2020) consider estimation and inference on average treatment effects under the assumption that the final treatment assignment is independent of potential outcomes conditional on observables. Their estimator is not applicable to the IV world we consider. Their method and our method also achieve different goals; their goal lies in finite-sample optimality and asymptotically valid inference while our goal is to obtain consistency and asymptotic normality.

\(^5\)This paper also has a conceptual connection to the heated conversation about whether algorithmic decisions
2 Framework

Our framework is a mix of the conditional independence, multidimensional RDD, and instrumental variable scenarios. In the setup in the introduction, we are interested in the effect of some binary treatment \( D_i \in \{0,1\} \) on some outcome of interest \( Y_i \in \mathbb{R} \). As is standard in the literature, we impose the exclusion restriction that the treatment recommendation \( Z_i \in \{0,1\} \) does not affect the observed outcome other than through the treatment assignment \( D_i \). This allows us to define the potential outcomes indexed against the treatment assignment \( D_i \) alone.\(^6\)

We consider algorithms that make treatment recommendations based solely on individual \( i \)'s predetermined, observable covariates \( X_i = (X_{i1}, ..., X_{ip})' \in \mathbb{R}^p \). Let the function \( A : \mathbb{R}^p \rightarrow [0,1] \) represent the decision algorithm, where \( A(X_i) = \Pr(Z_i = 1|X_i) \) is the probability that the treatment is recommended for individual \( i \) with covariates \( X_i \).\(^7\) The central assumption is that the analyst knows function \( A \) and is able to simulate it. That is, the analyst is able to compute the recommendation probability \( A(x) \) given any input value \( x \in \mathbb{R}^p \). The treatment recommendation \( Z_i \) for individual \( i \) is then randomly determined with probability \( A(X_i) \) independently of everything else. Consequently, the following conditional independence property holds.

**Property 1** (Conditional Independence). \( Z_i \perp \perp (Y_i(1), Y_i(0), D_i(1), D_i(0))|X_i \).

Note that the codomain of \( A \) contains 0 and 1, allowing for deterministic treatment assignments conditional on \( X_i \). Our framework therefore nests the RDD as a special case.\(^8\) Another special case of our framework is the classic conditional independence scenario with the common support condition (\( A(X_i) \in (0,1) \) almost surely). In addition to these simple settings, this framework nests many other situations, such as multidimensional RDDs and complex machine learning and market-design algorithms, as illustrated in Section 7.

In typical machine-learning scenarios, an algorithm first applies machine learning on \( X_i \) to make some prediction and then uses the prediction to output the recommendation probability \( A(X_i) \), as in the following example.

**Example.** Automated disease detection algorithms (Gulshan et al., 2016) use machine learning, in particular deep learning, to detect various diseases and to identify patients at risk. Using our framework described above, a detection algorithm predicts whether an individual \( i \) has a certain disease \( (Z_i = 1) \) or not \( (Z_i = 0) \) based on a digital image \( X_i \in \mathbb{R}^p \) of a part of the are better than human decisions. Here “better” is in terms of fairness and efficiency (Hoffman, Kahn and Li, 2017; Horton, 2017; Kleinberg et al., 2017). In this study, we take a complementary perspective in that we take a decision algorithm as given, no matter whether it is good or bad, and study how to use its produced data for impact evaluation.

\(^6\)Formally, let \( Y_i(d,z) \) denote the potential outcome that would be realized if \( i \)'s treatment assignment and recommendation were \( d \) and \( z \), respectively. The exclusion restriction assumes that \( Y_i(d,1) = Y_i(d,0) \) for \( d \in \{0,1\} \) (Imbens and Angrist, 1994).

\(^7\)We assume that the function \( A \) is supported on \( \mathbb{R}^p \) irrespective of the support of \( X_i \).

\(^8\)Most of the existing studies on RDDs define the potential treatment assignment indexed against the running variable like \( D_i(x) \), which represents the counterfactual treatment assignment the individual \( i \) would have received if her running variable had been set to \( x \). Unlike prior work, we define it indexed against the treatment recommendation \( z \).
individual’s body, where each \( X_{ij} \in \mathbb{R} \) denotes the intensity value of a pixel in the image. The algorithm uses training data to construct a binary classifier \( A : \mathbb{R}^p \rightarrow \{0, 1\} \). The classifier takes an image of individual \( i \) as input and makes a binary prediction of whether the individual has the disease:

\[
Z_i \equiv A(X_i).
\]

The algorithm’s diagnosis \( Z_i \) may influence the doctor’s treatment decision for the individual, denoted by \( D_i \in \{0, 1\} \). We are interested in how the treatment decision \( D_i \) affects the individual’s outcome \( Y_i \).

Let \( Y_{zi} \) be defined as

\[
Y_{zi} \equiv D_i(z)Y_i(1) + (1 - D_i(z))Y_i(0)
\]

for \( z \in \{0, 1\} \). \( Y_{zi} \) is the potential outcome when the treatment recommendation is \( Z_i = z \). It follows from Property 1 that \( Z_i \perp \perp (Y_i(1), Y_i(0)) | X_i \).

Assumption 1.

(a) (Almost Everywhere Continuity of \( A \)) \( A \) is continuous almost everywhere with respect to the Lebesgue measure.

(b) (Measure Zero Boundaries of \( X_0 \) and \( X_1 \)) \( \mathcal{L}^p(\mathcal{X}_k) = \mathcal{L}^p(\text{int}(\mathcal{X}_k)) \) for \( k = 0, 1 \).

Assumption 1 (a) allows the function \( A \) to be discontinuous on a set of points with the Lebesgue measure zero. For example, \( A \) is allowed to be a discontinuous step function as long as it is continuous almost everywhere. Assumption 1 (b) holds if the Lebesgue measures of the boundaries of \( X_0 \) and \( X_1 \) are zero.

3 Identification

What causal effects can be learned from data \( (Y_i, X_i, D_i, Z_i) \) generated by the algorithm \( A \)? A key step toward answering this question is what we call the Approximate Propensity Score (APS). To define it, let:

\[
p^A(x; \delta) = \frac{\int_{B(x, \delta)} A(x^*) dx^*}{\int_{B(x, \delta)} dx^*},
\]

where \( B(x, \delta) = \{x^* \in \mathbb{R}^p : \|x - x^*\| < \delta\} \) is the (open) \( \delta \)-ball around \( x \in \mathcal{X} \). Here, \( \|\cdot\| \) denotes the Euclidean norm on \( \mathbb{R}^p \). To make common \( \delta \) for all dimensions reasonable, we normalize \( X_{ij} \)

\[\text{---}
\]

\[9\]Whether we use an open ball or closed ball does not affect \( p^A(x; \delta) \). When we instead use a rectangle, ellipsoid, or any standard kernel function to define \( p^A(x; \delta) \), the limit \( \lim_{\delta \to 0} p^A(x; \delta) \) may be different at some points (e.g., at discontinuity points of \( A \)), but the same identification results hold under suitable conditions. We use a ball for simplicity and practicality.
to have mean zero and variance one for each $j = 1, ..., p$.\textsuperscript{10} We assume that $A$ is a $L^p$-measurable function so that the integrals exist. We then define APS as follows.

$$ p^A(x) \equiv \lim_{\delta \to 0} p^A(x; \delta). $$

APS at $x$ is the average probability of a treatment recommendation in a shrinking ball around $x$. We call this the \textit{Approximate} Propensity Score, since this score modifies the standard propensity score $A(X_i)$ to incorporate local variation in the score. APS exists for most covariate points and algorithms (see Appendix A.2).

Figure 1 illustrates APS. In the example, $X_i$ is two dimensional, and the support of $X_i$ is divided into three sets depending on the value of $A$. For the interior points of each set, APS is equal to $A$. On the border of any two sets, APS is the average of the $A$ values in the two sets. Thus, $p^A(x) = \frac{1}{2}(0+0.5) = 0.25$ for any $x$ in the open line segment $AB$, $p^A(x) = \frac{1}{2}(0.5+1) = 0.75$ for any $x$ in the open line segment $BC$, and $p^A(x) = \frac{1}{2}(0+1) = 0.5$ for any $x$ in the open line segment $BD$.

We say that a causal effect is \textit{identified} if it is uniquely determined by the joint distribution of $(Y_i, X_i, D_i, Z_i)$. Our identification analysis uses the following continuity condition.

**Assumption 2** (Local Mean Continuity). For $z \in \{0, 1\}$, the conditional expectation functions $E[Y_{zi}|X_i]$ and $E[D_i(z)|X_i]$ are continuous at any point $x \in \mathcal{X}$ such that $p^A(x) \in (0, 1)$ and $A(x) \in \{0, 1\}$.

Assumption 2 is a multivariate extension of the local mean continuity condition that is frequently assumed in the RDD.\textsuperscript{11} $A(x) \in \{0, 1\}$ means that the treatment recommendation $Z_i$ is deterministic conditional on $X_i = x$. If APS at the point $x$ is nondegenerate ($p^A(x) \in (0, 1)$), however, there exists a point close to $x$ that has a different value of $A$ from $x$’s, which creates variation in the treatment recommendation near $x$. For any such point $x$, Assumption 2 requires that the points close to $x$ have similar conditional means of the outcome $Y_{zi}$ and treatment assignment $D_i(z)$.\textsuperscript{12} Note that Assumption 2 does not require continuity of the conditional means at $x$ for which $A(x) \in (0, 1)$, since the identification of the conditional means at such points follows from Property 1 without continuity.

Under the above assumptions, APS provides an easy-to-check condition for whether an algorithm allows us to identify causal effects.

\textsuperscript{10}This normalization is without loss of generality in the following sense. Take a vector $X_i^*$ of any continuous random variables and $A^*: \mathbb{R}^p \to [0, 1]$. The normalization induces the random vector $X_i = T(X_i^* - E[X_i^*])$, where $T$ is a diagonal matrix with diagonal entries $\frac{1}{\sqrt{\text{var}(X_{i1})^{1/2}}}, ..., \frac{1}{\sqrt{\text{var}(X_{ip})^{1/2}}}$. Let $A(x) = A^*(T^{-1}x + E[X_i^*])$. Then $(X_i^*, A^*)$ is equivalent to $(X_i, A)$ in the sense that $A(X_i) = A^*(X_i^*)$ for any individual $i$.

\textsuperscript{11}In the RDD with a single running variable, the point for which $p^A(x) \in (0, 1)$ and $A(x) \in \{0, 1\}$ is the cutoff point at which the treatment probability discontinuously changes.

\textsuperscript{12}In the context of the RDD with a single running variable, one sufficient condition for continuity of $E[Y_{zi}|X_i]$ is a local independence condition in the spirit of Hahn, Todd and van der Klaauw (2001): $(Y_i(1), Y_i(0), D_i(1), D_i(0))$ is independent of $X_i$ near $x$. A weaker sufficient condition, which allows such dependence, is that $E[Y_i(d)|D_i(1) = d_1, D_i(0) = d_0, X_i]$ and $\text{Pr}(D_i(1) = d_1, D_i(0) = d_0|X_i)$ are continuous at $x$ for every $d \in \{0, 1\}$ and $(d_1, d_0) \in [0, 1]^2$ (Dong, 2018). This assumes that the conditional means of the potential outcomes for each of the four types determined based on the potential treatment assignment $D_i(z)$ and the conditional probabilities of those types are continuous at the cutoff. These two sets of conditions are sufficient for continuity of $E[Y_{zi}|X_i]$ regardless of the dimension of $X_i$, accommodating multidimensional RDDs.
Proposition 1 (Identification). Under Assumptions 1 and 2:

(a) $E[Y_{1i} - Y_{0i}|X_i = x]$ and $E[D_i(1) - D_i(0)|X_i = x]$ are identified for every $x \in \text{int}(\mathcal{X})$ such that $p^A(x) \in (0,1)$.\textsuperscript{13}

(b) Let $S$ be any open subset of $\mathcal{X}$ such that $p^A(x)$ exists for all $x \in S$. Then either $E[Y_{1i} - Y_{0i}|X_i \in S]$ or $E[D_i(1) - D_i(0)|X_i \in S]$ or both are identified only if $p^A(x) \in (0,1)$ for almost every $x \in S$ (with respect to the Lebesgue measure).\textsuperscript{14}

Proof. See Appendix C.1.

Proposition 1 characterizes a necessary and sufficient condition for identification. Part (a) says that the average effects of the treatment recommendation $Z_i$ on the outcome $Y_i$ and on the treatment assignment $D_i$ for the individuals with $X_i = x$ are both identified if APS at $x$ is neither 0 nor 1. Non-degeneracy of APS at $x$ implies that there are both types of individuals who receive $Z_i = 1$ and $Z_i = 0$ among those whose $X_i$ is close to $x$. Assumption 2 ensures that these individuals are similar in terms of average potential outcomes and treatment assignments. We can therefore identify the average effects conditional on $X_i = x$. In Figure 1, $p^A(x) \in (0,1)$ holds for any $x$ in the shaded region (the union of the minor circular segment made by the chord $AC$ and the line segment $BD$).

Part (b) provides a necessary condition for identification. It says that if the average effect of the treatment recommendation conditional on $X_i$, being in some open set $S$ is identified, then we must have $p^A(x) \in (0,1)$ for almost every $x \in S$. If, to the contrary, there is a subset of $S$ of nonzero measure for which $p^A(x) = 1$ (or $p^A(x) = 0$), then $Z_i$ has no variation in the subset, which makes it impossible to identify the average effect for the subset.

Proposition 1 concerns causal effects of treatment recommendation, not of treatment assignment. The proposition implies that the conditional average treatment effects and the conditional local average treatment effects (LATEs) are identified under additional assumptions.

Corollary 1 (Perfect and Imperfect Compliance). Under Assumptions 1 and 2:

(a) The average treatment effect conditional on $X_i = x$, $E[Y_i(1) - Y_i(0)|X_i = x]$, is identified for every $x \in \text{int}(\mathcal{X})$ such that $p^A(x) \in (0,1)$ and $\Pr(D_i(1) > D_i(0)|X_i = x) = 1$ (perfect compliance).

(b) The local average treatment effect conditional on $X_i = x$, $E[Y_i(1) - Y_i(0)|D_i(1) \neq D_i(0), X_i = x]$, is identified for every $x \in \text{int}(\mathcal{X})$ such that $p^A(x) \in (0,1)$, $\Pr(D_i(1) \geq D_i(0)|X_i = x) = 1$ (monotonicity), and $\Pr(D_i(1) \neq D_i(0)|X_i = x) > 0$ (existence of compliers).

Proof. See Appendix C.2.

\textsuperscript{13}The causal effects may not be identified at a boundary point $x$ of $\mathcal{X}$ for which $p^A(x) \in (0,1)$. For example, if $A(x') = 1$ for all $x' \in B(x, \delta) \cap \mathcal{X}$ and $A(x^*) = 0$ for all $x^* \in B(x, \delta) \setminus \mathcal{X}$ for any sufficiently small $\delta > 0$, $p^A(x) \in (0,1)$ but the causal effects are not identified at $x$ since $\Pr(Z_i = 0|X_i \in B(x, \delta)) = 0$.

\textsuperscript{14}We assume that $p^A$ is a $\mathcal{L}^\infty$-measurable function so that $\{x \in S : p^A(x) = 0\}$ and $\{x \in S : p^A(x) = 1\}$ are $\mathcal{L}^\infty$-measurable.
Non-degeneracy of APS $p^A(x)$ therefore summarizes what causal effects the data from $A$ identify. Note that the key condition $(p^A(x) \in (0, 1))$ holds for some points $x$ for every standard algorithm except trivial algorithms that always recommend a treatment with probability 0 or 1. Therefore, the data from almost every algorithm identify some causal effect.

4 Estimation

The sources of quasi-random assignment characterized in Proposition 1 suggest a way of estimating causal effects of the treatment. In view of Proposition 1, it is possible to nonparametrically estimate conditional average causal effects $E[Y_1_i - Y_0_i | X_i = x]$ and $E[D_i(1) - D_i(0) | X_i = x]$ for points $x$ such that $p^A(x) \in (0, 1)$. This approach is hard to use in practice, however, when $X_i$ has many elements.

We instead seek an estimator that aggregates conditional effects at different points into a single average causal effect. Proposition 1 suggests that conditioning on APS makes algorithm-based treatment recommendation quasi-randomly assigned. This motivates the use of an algorithm’s recommendation as an instrument conditional on APS, which we operationalize as follows.

4.1 Two-Stage Least Squares Meets APS

Suppose that we observe a random sample $\{(Y_i, X_i, D_i, Z_i)\}_{i=1}^n$ of size $n$ from the population whose data generating process is as described in the introduction and Section 2. Consider the following 2SLS regression using the observations with $p^A(X_i; \delta_n) \in (0, 1)$:

$$D_i = \gamma_0 + \gamma_1 Z_i + \gamma_2 p^A(X_i; \delta_n) + \nu_i$$

$$Y_i = \beta_0 + \beta_1 D_i + \beta_2 p^A(X_i; \delta_n) + \epsilon_i,$$

where bandwidth $\delta_n$ shrinks toward zero as the sample size $n$ increases. Let $I_{i,n} = 1\{p^A(X_i; \delta_n) \in (0, 1)\}$, $D_{i,n} = (1, D_i, p^A(X_i; \delta_n))^\prime$, and $Z_{i,n} = (1, Z_i, p^A(X_i; \delta_n))^\prime$. The 2SLS estimator $\hat{\beta}$ is then given by

$$\hat{\beta} = (\sum_{i=1}^n Z_{i,n} D_{i,n}^\prime I_{i,n})^{-1} \sum_{i=1}^n Z_{i,n} Y_i I_{i,n}.$$

Let $\hat{\beta}_1$ denote the 2SLS estimator of $\beta_1$ in the above regression.

The above regression uses true APS $p^A(X_i; \delta_n)$, but it may be difficult to analytically compute if $A$ is complex. In such a case, we propose to approximate $p^A(X_i; \delta_n)$ using brute force simulation. We draw a value of $x$ from the uniform distribution on $B(X_i, \delta_n)$ a number of times, compute $A(x)$ for each draw, and take the average of $A(x)$ over the draws.\footnote{See Appendix A.5 for how to efficiently sample from the uniform distribution on a $p$-dimensional ball.} Formally, let $X_{i,1}^*, ..., X_{i,S_n}^*$ be $S_n$ independent draws from the uniform distribution on $B(X_i, \delta_n)$, and calculate

$$p^A(x; \delta_n) = \frac{1}{S_n} \sum_{s=1}^{S_n} A(X_{i,s}^*).$$
We compute $p^s(X_i; \delta_n)$ for each $i = 1, ..., n$ independently across $i$ so that $p^s(X_1; \delta_n), ..., p^s(X_n; \delta_n)$ are independent of each other. For fixed $n$ and $X_i$, the approximation error relative to true $p^A(X_i; \delta_n)$ has a $1/\sqrt{S_n}$ rate of convergence. This rate does not depend on the dimension of $X_i$, so the simulation error can be made negligible even when $X_i$ has many elements.

Now consider the following simulation version of the 2SLS regression using the observations with $p^s(X_i; \delta_n) \in (0, 1)$:

\begin{align*}
D_i &= \gamma_0 + \gamma_1 Z_i + \gamma_2 p^s(X_i; \delta_n) + \nu_i \\
Y_i &= \beta_0 + \beta_1 D_i + \beta_2 p^s(X_i; \delta_n) + \epsilon_i.
\end{align*}

Let $\hat{\beta}^s_1$ denote the 2SLS estimator of $\beta_1$ in the simulation-based regression. This regression is the same as the 2SLS regression (1) and (2) except that it uses the simulated APS $p^s(X_i; \delta_n)$ in place of $p^A(X_i; \delta_n)$.

### 4.2 Consistency and Asymptotic Normality

We establish the consistency and asymptotic normality of the 2SLS estimators $\hat{\beta}_1$ and $\hat{\beta}^s_1$. Our consistency and asymptotic normality result uses the following assumptions.

**Assumption 3.**

(a) (Finite Moment) $E[Y_i^4] < \infty$.

(b) (Nonzero First Stage) There exists a constant $c > 0$ such that $E[D_i(1) - D_i(0)|X_i = x] > c$ for every $x \in \mathcal{X}$ such that $p^A(x) \in (0, 1)$.

(c) (Nonzero Conditional Variance) If $Pr(A(X_i) \in (0, 1)) > 0$, then $Var(A(X_i)|A(X_i) \in (0, 1)) > 0$.

If $Pr(A(X_i) \in (0, 1)) = 0$, then the following conditions (d)-(g) hold.

(d) (Nonzero Variance) $Var(A(X_i)) > 0$.

For a set $S \subset \mathbb{R}^p$, let $cl(S)$ denote the closure of $S$ and let $\partial S$ denote the boundary of $S$, i.e., $\partial S = cl(S) \setminus int(S)$.

(e) ($C^2$ Boundary of $\Omega^*$) There exists a partition $\{\Omega^*_1, ..., \Omega^*_M\}$ of $\Omega^* = \{x \in \mathbb{R}^p : A(x) = 1\}$ (the set of the covariate points whose $A$ value is one) such that

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16 More precisely, we have $|p^s(X_i; \delta_n) - p^A(X_i; \delta_n)| = O_p(1/\sqrt{S_n})$, where $O_p^\circ$ indicates the stochastic boundedness in terms of the probability distribution of the $S_n$ simulation draws.

17 In many industry and policy applications, the analyst is only able to change the algorithm’s recommendation $Z_i$ by redesigning the algorithm. In this case, the effect of recommendation $Z_i$ on outcome $Y_i$ may also be of interest. We can estimate the effect of recommendation by running the following ordinary least squares (OLS) regression using the observations with $p^s(X_i; \delta) \in (0, 1)$:

\[ Y_i = \alpha_0 + \alpha_1 Z_i + \alpha_2 p^s(X_i; \delta) + u_i. \]

The estimated coefficient on $Z_i$, $\hat{\alpha}_1$, is our preferred estimator of the recommendation effect.
(i) \( \text{dist}(\Omega_m^*,\Omega_{m'}^*) > 0 \) for any \( m, m' \in \{1, \ldots, M\} \) such that \( m \neq m' \). Here \( \text{dist}(S,T) = \inf_{x \in S,y \in T} \|x - y\| \) is the distance between two sets \( S \) and \( T \subset \mathbb{R}^p \);

(ii) \( \Omega_m^* \) is nonempty, bounded, open, connected and twice continuously differentiable for each \( m \in \{1, \ldots, M\} \). Here we say that a bounded open set \( S \subset \mathbb{R}^p \) is twice continuously differentiable if for every \( x \in S \), there exists a ball \( B(x,\epsilon) \) and a one-to-one mapping \( \psi \) from \( B(x,\epsilon) \) onto an open set \( D \subset \mathbb{R}^p \) such that \( \psi^{-1} \) are twice continuously differentiable, \( \psi(B(x,\epsilon) \cap S) \subset \{(x_1, \ldots, x_p) \in \mathbb{R}^p : x_p > 0\} \) and \( \psi(B(x,\epsilon) \cap \partial S) \subset \{(x_1, \ldots, x_p) \in \mathbb{R}^p : x_p = 0\} \).

Let \( f_X \) denote the probability density function of \( X_i \) and let \( \mathcal{H}^k \) denote the \( k \)-dimensional Hausdorff measure on \( \mathbb{R}^p \).

(f) (Regularity of Deterministic \( A \))

(i) \( \mathcal{H}^{p-1}(\partial \Omega^*) < \infty \), and \( \int_{\partial \Omega^*} f_X(x) d\mathcal{H}^{p-1}(x) > 0 \).

(ii) There exists \( \delta > 0 \) such that \( A(x) = 0 \) for almost every \( x \in N(\mathcal{X}, \delta) \setminus \Omega^* \), where \( N(S,\delta) = \{x \in \mathbb{R}^p : \|x - y\| < \delta \text{ for some } y \in S\} \) for a set \( S \subset \mathbb{R}^p \) and \( \delta > 0 \).

(g) (Conditional Moments and Density near \( \partial \Omega^* \)) There exists \( \delta > 0 \) such that

(i) \( E[Y_{1i} | X_i], E[Y_{0i} | X_i], E[D_i(1) | X_i], E[D_i(0) | X_i] \) and \( f_X \) are continuously differentiable and have bounded partial derivatives on \( \mathcal{N}(\partial \Omega^*, \delta) \);

(ii) \( E[Y_{1i}^2 | X_i], E[Y_{0i}^2 | X_i], E[Y_{1i}D_i(1) | X_i], E[Y_{0i}D_i(0) | X_i] \) are continuous on \( \mathcal{N}(\partial \Omega^*, \delta) \);

(iii) \( E[Y_{1i}^4 | X_i] \) is bounded on \( \mathcal{N}(\partial \Omega^*, \delta) \).

Assumption 3 is a set of conditions for establishing consistency. Assumption 3 (b) assumes that, conditional on each value of \( X_i \) for which APS is nondegenerate, more individuals would change their treatment assignment status from 0 to 1 in response to treatment recommendation than would change it from 1 to 0. Under this assumption, the estimated first-stage coefficient on \( Z_i \) converges to a positive quantity. Note that, if there exists \( c < 0 \) such that \( E[D_i(1) - D_i(0) | X_i = x] < c \) for every \( x \in \mathcal{X} \) with \( p^A(x) \in (0,1) \), changing the labels of treatment recommendation makes Assumption 3 (b) hold.

Assumption 3 (c) rules out potential multicollinearity. If the support of \( A(X_i) \) contains only one value in \( (0,1) \), \( p^A(X_i; \delta_n) \) is asymptotically constant and equal to \( A(X_i) \) conditional on \( p^A(X_i; \delta_n) \in (0,1) \), resulting in the multicollinearity between \( p^A(X_i; \delta_n) \) and the constant term. Although dropping the constant term from the 2SLS regression solves this issue, Assumption 3 (c) allows us to only consider the regression with a constant for the purpose of simplifying the presentation. In Appendix C.3, we provide 2SLS estimators that are consistent and asymptotically normal even if we do not know whether Assumption 3 (c) holds.

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18. The \( k \)-dimensional Hausdorff measure on \( \mathbb{R}^p \) is defined as follows. Let \( \Sigma \) be the Lebesgue \( \sigma \)-algebra on \( \mathbb{R}^p \) (the set of all Lebesgue measurable sets on \( \mathbb{R}^p \)). For \( S \in \Sigma \) and \( \delta > 0 \), let \( \mathcal{H}^k_S(S) = \inf\{ \sum_{j=1}^{\infty} d(E_j)^k : S \subset \bigcup_{j=1}^{\infty} E_j, d(E_j) < \delta, E_j \subset \mathbb{R}^p \text{ for all } j \} \), where \( d(E) = \sup\{\|x-y\| : x, y \in E\} \). The \( k \)-dimensional Hausdorff measure of \( S \) on \( \mathbb{R}^p \) is \( \mathcal{H}^k(S) = \lim_{\delta \to 0} \mathcal{H}^k_S(S) \).

19. At the cost of making the presentation more complex, the assumption can be relaxed so that the sign of \( E[D_i(1) - D_i(0) | X_i = x] \) is allowed to vary over \( x \) with \( p^A(x) \in (0,1) \).
Assumption 3 (d)–(g) are a set of conditions we require for proving consistency and asymptotic normality of $\hat{\beta}_1$ when $A$ is deterministic and produces only multidimensional regression-discontinuity variation. Assumption 3 (d) says that $A$ produces variation in the treatment recommendation.

Assumption 3 (e) imposes the differentiability of the boundary of $\Omega^* = \{x \in \mathbb{R}^p : A(x) = 1\}$. The conditions are satisfied if, for example, $\Omega^* = \{x \in \mathbb{R}^p : f(x) \geq 0\}$ for some twice continuously differentiable function $f : \mathbb{R}^p \to \mathbb{R}$ such that $\nabla f(x) = (\frac{\partial f(x)}{\partial x_1}, \ldots, \frac{\partial f(x)}{\partial x_p})' \neq 0$ for all $x \in \mathbb{R}^p$ with $f(x) = 0$. $\Omega^*$ takes this form, for example, when the conditional treatment effect $E[Y_i(1) - Y_i(0) | X]$ is predicted by supervised learning based on smooth models such as lasso and ridge regressions, and treatment is recommended to individuals who are estimated to experience nonnegative treatment effects.

In general, the differentiability of $\Omega^*$ may not hold. For example, if tree-based algorithms such as Classification And Regression Tree (CART) and random forests are used to predict the conditional treatment effect, the predicted conditional treatment effect function is not differentiable at some points. Although the resulting $\Omega^*$ does not exactly satisfy Assumption 3 (e), the assumptions approximately hold in that $\Omega^*$ is arbitrarily well approximated by a set that satisfies the differentiability condition.\(^\text{20}\)

Part (i) of Assumption 3 (f) says that the boundary of $\Omega^*$ is $(p-1)$ dimensional and that the boundary has nonzero density. Part (ii) puts a weak restriction on the values $A$ takes on outside the support of $X_i$. It requires that $A(x) = 0$ for almost every $x \notin \Omega^*$ that is outside $\mathcal{X}$ but is in the neighborhood of $\mathcal{X}$. $A(x)$ may take on any value if $x$ is not close to $\mathcal{X}$. These conditions hold in practice.\(^\text{21}\) Assumption 3 (g) imposes continuity, continuous differentiability and boundedness on the conditional moments of potential outcomes and the probability density near the boundary of $\Omega^*$.

When $A$ is stochastic, asymptotic normality requires additional assumptions. Let

$$C^* = \{x \in \mathbb{R}^p : A \text{ is continuously differentiable at } x\},$$

and let $D^* = \mathbb{R}^p \setminus C^*$ be the set of points at which $A$ is not continuously differentiable.

**Assumption 4.** If $\Pr(A(X_i) \in (0,1)) > 0$, then the following conditions (a)–(c) hold.

(a) (Probability of Neighborhood of $D^*$) $\Pr(X_i \in N(D^*, \delta)) = O(\delta)$.

(b) (Bounded Partial Derivatives of $A$) The partial derivatives of $A$ are bounded on $C^*$.

(c) (Bounded Conditional Mean) $E[Y_i | X_i]$ is bounded on $\mathcal{X}$.

\(^{20}\)For example, suppose that $p = 2$, $A(x) = 1$ if $x_1 > 0$ and $x_2 > 0$, and $A(x) = 0$ otherwise. In this case, $\Omega^* = \{x \in \mathbb{R}^2 : x_1 > 0, x_2 > 0\}$. Let $\{\Omega_k\}_{k=1}^\infty$ be a sequence of subsets of $\mathbb{R}^2$, where $\Omega_k = \{x \in \mathbb{R}^2 : x_2 \geq \frac{1}{k}, x_1 > 0\}$ for each $k$. $\Omega_k$ is twice continuously differentiable for all $k$, and well approximates $\Omega^*$ for a large $k$ in that $d_H(\Omega^*, \Omega_k) \to 0$ as $k \to \infty$, where $d_H(S,T) = \max\{\sup_{s \in S} \inf_{t \in T} ||x - y||, \sup_{t \in T} \inf_{s \in S} ||x - y||\}$ is the Hausdorff distance between two sets $S$ and $T \subset \mathbb{R}^p$.

\(^{21}\)The boundary of $\Omega^*$ fails to be $(p-1)$ dimensional, for example, when the covariate space is three dimensional ($p = 3$) and $\Omega^*$ is a straight line, not a set with nonzero volume nor even a plane. In this example, the boundary is the same as $\Omega^*$, and its two-dimensional Hausdorff measure is zero.
Theorem 1

normal estimators of a weighted average treatment effect. In which case
\( p^A(x_\delta) = A(x_\delta) \) (as formally implied by Proposition A.2 in Appendix A.2). However, \( p^A(x_\delta; \delta) \) does not necessarily get sufficiently close to \( A(x_\delta) \) even as \( \delta \to 0 \), since \( x_\delta \) is in the \( \delta \)-neighborhood of \( D^* \) and hence \( A \) may discontinuously change within the \( \delta \)-ball \( B(x_\delta, \delta) \). Assumption 4 (a) requires that the probability of \( X_i \) being in the \( \delta \)-neighborhood of \( D^* \) shrink to zero at the rate of \( \delta_i \), which makes the points in the neighborhood negligible.

Assumption 4 (a) often holds in practice. If \( A \) is continuously differentiable on \( \mathcal{X} \), then \( D^* \cap \mathcal{X} = \emptyset \), so this condition holds. If, for example, the treatment recommendation is randomly assigned based on a stratified randomized experiment or on the \( \epsilon \)-Greedy algorithm (see Example A.1 (a) in Appendix A.6), \( D^* \) is the boundary at which the recommendation probability changes discontinuously. For any boundary of standard shape, the probability of \( X_i \) being in the \( \delta \)-neighborhood of the boundary vanishes at the rate of \( \delta_i \), and the required condition is satisfied. We provide a sufficient condition for this condition in Appendix A.4.

Assumption 4 (b) and (c) are regularity conditions, imposing the boundedness of the partial derivatives of \( A \) and of the conditional mean of the outcome.

Assumption 5 (The Number of Simulation Draws). \( n^{-1/2} S_n \to \infty \), and \( \Pr(p^A(X_i; \delta_n) \in (0, \gamma \log n / S_n) \cup (1 - \gamma \log n / S_n, 1)) = o(n^{-1/2} \delta_n^{1/2}) \) for some \( \gamma > 1/2 \).

Assumption 5 is the key to proving asymptotic normality of the simulation-based estimator \( \tilde{\beta}_1^s \). Assumption 5 says that we need to choose the number of simulation draws \( S_n \) so that it grows to infinity faster than \( n^{1/2} \), and that the probability that \( p^A(X_i; \delta_n) \) lies on the tails \( (0, \gamma \log n / S_n) \cup (1 - \gamma \log n / S_n, 1) \) vanishes faster than \( n^{-1/2} \delta_n^{1/2} \). This condition makes the bias caused by using \( p^s(X_i; \delta_n) \) instead of \( p^A(X_i; \delta_n) \) asymptotically negligible. To illustrate how the second part of this assumption restricts the rate at which \( S_n \) goes to infinity, consider an example where \( \Pr(p^A(X_i; \delta_n) \in (0, 1)) = O(\delta_n) \), and \( p^A(X_i; \delta_n) \) is approximately uniformly distributed on the tails \( (0, \gamma \log n / S_n) \cup (1 - \gamma \log n / S_n, 1) \). In this case, \( \Pr(p^A(X_i; \delta_n) \in (0, \gamma \log n / S_n) \cup (1 - \gamma \log n / S_n, 1)) = O(\delta_n \log n / S_n) \), and the second part of Assumption 5 requires that \( S_n \) grow sufficiently fast so that \( n^{1/2} \delta_n^{1/2} \log n / S_n = o(1) \). One choice of \( S_n \) satisfying this is \( S_n = \alpha n^\kappa \delta_n^{1/2} \) for some \( \alpha > 0 \) and \( \kappa > 1/2 \), in which case \( n^{1/2} \delta_n^{1/2} \log n / S_n = \log n / \alpha n^{\kappa - 1/2} = o(1) \).

Under the above conditions, the 2SLS estimators \( \tilde{\beta}_1 \) and \( \tilde{\beta}_1^s \) are consistent and asymptotically normal estimators of a weighted average treatment effect.

Theorem 1 (Consistency and Asymptotic Normality). Suppose that Assumptions 1 and 3 hold, and that \( \delta_n \to 0 \), \( n \delta_n \to \infty \) and \( S_n \to \infty \) as \( n \to \infty \). Then the 2SLS estimators \( \tilde{\beta}_1 \) and \( \tilde{\beta}_1^s \) converge in probability to
\[
\beta_1 \equiv \lim_{\delta \to 0} E[\omega_i(\delta)(Y_i(1) - Y_i(0))],
\]
where
\[
\omega_i(\delta) = \frac{p^A(X_i; \delta)(1 - p^A(X_i; \delta))(D_i(1) - D_i(0))}{E[p^A(X_i; \delta)(1 - p^A(X_i; \delta))(D_i(1) - D_i(0))]}.
\]
Suppose, in addition, that Assumptions 4 and 5 hold and that \( n\sigma^2_{\text{A}} \to 0 \) as \( n \to \infty \). Then

\[
\hat{\sigma}_n^{-1}(\hat{\beta}_1 - \beta_1) \xrightarrow{d} \mathcal{N}(0, 1),
\]

\[
(\hat{\sigma}^2_n)^{-1}(\hat{\beta}_1 - \beta_1) \xrightarrow{d} \mathcal{N}(0, 1),
\]

where we define \( \hat{\sigma}_n^{-1} \) and \( (\hat{\sigma}^2_n)^{-1} \) as follows. Let

\[
\hat{\Sigma}_n = \left( \sum_{i=1}^{n} Z_{i,n}D_{i,n}I_{i,n} \right)^{-1}\left( \sum_{i=1}^{n} \hat{\epsilon}_{i,n}^2 Z_{i,n}Z_{i,n}'I_{i,n} \right)\left( \sum_{i=1}^{n} D_{i,n}Z_{i,n}Z_{i,n}'I_{i,n} \right)^{-1},
\]

where

\[
\hat{\epsilon}_{i,n} = Y_i - D_{i,n}\hat{\beta}.
\]

\( \hat{\Sigma}_n \) is the conventional heteroskedasticity-robust estimator for the variance of the 2SLS estimator. \( \hat{\sigma}_n^2 \) is the second diagonal element of \( \hat{\Sigma}_n \). \( (\hat{\sigma}^2_n)^2 \) is the analogously-defined estimator for the variance of \( \hat{\beta}_1 \) from the simulation-based regression.

**Proof.** See Appendix C.3.

Theorem 1 says that the 2SLS estimators converge to a weighted average of causal effects for the subpopulation whose APS is nondegenerate \( (p^A(X_i; \delta) \in (0, 1)) \) and who would switch their treatment status in response to the treatment recommendation \( (D_i(1) \neq D_i(0)) \). The limit \( \lim_{\delta \to 0} E[\delta_i(\delta)(Y_i(1) - Y_i(0))] \) always exists under the assumptions of Theorem 1. It also shows that inference based on the conventional 2SLS heteroskedasticity-robust standard errors is asymptotically valid if \( \delta_n \) goes to zero at an appropriate rate. The convergence rate of \( \hat{\beta}_1 \) is \( O_p(1/\sqrt{n}) \) if \( \Pr(A(X_i) \in (0, 1)) > 0 \) and is \( O_p(1/\sqrt{n\delta_n}) \) if \( \Pr(A(X_i) \in (0, 1)) = 0 \).

Our consistency result requires that \( \delta_n \) go to zero slower than \( n^{-1} \). The rate condition ensures that, when \( \Pr(A(X_i) \in (0, 1)) = 0 \), we have sufficiently many observations in the \( \delta_n \)-neighborhood of the boundary of \( \Omega^* \). Importantly, the rate condition does not depend on the dimension of \( X_i \), unlike other bandwidth-based estimation methods such as kernel methods. This is because we use all the observations in the \( \delta \)-neighborhood of the boundary, and the number of those observations is of order \( n\delta_n \) regardless of the dimension of \( X_i \) if the dimension of the boundary is one less than the dimension of \( X_i \), i.e., \( (p - 1) \).

The asymptotic normality result requires that \( \delta_n \) go to zero sufficiently quickly so that \( n\delta^2_n \to 0 \). When \( \Pr(A(X_i) \in (0, 1)) > 0 \), we need to use a small enough \( \delta_n \) so that \( p^A(X_i; \delta_n) \) converges to \( p^A(X_i) \) at a fast rate and \( \delta_n \)-neighborhood of \( D^* \) is asymptotically small enough. When \( \Pr(A(X_i) \in (0, 1)) = 0 \), the asymptotic normality is based on undersmoothing, which eliminates the asymptotic bias by using the observations sufficiently close to the boundary of \( \Omega^* \). In both

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22 In principle, it is possible to estimate other weighted averages and the unweighted average by reweighting different observations appropriately. For example, we can estimate the unweighted average treatment effect by weighting observations by the inverse of APS.

23 In the standard RDD with a single running variable \( X_i \) and cutoff \( c \), \( p^A(X_i; \delta) = \frac{\delta}{2\sigma_n} + \frac{1}{2} \) if \( X_i \in [c - \delta_n, c + \delta_n] \) and \( p^A(X_i; \delta_n) \in (0, 1] \) otherwise. Since \( p^A(X_i; \delta_n) \) is linear in the running variable \( X_i \), if \( p^A(X_i; \delta_n) \in (0, 1) \), the estimator \( \hat{\beta}_1 \) becomes a local regression estimator with the box kernel that places the same slope coefficient of \( X_i \) on both sides of the cutoff. Under our assumptions, \( \hat{\beta}_1 \) and standard local linear estimators are shown to have the same fastest possible convergence rate.
cases, the bias of our estimator is \(O(\delta_n)\). On the other hand, the standard deviation is \(O(1/\sqrt{n})\) when \(\Pr(A(X_i) \in (0, 1)) > 0\) and is \(O(1/\sqrt{n}\delta_n)\) when \(\Pr(A(X_i) \in (0, 1)) = 0\). The condition that \(n\delta_n^2 \to 0\) ensures that the bias converges to zero faster than the standard deviation in either case.

Whether or not \(\Pr(A(X_i) \in (0, 1)) = 0\), when we use simulated APS, the consistency result requires that the number of simulation draws \(S_n\) go to infinity as \(n\) increases. The asymptotic normality result requires a sufficiently fast growth rate of \(S_n\) given by Assumption 5 to make the bias caused by using \(p^s(X_i; \delta_n)\) negligible.

Finally, note that the weight \(\omega_i(\delta)\) given in Theorem 1 is negative if \(D_i(1) < D_i(0)\), so \(E[\omega_i(\delta)(Y_i(1) - Y_i(0))]\) may not be a causally interpretable convex combination of treatment effects \(Y_i(1) - Y_i(0)\). This can happen because the treatment effect of those whose treatment assignment switches from 1 to 0 in response to the treatment recommendation (defiers) negatively contributes to \(E[\omega_i(\delta)(Y_i(1) - Y_i(0))]\). Additional assumptions prevent this problem. If the treatment effect is constant, for example, the 2SLS estimators are consistent for the treatment effect.

**Corollary 2.** Suppose that Assumptions 1 and 3 hold, that the treatment effect is constant, i.e., \(Y_i(1) - Y_i(0) = b\) for some constant \(b\), and that \(\delta_n \to 0\), \(n\delta_n \to \infty\), and \(S_n \to \infty\) as \(n \to \infty\). Then the 2SLS estimators \(\hat{\beta}_1\) and \(\hat{\beta}_1^s\) converge in probability to \(b\).

Another approach is to impose monotonicity (Imbens and Angrist, 1994). Let \(LATE(x) = E[Y_i(1) - Y_i(0)|D_i(1) \neq D_i(0), X_i = x]\) be the local average treatment effect (LATE) conditional on \(X_i = x\).

**Corollary 3.** Suppose that Assumptions 1 and 3 hold, that \(\Pr(D_i(1) \geq D_i(0)|X_i = x) = 1\) for any \(x \in X\) with \(p^A(x) \in (0, 1)\), and that \(\delta_n \to 0\), \(n\delta_n \to \infty\) and \(S_n \to \infty\) as \(n \to \infty\). Then the 2SLS estimators \(\hat{\beta}_1\) and \(\hat{\beta}_1^s\) converge in probability to

\[
\lim_{\delta \to 0} E[\omega(X_i; \delta)LATE(X_i)],
\]

where

\[
\omega(x; \delta) = \frac{p^A(x; \delta)(1 - p^A(x; \delta))E[D_i(1) - D_i(0)|X_i = x]}{E[p^A(X_i; \delta)(1 - p^A(X_i; \delta))(D_i(1) - D_i(0))]|D_i(1) - D_i(0)\}
\]

The 2SLS estimators are consistent for a weighted average of conditional LATEs over all values of \(X_i\) with nondegenerate APS \(p^A(X_i; \delta_n)\). The weights are proportional to \(p^A(X_i; \delta_n)(1 - p^A(X_i; \delta_n))\), and to the proportion of compliers, \(E[D_i(1) - D_i(0)|X_i]\).

### 4.3 Special Cases

The result in Theorem 1 holds whether \(A\) is stochastic (\(\Pr(A(X_i) \in (0, 1)) > 0\)) or deterministic (\(\Pr(A(X_i) \in (0, 1)) = 0\)). As shown in the proof of Theorem 1 in Appendix C.3, if we consider these two underlying cases separately, the probability limit of the 2SLS estimators has a more specific expression. If \(\Pr(A(X_i) \in (0, 1)) > 0\),

\[
\text{plim} \hat{\beta}_1 = \text{plim} \hat{\beta}_1^s = \frac{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))(Y_i(1) - Y_i(0))]}{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))]}.
\]

(5)
The 2SLS estimators converge to a weighted average of treatment effects for the subpopulation with nondegenerate \( A(X_i) \). To relate this result to existing work, consider the following 2SLS regression with the (standard) propensity score \( A(X_i) \) control:

\[
D_i = \gamma_0 + \gamma_1 Z_i + \gamma_2 A(X_i) + \nu_i
\]

\[
Y_i = \beta_0 + \beta_1 D_i + \beta_2 A(X_i) + \epsilon_i.
\]

Existing results show that under conditional independence, the 2SLS estimator from this regression converges in probability to the treatment-variance weighted average of treatment effects in (5) (Angrist and Pischke, 2008; Hull, 2018).24 Not surprisingly, for this selection-on-observables case, our result shows that the 2SLS estimator is consistent for the same treatment effects in (5) (Angrist and Pischke, 2008; Hull, 2018).

Importantly, using APS as a control allows us to consistently estimate a causal effect even if \( A \) is deterministic and produces multidimensional regression-discontinuity variation. If we use as a control the propensity score, APS, or simulated APS.

Precisely speaking, Angrist and Pischke (2008) consider the OLS regression of \( Y_i \) (or \( D_i \)) on \( Z_i \) controlling a dummy variable for every value taken on by \( X_i \) (i.e., the model is saturated in \( X_i \)) when \( X_i \) is a discrete variable:

\[
Y_i = \alpha_1 Z_i + \sum_{x \in \mathcal{X}} \alpha_{2,x} 1\{X_i = x\} + u_i.
\]

By the Frisch-Waugh Theorem, the population coefficient on \( Z_i \) from (8) is given by \( \alpha_1 = \frac{E[(Z_i - E[Z_i|X_i])Y_i]}{E[(Z_i - E[Z_i|X_i])^2]} \). Angrist and Pischke (2008) show that this expression is reduced to the treatment-variance weighted average of treatment effects \( \frac{E[A(X_i)(1-A(X_i))(Y_{i1} - Y_{i0})]}{E[A(X_i)(1-A(X_i))^2]} \) under the conditional independence assumption. Their derivation follows even when \( X_i \) is continuous and we control the propensity score linearly.

25Our approach using geometric theory shows that \( \hat{\beta}_1 \) converges to an integral of the conditional treatment effect over boundary points with respect to the Hausdorff measure. In contrast, prior studies on multidimensional RDDs express treatment effect estimates in terms of expectations conditional on \( X_i \) being in the boundary like \( E[Y_{i1} - Y_{i0}|X_i \in \partial \mathcal{Y}] \) (Zajonc, 2012). However, those conditional expectations are, formally, not well-defined, since \( \mathcal{L}^p(\partial \mathcal{Y}) = 0 \) and hence \( \Pr(X_i \in \partial \mathcal{Y}) = 0 \). We therefore prefer our expression in terms of an integral with respect to the Hausdorff measure to any expressions in terms of conditional expectations on the boundary. Arias, Rubio-Ramírez and Waggoner (2018), Bornn, Shephard and Solgi (2019), and Qiao (2021) use similar tools from differential geometry and geometric measure theory, but for different purposes.
over \( N(\partial \Omega^*, \delta) \) in terms of the iterated integral over the levels sets of the signed distance function of \( \Omega^* \):

\[
\int_{N(\partial \Omega^*, \delta)} g(x)dx = \int_{-\delta}^{\delta} \int_{\{x' \in \mathbb{R}^p : d_{\Omega^*}^*(x') = \lambda\}} g(x)d\mathcal{H}^{p-1}(x)d\lambda,
\]

where \( d_{\Omega^*}^* \) is the signed distance function of \( \Omega^* \) (see Appendix B.2 for the definition). The set \( \{x' \in \mathbb{R}^p : d_{\Omega^*}^*(x') = \lambda\} \) is a level set of \( d_{\Omega^*}^* \), which collects the points in \( \Omega^* \) when \( \lambda > 0 \) and the points in \( \mathbb{R}^p \setminus \Omega^* \) when \( \lambda < 0 \) whose distance to the boundary \( \partial \Omega^* \) is \( |\lambda| \). Figure 2a shows a visual illustration.

We then use the area formula (Lemma B.4 in Appendix B.3) to write the integral over each level set in terms of the integral over the boundary \( \partial \Omega^* \):

\[
\int_{\{x' \in \mathbb{R}^p : d_{\Omega^*}^*(x') = \lambda\}} g(x)d\mathcal{H}^{p-1}(x) = \int_{\partial \Omega^*} g(x^* + \lambda \nu_{\Omega^*}(x^*)) J_{\mathbb{R}^{p-1}}^\partial \psi_{\Omega^*}(x^*, \lambda) d\mathcal{H}^{p-1}(x^*),
\]

where \( \nu_{\Omega^*}(x^*) \) is the inward unit normal vector of \( \partial \Omega^* \) at \( x^* \) (the unit vector orthogonal to all vectors in the tangent space of \( \partial \Omega^* \) at \( x^* \) that points toward the inside of \( \Omega^* \)), and \( J_{\mathbb{R}^{p-1}}^\partial \psi_{\Omega^*}(x^*, \lambda) \) is the Jacobian of the transformation \( \psi_{\Omega^*}(x^*, \lambda) = x^* + \lambda \nu_{\Omega^*}(x^*) \). Figure 2b illustrates this change of variables formula. Finally, combining (10) and (11) and proceeding with further analysis, we prove in Appendix C.3.3 that when \( g \) is continuous,

\[
\int_{N(\partial \Omega^*, \delta)} g(x)dx = \delta \left( \int_{\partial \Omega^*} g(x)d\mathcal{H}^{p-1}(x) + o(1) \right).
\]

Thus, the integral over the \( \delta \)-neighborhood of \( \partial \Omega^* \) scaled up by \( \delta^{-1} \) converges to the integral over boundary points with respect to the \( (p - 1) \)-dimensional Hausdorff measure. This result is used to derive the expression of the probability limit of \( \hat{\beta}_1 \) given by (9).

## 5 Machine Learning Simulation

This section conducts a Monte Carlo experiment to assess the feasibility and performance of our method. Consider a tech company that conducts a randomized controlled trial (RCT) using a small segment of the population. At the same time, the company applies a deterministic decision algorithm to the rest of the population. This mimics a situation in which the decision maker first conducts an experiment that randomly assigns \( Z_i \) to predict the conditional average effect of \( Z_i \) and then constructs an algorithm that greedily chooses the treatment predicted to perform better based on the predicted effect.

We simulate 1,000 hypothetical samples from the following data-generating process. Each sample \( \{(Y_i, X_i, D_i, Z_i)\}_{i=1}^n \) is of size \( n = 10,000 \). There are 100 covariates \( (p = 100) \), and \( X_i \sim \mathcal{N}(0, \Sigma) \). \( Y_i(0) \) is generated as \( Y_i(0) = 0.75X_i'\alpha_0 + 0.25\epsilon_{0i} \), where \( \alpha_0 \in \mathbb{R}^{100} \), and \( \epsilon_{0i} \sim \mathcal{N}(0, 1) \). We consider two models for \( Y_i(1) \), one in which the treatment effect \( Y_i(1) - Y_i(0) \) does not depend on \( X_i \) and one in which the treatment effect depends on \( X_i \).

Model A. \( Y_i(1) = Y_i(0) + \epsilon_{1i} \), where \( \epsilon_{1i} \sim \mathcal{N}(0, 1) \).
Model B. $Y_i(1) = Y_i(0) + X_i'\alpha_1$, where $\alpha_1 \in \mathbb{R}^{100}$.

The choice of parameters $\Sigma$, $\alpha_0$ and $\alpha_1$ is explained in Appendix D. $D_i(0)$ and $D_i(1)$ are generated as $D_i(0) = 0$ and $D_i(1) = 1\{Y_i(1) - Y_i(0) > u_i\}$, where $u_i \sim \mathcal{N}(0,1)$. To generate $Z_i$, let $q_{0.495}$ and $q_{0.505}$ be the 49.5th and 50.5th (empirical) quantiles of the first covariate $X_{i1}$.

Let $\tau_{pred}(X_i)$ be a real-valued function of $X_i$, which we regard as a prediction of the effect of recommendation on the outcome for individual $i$ obtained from past data. We construct $\tau_{pred}$ using an independent sample $\{(\tilde{Y}_i, \tilde{X}_i, \tilde{D}_i, \tilde{Z}_i)\}_{i=1}^{n}$ of size $n = 2,000$. The distribution of $(\tilde{Y}_i, \tilde{X}_i, \tilde{D}_i, \tilde{Z}_i)$ is the same as that of $(Y_i, X_i, D_i, Z_i)$ except (1) that $\tilde{Y}_i(1)$ is generated as $\tilde{Y}_i(1) = \tilde{Y}_i(0) + 0.5\tilde{X}_i'\alpha_1 + 0.5\epsilon_{1i}$, where $\epsilon_{1i} \sim \mathcal{N}(0,1)$ and (2) that $\tilde{Z}_i \sim \text{Bernoulli}(0.5)$. This can be viewed as data from a past randomized experiment conducted to construct the algorithm.

We then use random forests separately for the subsamples with $\tilde{Z}_i = 1$ and $\tilde{Z}_i = 0$ to predict $\tilde{Y}_i$ from $\tilde{X}_i$. Let $\mu_2(x)$ be the trained prediction model. Set $\tau_{pred}(x) = \mu_1(x) - \mu_0(x)$. We generate the sample $\{(\tilde{Y}_i, \tilde{X}_i, \tilde{D}_i, \tilde{Z}_i)\}_{i=1}^{n}$ and construct $\tau_{pred}$ only once, and we use it for all of the 1,000 samples. The distribution of the sample $\{(Y_i, X_i, D_i, Z_i)\}_{i=1}^{n}$ is thus held fixed for all simulations.

$Z_i$ is then generated as

$$Z_i = \begin{cases} Z^*_i \sim \text{Bernoulli}(0.5) & \text{if } X_{i1} \in [q_{0.495}, q_{0.505}] \\ 1 & \text{if } X_{i1} \notin [q_{0.495}, q_{0.505}] \text{ and } \tau_{pred}(X_i) \geq 0 \\ 0 & \text{if } X_{i1} \notin [q_{0.495}, q_{0.505}] \text{ and } \tau_{pred}(X_i) < 0. \end{cases}$$

The first case corresponds to the RCT segment while the latter two cases to the deterministic algorithm segment. The function $A$ is given by

$$A(x) = \begin{cases} 0.5 & \text{if } x_1 \in [q_{0.495}, q_{0.505}] \\ 1 & \text{if } x_1 \notin [q_{0.495}, q_{0.505}] \text{ and } \tau_{pred}(x) \geq 0 \\ 0 & \text{if } x_1 \notin [q_{0.495}, q_{0.505}] \text{ and } \tau_{pred}(x) < 0. \end{cases}$$

Finally, $D_i$ and $Y_i$ are generated as $D_i = Z_iD_i(1) + (1-Z_i)D_i(0)$ and $Y_i = D_iY_i(1) + (1-D_i)Y_i(0)$, respectively.

**Estimators and Estimands.** We use the data $\{(Y_i, X_i, D_i, Z_i)\}_{i=1}^{n}$ to estimate treatment effect parameters. Our main approach is 2SLS with APS controls in Theorem 1. To compute APS, we use $S = 400$ simulation draws for each observation.

We compare our approach with two alternatives. The first alternative is OLS of $Y_i$ on a constant and $D_i$ (i.e., the difference in the sample mean of $Y_i$ between the treated group and untreated group) using all observations. The second alternative is 2SLS with $A$ controls. This method uses the observations with $A(X_i) \in (0,1)$ to run the 2SLS regression of $Y_i$ on a constant, $D_i$, and $A(X_i)$ using $Z_i$ as an instrument for $D_i$ (see (6) and (7) in Section 4.3) and reports the coefficient on $D_i$.

We consider four parameters as target estimands: $\text{ATE} \equiv E[Y_i(1) - Y_i(0)]$, $\text{ATE(RCT)} \equiv E[Y_i(1) - Y_i(0)|X_{i1} \in [q_{0.495}, q_{0.505}]]$, $\text{LATE} \equiv E[Y_i(1) - Y_i(0)|D_i(1) \neq D_i(0)]$, and $\text{LATE(RCT)} \equiv E[Y_i(1) - Y_i(0)|D_i(1) \neq D_i(0), X_{i1} \in [q_{0.495}, q_{0.505}]]$. In the case where the treatment effect does not depend on $X_i$ (Model A), the conditional effects are homogeneous, so that $\text{ATE}$ and $\text{LATE}$
are the same as ATE(RCT) and LATE(RCT), respectively. In the case where the treatment effect depends on $X_i$ (Model B), the conditional effects are heterogeneous. However, since the RCT segment consists of those in the middle of the distribution of $X_{i1}$, the average effect for the RCT segment is close to the unconditional average effect. As a result, ATE is equal to ATE(RCT) and LATE is similar to LATE(RCT) under this data-generating process.

For both models, the 2SLS estimator converges in probability to LATE(RCT) whether we control for APS or $A$. However, 2SLS with $A$ controls uses only the individuals for the RCT segment while 2SLS with APS controls additionally uses the individuals near the decision boundary of the deterministic algorithm (i.e., the boundary of the region for which $\tau_{pred}(x) \geq 0$). Therefore, 2SLS with APS controls is expected to produce a more precise estimate than 2SLS with $A$ controls if the conditional effects for those near the boundary are not far from the target estimand.

Results. Table 1 reports the bias, standard deviation (SD), and root mean squared error (RMSE) of each estimator. Panels A and B present the results for the cases where the conditional effects are homogeneous and heterogeneous, respectively. Note first that OLS with no controls is significantly biased, showing the importance of correcting for omitted variable bias. 2SLS with APS achieves this goal, as demonstrated by its smaller biases across all possible treatment effect models, target parameters, and values of the bandwidth $\delta$. 2SLS with APS controls shows a consistent pattern; as the bandwidth $\delta$ grows, the bias increases while the variance declines. For several values of $\delta$, 2SLS with APS controls outperforms 2SLS with $A$ controls in terms of the RMSE. This finding implies that exploiting individuals near the multidimensional decision boundary of the deterministic algorithm can lead to better performance than using only the individuals in the RCT segment.

We also evaluate our inference procedure based on Theorem 1. Table 1 reports the coverage probabilities of the 95% confidence intervals for LATE(RCT) constructed from the 2SLS estimates and their heteroskedasticity-robust standard errors. The confidence intervals offer nearly correct coverage when $\delta$ is small, which supports the implication of Theorem 1 that the inference procedure is valid when we use a sufficiently small $\delta$. Overall, Table 1 shows that our estimator is feasible in a high-dimensional setting and performs better than alternative estimators.

## 6 Empirical Policy Application

### 6.1 Hospital Relief Funding during the COVID-19 Pandemic

Here we provide our real-world empirical application. As part of the 3-phase Coronavirus Aid, Relief, and Economic Security (CARES) Act, the government has distributed tens of billions of dollars of relief funding to hospitals since April 2020. This funding intended to help health care providers hit hardest by the COVID-19 outbreak and at a high risk of closing. The bill specified that providers may (but are not required to) use the funds for COVID-19-related expenses, such as construction of temporary structures, leasing of properties, purchasing medical supplies and

\[26\] The 2SLS estimators converge in probability to the right-hand side of equation (5), which is the same as LATE(RCT) under the data-generating process of this simulation.
equipment (including personal protective equipment and testing supplies), increased workforce utilization and training, establishing emergency operation centers, retrofitting facilities and managing the surge in capacity, among others. We are interested in whether this funding had a causal impact on hospital operation and activities in dealing with COVID-19 cases.

We focus on an initial portion of this funding ($10 billion), which was allocated to hospitals that qualified as “safety net hospitals” according to a specific eligibility criterion. This eligibility criterion intends to direct funding towards hospitals that “disproportionately provide care to the most vulnerable, and operate on thin margins.” Specifically, an acute care hospital was deemed eligible for funding if the following conditions hold:

- Medicare Disproportionate Patient Percentage (DPP) of 20.2% or greater. DPP is equal to the sum of the percentage of Medicare inpatient days attributable to patients eligible for both Medicare Part A and Supplemental Security Income (SSI), and the percentage of total inpatient days attributable to patients eligible for Medicaid but not Medicare Part A.\(^\text{27}\)

- Annual Uncompensated Care (UCC) of at least $25,000 per bed. UCC is a measure of hospital care provided for which no payment was received from the patient or insurer. It is the sum of a hospital’s bad debt and the financial assistance it provides.\(^\text{28}\)

- Profit Margin (Net income/(Net patient revenue + Total other income)) of 3.0% or less.

Hospitals that do not qualify on any of the three dimensions are funding ineligible. Figure 4 visualizes how the three dimensions determine safety net eligibility. As the bottom two-dimensional planes show, eligibility discontinuously changes as hospitals cross the eligibility boundary in the space of the three characteristics. This setting is a three-dimensional RDD, falling under our framework.

The final funding amount is calculated as follows. Each eligible hospital is assigned an individual facility score, which is calculated as the product of DPP and the number of beds in that hospital. This facility score determines the share of funding allocated to the hospital, out of the total $10 billion. The share received by each hospital is determined by the ratio of the hospital’s facility score to the sum of facility scores across all eligible hospitals. The amount of funding that can be received is bounded below at $5 million and capped above at $50 million. Figure 3 shows the distribution of funding amounts received by eligible hospitals. A majority of eligible hospitals receive the minimum amount of $5 million.

We use publicly available data from the Healthcare Cost Report Information System (HCRIS) for the 2018 financial year, to replicate the funding eligibility status as well as the amount of funding received.\(^\text{29}\) To obtain outcome measures of interest, we use the publicly available COVID-19 Reported Patient Impact and Hospital Capacity by Facility dataset. This provides \(^\text{27}\)Source: https://www.cms.gov/Medicare/Medicare-Fee-for-Service-Payment/AcuteInpatientPPS/dsh
\(^\text{28}\)Source: https://www.aha.org/fact-sheets/2020-01-06-fact-sheet-uncompensated-hospital-care-cost
\(^\text{29}\)We use the methodology detailed in the CARES ACT website to project funding based on 2018 financial year cost reports. We use the RAND cleaned version of the dataset which can be accessed at https://www.hospitaldatasets.org/
facility-level data on hospital utilization aggregated on a weekly basis, from July 31st 2020 onwards. Summary statistics about hospital outcomes and characteristics are documented in Table 2. Safety net hospitals have higher fractions of inpatient and ICU beds occupied by COVID-19 patients. Safety net hospitals also have a higher disproportionate patient percentage, higher uncompensated care per bed, lower profit margins, more employees and beds, and shorter lengths of inpatient stay.

6.2 Covariate Balance Estimates

Using the above data, we study the effect of safety net funding on relevant hospital outcomes, such as the total number of inpatient beds and the number of staffed ICU beds occupied by adult COVID patients reported between July 31st 2020 and August 6th 2020.

We first evaluate the balancing property of APS conditioning using APS-controlled differences in covariate means for hospitals who are and are not deemed eligible for safety net funding. Specifically, we run the following OLS regression of hospital-level characteristics on the eligibility status using observations with $p^s(X_i; \delta_n) \in (0, 1)$:

$$W_i = \gamma_0 + \gamma_1 Z_i + \gamma_2 p^s(X_i; \delta_n) + \eta_i,$$

where $W_i$ is one of the predetermined characteristics of the hospital, $Z_i$ is a funding eligibility dummy, $X_i$ is a vector of three input variables (DPP, UCC, and profit margin) that determine the funding eligibility, and $p^s(X_i; \delta_n)$ is the simulated APS. We compute APS using $S = 10,000$ simulation draws. The estimated coefficient on $Z_i$ is the APS-controlled difference in the mean of the covariate between eligible and ineligible hospitals. For comparison, we also run the OLS regression of hospital characteristics on the eligibility status with no controls using the whole sample.

Table 3 reports the covariate balance estimates. Column 1 shows that, without controlling for APS, eligible hospitals are significantly different from ineligible hospitals. We find that all the relevant hospital eligibility characteristics are strongly associated with eligibility. Once we control for APS with small enough bandwidth $\delta$, eligible and ineligible hospitals have similar financial and utilization characteristics, as reported in columns 2–6 of Table 3. These estimates are consistent with our theoretical results, establishing the empirical relevance of APS controls.

6.3 2SLS Estimates

The balancing performance of APS motivates us to estimate causal effects of safety net funding by 2SLS using funding eligibility as an instrument for the amount of funding received. We run the following 2SLS regression on four different hospital-level outcome variables, using hospitals

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30 Source: https://healthdata.gov/Hospital/COVID-19-Reported-Patient-Impact-and-Hospital-Capa/anag-cv7u

31 Figure A.2 in Appendix E.4 reports APS for several hospitals with varying numbers of simulation draws. We find that $S = 10,000$ is sufficient for well stabilizing APS simulation.
with \( p^s(X_i; \delta) \in (0, 1) \):

\[
D_i = \gamma_0 + \gamma_1 Z_i + \gamma_2 p^s(X_i; \delta) + v_i
\]

\[
Y_i = \beta_0 + \beta_1 D_i + \beta_2 p^s(X_i; \delta) + \epsilon_i,
\]

where \( Y_i \) is a hospital-level outcome and \( D_i \) is the amount of relief funding received.\(^{32}\) We also run the OLS and 2SLS regressions with no controls, as well as OLS regression controlling for the three eligibility determinants (disproportionate patient percentage, uncompensated care per bed and profit margin).\(^{33}\) These alternative regressions are computed using the sample of all hospitals, as benchmark estimators.

The first stage effects of safety net eligibility on funding amount (in millions), shown in columns 3–10 of Table 4, suggest that safety net eligibility boosts funding significantly. For example, in column 3 of Table 4, we can see that funding eligibility increases the safety net funding by approximately 15 million dollars.

OLS estimates of funding effects, reported as the benchmark in column 1 of Table 4, indicate that safety net funding is associated with a higher number of adult inpatient beds and higher number of staffed ICU beds utilized by patients who have lab-confirmed or suspected COVID. The estimates indicate that a million dollar increase in funding is associated with 5.58 more adult inpatient beds occupied by patients with lab-confirmed or suspected COVID. The corresponding increase in total adult inpatient beds occupied by those who have lab-confirmed COVID is 4.53 and the increase in staffed ICU beds occupied by those who have lab-confirmed or suspected COVID is 1.67. The estimated increase in staffed ICU beds occupied by lab-confirmed COVID patients is 1.51. Naive 2SLS estimates with no controls and OLS with covariate controls produce similar significantly positive results.

In contrast with the OLS or uncontrolled 2SLS estimates, the 2SLS estimates with APS controls in columns 4–10 show a different picture. The gains in number of inpatient beds and staffed ICU beds occupied by suspected and lab-confirmed COVID patients become much smaller and lose significance across all bandwidth specifications. These results suggest that APS reveals important selection bias in the estimated effects of safety net funding. Once we control for APS

\(^{32}\)This specification uses a continuous treatment, unlike our theoretical framework with a binary treatment. We obtain similar results when the treatment is a binary transformation of the amount of relief funding received (e.g., a dummy indicating whether the amount exceeds a certain value). Results are available upon request.

\(^{33}\)Precisely speaking, we run the following specification for each hospital-level outcome variable \( Y_i \). For the OLS regression without any controls, we estimate:

\[
Y_i = \beta_0 + \beta_1 D_i + \epsilon_i.
\]

For the 2SLS regression without any controls, we run:

\[
D_i = \gamma_0 + \gamma_1 Z_i + v_i
\]

\[
Y_i = \beta_0 + \beta_1 D_i + \epsilon_i.
\]

For the OLS regression controlling for disproportionate patient percentage, uncompensated care per bed and profit margin, we estimate:

\[
Y_i = \beta_0 + \beta_1 D_i + \beta_2 X_{i1} + \beta_3 X_{i2} + \beta_4 X_{i3} + \epsilon_i,
\]

where \( X_{i1} \) is disproportionate patient percentage, \( X_{i2} \) is uncompensated care per bed, and \( X_{i3} \) is profit margin.
to eliminate the bias, the safety net relief funding has little to no effect on the hospital utilization level by COVID-19 patients.\textsuperscript{34}

The above analysis looks at the immediate effects of relief funding. However, the effects of relief funding might kick in after a time lag, given that expansion in capacity and staff takes time. To investigate the relevance of this concern, we finally measure the evolving effects of relief funding. We estimate our main 2SLS specification on the 7-day average of each hospital outcome for each week from July 31st, 2020 to April 2nd, 2021. We plot the results in Figure 5. The estimated effects are similar to the initial null effects in Table 4, even several months after the distribution of relief funding. The overall insignificance of the estimates suggests that CARES funding had largely no effect on hospital utilization trends during the pandemic. This finding is consistent with policy and media arguments that CARES Act funding was not well targeted toward needy providers. Unlike the previous arguments, the analysis here provide causal evidence.

7 Other Examples

Here we give real-world examples of other algorithms and discuss the applicability of our framework.

**Example 1** (Bandit Algorithms). We are constantly exposed to digital information (movie, music, news, search results, advertisements, and recommendations) through a variety of devices and platforms. Tech companies allocate these pieces of content by using bandit algorithms. Our method is applicable to many popular bandit algorithms. For simplicity, assume a perfect-compliance scenario where the company perfectly controls the treatment assignment ($D_i = Z_i$). The algorithms below first use past data and supervised learning to estimate the conditional means and variances of potential outcomes, $E[Y_i(z)|X_i]$ and $\text{Var}(Y_i(z)|X_i)$, for each $z \in \{0, 1\}$. Let $\mu_z$ and $\sigma_z^2$ denote the estimated functions. The algorithms use $\mu_z(X_i)$ and $\sigma_z^2(X_i)$ to determine the treatment assignment for individual $i$.

(a) (Thompson Sampling Using Gaussian Priors) The algorithm first samples potential outcomes from the normal distribution with mean $(\mu_0(X_i), \mu_1(X_i))$ and variance-covariance matrix diag$(\sigma_0^2(X_i), \sigma_1^2(X_i))$. The algorithm then chooses the treatment with the highest sampled potential outcome:

$$Z_{i}^{TS} \equiv \arg \max_{z \in \{0, 1\}} y(z), \quad A^{TS}(X_i) = E[\arg \max_{z \in \{0, 1\}} y(z)|X_i],$$

where $y(z) \sim \mathcal{N}(\mu_z(X_i), \sigma_z^2(X_i))$ independently across $z$. These algorithms often induce quasi-experimental variation in treatment assignment, as a strand of the computer science literature has observed (Precup, 2000; Li et al., 2010; Narita, Yasui and Yata, 2019; Saito, 2019).

\textsuperscript{34}Estimates reported in Table A.1 in Appendix E.4 show little difference in outcome availability rates between eligible and ineligible hospitals once we control for APS. The 2SLS estimates in Table 4 are therefore unlikely to be compromised by differential attrition.
Aihara, Matsutani and Narita, 2021). The function $A$ has an analytical expression:

$$A^{TS}(x) = 1 - \Phi \left( \frac{\mu_0(x) - \mu_1(x)}{\sqrt{\sigma_0^2(x) + \sigma_1^2(x)}} \right),$$

where $\Phi$ is the cumulative distribution function of a standard normal distribution. Suppose that the functions $\mu_0$, $\mu_1$, $\sigma_0^2$ and $\sigma_1^2$ are continuous. APS for this case is given by

$$p^{TS}(x) = 1 - \Phi \left( \frac{\mu_0(x) - \mu_1(x)}{\sqrt{\sigma_0^2(x) + \sigma_1^2(x)}} \right).$$

This APS is nondegenerate, meaning that the data from the algorithm allow for causal-effect identification.

(b) (Upper Confidence Bound, UCB) Unlike the above stochastic algorithm, the UCB algorithm (Li et al., 2010) is a deterministic algorithm, producing a less obvious example of our framework. This algorithm chooses the treatment with the highest upper confidence bound for the potential outcome:

$$Z^{UCB}_i \equiv \arg \max_{z=0,1} \{\mu_z(X_i) + \alpha \sigma_z(X_i)\}, \quad A^{UCB}(x) = \arg \max_{z=0,1} \{\mu_z(x) + \alpha \sigma_z(x)\},$$

where $\alpha$ is chosen so that $|\mu_z(x) - E[Y_i(z)|X_i = x]| \leq \alpha \sigma_z(x)$ at least with some probability, for example, 0.95, for every $x$. Suppose that the function $g = \mu_1 - \mu_0 + \alpha (\sigma_1 - \sigma_0)$ is continuous on $\mathcal{X}$ and is continuously differentiable in a neighborhood of $x$ with $\nabla g(x) \neq 0$ for any $x \in \mathcal{X}$ such that $g(x) = 0$. APS for this case is given by

$$p^{UCB}(x) = \begin{cases} 
0 & \text{if } \mu_1(x) + \alpha \sigma_1(x) < \mu_0(x) + \alpha \sigma_0(x) \\
0.5 & \text{if } \mu_1(x) + \alpha \sigma_1(x) = \mu_0(x) + \alpha \sigma_0(x) \\
1 & \text{if } \mu_1(x) + \alpha \sigma_1(x) > \mu_0(x) + \alpha \sigma_0(x). 
\end{cases}$$

This means that the UCB algorithm produces potentially complicated quasi-experimental variation along the boundary in the covariate space where the algorithm’s treatment recommendation changes from one to the other. It is possible to identify and estimate causal effects across the boundary.

Example 2 (Unsupervised Learning). Customer segmentation is a core marketing practice that divides a company’s customers into groups based on their characteristics and behavior so that the company can effectively target marketing activities at each group. Many businesses today use unsupervised learning algorithms, clustering algorithms in particular, to perform customer segmentation. Using our notation, assume that a company decides whether it targets a campaign at customer $i$ ($Z_i = 1$) or not ($Z_i = 0$). The company first uses a clustering algorithm such as $K$-means clustering or Gaussian mixture model clustering to divide customers into $K$ groups, making a partition $\{S_1, \ldots, S_K\}$ of the covariate space $\mathbb{R}^p$. The company then conducts the campaign targeted at some of the groups:

$$Z^{CL}_i \equiv 1\{X_i \in \cup_{k \in T} S_k\}, \quad A^{CL}(x) = 1\{x \in \cup_{k \in T} S_k\},$$
where $T \subset \{1, \ldots, K\}$ is the set of the indices of the target groups.

For example, suppose that the company uses $K$-means clustering, which creates a partition in which a covariate value $x$ belongs to the group with the nearest centroid. Let $c_1, \ldots, c_K$ be the centroids of the $K$ groups. Define a set-valued function $C : \mathbb{R}^p \to 2^{\{1, \ldots, K\}}$, where $2^{\{1, \ldots, K\}}$ is the power set of $\{1, \ldots, K\}$, as $C(x) \equiv \arg \min_{k \in \{1, \ldots, K\}} \|x - c_k\|$. If $C(x)$ is a singleton, $x$ belongs to the unique group in $C(x)$. If $C(x)$ contains more than one indices, the group to which $x$ belongs is arbitrarily determined. APS for this case is given by

$$p^{CL}(x) = \begin{cases} 
0 & \text{if } C(x) \cap T = \emptyset \\
0.5 & \text{if } |C(x)| = 2, x \in \partial(\cup_{k \in T} S_k) \\
1 & \text{if } C(x) \subset T
\end{cases}$$

and $p^{CL}(x) \in (0, 1)$ if $|C(x)| \geq 3$ and $x \in \partial(\cup_{k \in T} S_k)$, where $|C(x)|$ is the number of elements in $C(x)$.\textsuperscript{35} Thus, it is possible to identify causal effects across the boundary $\partial(\cup_{k \in T} S_k)$.

**Example 3** (Supervised Learning). Millions of times each year, judges make jail-or-release decisions that hinge on a prediction of what a defendant would do if released. Many judges now use proprietary algorithms (like COMPAS criminal risk score) to make such predictions and use the predictions to support jail-or-release decisions. Using our notation, assume that a criminal risk algorithm recommends jailing ($Z_i = 1$) or releasing ($Z_i = 0$) for each defendant $i$. The algorithm uses defendant $i$’s observable characteristics $X_i$, including criminal history and demographics. The algorithm first translates $X_i$ into a risk score $r(X_i)$, where $r : \mathbb{R}^p \to \mathbb{R}$ is a function estimated by supervised learning based on past data and assumed to be fixed. For example, Kleinberg et al. (2017) construct a version of $r(X_i)$ using gradient boosted decision trees. The algorithm then uses the risk score to make the final recommendation:

$$Z_i^{SL} \equiv 1\{r(X_i) > c\}, \quad A^{SL}(x) = 1\{r(x) > c\},$$

where $c \in \mathbb{R}$ is a constant threshold that is set ex ante.\textsuperscript{36} A similar procedure applies to the screening of potential borrowers by banks and insurance companies based on credit scores estimated by supervised learning (Agarwal, Chomsisengphet, Mahoney and Stroebel, 2017).

A widely-used approach to identifying and estimating treatment effects in these settings is to use the score $r(X_i)$ as a continuous univariate running variable and apply a univariate RDD method (Cowgill, 2018). However, whether $r(X_i)$ is continuously distributed or not depends on how the function $r$ is constructed. For example, suppose that $r$ is constructed by a tree-based algorithm and is the following simple regression tree with three terminal nodes:

$$r(x) = \begin{cases} 
1 & \text{if } x_1 \leq 0 \\
2 & \text{if } x_1 > 0, x_2 \leq 0 \\
3 & \text{if } x_1 > 0, x_2 > 0,
\end{cases}$$

\textsuperscript{35}If $|C(x)| = 2$ and $x \in \partial(\cup_{k \in T} S_k)$, $x$ is on a linear boundary between one target group and one non-target group, and hence APS is 0.5. If $|C(x)| \geq 3$ and $x \in \partial(\cup_{k \in T} S_k)$, $x$ is a common endpoint of several group boundaries, and APS is determined by the angles at which the boundaries intersect.

\textsuperscript{36}The algorithm sometimes discretizes the original risk score $r(X_i)$ into $d(r(X_i))$, where $d : \mathbb{R} \to \mathbb{N}$ (Cowgill, 2018). In this case, the algorithm uses the discretized risk score to make the final recommendation: $Z_i^{SL} \equiv 1\{d(r(X_i)) > c\}$.\textsuperscript{35}
where \( r_1 < r_2 < c < r_3 \). In this case, the score \( r(X_i) \) is a discrete variable, and hence it may not be suitable to apply a standard univariate RDD method.

Our approach is applicable to this case as long as at least one of the original multi-dimensional covariates \( X_i \) are continuously distributed. Since \( A_{SL}(x) = 1\{r(x) > c\} = 1\{x_1 > 0, x_2 > 0\} \), APS for this case is given by

\[
p_{SL}(x) = \begin{cases} 
0 & \text{if } x_1 < 0 \text{ or } x_2 < 0 \\
0.25 & \text{if } x_1 = x_2 = 0 \\
0.5 & \text{if } (x_1 = 0, x_2 > 0) \text{ or } (x_1 > 0, x_2 = 0) \\
1 & \text{if } x_1 > 0, x_2 > 0.
\end{cases}
\]

It is therefore possible to identify causal effects across the boundary \( \{x \in X : (x_1 = 0, x_2 \geq 0) \text{ or } (x_1 > 0, x_2 = 0)\} \).

**Example 4** (Policy Eligibility Rules). Medicaid and other welfare policies often decide who are eligible based on algorithmic rules, as studied by Currie and Gruber (1996) and Brown, Kowalski and Lurie (2020).\(^{37}\) Using our notation, the state government determines whether each individual \( i \) is eligible \( (Z_i = 1) \) or not \( (Z_i = 0) \) for Medicare. The state government’s eligibility rule \( A_{Medicaid} \) maps individual characteristics \( X_i \) (e.g. income, family composition) into an eligibility decision \( Z_i^{Medicare} \). A similar procedure also applies to bankruptcy laws (Mahoney, 2015). These policy eligibility rules produce quasi-experimental variation as in Example 3.

**Example 5** (Mechanism Design: Matching and Auction). Centralized economic mechanisms such as matching and auction are also suitable examples, as summarized below:

<table>
<thead>
<tr>
<th>( i )</th>
<th>Matching (e.g., School Choice)</th>
<th>Auction</th>
</tr>
</thead>
<tbody>
<tr>
<td>( X_i )</td>
<td>Preference/Priority/Tie-breaker</td>
<td>Bid</td>
</tr>
<tr>
<td>( Z_i )</td>
<td>Whether student ( i ) is assigned treatment school</td>
<td>Whether bidder ( i ) wins the good</td>
</tr>
<tr>
<td>( D_i )</td>
<td>Whether student ( i ) attends treatment school</td>
<td>Same as ( Z_i )</td>
</tr>
<tr>
<td>( Y_i )</td>
<td>Student ( i )’s future test score</td>
<td>Bidder ( i )’s future economic performance</td>
</tr>
</tbody>
</table>

In mechanism design and other algorithms with capacity constraints, the treatment recommendation for individual \( i \) may depend not only on \( X_i \) but also on the characteristics of others. These interactive situations can be accommodated by our framework if we consider the following large market setting.\(^{38}\) Suppose that there is a continuum of individuals \( i \in [0, 1] \) and that the

\(^{37}\)These papers estimate the effect of Medicaid eligibility by exploiting variation in the eligibility rule across states and over time (simulated instrumental variable method). In contrast, our method exploits local variation in the eligibility status across different individuals given a fixed eligibility rule.

\(^{38}\)The approach proposed by Borusyak and Hull (2020) is applicable to finite-sample settings if the treatment recommendation probability, which may depend on all individuals’ characteristics, is nondegenerate for multiple individuals.
recommendation probability for individual $i$ with covariate $X_i$ is determined by a function $M$ as follows:

$$ Pr(Z_i = 1 | X_i; F_{X_{-i}}) = M(X_i; F_{X_{-i}}). $$

Here $F_{X_{-i}} = \Pr(\{j \in [0,1] \setminus \{i\} : X_j \leq x\})$ is the distribution of $X$ among all individuals $j \in [0,1] \setminus \{i\}$. The function $M : \mathbb{R}^p \times \mathcal{F} \rightarrow [0,1]$, where $\mathcal{F}$ is a set of distributions on $\mathbb{R}^p$, gives the recommendation probability for each individual in the market. With a continuum of individuals, for any $i \in [0,1]$, $F_{X_{-i}}$ is the same as the distribution of $X$ in the whole market, denoted by $F_X$. Therefore, the data generated by the mechanism $M$ are equivalent to the data generated by the algorithm $A : \mathbb{R}^p \rightarrow [0,1]$ such that $A(x) \equiv M(x; F_{X})$ for all $x \in \mathbb{R}^p$. Our framework is applicable to this large-market interactive setting.

The above discussions can be summarized as follows.

**Corollary 4.** In all the above examples, there exists $x \in \text{int}(\mathcal{X})$ such that $p^A(x) \in (0,1)$. Therefore, a causal effect is identified under Assumptions 1 and 2.

## 8 Conclusion

As algorithmic decisions become the new norm, the world becomes a mountain of natural experiments and instruments. We develop a general method to use these algorithm-produced instruments to identify and estimate causal treatment effects. Our analysis of the CARES Act hospital relief funding uses the proposed method to find that relief funding has little effect on COVID-19-related hospital activities. OLS or uncontrolled 2SLS estimates, by contrast, show considerably larger and more significant effects. The large estimates appear to be an artifact of selection bias; relief funding just went to hospitals with more COVID-19 patients, without helping hospitals accommodate additional patients.

Our analysis clarifies a few implications for policy and management practices around algorithmic decision-making. It is important to record the implementation of algorithms in a replicable, simulatable way, including what input variables $X_i$ are used to make algorithmic recommendation $Z_i$. Another key lesson is the importance of recording an algorithm’s recommendation $Z_i$ even if they are superseded by a human decision $D_i$. These data retention efforts would go a long way to exploit the full potential of algorithms as natural experiments.

Our agenda for future research includes a formalization of such optimal policy (algorithm) learning. Another important topic is estimation details, such as data-driven bandwidth selection. This work needs to extend Imbens and Kalyanaraman (2012) and Calonico et al. (2014)’s bandwidth selection methods in the univariate RDD to our setting. Inference on treatment effects in our framework relies on conventional large sample reasoning. It seems natural to additionally consider permutation or randomization inference as in Imbens and Rosenbaum (2005). It will also be challenging but interesting to develop finite-sample optimal estimation and inference strategies such as those recently introduced by Armstrong and Kolesár (2018, 2020) and Imbens and Wager (2019). Finally, we look forward to empirical applications of our method in a variety of business, policy, and scientific domains.
References


Figure 1: Example of the Approximate Propensity Score

Figure 2: Illustration of the Change of Variables Techniques

(a) 

(b)
Table 1: Bias, SD, and RMSE of Estimators and Coverage of 95% Confidence Intervals

<table>
<thead>
<tr>
<th></th>
<th>Our Method: 2SLS with Approximate Propensity Score Controls</th>
<th>2SLS with A Controls</th>
<th>OLS with No Controls</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$\delta = 0.01$</td>
<td>$\delta = 0.05$</td>
<td>$\delta = 0.1$</td>
</tr>
<tr>
<td>Estimand: ATE $= \text{ATE(RCT)} = 0$</td>
<td>Bias</td>
<td>.603</td>
<td>.634</td>
</tr>
<tr>
<td></td>
<td>SD</td>
<td>.304</td>
<td>.205</td>
</tr>
<tr>
<td></td>
<td>RMSE</td>
<td>.675</td>
<td>.667</td>
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<tr>
<td>Coverage</td>
<td>94.8%</td>
<td>92.8%</td>
<td>92.9%</td>
</tr>
<tr>
<td>Avg $N$</td>
<td>235</td>
<td>727</td>
<td>1275</td>
</tr>
</tbody>
</table>

Panel A: Homogeneous Conditional Effects

<table>
<thead>
<tr>
<th></th>
<th>Our Method: 2SLS with Approximate Propensity Score Controls</th>
<th>2SLS with A Controls</th>
<th>OLS with No Controls</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$\delta = 0.01$</td>
<td>$\delta = 0.05$</td>
<td>$\delta = 0.1$</td>
</tr>
<tr>
<td>Estimand: LATE $= \text{LATE(RCT)} = 0.564$</td>
<td>Bias</td>
<td>.039</td>
<td>.070</td>
</tr>
<tr>
<td></td>
<td>SD</td>
<td>.304</td>
<td>.205</td>
</tr>
<tr>
<td></td>
<td>RMSE</td>
<td>.306</td>
<td>.217</td>
</tr>
<tr>
<td>Coverage</td>
<td>94.8%</td>
<td>92.8%</td>
<td>92.9%</td>
</tr>
<tr>
<td>Avg $N$</td>
<td>235</td>
<td>723</td>
<td>1274</td>
</tr>
</tbody>
</table>

Panel B: Heterogeneous Conditional Effects

<table>
<thead>
<tr>
<th></th>
<th>Our Method: 2SLS with Approximate Propensity Score Controls</th>
<th>2SLS with A Controls</th>
<th>OLS with No Controls</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$\delta = 0.01$</td>
<td>$\delta = 0.05$</td>
<td>$\delta = 0.1$</td>
</tr>
<tr>
<td>Estimand: LATE $= \text{LATE(RCT)} = 0.564$</td>
<td>Bias</td>
<td>.004</td>
<td>.023</td>
</tr>
<tr>
<td></td>
<td>SD</td>
<td>.331</td>
<td>.222</td>
</tr>
<tr>
<td></td>
<td>RMSE</td>
<td>.331</td>
<td>.223</td>
</tr>
<tr>
<td>Coverage</td>
<td>95.9%</td>
<td>94.8%</td>
<td>95.0%</td>
</tr>
<tr>
<td>Avg $N$</td>
<td>235</td>
<td>723</td>
<td>1274</td>
</tr>
</tbody>
</table>

Notes: This table shows the bias, the standard deviation (SD) and the root mean squared error (RMSE) of 2SLS with Approximate Propensity Score controls, 2SLS with $A$ controls, and OLS with no controls. These statistics are computed with the estimand set to ATE, ATE(RCT), LATE, or LATE(RCT). The row “Coverage” in each panel shows the probabilities that the 95% confidence intervals of the form $[\hat{\beta}_1 - 1.96\hat{\sigma}_n, \hat{\beta}_1 + 1.96\hat{\sigma}_n]$ contains LATE(RCT), where $\hat{\beta}_1$ is the 2SLS estimate with Approximate Propensity Score controls and $\hat{\sigma}_n$ is its heteroskedasticity-robust standard error. We use 1,000 replications of a size 10,000 simulated sample to compute these statistics. We use several possible values of $\delta$ to compute the Approximate Propensity Score. All Approximate Propensity Scores are computed by averaging 400 simulation draws of the $A$ value. Panel A reports the results under the model in which the treatment effect does not depend on $X_i$. Panel B reports the results under the model in which the treatment effect depends on $X_i$. The bottom row “Avg $N$” in each panel shows the average number of observations used for estimation (i.e., the average number of observations for which the Approximate Propensity Score or the $A$ value is strictly between 0 and 1).
Table 2: Hospital Characteristics and Outcomes

<table>
<thead>
<tr>
<th></th>
<th>All</th>
<th>Ineligible</th>
<th>Eligible</th>
<th>Hospitals with APS ∈ (0,1)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>105.59</td>
<td>98.41</td>
<td>136.61</td>
<td>125.19</td>
</tr>
<tr>
<td># Confirmed/Suspected Covid Patients</td>
<td>80.1</td>
<td>73.86</td>
<td>107.83</td>
<td>86.78</td>
</tr>
<tr>
<td># Confirmed Covid Patients</td>
<td>31.37</td>
<td>28.92</td>
<td>42.1</td>
<td>36.84</td>
</tr>
<tr>
<td># Confirmed/Suspected Covid Patients in ICU</td>
<td>26.62</td>
<td>24.41</td>
<td>36.56</td>
<td>31.41</td>
</tr>
<tr>
<td>Observations</td>
<td>4,008</td>
<td>3,293</td>
<td>715</td>
<td>429</td>
</tr>
</tbody>
</table>

Panel A: Outcome Variable Means

<table>
<thead>
<tr>
<th></th>
<th>All</th>
<th>Ineligible</th>
<th>Eligible</th>
<th>Hospitals with APS ∈ (0,1)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Beds</td>
<td>143.66</td>
<td>134.6</td>
<td>188.35</td>
<td>206.47</td>
</tr>
<tr>
<td>Interns and residents (full-time equivalents) per bed</td>
<td>.06</td>
<td>.05</td>
<td>.11</td>
<td>.09</td>
</tr>
<tr>
<td>Adult and pediatric hospital beds</td>
<td>120.26</td>
<td>113.29</td>
<td>154.66</td>
<td>170.49</td>
</tr>
<tr>
<td>Ownership: Proprietary (for-profit)</td>
<td>.19</td>
<td>.2</td>
<td>.18</td>
<td>.15</td>
</tr>
<tr>
<td>Ownership: Governmental</td>
<td>.22</td>
<td>.22</td>
<td>.23</td>
<td>.16</td>
</tr>
<tr>
<td>Ownership: Voluntary (non-profit)</td>
<td>.58</td>
<td>.58</td>
<td>.59</td>
<td>.68</td>
</tr>
<tr>
<td>Inpatient length of stay</td>
<td>9.21</td>
<td>10.14</td>
<td>4.66</td>
<td>4.38</td>
</tr>
<tr>
<td>Employees on payroll (full-time equivalents)</td>
<td>973.9</td>
<td>897.31</td>
<td>1351.57</td>
<td>1525.06</td>
</tr>
<tr>
<td>Disproportionate patient percentage</td>
<td>.21</td>
<td>.18</td>
<td>.38</td>
<td>.36</td>
</tr>
<tr>
<td>Uncompensated care per bed</td>
<td>59,850</td>
<td>56,556.03</td>
<td>76,096.31</td>
<td>45,996.48</td>
</tr>
<tr>
<td>Profit margin</td>
<td>.02</td>
<td>.04</td>
<td>-.07</td>
<td>-.03</td>
</tr>
<tr>
<td>Observations</td>
<td>4,633</td>
<td>3,582</td>
<td>781</td>
<td>485</td>
</tr>
</tbody>
</table>

Panel B: Hospital Characteristics

Notes: This table reports averages of outcome variables and hospital characteristics by safety net eligibility. A safety net hospital is defined as any acute care hospital with disproportionate patient percentage of 20.2% or greater, annual uncompensated care of at least $25,000 per bed and profit margin of 3.0% or less. Panel A reports the outcome variable means. Outcome variable estimates are 7 day sums for the week spanning July 31st 2020 to August 6th 2020. Confirmed or Suspected COVID patients refer to the sum of patients in inpatient beds with lab-confirmed/suspected COVID-19. Confirmed COVID patients refer to the sum of patients in inpatient beds with lab-confirmed COVID-19, including those with both lab-confirmed COVID-19 and influenza. Inpatient bed totals also include observation beds. Similarly, Confirmed/ Suspected COVID patients in ICU refer to the sum of patients in ICU beds with lab-confirmed or suspected COVID-19. Confirmed COVID patients in ICU refers to the sum of patients in ICU beds with lab-confirmed COVID-19, including those with both lab-confirmed COVID-19 and influenza. Panel B reports the means for hospital characteristics for the financial year 2018. Column 1 shows the means for All Hospitals. Columns 2 and 3 show the means for hospitals that are ineligible and eligible to receive safety net funding respectively. Column 4 shows the means for the hospitals with non-degenerate Approximate Propensity Score with bandwidth δ = 0.05. Approximate Propensity Score is computed by averaging 10,000 simulation draws.
Figure 3: Funding Distribution for Eligible Hospitals

Notes: The figure shows the distribution of funding amount for eligible hospitals. Each eligible hospital is assigned an individual facility score, which is the product of Disproportionate Patient Percentage and number of beds in the hospital. The share of $10 billion received by an eligible hospital is determined by the ratio of the individual facility score of that hospital to the sum of facility scores across all eligible hospitals. The amount of funding that can be received by an eligible hospital is calculated as the product of this ratio and $10 billion, and is bounded below at $5 million and bounded above at $50 million.
Notes: The top figure visualizes the three dimensions that determine safety net eligibility. The bottom figures show the data points plotted along 2 out of 3 dimensions. The bottom left panel plots disproportionate patient percentage against profit margin, while the bottom right panel plots uncompensated care per bed against profit margin. We remove hospitals above the 99th percentile of disproportionate patient percentage and uncompensated care per bed, for visibility purposes.
<table>
<thead>
<tr>
<th>No Controls</th>
<th>Our Method: 2SLS with Approximate Propensity Score Controls</th>
<th>Mean (Ineligible)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(\delta = 0.01)</td>
<td>(\delta = 0.025)</td>
</tr>
<tr>
<td>Beds</td>
<td>(7.05)</td>
<td>(105.74)</td>
</tr>
<tr>
<td></td>
<td>53.75***</td>
<td>198.67</td>
</tr>
<tr>
<td></td>
<td>(7.05)</td>
<td>(105.74)</td>
</tr>
<tr>
<td>Costs per discharge</td>
<td>(17.93)</td>
<td>(2.18)</td>
</tr>
<tr>
<td>(in thousands)</td>
<td>-49.95**</td>
<td>3.83</td>
</tr>
<tr>
<td></td>
<td>(17.93)</td>
<td>(2.18)</td>
</tr>
<tr>
<td>Disproportionate</td>
<td>0.21***</td>
<td>-0.09</td>
</tr>
<tr>
<td></td>
<td>(0.01)</td>
<td>(0.09)</td>
</tr>
<tr>
<td>patient percentage</td>
<td>N=4633</td>
<td>N=89</td>
</tr>
<tr>
<td>Full time employees</td>
<td>454.26***</td>
<td>2,670.64</td>
</tr>
<tr>
<td>(in millions)</td>
<td>(69.23)</td>
<td>(1,652.55)</td>
</tr>
<tr>
<td></td>
<td>N=4626</td>
<td>N=89</td>
</tr>
<tr>
<td>Medicare net revenue</td>
<td>18.36***</td>
<td>35.34</td>
</tr>
<tr>
<td></td>
<td>N=4511</td>
<td>N=88</td>
</tr>
<tr>
<td>Occupancy</td>
<td>0.07***</td>
<td>0.19*</td>
</tr>
<tr>
<td></td>
<td>(0.01)</td>
<td>(0.09)</td>
</tr>
<tr>
<td>Operating margin</td>
<td>-0.11***</td>
<td>-0.03</td>
</tr>
<tr>
<td></td>
<td>(0.01)</td>
<td>(0.06)</td>
</tr>
<tr>
<td>Profit margin</td>
<td>-0.11***</td>
<td>-0.03</td>
</tr>
<tr>
<td>Uncompensated</td>
<td>19,540.28***</td>
<td>4,905.22</td>
</tr>
<tr>
<td>care per bed</td>
<td>(3,827.22)</td>
<td>(12,161.31)</td>
</tr>
<tr>
<td></td>
<td>N=4633</td>
<td>N=89</td>
</tr>
<tr>
<td>p-value for joint significance</td>
<td>0</td>
<td>.74</td>
</tr>
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</table>

Notes: This table shows the results of the covariate balance regressions at the hospital level. The dependent variables for these regressions are drawn from the Healthcare Cost Report Information System for the financial year 2018. Disproportionate patient percentage, profit margin and uncompensated care per bed are used to determine the hospital’s safety net funding eligibility. Other dependent variables shown indicate the financial health and utilization of the hospitals. In column 1, we regress the dependent variables on the safety net eligibility of the hospital with no controls. In columns 2–8, we regress the dependent variables on funding eligibility controlling for the Approximate Propensity Score with different values of bandwidth \(\delta\). All Approximate Propensity Scores are computed by averaging 10,000 simulation draws. Column 9 shows the mean of dependent variables for hospitals that are ineligible to receive safety net funding. Robust standard errors are reported in the parenthesis and number of observations are reported separately for each regression. The last row reports the p-value of the joint significance test.
Table 4: Estimated Effects of Funding on Hospital Utilization

<table>
<thead>
<tr>
<th></th>
<th>OLS with No Controls</th>
<th>OLS with Covariate Controls</th>
<th>2SLS with No Controls</th>
<th>2SLS with Covariate Controls</th>
<th>Our Method: 2SLS with Approximate Propensity Score Controls</th>
<th>(\delta = )</th>
<th>(\delta = )</th>
<th>(\delta = )</th>
<th>(\delta = )</th>
<th>(\delta = )</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
<td>(4)</td>
<td>(5)</td>
<td>(6)</td>
<td>(7)</td>
<td>(8)</td>
<td>(9)</td>
<td>(10)</td>
</tr>
<tr>
<td># Confirmed/Suspected COVID Patients</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
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<tr>
<td>First stage</td>
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<td></td>
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<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(in millions)</td>
<td>(0.49)</td>
<td>(5.83)</td>
<td>(3.54)</td>
<td>(2.27)</td>
<td>(1.87)</td>
<td>(1.61)</td>
<td>(1.03)</td>
<td>(0.74)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$1\text{mm of funding}$</td>
<td>5.58***</td>
<td>3.25***</td>
<td>2.77***</td>
<td>-1.03</td>
<td>-1.86</td>
<td>-4.08</td>
<td>-2.91</td>
<td>0.15</td>
<td>-0.31</td>
<td></td>
</tr>
<tr>
<td>Observations</td>
<td>3532</td>
<td>3532</td>
<td>3532</td>
<td>73</td>
<td>195</td>
<td>547</td>
<td>719</td>
<td>1389</td>
<td>1947</td>
<td></td>
</tr>
<tr>
<td># Confirmed COVID Patients</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
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<td></td>
<td></td>
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</tr>
<tr>
<td>First stage</td>
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<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(in millions)</td>
<td>(0.50)</td>
<td>(6.11)</td>
<td>(3.66)</td>
<td>(2.33)</td>
<td>(1.91)</td>
<td>(1.64)</td>
<td>(1.04)</td>
<td>(0.75)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$1\text{mm of funding}$</td>
<td>4.53***</td>
<td>2.50**</td>
<td>2.44***</td>
<td>0.05</td>
<td>-2.14</td>
<td>1.42</td>
<td>0.13</td>
<td>-0.03</td>
<td>-0.09</td>
<td>-0.63</td>
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<tr>
<td>Observations</td>
<td>3558</td>
<td>3558</td>
<td>3558</td>
<td>70</td>
<td>191</td>
<td>385</td>
<td>539</td>
<td>709</td>
<td>1366</td>
<td>1923</td>
</tr>
<tr>
<td># Confirmed COVID Patients in ICU</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>First stage</td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(in millions)</td>
<td>(0.51)</td>
<td>(5.59)</td>
<td>(3.49)</td>
<td>(2.32)</td>
<td>(1.93)</td>
<td>(1.67)</td>
<td>(1.06)</td>
<td>(0.76)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$1\text{mm of funding}$</td>
<td>1.67***</td>
<td>0.91**</td>
<td>0.95***</td>
<td>0.93</td>
<td>0.71</td>
<td>0.36</td>
<td>-0.05</td>
<td>0.16</td>
<td>-0.03</td>
<td>-0.32</td>
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<tr>
<td>Observations</td>
<td>3445</td>
<td>3445</td>
<td>3445</td>
<td>72</td>
<td>186</td>
<td>374</td>
<td>520</td>
<td>678</td>
<td>1314</td>
<td>1846</td>
</tr>
<tr>
<td># Confirmed COVID Patients in ICU</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
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<tr>
<td>First stage</td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(in millions)</td>
<td>(0.50)</td>
<td>(6.15)</td>
<td>(3.73)</td>
<td>(2.41)</td>
<td>(2.02)</td>
<td>(1.73)</td>
<td>(1.09)</td>
<td>(0.77)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$1\text{mm of funding}$</td>
<td>1.51***</td>
<td>0.82**</td>
<td>0.88***</td>
<td>0.50</td>
<td>-0.11</td>
<td>0.18</td>
<td>0.04</td>
<td>0.12</td>
<td>-0.13</td>
<td>-0.35</td>
</tr>
<tr>
<td>Observations</td>
<td>3503</td>
<td>3503</td>
<td>3503</td>
<td>67</td>
<td>181</td>
<td>370</td>
<td>514</td>
<td>671</td>
<td>1321</td>
<td>1868</td>
</tr>
</tbody>
</table>

Notes: In this table we regress relevant outcomes at the hospital level on safety net funding. Column 1 presents the results of OLS regression of the outcome variables on safety net funding without any controls. Column 2 presents the results of OLS regression of the outcome variables on safety net funding controlling for disproportionate patient percentage, uncompensated care per bed and profit margin. In columns 3–10, we instrument safety net funding with eligibility to receive this funding and present the results of 2SLS regressions. In columns 3–10, the first stage shows the effect of being deemed eligible on the amount of relief funding received by hospitals, in millions of dollars. Column 3 shows the results of a 2SLS regression with no controls. In columns 4–10, we run this regression controlling for the Approximate Propensity Score with different values of bandwidth \(\delta\) on the sample with nondegenerate Approximate Propensity Score. All Approximate Propensity Scores are computed by averaging 10,000 simulation draws. The outcome variables are the 7 day totals for the week spanning July 31st, 2020 to August 6th, 2020. Confirmed or Suspected COVID patients refer to the sum of patients in inpatient beds with lab-confirmed/suspected COVID-19. Confirmed COVID patients refer to the sum of patients in inpatient beds with lab-confirmed COVID-19, including those with both lab-confirmed COVID-19 and influenza. Inpatient bed totals also include observation beds. Similarly, Confirmed/ Suspected COVID patients in ICU refer to the sum of patients in ICU beds with lab-confirmed or suspected COVID-19. Confirmed COVID patients in ICU refer to the sum of patients in ICU beds with lab-confirmed COVID-19, including those with both lab-confirmed COVID-19 and influenza. Robust standard errors are reported in parentheses.
Figure 5: Dynamic Effects of Funding on Weekly Hospital Outcomes

(a) # Confirmed/Suspected COVID Patients
(b) # Confirmed COVID Patients
(c) # Confirmed/Suspected COVID Patients in ICU
(d) # Confirmed COVID Patients in ICU

Notes: The figure shows the results of estimating our main 2SLS specification about the effect of $1mm of relief funding on weekly hospital outcomes from 07/31/2020 to 04/02/2021. The outcomes record the 7-day sum of the number of hospitalized patients with the specified condition. We compute the Approximate Propensity Score with $S = 10,000$ and $\delta = 0.05$. The estimates from the uncontrolled OLS, uncontrolled 2SLS, and 2SLS with the Approximate Propensity Score controls are plotted on the y-axis. Standard error ribbons are given in grey.
A Extensions and Discussions

A.1 Related Literature: Details

In this section, we discuss the related methodological literature on the multidimensional RDD in detail. Imbens and Wager (2019) propose the finite-sample-minimax linear estimator of the form \[ \sum_{i=1}^{n} \gamma_i Y_i \] and uniform confidence intervals for treatment effects in the multidimensional RDD. One version of their approach constructs a linear estimator by choosing the weight \( (\gamma_i)_{i=1}^{n} \) greedily to make the inference as precise as possible. Although their estimator is favorable in terms of precision, it is not obvious what estimand the estimator estimates, without assuming a constant treatment effect. The other version of Imbens and Wager (2019)’s approach and some other existing approaches (Zajonc, 2012; Keele and Titiunik, 2015) consider nonparametric estimation of the conditional average treatment effect \( E[Y_i(1) - Y_i(0)|X_i = x] \) for a specified boundary point \( x \). The estimand has a clear interpretation, but “when curvature is nonnegligible, equation (6) can effectively make use of only data near the specified focal point \( c \), thus resulting in relatively long confidence intervals” (Imbens and Wager, 2019, p. 268), where equation (6) defines their estimator.

To obtain more precise estimates while keeping interpretability, several papers studying a two-dimensional RDD, including Zajonc (2012) and Keele and Titiunik (2015), propose to estimate an integral of conditional average treatment effects over the boundary. Their approach first nonparametrically estimates \( E[Y_i(1) - Y_i(0)|X_i = x] \) and the density of \( X_i \) for a large number of points \( x \) in the boundary and then computes the weighted average of the estimated conditional average treatment effects with the weight set to the estimated density.

The above approach is difficult to implement, however, when \( X_i \) is high dimensional or the decision algorithm is a complex, black box function of \( X_i \), for the following reasons. First, it is computationally demanding to estimate \( E[Y_i(1) - Y_i(0)|X_i = x] \) for numerous points in the boundary such that the weighted average well approximates the integral of \( E[Y_i(1) - Y_i(0)|X_i = x] \) over the boundary. Second, identifying boundary points from a general decision algorithm itself is hard unless it has a known analytical form. By contrast, we develop an estimator that uses observations near all the boundary points without tracing out the boundary or knowing its analytical form, thus alleviating the limitations of existing estimators.

A.2 Existence of the Approximate Propensity Score

Proposition 1 assumes that APS exists, but is it fair to assume so? In general, APS may fail to exist. Figure A.1 shows such an example. In this example, \( X_i \) is two dimensional, and

\[
A(x) = \begin{cases} 
1 & \text{if } 3(\frac{1}{2})^{k-1} < \|x\| \leq 4(\frac{1}{2})^{k-1} \text{ for some } k = 1, 2, \cdots \\
0 & \text{if } 2(\frac{1}{2})^{k-1} < \|x\| \leq 3(\frac{1}{2})^{k-1} \text{ for some } k = 1, 2, \cdots.
\end{cases}
\]
It is shown that
\[ p_A(0; \delta) = \begin{cases} \frac{7}{12} & \text{if } \delta = 4\left(\frac{1}{2}\right)^k - 1 \text{ for some } k = 1, 2, \ldots \\ \frac{7}{27} & \text{if } \delta = 3\left(\frac{1}{2}\right)^k - 1 \text{ for some } k = 1, 2, \ldots \end{cases} \]

Therefore, \( \lim_{\delta \to 0} p_A(0; \delta) \) does not exist.

Nevertheless, APS exists for almost every \( x \), as shown in the following proposition.

**Proposition A.1.** \( p_A(x) \) exists and is equal to \( A(x) \) for almost every \( x \in \mathcal{X} \) (with respect to the Lebesgue measure).

**Proof.** See Appendix C.6.

Does APS exist at a specific point \( x \)? What is the value of APS at \( x \) if it is not equal to \( A(x) \)? We show that APS exists and is of a particular form for most covariate points and typical algorithms. For each \( x \in \mathcal{X} \) and each \( q \in \text{Supp}(A(X_i)) \), define
\[ U_{x,q} \equiv \{ u \in B(0, 1) : \lim_{\delta \to 0} A(x + \delta u) = q \}. \]

\( U_{x,q} \) is the set of vectors in \( B(0, 1) \) such that the value of \( A \) approaches \( q \) as we approach \( x \) from the direction of the vector. With this notation, we obtain a sufficient condition for the existence of APS at a point \( x \).

**Proposition A.2.** Take any \( x \in \mathcal{X} \). If there exists a countable set \( Q \subset \text{Supp}(A(X_i)) \) such that \( \mathcal{L}^p(\bigcup_{q \in Q} U_{x,q}) = \mathcal{L}^p(B(0, 1)) \) and \( U_{x,q} \) is \( \mathcal{L}^p \)-measurable for all \( q \in Q \), then \( p_A(x) \) exists and is given by
\[ p_A(x) = \frac{\sum_{q \in Q} q \mathcal{L}^p(U_{x,q})}{\mathcal{L}^p(B(0, 1))}. \]

**Proof.** See Appendix C.4.
If almost every point in $B(0, 1)$ is contained by one of countably many $\mathcal{U}_{x,q}$’s, therefore, APS exists and is equal to the weighted average of the values of $q$ with the weight proportional to the hypervolume of $\mathcal{U}_{x,q}$. This result implies that APS exists in practically important cases.

Corollary A.1.

1. (Continuity points) If $A$ is continuous at $x \in \mathcal{X}$, then $p^A(x)$ exists and $p^A(x) = A(x)$.

2. (Interior points) Let $\mathcal{X}_q = \{x \in \mathcal{X} : A(x) = q\}$ for some $q \in [0, 1]$. Then, for any interior point $x \in \text{int}(\mathcal{X}_q)$, $p^A(x)$ exists and $p^A(x) = q$.

3. (Smooth boundary points) Suppose that $\{x \in \mathcal{X} : A(x) = q_1\} = \{x \in \mathcal{X} : f(x) \geq 0\}$ and $\{x \in \mathcal{X} : A(x) = q_2\} = \{x \in \mathcal{X} : f(x) < 0\}$ for some $q_1, q_2 \in [0, 1]$, where $f : \mathbb{R}^p \to \mathbb{R}$.

Let $x \in \mathcal{X}$ be a boundary point such that $f(x) = 0$, and suppose that $f$ is continuously differentiable in a neighborhood of $x$ with $\nabla f(x) \neq 0$. In this case, $p^A(x)$ exists and $p^A(x) = \frac{1}{2}(q_1 + q_2)$.

4. (Intersection points under CART and random forests) Let $p = 2$, and suppose that $\{x \in \mathcal{X} : A(x) = q_1\} = \{(x_1, x_2) \in \mathcal{X} : x_1 \leq 0 \text{ or } x_2 \leq 0\}$, $\{x \in \mathcal{X} : A(x) = q_2\} = \{(x_1, x_2) \in \mathcal{X} : x_1 > 0, x_2 > 0\}$, and $0 = (0, 0)' \in \mathcal{X}$. This is an example in which tree-based algorithms such as Classification And Regression Tree (CART) and random forests are used to create $A$. In this case, $p^A(0)$ exists and $p^A(0) = \frac{3}{4}q_1 + \frac{1}{4}q_2$.

Proof. See Appendix C.5.

A.3 Discrete Covariates

In this section, we provide the definition of APS and identification and consistency results when $X_i$ includes discrete covariates. Suppose that $X_i = (X_{di}, X_{ci})$, where $X_{di} \in \mathbb{R}^{p_d}$ is a vector of discrete covariates, and $X_{ci} \in \mathbb{R}^{p_c}$ is a vector of continuous covariates. Let $\mathcal{X}_d$ denote the support of $X_{di}$ and be assumed to be finite. We also assume that $X_{ci}$ is continuously distributed conditional on $X_{di}$, and let $\mathcal{X}_{ci}(x_d)$ denote the support of $X_{ci}$ conditional on $X_{di} = x_d$ for each $x_d \in \mathcal{X}_d$. Let $\mathcal{X}_{c,0}(x_d) = \{x_c \in \mathcal{X}_{ci}(x_d) : A(x_d, x_c) = 0\}$ and $\mathcal{X}_{c,1}(x_d) = \{x_c \in \mathcal{X}_{ci}(x_d) : A(x_d, x_c) = 1\}$.

Define APS as follows: for each $x = (x_d, x_c) \in \mathcal{X}$,

$$p^A(x; \delta) \equiv \frac{\int_{B(x_c, \delta)} A(x_d, x_c^*) dx_c^*}{\int_{B(x_c, \delta)} dx_c^*},$$

$$p^A(x) \equiv \lim_{\delta \to 0} p^A(x; \delta),$$

where $B(x_c, \delta) = \{x_c^* \in \mathbb{R}^{p_c} : \|x_c - x_c^*\| \leq \delta\}$ is the $\delta$-ball around $x_c \in \mathbb{R}^{p_c}$. In other words, we take the average of the $A(x_d, x_c^*)$ values when $x_c^*$ is uniformly distributed on $B(x_c, \delta)$ holding $x_d$ fixed, and let $\delta \to 0$. Below, we assume that Assumptions 1, 2, 3 and 4 hold conditional on $X_{di}$.

Assumption A.1 (Almost Everywhere Continuity of $A$).
(a) For every \( x_d \in X_d \), \( A(x_d, \cdot) \) is continuous almost everywhere with respect to the Lebesgue measure \( \mathcal{L}^{p_c} \).

(b) For every \( x_d \in X_d \), \( \mathcal{L}^{p_c}(\mathcal{X}_{c,k}(x_d)) = \mathcal{L}^{p_c}(\text{int}(\mathcal{X}_{c,k}(x_d))) \) for \( k = 0, 1 \).

### A.3.1 Identification

**Assumption A.2** (Local Mean Continuity). For every \( x_d \in X_d \) and \( z \in \{0, 1\} \), the conditional expectation functions \( E[Y_{zi}|X_i = (x_d, x_c)] \) and \( E[D_i(z)|X_i = (x_d, x_c)] \) are continuous in \( x_c \) at any point \( x_c \in \mathcal{X}_c(x_d) \) such that \( p^A(x_d, x_c) \in (0, 1) \) and \( A(x_d, x_c) \in \{0, 1\} \).

Let \( \text{int}_c(\mathcal{X}) = \{(x_d, x_c) \in \mathcal{X} : x_c \in \text{int}(\mathcal{X}_c(x_d))\} \). We say that a set \( S \subset \mathbb{R}^p \) is open relative to \( \mathcal{X} \) if there exists an open set \( U \subset \mathbb{R}^p \) such that \( S = U \cap \mathcal{X} \). For a set \( S \subset \mathbb{R}^p \), let \( \mathcal{X}_d^S = \{x_d \in X_d : (x_d, x_c) \in S \text{ for some } x_c \in \mathbb{R}^{p_c}\} \) and \( \mathcal{X}_c^S(x_d) = \{x_c \in \mathcal{X}_c : (x_d, x_c) \in S\} \) for each \( x_d \in \mathcal{X}_d^S \).

**Proposition A.3.** Under Assumptions A.1 and A.2:

(a) \( E[Y_{zi} - Y_{0i}|X_i = x] \) and \( E[D_i(1) - D_i(0)|X_i = x] \) are identified for every \( x \in \text{int}_c(\mathcal{X}) \) such that \( p^A(x) \in (0, 1) \).

(b) Let \( S \) be any subset of \( \mathcal{X} \) open relative to \( \mathcal{X} \) such that \( p^A(x) \) exists for all \( x \in S \). Then either \( E[Y_{zi} - Y_{0i}|X_i \in S] \) or \( E[D_i(1) - D_i(0)|X_i \in S] \), or both are identified only if \( p^A(x) \in (0, 1) \) for almost every \( x_c \in \mathcal{X}_c^S(x_d) \) for every \( x_d \in \mathcal{X}_d^S \).

**Proof.** See Appendix C.7.

### A.3.2 Estimation

For each \( x_d \in X_d \), let \( \Omega^*\hat{(x_d)} = \{x_c \in \mathbb{R}^{p_c} : A(x_d, x_c) = 1\} \). Also, let \( \mathcal{X}_d^* = \{x_d \in X_d : \text{Var}(A(X_i)|X_{di} = x_d) > 0\} \), and let \( f_{X_c|X_d} \) denote the probability density function of \( X_{ci} \) conditional on \( X_{di} \). In addition, for each \( x_d \in \mathcal{X}_d \), let

\[
C^*(x_d) = \{x_c \in \mathbb{R}^{p_c} : A(x_d, \cdot) \text{ is continuously differentiable at } x_c\},
\]

and let \( D^*(x_d) = \mathbb{R}^{p_c} \setminus C^*(x_d) \).

**Assumption A.3.**

(a) (Finite Moments) \( E[Y_i^4] < \infty \).

(b) (Nonzero First Stage) There exists a constant \( c > 0 \) such that \( E[D_i(1) - D_i(0)|X_i = x] > c \) for every \( x \in \mathcal{X} \) such that \( p^A(x) \in (0, 1) \).

(c) (Nonzero Conditional Variance) If \( \Pr(A(X_i) \in (0, 1)) > 0 \), then \( \text{Var}(A(X_i)|A(X_i) \in (0, 1)) > 0 \).

If \( \Pr(A(X_i) \in (0, 1)) = 0 \), then the following conditions (d)–(g) hold.
(d) (Nonzero Variance) $\mathcal{X}_d^* \neq \emptyset$.

(e) ($C^2$ Boundary of $\Omega^*(x_d)$) For each $x_d \in \mathcal{X}_d^*$, there exists a partition $\{\Omega_1^*(x_d), \ldots, \Omega_M^*(x_d)\}$ of $\Omega^*(x_d)$ such that

(i) $\text{dist}(\Omega_m^*(x_d), \Omega_{m'}^*(x_d)) > 0$ for any $m, m' \in \{1, \ldots, M\}$ such that $m \neq m'$;

(ii) $\Omega_m^*(x_d)$ is nonempty, bounded, open, connected and twice continuously differentiable for each $m \in \{1, \ldots, M\}$.

(f) (Regularity of Deterministic $A$)

(i) For each $x_d \in \mathcal{X}_d^*$, $\mathcal{H}^{\rho - 1}(\partial \Omega^*(x_d)) < \infty$, and $\int_{\partial \Omega^*(x_d)} f_{X_c | x_d}(x_c | x_d) d\mathcal{H}^{\rho - 1}(x_c) > 0$.

(ii) There exists $\delta > 0$ such that $A(x_d, x_c) = 0$ for almost every $x_c \in N(\mathcal{X}_c(x_d), \delta) \setminus \Omega^*(x_d)$.

(g) (Conditional Means and Density near $\partial \Omega^*(x_d)$) For each $x_d \in \mathcal{X}_d^*$, there exists $\delta > 0$ such that

(i) $E[Y_i | X_i = (x_d, \cdot)]$, $E[Y_0 | X_i = (x_d, \cdot)]$, $E[D_i(1) | X_i = (x_d, \cdot)]$, $E[D_i(0) | X_i = (x_d, \cdot)]$ and $f_{X_c | x_d}(\cdot | x_d)$ are continuously differentiable and have bounded partial derivatives on $N(\partial \Omega^*(x_d), \delta)$;

(ii) $E[Y_i^2 | X_i = (x_d, \cdot)]$, $E[Y_0^2 | X_i = (x_d, \cdot)]$, $E[Y_i D_i(1) | X_i = (x_d, \cdot)]$ and $E[Y_0 D_i(0) | X_i = (x_d, \cdot)]$ are continuous on $N(\partial \Omega^*(x_d), \delta)$;

(iii) $E[Y_i^4 | X_i = (x_d, \cdot)]$ is bounded on $N(\partial \Omega^*(x_d), \delta)$.

Assumption A.4. If $\Pr(A(X_i) \in (0, 1)) > 0$, then the following conditions (a)–(c) hold.

(a) (Probability of Neighborhood of $D^*(x_d)$) For each $x_d \in \mathcal{X}_d^*$, $\Pr(X_i \in N(D^*(x_d), \delta)) = O(\delta)$.

(b) (Bounded Partial Derivatives of $A$) For each $x_d \in \mathcal{X}_d^*$, the partial derivatives of $A(x_d, \cdot)$ are bounded on $C^* (x_d)$.

(c) (Bounded Conditional Mean) For each $x_d \in \mathcal{X}_d^*$, $E[Y_i | X_i = (x_d, \cdot)]$ is bounded on $\mathcal{X}_c(x_d)$.

Theorem A.1. Suppose that Assumptions A.1 and A.3 hold, and that $\delta_n \to 0$, $n \delta_n \to \infty$ and $S_n \to \infty$ as $n \to \infty$. Then the 2SLS estimators $\hat{\beta}_1$ and $\hat{\beta}_1^*$ converge in probability to

$$\beta_1 \equiv \lim_{\delta \to 0} E[\omega_1(\delta)(Y_i(1) - Y_i(0))]$$

where

$$\omega_1(\delta) = \frac{p^A(X_i; \delta)(1 - p^A(X_i; \delta))(D_i(1) - D_i(0))}{E[p^A(X_i; \delta)(1 - p^A(X_i; \delta))(D_i(1) - D_i(0))]}.$$ 

Suppose, in addition, that Assumptions A.4 and 5 hold and that $n \delta_n^2 \to 0$ as $n \to \infty$. Then

$$\hat{\sigma}_n^{-1}(\hat{\beta}_1 - \beta_1) \xrightarrow{d} \mathcal{N}(0, 1),$$

$$(\hat{\sigma}_n^*)^{-1}(\hat{\beta}_1^* - \beta_1) \xrightarrow{d} \mathcal{N}(0, 1).$$
Assumption A.5. We provide a sufficient condition for Assumption 4 (a).

A.4 A Sufficient Condition for Assumption 4 (a)

We provide a sufficient condition for Assumption 4 (a).

Assumption A.5.

(a) (Twice Continuous Differentiability of $D^*$) There exist $C_1^*, ..., C_M^* \subset \mathbb{R}^p$ such that
   
   (i) $\partial(\bar{C}^*) = D^*$, where $\bar{C}^* = \bigcup_{m=1}^{M} C_m^*$;
   
   (ii) dist$(C_m^*, C_{m'}^*) > 0$ for any $m, m' \in \{1, ..., M\}$ such that $m \neq m'$;
   
   (iii) $C_m^*$ is nonempty, bounded, open, connected and twice continuously differentiable for each $m \in \{1, ..., M\}$.

(b) (Regularity of $D^*$) $\mathcal{H}^{p-1}(D^*) < \infty$.

(c) (Bounded Density near $D^*$) There exists $\delta > 0$ such that $f_X$ is bounded on $N(D^*, \delta)$.

The key condition is the twice continuous differentiability of $D^*$. This condition holds if, for example, the $\epsilon$-Greedy algorithm described in Example A.1 (a) in Appendix A.6 uses an estimated $Q$-function that is twice continuously differentiable in $x$.

Under Assumption A.5 (a), by Lemma B.4 in Appendix B.3 and with change of variables $v = \frac{\lambda}{\delta}$, for any sufficiently small $\delta > 0$,

$$\Pr(X_i \in N(D^*, \delta)) = \int_{-\delta}^{\delta} \int_{D^*} f_X(u + \lambda \nu_{\bar{C}^*}(u))J_{p-1}D^*\psi_{\bar{C}^*}(u, \lambda)d\mathcal{H}^{p-1}(u)d\lambda$$

$$= \delta \int_{-1}^{1} \int_{D^*} f_X(u + \delta \nu_{\bar{C}^*}(u))J_{p-1}D^*\psi_{\bar{C}^*}(u, \delta v)d\mathcal{H}^{p-1}(u)dv.$$ (See Appendix B for the notation.) If $f_X$ is bounded on $N(D^*, \delta)$ and $\mathcal{H}^{p-1}(D^*) < \infty$, the right-hand side is $O(\delta)$. 

Proof. See Appendix C.8.
A.5 Sampling from Uniform Distribution on \( p \)-Dimensional Ball

When we calculate APS by simulation, we need to uniformly sample from \( B(X_i, \delta) \). We introduce three existing methods to uniformly sample from a \( p \)-dimensional unit ball \( B(0,1) \). By multiplying the sampled vector by \( \delta \) and adding \( X_i \) to it, we can sample from a uniform distribution on \( B(X_i, \delta) \).

**Method 1.**
1. Sample \( x_1, \ldots, x_p \) independently from the uniform distribution on \([-1,1]\).
2. Accept the vector \( x = (x_1, \ldots, x_p) \) if \( \sum_{k=1}^{p} x_k^2 \leq 1 \) and reject it otherwise.

Method 1 is a practical choice when \( p \) is small (e.g. \( p = 2, 3 \)), but is inefficient for higher dimensions, since the acceptance rate decreases to zero quickly as \( p \) increases. The conventional method used for higher dimensions is the following.

**Method 2.**
1. Sample \( x^*_1, \ldots, x^*_p \) independently from the standard normal distribution, and compute the vector \( s = (x^*_1, \ldots, x^*_p)/\sqrt{\sum_{k=1}^{p} (x^*_k)^2} \).
2. Sample \( u \) from the uniform distribution on \([0,1]\).
3. Return the vector \( x = u^{1/p} s \).

There is yet another method efficient for higher dimensions, which is recently proposed by Voelker, Gosmann and Stewart (2017).

**Method 3.**
1. Sample \( x^*_1, \ldots, x^*_{p+2} \) independently from the standard normal distribution, and compute the vector \( s = (x^*_1, \ldots, x^*_{p+2})/\sqrt{\sum_{k=1}^{p+2} (x^*_k)^2} \).
2. Return the vector \( x = (s_1, \ldots, s_p) \).

A.6 Additional Examples

**Example A.1 (Reinforcement Learning Algorithms).** Extending bandit algorithms to dynamically changing environments, reinforcement learning algorithms optimize decisions in dynamic environments, where the state (the set of observables that the agent receives from the environment) and action in the current period can affect the future states and outcomes. Let \( \{(X_{ti}, Z_{ti}, Y_{ti})\}_{t=0}^{\infty} \) denote the trajectory of the states, treatment assignments, and outcomes in periods \( t = 0, 1, 2, \cdots \) for individual \( i \). For simplicity, we assume that the trajectory follows a Markov decision process.\(^{39}\) Let \( Y_{ti}(1) \) and \( Y_{ti}(0) \) represent the potential outcomes in period \( t \). Let \( Q : \mathcal{X} \times [0,1] \rightarrow \mathbb{R} \) be the optimal state-action value function, called the \( Q \)-function: for \((x, z) \in \mathcal{X} \times [0,1] \),

\[
Q(x, z) = \max_{\pi : \mathcal{X} \rightarrow [0,1]} \mathbb{E} \left[ \sum_{t=0}^{\infty} \gamma^t (Y_{ti}(1) \pi(X_{ti}) + Y_{ti}(0)(1 - \pi(X_{ti}))) | X_{0i} = x, Z_{0i} = z \right],
\]

\(^{39}\)Under a Markov decision process, the distribution of the state \( X_{ti} \) only depends on the last state and treatment assignment \( (X_{t-1,i}, Z_{t-1,i}) \), the distribution of the outcome \( Y_{ti} \) only depends on the current state and treatment assignment \( (X_{ti}, Z_{ti}) \), and these distributions are stationary over periods.
where $\gamma \in [0, 1)$ is a discount factor, and $\pi$ is a policy function that assigns the probability of treatment to each possible state.

(a) (\(\epsilon\)-Greedy) This algorithm first uses past data to yield $\hat{Q}$, an estimate of the $Q$-function. For example, the fitted $Q$ iteration (Ernst, Geurts and Wehenkel, 2005) is used to estimate $Q$.40

B Notation and Lemmas

B.1 Basic Notations

For a scalar-valued differentiable function $f : S \subset \mathbb{R}^n \to \mathbb{R}$, let $\nabla f : S \to \mathbb{R}^n$ be a gradient of $f$:

$$\nabla f(x) = \left( \frac{\partial f(x)}{\partial x_1}, \ldots, \frac{\partial f(x)}{\partial x_n} \right).$$

Also, when the second-order partial derivatives of $f$ exist, let $D^2 f(x)$ be the Hessian matrix:

$$D^2 f(x) = \begin{bmatrix} \frac{\partial^2 f(x)}{\partial x_1^2} & \cdots & \frac{\partial^2 f(x)}{\partial x_1 \partial x_n} \\ \vdots & \ddots & \vdots \\ \frac{\partial^2 f(x)}{\partial x_n \partial x_1} & \cdots & \frac{\partial^2 f(x)}{\partial x_n^2} \end{bmatrix}$$

for each $x \in S$.

Let $f : S \subset \mathbb{R}^m \to \mathbb{R}^n$ be a function such that its first-order partial derivatives exist. For each $x \in S$, let $Jf(x)$ be the Jacobian matrix of $f$ at $x$:

$$Jf(x) = \begin{bmatrix} \frac{\partial f_1(x)}{\partial x_1} & \cdots & \frac{\partial f_1(x)}{\partial x_m} \\ \vdots & \ddots & \vdots \\ \frac{\partial f_n(x)}{\partial x_1} & \cdots & \frac{\partial f_n(x)}{\partial x_m} \end{bmatrix}.$$

For a positive integer $n$, let $I_n$ denote the $n \times n$ identity matrix.

B.2 Differential Geometry

We provide some concepts and facts from differential geometry of twice continuously differentiable sets, following Crasta and Malusa (2007). Let $S \subset \mathbb{R}^p$ be a twice continuously differentiable set. For each $x \in \partial S$, we denote by $\nu_S(x) \in \mathbb{R}^p$ the inward unit normal vector of $\partial S$ at $x$, that is,

40Suppose that we have collected a set of $L$ four-tuples $\{(x_l^t, z_l^t, y_l^{t+1}, x_{l+1}^{t+1})\}_{l=1}^L$ as a result of the agent interacting with the dynamic environment. Given the dataset and an initial approximation $\hat{Q}$ of $Q$ (e.g., $\hat{Q}(x, z) = 0$ for all $(x, z)$), we repeat the following steps until some stopping condition is reached: 1. For each $l = 1, \ldots, L$, calculate $q_l = y_l + \gamma \max_{z \in \{0, 1\}} \hat{Q}(x_{l+1}^{t+1}, z); 2$. Use $\{(x_l^t, z_l^{t+1}, q_l)\}_{l=1}^L$ and a supervised learning method to train a model that predicts $q$ from $(x, z)$. Let the model be a new approximation $\hat{Q}$ of $Q$. Possible supervised learning methods used in the second step include tree-based methods, neural networks (Neural Fitted $Q$ Iteration) and deep neural networks (Deep Fitted $Q$ Iteration).
We apply results from Crasta and Malusa (2007). Let \( t \nu \) be the unit vector orthogonal to all vectors in the tangent space of \( \partial S \) at \( x \) that points toward the inside of \( A \). For a set \( S \subset \mathbb{R}^p \), let \( d_S : \mathbb{R}^p \to \mathbb{R} \) be the signed distance function of \( S \), defined by

\[
d_S(x) = \begin{cases} 
    d(x, \partial S) & \text{if } x \in \text{cl}(S) \\
    -d(x, \partial S) & \text{if } x \in \mathbb{R}^p \setminus \text{cl}(S),
\end{cases}
\]

where \( d(x, B) = \inf_{y \in B} \|y - x\| \) for any \( x \in \mathbb{R}^p \) for a set \( B \subset \mathbb{R}^p \). Note that we can write \( N(\partial S, \delta) = \{ x \in \mathbb{R}^p : -\delta < d_S(x) < \delta \} \) for \( \delta > 0 \). Lastly, let \( \Pi_{\partial S}(x) = \{ y \in \partial S : \|y - x\| = d(x, \partial S) \} \) be the set of projections of \( x \) on \( \partial S \).

**Lemma B.1** (Corollary of Theorem 4.16, Crasta and Malusa (2007)). Let \( S \subset \mathbb{R}^p \) be nonempty, bounded, open, connected and twice continuously differentiable. Then the function \( d_S \) is twice continuously differentiable on \( N(\partial S, \mu) \) for some \( \mu > 0 \). In addition, for every \( x_0 \in \partial S \), \( \Pi_{\partial S}(x_0 + t\nu_S(x_0)) = \{ x_0 \} \) for every \( t \in (-\mu, \mu) \). Furthermore, for every \( x \in N(\partial S, \mu) \), \( \Pi_{\partial S}(x) \) is a singleton, \( \nabla d_S(x) = \nu_S(y) \) and \( x = y + d_S(x)\nu_S(y) \) for \( y \in \Pi_{\partial S}(x) \), and \( \|\nabla d_S(x)\| = 1 \).

**Proof.** We apply results from Crasta and Malusa (2007). Let \( K = \{ x \in \mathbb{R}^p : \|x\| \leq 1 \} \). \( K \) is nonempty, compact, convex subset of \( \mathbb{R}^p \) with the origin as an interior point. The polar body of \( K \), defined as \( K_0 = \{ y \in \mathbb{R}^p : y \cdot x \leq 1 \text{ for all } x \in K \} \), is \( K \) itself. The gauge functions \( \rho_K, \rho_{K_0} : \mathbb{R}^p \to [0, \infty] \) of \( K \) and \( K_0 \) are given by

\[
\rho_K(x) = \inf \{ t \geq 0 : x \in tK \} = \|x\|,
\rho_{K_0}(x) = \inf \{ t \geq 0 : x \in tK_0 \} = \|x\|.
\]

Given \( \rho_{K_0} \), the Minkowski distance from a set \( S \subset \mathbb{R}^p \) is defined as

\[
\delta_S(x) = \inf_{y \in S} \rho_{K_0}(x - y), \quad x \in \mathbb{R}^p.
\]

Note that we can write

\[
d_S(x) = \begin{cases} 
    \delta_{\partial S}(x) & \text{if } x \in \text{cl}(S) \\
    -\delta_{\partial S}(x) & \text{if } x \in \mathbb{R}^p \setminus \text{cl}(S),
\end{cases}
\]

It then follows from Theorem 4.16 of Crasta and Malusa (2007) that \( d_S \) is twice continuously differentiable on \( N(\partial S, \mu) \) for some \( \mu > 0 \), and for every \( x_0 \in \partial S \),

\[
\nabla d_S(x_0) = \frac{\nu_S(x_0)}{\rho_K(\nu_S(x_0))} \nu_S(x_0) = \frac{\nu_S(x_0)}{\|\nu_S(x_0)\|} \nu_S(x_0),
\]

where the last equality follows since \( \nu_S(x_0) \) is a unit vector. It then follows that \( \|\nabla d_S(x_0)\| = \|\nu_S(x_0)\| = 1 \) for every \( x_0 \in \partial S \). Also, it is obvious that, for every \( x_0 \in \partial S \), \( \Pi_{\partial S}(x_0) = \{ x_0 \} \) and \( x_0 = x_0 + d_S(x_0)\nu_S(x_0) \), since \( d_S(x_0) = 0 \). In addition, as stated in the proof of Theorem 4.16 of Crasta and Malusa (2007), \( \mu \) is chosen so that (4.7) in Proposition 4.6 of Crasta and Malusa (2007) holds for every \( x_0 \in \partial S \) and every \( t \in (-\mu, \mu) \). That is, \( \Pi_{\partial S}(x_0 + t\nu_K(\nu_S(x_0))) = \{ x_0 \} \) for every \( x_0 \in \partial S \) and every \( t \in (-\mu, \mu) \). Since \( \nabla \rho_K(\nu_S(x_0)) = \frac{\nu_S(x_0)}{\|\nu_S(x_0)\|} = \nu_S(x_0) \), \( \Pi_{\partial S}(x_0 + t\nu_S(x_0)) = \{ x_0 \} \) for every \( x_0 \in \partial S \) and every \( t \in (-\mu, \mu) \).
Furthermore, for every \( x \in N(\partial S, \mu) \setminus \partial S \), \( \Pi_{\partial S}(x) \) is a singleton as shown in the proof of Theorem 4.16 of Crasta and Malusa (2007). Let \( \pi_{\partial S}(x) \) be the unique element in \( \Pi_{\partial S}(x) \). By Lemma 4.3 of Crasta and Malusa (2007), for every \( x \in N(\partial S, \mu) \setminus \partial S \),

\[
\nabla d^s_S(x) = \frac{\nu_S(\pi_{\partial S}(x))}{\rho_K(\nu_S(\pi_{\partial S}(x)))} = \frac{\nu_S(\pi_{\partial S}(x))}{\|\nu_S(\pi_{\partial S}(x))\|} = \nu_S(\pi_{\partial S}(x)),
\]

where the last equality follows since \( \nu_S(\pi_{\partial S}(x)) \) is a unit vector. It then follows that \( \|\nabla d^s_S(x)\| = \|\nu_S(\pi_{\partial S}(x))\| = 1 \) for every \( x \in N(\partial S, \mu) \setminus \partial S \).

Lastly, note that

\[
\delta_{\partial S}(x) = \begin{cases} 
d^s_S(x) & \text{if } x \in N(\partial S, \mu) \cap \text{int}(S) \\
d^s_S(x) & \text{if } x \in N(\partial S, \mu) \setminus \text{cl}(S), 
\end{cases}
\]

and

\[
\nabla \delta_{\partial S}(x) = \begin{cases} 
\nabla d^s_S(x) & \text{if } x \in N(\partial S, \mu) \cap \text{int}(S) \\
-\nabla d^s_S(x) & \text{if } x \in N(\partial S, \mu) \setminus \text{cl}(S), 
\end{cases}
\]

so \( \delta_{\partial S}(x) \nabla \delta_{\partial S}(x) = d^s_S(x) \nabla d^s_S(x) = d^s_S(x) \nu_S(\pi_{\partial S}(x)) \) for every \( x \in N(\partial S, \mu) \setminus \partial S \). By Proposition 3.3 (i) of Crasta and Malusa (2007), for every \( x \in N(\partial S, \mu) \setminus \partial S \),

\[
\nabla \rho_K(\nabla \delta_{\partial S}(x)) = \frac{x - \pi_{\partial S}(x)}{\delta_{\partial S}(x)},
\]

which implies that

\[
x = \pi_{\partial S}(x) + \delta_{\partial S}(x) \nabla \rho_K(\nabla \delta_{\partial S}(x)) \\
= \pi_{\partial S}(x) + \delta_{\partial S}(x) \frac{\nabla \delta_{\partial S}(x)}{\|\nabla \delta_{\partial S}(x)\|} = \pi_{\partial S}(x) + d^s_S(x) \nu_S(\pi_{\partial S}(x)).
\]

\[\square\]

We say that a set \( S \subset \mathbb{R}^n \) is a \( m \)-dimensional \( C^1 \) submanifold of \( \mathbb{R}^n \) if for every point \( x \in S \), there exist an open neighborhood \( V \subset \mathbb{R}^n \) of \( x \) and a one-to-one continuously differentiable function \( \phi \) from an open set \( U \subset \mathbb{R}^m \) to \( \mathbb{R}^n \) such that the Jacobian matrix \( J\phi(u) \) is of rank \( m \) for all \( u \in U \), and \( \phi(U) = V \cap S \).

**Lemma B.2.** Let \( S \subset \mathbb{R}^p \) be nonempty, bounded, open, connected and twice continuously differentiable. Then \( \partial S \) is a \( (p-1) \)-dimensional \( C^1 \) submanifold of \( \mathbb{R}^p \).

**Proof.** Fix any \( x^* \in \partial S \). By Lemma B.1, \( \nabla d^s_S(x^*) \) is nonzero. Without loss of generality, let \( \frac{\partial d^s_S(x^*)}{\partial x^p} \neq 0 \). Let \( \psi : \mathbb{R}^p \to \mathbb{R}^p \) be the function such that \( \psi(x) = (x_1, \ldots, x_{p-1}, d^s_S(x)) \). \( \psi \) is continuously differentiable, and the Jacobian matrix of \( \psi \) at \( x^* \) is given by

\[
J\psi(x^*) = \begin{pmatrix}
\frac{\partial \psi_1}{\partial x_1}(x^*) & \ldots & \frac{\partial \psi_1}{\partial x_p}(x^*) \\
\vdots & \ddots & \vdots \\
\frac{\partial \psi_{p-1}}{\partial x_1}(x^*) & \ldots & \frac{\partial \psi_{p-1}}{\partial x_p}(x^*) \\
\frac{\partial d^s_S}{\partial x_1}(x^*) & \ldots & \frac{\partial d^s_S}{\partial x_p}(x^*) \\
\frac{\partial d^s_S}{\partial x_{p-1}}(x^*) & \ldots & \frac{\partial d^s_S}{\partial x_p}(x^*)
\end{pmatrix} = \begin{pmatrix} I_{p-1} & 0 \\
0 & \frac{\partial d^s_S(x^*)}{\partial x_1} & \ldots & \frac{\partial d^s_S(x^*)}{\partial x_p}
\end{pmatrix}.
\]

50
Since \( \frac{\partial d^p_s(x^*)}{\partial x_p} \neq 0 \), the Jacobian matrix is invertible. By the Inverse Function Theorem, there exist an open set \( V \) containing \( x^* \) and an open set \( W \) containing \( \psi(x^*) \) such that \( \psi: V \to W \) has an inverse function \( \psi^{-1}: W \to V \) that is continuously differentiable. We make \( V \) small enough so that \( \frac{\partial d^p_s(x)}{\partial x_p} \neq 0 \) for every \( x \in V \). The Jacobian matrix of \( \psi^{-1} \) is given by \( J\psi^{-1}(y) = J\psi(\psi^{-1}(y))^{-1} \) for all \( y \in W \).

Now note that \( \psi(x) = (x_1, \ldots, x_{p-1}, 0) \) for all \( x \in V \cap \partial S \) by the definition of \( d^p_s \). Let \( U = \{(x_1, \ldots, x_{p-1}) \in \mathbb{R}^{p-1} : x \in V \cap \partial S \} \) and \( \phi: U \to \mathbb{R}^p \) be a function such that \( \phi(u) = \psi^{-1}((u, 0)) \) for all \( u \in U \). Below we verify that \( \phi \) is one-to-one and continuously differentiable, that \( J\phi(u) \) is of rank \( p-1 \) for all \( u \in U \), that \( \phi(U) = V \cap \partial S \), and that \( U \) is open.

First, \( \phi \) is one-to-one, since \( \psi^{-1} \) is one-to-one, and \( (u, 0) \neq (u', 0) \) if \( u \neq u' \). Second, \( \phi \) is continuously differentiable, since \( \psi^{-1} \) is so. The Jacobian matrix of \( \phi \) at \( u \in U \) is by definition

\[
J\phi(u) = \begin{pmatrix}
\frac{\partial \psi_{-1}^{-1}((u, 0))}{\partial y_1} & \cdots & \frac{\partial \psi_{-1}^{-1}((u, 0))}{\partial y_{p-1}} \\
\vdots & \ddots & \vdots \\
\frac{\partial \psi_{-1}^{-1}((u, 0))}{\partial y_1} & \cdots & \frac{\partial \psi_{-1}^{-1}((u, 0))}{\partial y_{p-1}}
\end{pmatrix}.
\]

Note that this is the left \( p \times (p-1) \) submatrix of \( J\psi^{-1}((u, 0)) \). Since \( J\psi^{-1}((u, 0)) \) has full rank, \( J\phi(u) \) is of rank \( p-1 \). Moreover,

\[
\phi(U) = \{ \psi^{-1}((u, 0)) : u \in U \} = \{ \psi^{-1}((x_1, \ldots, x_{p-1}, 0)) : x \in V \cap \partial S \} = \{ \psi^{-1}(\psi(x)) : x \in V \cap \partial S \} = V \cap \partial S.
\]

Lastly, we show that \( U \) is open. Pick any \( \bar{u} \in U \). Then, there exists \( \bar{x}_p \in \mathbb{R} \) such that \( (\bar{u}, \bar{x}_p) \in V \cap \partial S \). As \( (\bar{u}, \bar{x}_p) \in V \cap \partial S \), \( d^p_s((\bar{u}, \bar{x}_p)) = 0 \). Since \( \frac{\partial d^p_s((\bar{u}, \bar{x}_p))}{\partial x_p} \neq 0 \), it follows by the Implicit Function Theorem that there exist an open set \( S \subset \mathbb{R}^{p-1} \) containing \( \bar{u} \) and a continuously differentiable function \( g: S \to \mathbb{R} \) such that \( g(\bar{u}) = \bar{x}_p \) and \( d^p_s(u, g(u)) = 0 \) for all \( u \in S \). Since \( g \) is continuous, \( (\bar{u}, g(\bar{u})) \in V \) and \( V \) is open, there exists an open set \( S' \subset S \) containing \( \bar{u} \) such that \( (u, g(u)) \in V \) for all \( u \in S' \). By the definition of \( d^p_s \), \( d^p_s(x) = 0 \) if and only if \( x \in \partial S \). Therefore, if \( u \in S' \), \( (u, g(u)) \) must be contained by \( \partial S \), for otherwise \( d^p_s(u, g(u)) \neq 0 \), which is a contradiction. Thus, \( (u, g(u)) \in V \cap \partial S \) and hence \( u \in U \) for all \( u \in S' \). This implies that \( S' \) is an open subset of \( U \) containing \( \bar{u} \), which proves that \( U \) is open.

\[ \Box \]

### B.3 Geometric Measure Theory

We provide some concepts and facts from geometric measure theory, following Krantz and Parks (2008). Recall that for a function \( f: S \subset \mathbb{R}^m \to \mathbb{R}^n \) and a point \( x \in S \) at which \( f \) is differentiable, \( Jf(x) \) denotes the Jacobian matrix of \( f \) at \( x \).

**Lemma B.3** (Coarea Formula, Lemma 5.1.4 and Corollary 5.2.6 of Krantz and Parks (2008)). If \( f: \mathbb{R}^m \to \mathbb{R}^n \) is a Lipschitz function and \( m \geq n \), then

\[
\int_S g(x)J_n f(x)d\mathcal{L}^m(x) = \int_{\mathbb{R}^n} \int_{\{x' \in S : f(x') = y\}} g(x)d\mathcal{H}^{m-n}(x)d\mathcal{L}^n(y)
\]

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for every Lebesgue measurable subset $S$ of $\mathbb{R}^m$ and every $\mathcal{L}^m$-measurable function $g : S \to \mathbb{R}$, where for each $x \in \mathbb{R}^m$ at which $f$ is differentiable,

$$J_n f(x) = \sqrt{\det((Jf(x))(Jf(x))^T)}.$$

Let $S$ be an $m$-dimensional $C^1$ submanifold of $\mathbb{R}^n$. Let $x \in S$ and let $\phi : U \subset \mathbb{R}^m \to \mathbb{R}^n$ be as in the definition of $m$-dimensional $C^1$ submanifold. We denote by $T_S(x)$ the tangent space of $S$ at $x$, $\{J\phi(u) : u \in \mathbb{R}^m\}$, where $u = \phi^{-1}(x)$.

**Lemma B.4** (Area Formula, Lemma 5.3.5 and Theorem 5.3.7 of Krantz and Parks (2008)). Suppose $m \leq \nu$ and $f : \mathbb{R}^n \to \mathbb{R}^\nu$ is Lipschitz. If $S$ is an $m$-dimensional $C^1$ submanifold of $\mathbb{R}^n$, then

$$\int_S g(x) J^S_m f(x) d\mathcal{H}^m(x) = \int_{\mathbb{R}^\nu} \sum_{x \in S \cap \{ f(y) = y \}} g(x) d\mathcal{H}^m(y)$$

for every $\mathcal{H}^m$-measurable function $g : S \to \mathbb{R}$, where for each $x \in \mathbb{R}^n$ at which $f$ is differentiable,

$$J^S_m f(x) = \frac{\mathcal{H}^m(\{ Jf(x) y : y \in P \})}{\mathcal{H}^m(P)}$$

for an arbitrary $m$-dimensional parallelepiped $P$ contained in $T_S(x)$.

Let $S \subset \mathbb{R}^p$. For each $x \in \mathbb{R}^p$ at which $d_S^a$ is differentiable and for each $\lambda \in \mathbb{R}$, let $\psi_S(x, \lambda) = x + \lambda \nabla d_S^a(x)$.

**Lemma B.5.** Let $\Omega \subset \mathbb{R}^p$, and suppose that there exists a partition $\{\Omega_1, ..., \Omega_M\}$ of $\Omega$ such that

(i) $\text{dist}(\Omega_m, \Omega_{m'}) > 0$ for any $m, m' \in \{1, ..., M\}$ such that $m \neq m'$;

(ii) $\Omega_m$ is nonempty, bounded, open, connected and twice continuously differentiable for each $m \in \{1, ..., M\}$.

Then there exists $\mu > 0$ such that $d^a_{\Omega}$ is twice continuously differentiable on $N(\partial \Omega, \mu)$ and that

$$\int_{N(\partial \Omega, \delta)} g(x) dx = \int_{-\delta}^{\delta} \int_{\partial \Omega} g(u + \lambda \nu_{\Omega}(u)) J^\partial_{p-1} \psi_{\Omega}(u, \lambda) d\mathcal{H}^{p-1}(u) d\lambda$$

for every $\delta \in (0, \mu)$ and every function $g : \mathbb{R}^p \to \mathbb{R}$ that is integrable on $N(\partial \Omega, \delta)$, where for each fixed $\lambda \in (-\mu, \mu)$, $J^{\partial_{p-1}}_{\Omega} \psi_{\Omega}(\cdot, \lambda)$ is calculated by applying the operation $J^{\partial_{p-1}}_{\Omega}$ to the function $\psi_{\Omega}(\cdot, \lambda)$. Furthermore, $J^{\partial_{p-1}}_{\Omega} \psi_{\Omega}(x, \cdot)$ is continuously differentiable in $\lambda$ and $J^{\partial_{p-1}}_{\Omega} \psi_{\Omega}(x, 0) = 1$ for every $x \in \partial \Omega$, and $J^{\partial_{p-1}}_{\Omega} \psi_{\Omega}(\cdot, \cdot)$ and $\frac{\partial J^{\partial_{p-1}}_{\Omega} \psi_{\Omega}(\cdot, \cdot)}{\partial \lambda}$ are bounded on $\partial \Omega \times (-\mu, \mu)$.

Proof. Let $\tilde{\mu} = \frac{1}{2} \min_{m, m' \in \{1, ..., M\}, m \neq m'} \text{dist}(\Omega^*_{m}, \Omega_{m'})$ so that $\{N(\partial \Omega_m, \tilde{\mu})\}_{m=1}^M$ is a partition of $N(\partial \Omega, \mu)$. Note that for every $m \in \{1, ..., M\}$, $d^a_{\Omega}(x) = d^a_{\Omega_m}(x)$ for every $x \in N(\partial \Omega_m, \tilde{\mu})$. By Lemma B.1, for every $m \in \{1, ..., M\}$, there exists $\bar{\mu}_m > 0$ such that $d^a_{\Omega_m}$ is twice continuously differentiable on $N(\partial \Omega_m, \bar{\mu}_m)$. Letting $\bar{\mu} \in (0, \min\{\bar{\mu}_1, \bar{\mu}_2, ..., \bar{\mu}_M\})$, we have that $d^a_{\Omega}$ is twice
continuously differentiable on $N(\partial \Omega, \mu)$. This implies that $d^s_\Omega$ is Lipschitz on $N(\partial \Omega, \mu)$. For every $\delta \in (0, \mu)$ and every function $g : \mathbb{R}^p \to \mathbb{R}$ that is integrable on $N(\partial \Omega, \delta)$,

$$
\int_{N(\partial \Omega, \delta)} g(x) dx = \int_{\{x' \in \mathbb{R}^p : d^s_\Omega(x') \in (-\delta, \delta)\}} g(x) \sqrt{\det(\|\nabla d^s_\Omega(x)\|)} dx
$$

$$
= \int_{\{x' \in \mathbb{R}^p : d^s_\Omega(x') \in (-\delta, \delta)\}} g(x) \sqrt{\det(\nabla d^s_\Omega(x)'\nabla d^s_\Omega(x))} dx
$$

$$
= \int_{\{x' \in \mathbb{R}^p : d^s_\Omega(x') \in (-\delta, \delta)\}} g(x) \sqrt{\det((Jd^s_\Omega(x))(Jd^s_\Omega(x)'))} dx
$$

$$
= \int_{\mathbb{R}} \int_{\{x' \in \mathbb{R}^p : d^s_\Omega(x') \in (-\delta, \delta), d^s_\Omega(x') = \lambda\}} g(x) d \mathcal{H}^{p-1}(x) d \lambda
$$

$$
= \int_{-\delta}^\delta \int_{\{x' \in \mathbb{R}^p : d^s_\Omega(x') = \lambda\}} g(x) d \mathcal{H}^{p-1}(x) d \lambda,
$$

(12)

where the first equality follows since $\|\nabla d^s_\Omega(x)\| = 1$ for every $x \in N(\partial \Omega, \delta)$ by Lemma B.1, the third equality follows from the definition of the Jacobian matrix, and the fourth equality follows from Lemma B.3.

Let $\Gamma(\lambda) = \{x \in \mathbb{R}^p : d^s_\Omega(x) = \lambda\}$ for each $\lambda \in (-\mu, \mu)$. Since $\nabla d^s_\Omega$ is differentiable on $N(\partial \Omega, \mu)$, $\psi_\Omega(x, \lambda)$ is defined on $N(\partial \Omega, \mu) \times \mathbb{R}$. We show that $\{\psi_\Omega(x_0, \lambda) : x_0 \in \partial \Omega\} \subset \Gamma(\lambda)$ for every $\lambda \in (-\mu, \mu)$. By Lemma B.1, for every $x_0 \in \partial \Omega$, $\psi_\Omega(x_0, \lambda) = x_0 + \lambda \nu_\Omega(x_0)$ and

$$
\Pi_{\partial \Omega}(\psi_\Omega(x_0, \lambda)) = \Pi_{\partial \Omega}(x_0 + \lambda \nu_\Omega(x_0)) = \{x_0\}.
$$

Hence,

$$
d(\psi_\Omega(x_0, \lambda), \partial \Omega) = \|\psi_\Omega(x_0, \lambda) - x_0\| = \|\lambda \nu_\Omega(x_0)\| = |\lambda|.
$$

Since $\nu_\Omega(x_0)$ is an inward normal vector, $\psi_\Omega(x_0, \lambda) \in \text{cl}(\Omega)$ if $0 \leq \lambda < \mu$, and $\psi_\Omega(x, \lambda_0) \in \mathbb{R}^p \setminus \text{cl}(\Omega)$ if $-\mu < \lambda < 0$. It follows that

$$
d^s_\Omega(\psi_\Omega(x_0, \lambda)) = \begin{cases} |\lambda| & \text{if } 0 \leq \lambda < \mu \\
-|\lambda| & \text{if } \mu < \lambda < 0 \end{cases}
$$

= $\lambda$,

so $\{\psi_\Omega(x_0, \lambda) : x_0 \in \partial \Omega\} \subset \Gamma(\lambda)$. It also holds that $\Gamma(\lambda) \subset \{\psi_\Omega(x_0, \lambda) : x_0 \in \partial \Omega\}$, since by Lemma B.1, for every $x \in \Gamma(\lambda)$,

$$
\psi_\Omega(\pi_{\partial \Omega}(x, \lambda) = \pi_{\partial \Omega}(x) + \lambda \nabla d^s_\Omega(\pi_{\partial \Omega}(x)) = \pi_{\partial \Omega}(x) + d^s_\Omega(x) \nu_\Omega(\pi_{\partial \Omega}(x)) = x,
$$

where $\pi_{\partial \Omega}(x)$ is the unique element in $\Pi_{\partial \Omega}(x)$. Thus, $\{\psi_\Omega(x_0, \lambda) : x_0 \in \partial \Omega\} = \Gamma(\lambda)$.

Now note that $\{\partial \Omega_m\}_{m=1}^M$ is a partition of $\partial \Omega$, since $\text{dist}(\Omega_m, \Omega_{m'}) > 0$ for any $m, m' \in \{1, \ldots, M\}$ such that $m \neq m'$. By Lemma B.2, $\partial \Omega_m$ is a $(p-1)$-dimensional $C^1$ submanifold of $\mathbb{R}^p$ for every $m \in \{1, \ldots, M\}$, and hence $\partial \Omega$ is a $(p-1)$-dimensional $C^1$ submanifold of $\mathbb{R}^p$. Furthermore, since $\nabla d^s_\Omega$ is continuously differentiable on $N(\partial \Omega, \mu)$, $\psi_\Omega(\cdot, \lambda)$ is continuously
differentiable on $N(\partial \Omega, \mu)$, which implies that $\psi_\Omega(\cdot, \lambda)$ is Lipschitz on $N(\partial \Omega, \mu)$ for every $\lambda \in \mathbb{R}$. Applying Lemma B.4, we have that for every $\lambda \in (-\mu, \mu)$,

$$\int_{\partial \Omega} g(u + \lambda \nu_\Omega(u)) J^{\partial \Omega}_{p-1} \psi_\Omega(u, \lambda) dH^{p-1}(u) = \int_{\partial \Omega} g(\psi_\Omega(u, \lambda)) J^{\partial \Omega}_{p-1} \psi_\Omega(u, \lambda) dH^{p-1}(u)$$

$$= \int_{\mathbb{R}^p} \sum_{u \in \partial \Omega : \psi_\Omega(u, \lambda) = x} g(\psi_\Omega(u, \lambda)) dH^{p-1}(x). \quad (13)$$

If $x \notin \{\psi_\Omega(u, \lambda) : u \in \partial \Omega\}$, $\{u \in \partial \Omega : \psi_\Omega(u, \lambda) = x\} = \emptyset$. If $x \in \{\psi_\Omega(u, \lambda) : u \in \partial \Omega\}$, there exists $u \in \partial \Omega$ such that $x = \psi_\Omega(u, \lambda)$. Since $P_{\partial \Omega}(x) = P_{\partial \Omega}(u + \lambda \nabla d^p_{\Omega}(u)) = P_{\partial \Omega}(u + \lambda \nu_\Omega(u)) = \{u\}$ by Lemma B.1, such $u$ is unique, and hence $\{u \in \partial \Omega : \psi_\Omega(u, \lambda) = x\}$ is a singleton. It follow that

$$\int_{\mathbb{R}^p} \sum_{u \in \partial \Omega : \psi_\Omega(u, \lambda) = x} g(\psi_\Omega(u, \lambda)) dH^{p-1}(x) = \int_{\{\psi_\Omega(u, \lambda) : u \in \partial \Omega\}} g(x) dH^{p-1}(x)$$

$$= \int_{\Gamma(\lambda)} g(x) dH^{p-1}(x), \quad (14)$$

where the last equality holds since $\{\psi_\Omega(u, \lambda) : u \in \partial \Omega\} = \Gamma(\lambda)$. Combining (12), (13) and (14), we obtain

$$\int_{N(\partial \Omega, \delta)} g(x) dx = \int \sum_{\delta} \int_{\partial \Omega} g(u + \lambda \nu_\Omega(u)) J^{\partial \Omega}_{p-1} \psi_\Omega(u, \lambda) dH^{p-1}(u) d\lambda. $$

We next show that $J^{\partial \Omega}_{p-1} \psi_\Omega(x, \cdot)$ is continuously differentiable in $\lambda$ and $J^{\partial \Omega}_{p-1} \psi_\Omega(x, 0) = 1$ for every $x \in \partial \Omega$. Fix an $x \in \partial \Omega$, and let $V_\Omega(x)$ be an arbitrary $p \times (p-1)$ matrix whose columns $v_1(x), \ldots, v_{p-1}(x) \in \mathbb{R}^p$ form an orthonormal basis of $T_{\partial \Omega}(x)$. Let $P(x) \subset T_{\partial \Omega}(x)$ be a parallellepipiped determined by $v_1(x), \ldots, v_{p-1}(x)$, that is, let $P(x) = \{\sum_{k=1}^{p-1} c_k v_k(x) : 0 \leq c_k \leq 1 \text{ for } k = 1, \ldots, p-1\}$. Since $v_1(x), \ldots, v_{p-1}(x)$ are linearly independent, $P(x)$ is a $(p-1)$-dimensional parallellepipeded. It follows that for each fixed $\lambda \in \mathbb{R}$,

$$\{J_{\psi_\Omega}(x, \lambda) y : y \in P(x)\} = \{J_{\psi_\Omega}(x, \lambda) \sum_{k=1}^{p-1} c_k v_k(x) : 0 \leq c_k \leq 1 \text{ for } k = 1, \ldots, p-1\}$$

$$= \{\sum_{k=1}^{p-1} c_k J_{\psi_\Omega}(x, \lambda) v_k(x) : 0 \leq c_k \leq 1 \text{ for } k = 1, \ldots, p-1\}$$

$$= \{\sum_{k=1}^{p-1} c_k w_k(x, \lambda) : 0 \leq c_k \leq 1 \text{ for } k = 1, \ldots, p-1\},$$

where $w_k(x, \lambda) = J_{\psi_\Omega}(x, \lambda) v_k(x)$ for $k = 1, \ldots, p-1$. Since $J_{\psi_\Omega}(x, \lambda)v_k(x)$ is the $k$-th column of $J_{\psi_\Omega}(x, \lambda)V_\Omega(x)$, $\{J_{\psi_\Omega}(x, \lambda) y : y \in P(x)\}$ is the parallelepipeded determined by the columns of
\[ J \psi(x, \lambda) V(x) \]. By Proposition 5.1.2 of Krantz and Parks (2008), we have that
\[
J_{p-1}^{\partial \Omega} \psi(x, \lambda) = \frac{\mathcal{H}^{-1}\left(\{\sum_{k=1}^{p-1} c_k w_k(x, \lambda) : 0 \leq c_k \leq 1 \text{ for } k = 1, \ldots, p-1\}\right)}{\mathcal{H}^{-1}(P(x))}
\]
\[
= \sqrt{\text{det}\left((J \psi(x, \lambda) V(x))' (J \psi(x, \lambda) V(x))\right)}
\]
\[
= \sqrt{\text{det}\left((V(x) + \lambda D^2 d_{ij}^s(x) V(x))' (V(x) + \lambda D^2 d_{ij}^s(x) V(x))\right)}
\]
\[
= \sqrt{\text{det}(I - I_{p-1} + \lambda V(x)' (2D^2 d_{ij}^s(x) + \lambda (D^2 d_{ij}^s(x))^2) V(x))}
\]
\[
= \sqrt{\text{det}(I + \lambda V(x)' (2D^2 d_{ij}^s(x) + \lambda (D^2 d_{ij}^s(x))^2))}
\]
where we use the fact that \( V(x)' V(x) = I_{p-1} \) and the fact that \( \text{det}(I_m + AB) = \text{det}(I_n + BA) \) for an \( m \times n \) matrix \( A \) and an \( n \times m \) matrix \( B \) (the Weinstein-Aronszajn identity). For every \( x \in \partial \Omega \), \( J_{p-1}^{\partial \Omega} \psi(x, \cdot) \) is continuously differentiable in \( \lambda \), and \( J_{p-1}^{\partial \Omega} \psi(x, 0) = \sqrt{\text{det}(I_p)} = 1 \).

Lastly, we show that \( J_{p-1}^{\partial \Omega} \psi(x, \cdot) \) and \( \frac{\partial J_{p-1}^{\partial \Omega} \psi(x, \cdot)}{\partial \lambda} \) are bounded on \( \partial \Omega \times (-\mu, \mu) \). Let \( f, h : \partial \Omega \times \mathbb{R}^{p \times (p-1)} \to \mathbb{R}^{p \times p} \) be functions such that
\[
f(x, A) = 2AA' D^2 d_{ij}^s(x),
\]
\[
h(x, A) = AA'(D^2 d_{ij}^s(x))^2.
\]
Also, let \( k : \partial \Omega \times \mathbb{R} \times \mathbb{R}^{p \times (p-1)} \to \mathbb{R} \) be a function such that
\[
k(x, \lambda, A) = \sqrt{\text{det}(I_p + \lambda f(x, A) + \lambda^2 h(x, A))}.
\]
Observe that
\[
J_{p-1}^{\partial \Omega} \psi(x, \lambda) = k(x, \lambda, V(x))
\]
and that
\[
\frac{\partial J_{p-1}^{\partial \Omega} \psi(x, \lambda)}{\partial \lambda} = \frac{\partial k(x, \lambda, A)}{\partial \lambda} \bigg|_{A = V(x)}
\]
\[
= \frac{1}{2k(x, \lambda, A)} \sum_{i,j} \frac{\partial \text{det}(I_p + \lambda f(x, A) + \lambda^2 h(x, A))}{\partial b_{ij}} (f_{ij}(x, A) + 2\lambda h_{ij}(x, A)) \bigg|_{A = V(x)}
\]
where \( \frac{\partial \text{det}(B)}{\partial b_{ij}} \) denotes the partial derivative of the function \( \text{det} : \mathbb{R}^{p \times p} \to \mathbb{R} \) with respect to the \((i, j)\) entry of \( B \).

Note that \( k(\cdot, \cdot, \cdot) \) and \( \frac{\partial k(\cdot, \cdot, \cdot)}{\partial \lambda} \) are continuous on \( \partial \Omega \times \mathbb{R} \times \mathbb{R}^{p \times (p-1)} \) (except at the points for which \( k(x, \lambda, A) = 0 \)), since \( \text{det} \) is infinitely differentiable, and \( f \) and \( h \) are continuous on
Lemma B.6. Let \( \{V_i\}_{i=1}^{\infty} \) be i.i.d. random variables such that \( E[V_i^2] < \infty \). If Assumption 1 holds, then for \( l \geq 0 \) and \( m = 0, 1 \),
\[
E[V_i p^A(X_i; \delta)]^l 1\{p^A(X_i; \delta) \in (0,1)\}^m \rightarrow E[V_i A(X_i)]^l 1\{A(X_i) \in (0,1)\}^m
\]
as \( \delta \rightarrow 0 \). Moreover, if, in addition, \( \delta_n \rightarrow 0 \) as \( n \rightarrow \infty \), then for \( l \geq 0 \),
\[
\frac{1}{n} \sum_{i=1}^{n} V_i p^A(X_i; \delta_n)^l I_{i,n} \xrightarrow{p} E[V_i A(X_i)]^l 1\{A(X_i) \in (0,1)\}
\]
as \( n \rightarrow \infty \).

Proof. Note that \( E[\frac{1}{n} \sum_{i=1}^{n} V_i p^A(X_i; \delta_n)^l I_{i,n}] = E[V_i p^A(X_i; \delta_n)^l 1\{p^A(X_i; \delta_n) \in (0,1)\}] \). We show that
\[
E[V_i p^A(X_i; \delta)]^l 1\{p^A(X_i; \delta) \in (0,1)\}^m \rightarrow E[V_i A(X_i)]^l 1\{A(X_i) \in (0,1)\}^m
\]
for \( l \geq 0 \) and \( m = 0, 1 \) as \( \delta \rightarrow 0 \), and that
\[
\text{Var}\left(\frac{1}{n} \sum_{i=1}^{n} V_i p^A(X_i; \delta_n)^l I_{i,n}\right) \rightarrow 0
\]
for \( l \geq 0 \) as \( n \rightarrow \infty \). For the first part, we have
\[
E[V_i p^A(X_i; \delta)]^l 1\{p^A(X_i; \delta) \in (0,1)\}^m = \int_{\mathcal{X}} E[V_i |X_i = x] p^A(x; \delta)]^l 1\{p^A(x; \delta) \in (0,1)\}^m f_X(x) dx.
\]
Suppose \( A \) is continuous at \( x \) and \( A(x) \in (0,1) \). Then \( \lim_{\delta \rightarrow 0} p^A(x; \delta) = A(x) \) by Part 1 of Corollary A.1, and hence \( p^A(x; \delta) \in (0,1) \) for sufficiently small \( \delta > 0 \). It follows that \( 1\{p^A(x; \delta) \in (0,1)\} \rightarrow 1 = 1\{A(x) \in (0,1)\} \) as \( \delta \rightarrow 0 \). Suppose \( x \in \text{int}(\mathcal{X}_0) \cup \text{int}(\mathcal{X}_1) \). Then \( B(x, \delta) \subset \mathcal{X}_0 \) or \( B(x, \delta) \subset \mathcal{X}_1 \) for sufficiently small \( \delta > 0 \) by the fact that \( \text{int}(\mathcal{X}_0) \) and \( \text{int}(\mathcal{X}_1) \) are open, and hence \( 1\{p^A(x; \delta) \in (0,1)\} \rightarrow 0 = 1\{A(x) \in (0,1)\} \) as \( \delta \rightarrow 0 \). Therefore, \( \lim_{\delta \rightarrow 0} p^A(x; \delta) = A(x) \) and \( \lim_{\delta \rightarrow 0} 1\{p^A(x; \delta) \in (0,1)\} = 1\{A(x) \in (0,1)\} \) for almost every \( x \in \mathcal{X} \), since \( A \) is continuous at \( x \) for almost every \( x \in \mathcal{X} \) by Assumption 1 (a), and either \( A(x) \in (0,1) \) or \( x \in \text{int}(\mathcal{X}_0) \cup \text{int}(\mathcal{X}_1) \) for almost every \( x \in \mathcal{X} \) by Assumption 1 (b). By the Dominated Convergence Theorem,
\[
E[V_i p^A(X_i; \delta)]^l 1\{p^A(X_i; \delta) \in (0,1)\}^m \rightarrow \int_{\mathcal{X}} E[V_i |X_i = x] A(x)^l 1\{A(x) \in (0,1)\}^m f_X(x) dx
\]
\[
= E[V_i A(X_i)]^l 1\{A(X_i) \in (0,1)\}^m
\]
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as $\delta \to 0$. As for variance,

$$\text{Var}\left(\frac{1}{n} \sum_{i=1}^{n} V_i p^A(X_i; \delta_n) I_{i,n}\right) \leq \frac{1}{n} E[V_i^2 p^A(X_i; \delta_n)^2 (I_{i,n})^2]$$

$$\leq \frac{1}{n} E[V_i^2]$$

$$\to 0$$

as $n \to \infty$. \hfill \Box$

**Lemma B.7.** Let $\{(\delta_n, S_n)\}_{n=1}^{\infty}$ be any sequence of positive numbers and positive integers. Fix $x \in \mathcal{X}$, and let $X_1^*, \ldots, X_{S_n}^*$ be $S_n$ independent draws from the uniform distribution on $B(x, \delta_n)$ so that

$$p^s(x; \delta_n) = \frac{1}{S_n} \sum_{s=1}^{S_n} A(X_s^*).$$

Then,

$$E[p^s(x; \delta_n) - p^A(x; \delta_n)] = 0,$$

$$E[(p^s(x; \delta_n) - p^A(x; \delta_n))^2] \leq \frac{1}{S_n},$$

$$|E[p^s(x; \delta_n)^2 - p^A(x; \delta_n)^2]| \leq \frac{1}{S_n},$$

$$E[(p^s(x; \delta_n)^2 - p^A(x; \delta_n)^2)^2] \leq \frac{4}{S_n},$$

$$\Pr(p^s(x; \delta_n) \in \{0, 1\}) \leq (1 - p^A(x; \delta_n))^{S_n} + p^A(x; \delta_n)^{S_n}.$$

Moreover, for any $\epsilon > 0$,

$$E[|p^s(x; \delta_n) - p^A(x; \delta_n)|] \leq \frac{1}{S_n \epsilon^2} + \epsilon,$$

and if $S_n \to \infty$, then

$$E[|p^s(x; \delta_n) - p^A(x; \delta_n)|] \to 0$$

as $n \to \infty$.

**Proof.** By construction, $E[A(X_s^*)] = p^A(x; \delta_n)$, so

$$E[p^s(x; \delta_n) - p^A(x; \delta_n)] = E\left[\frac{1}{S_n} \sum_{s=1}^{S_n} A(X_s^*)\right] - p^A(x; \delta_n)$$

$$= E[A(X_s^*)] - p^A(x; \delta_n)$$

$$= 0.$$
We have

\[ E[(p^s(x; \delta_n) - p^A(x; \delta_n))^2] = \text{Var}(p^s(x; \delta_n)) \]

\[ = \text{Var}\left(\frac{1}{S_n} \sum_{s=1}^{S_n} A(X^*_s)\right) \]

\[ = \frac{1}{S_n} \text{Var}(A(X^*_s)) \]

\[ \leq \frac{1}{S_n} E[A(X^*_s)^2] \]

\[ \leq \frac{1}{S_n}, \]

\[ |E[p^s(x; \delta_n)^2 - p^A(x; \delta_n)^2]| = |\text{Var}(p^s(x; \delta_n)) + (E[p^s(x; \delta_n)])^2 - p^A(x; \delta_n)^2| \]

\[ \leq \frac{1}{S_n} + |(p^A(x; \delta_n))^2 - p^A(x; \delta_n)^2| \]

\[ = \frac{1}{S_n}, \]

and

\[ E[(p^s(x; \delta_n)^2 - p^A(x; \delta_n)^2)^2] \]

\[ = E[(p^s(x; \delta_n) + p^A(x; \delta_n))^2(p^s(x; \delta_n) - p^A(x; \delta_n))^2] \]

\[ \leq 4E[(p^s(x; \delta_n) - p^A(x; \delta_n))^2] \]

\[ \leq \frac{4}{S_n}. \]

Now note that we have the following bounds on \( \Pr(A(X^*_s) = 0) \) and \( \Pr(A(X^*_s) = 1) \):

\[ 0 \leq \Pr(A(X^*_s) = 0) \leq 1 - p^A(x; \delta_n), \]

\[ 0 \leq \Pr(A(X^*_s) = 1) \leq p^A(x; \delta_n). \]

It follows that

\[ 0 \leq \Pr(p^s(x; \delta_n) \in \{0, 1\}) \]

\[ = \Pr(A(X^*_s) = 0)^{S_n} + \Pr(A(X^*_s) = 1)^{S_n} \]

\[ \leq (1 - p^A(x; \delta_n))^{S_n} + p^A(x; \delta_n)^{S_n}. \]

Lastly, for any \( \epsilon > 0 \),

\[ E[|p^s(x; \delta_n) - p^A(x; \delta_n)|] \]

\[ = E[|p^s(x; \delta_n) - p^A(x; \delta_n)||p^s(x; \delta_n) - p^A(x; \delta_n)| \geq \epsilon \Pr(|p^s(x; \delta_n) - p^A(x; \delta_n)| \geq \epsilon) \]

\[ + E[|p^s(x; \delta_n) - p^A(x; \delta_n)||p^s(x; \delta_n) - p^A(x; \delta_n)| < \epsilon \Pr(|p^s(x; \delta_n) - p^A(x; \delta_n)| < \epsilon) \]

\[ < 1 \cdot \frac{\text{Var}(p^s(x; \delta_n))}{\epsilon^2} + \epsilon \cdot 1 \]

\[ \leq \frac{1}{S_n\epsilon^2} + \epsilon, \]
where we use Chebyshev’s inequality for the first inequality. We can make \( E[|p^s(x; \delta_n) - p^A(x; \delta_n)|] \) arbitrarily close to zero by taking sufficiently small \( \epsilon > 0 \) and sufficiently large \( S_n \), which implies that \( E[|p^s(x; \delta_n) - p^A(x; \delta_n)|] = o(1) \) if \( S_n \to \infty \).

\[ \square \]

**Lemma B.8.** Let \( I_{i,n}^s = 1\{p^s(X_i; \delta_n) \in (0,1)\} \), and let \( \{V_i\}_{i=1}^\infty \) be i.i.d. random variables such that \( E[V_i^2] < \infty \). If Assumption 1 holds, \( S_n \to \infty \), and \( \delta_n \to 0 \), then

\[
\frac{1}{n} \sum_{i=1}^{n} V_ip^s(X_i; \delta_n)^l I_{i,n}^s - \frac{1}{n} \sum_{i=1}^{n} V_ip^A(X_i; \delta_n)^l I_{i,n} = o_p(1)
\]

for \( l = 0, 1, 2, 3, 4 \). If, in addition, Assumption 5 holds, and \( E[V_i|X_i] \) is bounded, then

\[
\frac{1}{\sqrt{n}} \sum_{i=1}^{n} V_ip^s(X_i; \delta_n)^l I_{i,n}^s - \frac{1}{\sqrt{n}} \sum_{i=1}^{n} V_ip^A(X_i; \delta_n)^l I_{i,n} = o_p(1)
\]

for \( l = 0, 1, 2 \).

**Proof.** We have

\[
\frac{1}{n} \sum_{i=1}^{n} V_ip^s(X_i; \delta_n)^l I_{i,n}^s - \frac{1}{n} \sum_{i=1}^{n} V_ip^A(X_i; \delta_n)^l I_{i,n} = \frac{1}{n} \sum_{i=1}^{n} V_i(p^s(X_i; \delta_n)^l - p^A(X_i; \delta_n)^l) I_{i,n}.
\]

We first consider \( \frac{1}{n} \sum_{i=1}^{n} V_i(p^s(X_i; \delta_n)^l - p^A(X_i; \delta_n)^l) I_{i,n} \). By Lemma B.7, for \( l = 0, 1, 2 \),

\[
|E[\frac{1}{n} \sum_{i=1}^{n} V_i(p^s(X_i; \delta_n)^l - p^A(X_i; \delta_n)^l) I_{i,n}]| \\
= |E[V_i(p^s(X_i; \delta_n)^l - p^A(X_i; \delta_n)^l) I_{i,n})]| \\
= E[|E[V_i|X_i]|E[p^s(X_i; \delta_n)^l - p^A(X_i; \delta_n)^l|X_i]|I_{i,n}] \\
\leq \frac{1}{S_n} E[|E[V_i|X_i]|I_{i,n}] \\
= O(S_n^{-1}).
\]

Also, by Lemma B.7,

\[
|E[\frac{1}{n} \sum_{i=1}^{n} V_i(p^s(X_i; \delta_n)^3 - p^A(X_i; \delta_n)^3) I_{i,n}]| \\
= |E[V_i(p^s(X_i; \delta_n) - p^A(X_i; \delta_n))(p^s(X_i; \delta_n) - p^A(X_i; \delta_n)p^A(X_i; \delta_n) + p^A(X_i; \delta_n)^2) I_{i,n}]| \\
\leq E[|E[V_i|X_i]|E[(p^s(X_i; \delta_n) - p^A(X_i; \delta_n))(p^s(X_i; \delta_n) - p^A(X_i; \delta_n)p^A(X_i; \delta_n) + p^A(X_i; \delta_n)^2)|X_i]|I_{i,n}] \\
\leq 3E[|E[V_i|X_i]|E[(p^s(X_i; \delta_n) - p^A(X_i; \delta_n))|X_i]|I_{i,n}] \\
= o(1),
\]
and

\[
|E[\frac{1}{n} \sum_{i=1}^{n} V_i(p^s(X_i; \delta_n)^4 - p^4(X_i; \delta_n)^4)I_{i,n}]|
\]

\[
=E[|V_i(p^s(X_i; \delta_n)^4 + p^4(X_i; \delta_n)^4)(p^4(X_i; \delta_n) - p^4(X_i; \delta_n))I_{i,n}|]
\]

\[
\leq E[|E[V_i|X_i]|E[(p^s(X_i; \delta_n)^4 + p^4(X_i; \delta_n)^4)(p^4(X_i; \delta_n) - p^4(X_i; \delta_n))|X_i]|I_{i,n}]
\]

\[
\leq 4E[|E[V_i|X_i]|E[|p^s(X_i; \delta_n) - p^4(X_i; \delta_n)||X_i|I_{i,n}]
\]

\[
= o(1).
\]

As for variance, for \( l = 0, 1, 2, \)

\[
\text{Var}\left(\frac{1}{n} \sum_{i=1}^{n} V_i(p^s(X_i; \delta_n)^l - p^4(X_i; \delta_n)^l)I_{i,n}\right) \leq \frac{1}{n} E[|V_i^2(p^s(X_i; \delta_n)^l - p^4(X_i; \delta_n)^l)^2I_{i,n}|]
\]

\[
\leq \frac{1}{n} E[E[V_i^2|X_i]E[(p^s(X_i; \delta_n)^l - p^4(X_i; \delta_n)^l)^2|X_i]|I_{i,n}]
\]

\[
\leq \frac{4}{nS_n} E[E[V_i^2|X_i]I_{i,n}]
\]

\[
= O((nS_n)^{-1}),
\]

and for \( l = 3, 4, \)

\[
\text{Var}\left(\frac{1}{n} \sum_{i=1}^{n} V_i(p^s(X_i; \delta_n)^l - p^4(X_i; \delta_n)^l)I_{i,n}\right) \leq \frac{1}{n} E[|V_i^2(p^s(X_i; \delta_n)^l - p^4(X_i; \delta_n)^l)^2I_{i,n}|]
\]

\[
= o(1).
\]

Therefore, \( \frac{1}{n} \sum_{i=1}^{n} V_i(p^s(X_i; \delta_n)^l - p^4(X_i; \delta_n)^l)I_{i,n} = o_p(1) \) if \( S_n \to \infty \) for \( l = 0, 1, 2, 3, 4, \) and \( \frac{1}{\sqrt{n}} \sum_{i=1}^{n} V_i(p^s(X_i; \delta_n)^l - p^4(X_i; \delta_n)^l)I_{i,n} = o_p(1) \) if \( n^{-1/2}S_n \to \infty \) for \( l = 0, 1, 2. \)

We next show that \( \frac{1}{n} \sum_{i=1}^{n} V_i p^s(X_i; \delta_n)^l(I_{i,n} - 1) = o_p(1) \) if \( S_n \to \infty \) and \( \delta_n \to 0 \) for \( l \geq 0. \)

We have

\[
|E[\frac{1}{n} \sum_{i=1}^{n} V_i p^s(X_i; \delta_n)^l(I_{i,n} - 1)]]| = |E[V_i p^s(X_i; \delta_n)^l(I_{i,n} - 1)]|
\]

\[
\leq E[|E[V_i|X_i]||E[p^s(X_i; \delta_n)^l(I_{i,n} - 1)||X_i]|]
\]

\[
\leq E[|E[V_i|X_i]|E[|I_{i,n} - 1||X_i]].
\]

Note that by construction, \( I\{p^s(X_i; \delta_n) \in (0, 1)\} \leq I\{p^4(X_i; \delta_n) \in (0, 1)\} \) with probability one conditional on \( X_i = x, \) so that

\[
E[|I_{i,n} - 1||X_i = x] = -E[I_{i,n} - 1|X_i = x].
\]

Suppose \( A \) is continuous at \( x \) and \( A(x) \in (0, 1). \) Then \( \lim_{\delta \to 0} p^A(x; \delta) = A(x) \) by Part 1 of Corollary A.1, and hence \( p^A(x; \delta_n) \in [\epsilon, 1 - \epsilon] \) for sufficiently small \( \delta_n > 0 \) for some constant
\( \epsilon \in (0, 1/2) \). It follows that
\[
E[I_{i,n}^s|X_i = x] = 1 - \Pr(p^s(x; \delta_n) \in \{0, 1\}) \\
\geq 1 - (1 - p^A(x; \delta_n))^{S_n} - p^A(x; \delta_n)^{S_n} \\
\geq 1 - 2(1 - \epsilon)^{S_n} \\
\to 1
\]
as \( S_n \to \infty \), where the first inequality follows from Lemma B.7. This implies that \( E[I_{i,n}^s|X_i = x] \to 0 \) as \( n \to \infty \). Suppose \( x \in \text{int}(X_0) \cup \text{int}(X_1) \). Then \( B(x, \delta_n) \subset X_0 \) or \( B(x, \delta_n) \subset X_1 \) for sufficiently small \( \delta_n > 0 \) by the fact that \( \text{int}(X_0) \) and \( \text{int}(X_1) \) are open, and hence \( p^A(x; \delta_n) \in \{0, 1\} \) and \( p^s(x; \delta_n) \in \{0, 1\} \) for sufficiently small \( \delta_n > 0 \), so that \( E[I_{i,n}^s - I_{i,n}|X_i = x] \to 0 \) as \( n \to \infty \). Therefore, \( E[I_{i,n}^s - I_{i,n}|X_i = x] \to 0 \) for almost every \( x \in \mathcal{X} \), since \( A \) is continuous at \( x \) for almost every \( x \in \mathcal{X} \) by Assumption 1 (a), and either \( A(x) \in (0, 1) \) or \( x \in \text{int}(X_0) \cup \text{int}(X_1) \) for almost every \( x \in \mathcal{X} \) by Assumption 1 (b). By the Dominated Convergence Theorem,
\[
-E[|E[V_i|X_i]|E[I_{i,n}^s - I_{i,n}|X_i]] \to 0
\]
as \( n \to \infty \).

As for variance,
\[
\text{Var}\left( \frac{1}{n} \sum_{i=1}^{n} V_i p^s(X_i; \delta_n)^l(I_{i,n}^s - I_{i,n}) \right) \leq \frac{1}{n} E[V_i^2 p^s(X_i; \delta_n)^{2l}(I_{i,n}^s - I_{i,n})^2] \\
\leq \frac{1}{n} E[V_i^2] \\
\to 0.
\]

Lastly, we show that, for \( l \geq 0 \), \( \frac{1}{\sqrt{n}} \sum_{i=1}^{n} V_i p^s(X_i; \delta_n)^l(I_{i,n}^s - I_{i,n}) = o_p(1) \) if Assumption 5 holds, and \( E[V_i|X_i] \) is bounded. Let \( \eta_n = \gamma \frac{\log n}{S_n} \), where \( \gamma \) is the one satisfying Assumption 5. We have
\[
|E[\frac{1}{\sqrt{n}} \sum_{i=1}^{n} V_i p^s(X_i; \delta_n)^l(I_{i,n}^s - I_{i,n})]| \\
\leq \sqrt{n}E[|E[V_i|X_i]|E[|I_{i,n}^s - I_{i,n}|]|X_i] \\
= -\sqrt{n}E[|E[V_i|X_i]|E[I_{i,n}^s - 1|X_i]I_{i,n}] \\
\leq \sqrt{n}E[|E[V_i|X_i]|((1 - p^A(X_i; \delta_n))^{S_n} + p^A(X_i; \delta_n)^{S_n}))I_{i,n}] \\
= \sqrt{n}E[|E[V_i|X_i]|((1 - p^A(X_i; \delta_n))^{S_n} + p^A(X_i; \delta_n)^{S_n})1\{p^A(X_i; \delta_n) \in (0, \eta_n) \cup (1 - \eta_n, 1)\}] \\
+ \sqrt{n}E[|E[V_i|X_i]|((1 - p^A(X_i; \delta_n))^{S_n} + p^A(X_i; \delta_n)^{S_n})1\{p^A(X_i; \delta_n) \in [\eta_n, 1 - \eta_n]\}] \\
\leq (\sup_{x \in \mathcal{X}} |E[V_i|X_i = x]|)(\sqrt{n} \Pr(p^A(X_i; \delta_n) \in (0, \eta_n) \cup (1 - \eta_n, 1)) + 2\sqrt{n}(1 - \eta_n)^{S_n}),
\]
where the second equality follows from the fact that \( I_{i,n}^s \leq I_{i,n} \) with strict inequality only if \( I_{i,n} = 1 \). By Assumption 5, \( \sqrt{n} \Pr(p^A(X_i; \delta_n) \in (0, \eta_n) \cup (1 - \eta_n, 1)) = o(1) \). As for \( \sqrt{n}(1 - \eta_n)^{S_n} \), first observe that \( \eta_n = \gamma \frac{\log n}{S_n} = \frac{\log n}{n^{1/2} (n^{1/2} + 2)} \to 0 \), since \( n^{-1/2} S_n \to \infty \) and \( \frac{\log n}{n^{3/2}} \to 0 \). Using the
Since \( \gamma > 1/2 \). As for variance,

\[
\text{Var}\left( \frac{1}{\sqrt{n}} \sum_{i=1}^{n} V_i p^*(X_i; \delta_n)(I_{i,n}^s - I_{i,n}) \right) \leq E[V_i^2 p^*(X_i; \delta_n)^2 (I_{i,n}^s - I_{i,n})^2] \\
\leq E[V_i^2 |I_{i,n}^s - I_{i,n}|] \\
= E[E[V_i^2 |X_i|]E[|I_{i,n}^s - I_{i,n}| |X_i]] \\
= o(1).
\]

\[
\Box
\]

\section{Proofs}

\subsection{Proof of Proposition 1}

Suppose that Assumptions 1 and 2 hold. Here, we only show that

(a) \( E[Y_{1i} - Y_{0i} | X_i = x] \) is identified for every \( x \in \text{int}(\mathcal{X}) \) such that \( p^A(x) \in (0, 1) \).

(b) Let \( S \) be any open subset of \( \mathcal{X} \) such that \( p^A(x) \) exists for all \( x \in S \). Then \( E[Y_{1i} - Y_{0i} | X_i \in S] \) is identified only if \( p^A(x) \in (0, 1) \) for almost every \( x \in S \).

The results for \( E[D_i(1) - D_i(0) | X_i = x] \) and \( E[D_i(1) - D_i(0) | X_i \in S] \) are obtained by a similar argument.

**Proof of Part (a).** Pick an \( x \in \text{int}(\mathcal{X}) \) such that \( p^A(x) \in (0, 1) \). If \( A(x) \in (0, 1) \), \( E[Y_{1i} - Y_{0i} | X_i = x] \) and \( E[D_i(1) - D_i(0) | X_i = x] \) are trivially identified by Property 1:

\[
E[Y_i | X_i = x, Z_i = 1] - E[Y_i | X_i = x, Z_i = 0] = E[Y_{1i} - Y_{0i} | X_i = x].
\]

We next consider the case where \( A(x) \in \{0, 1\} \). Since \( x \in \text{int}(\mathcal{X}) \), \( B(x, \delta) \subset \mathcal{X} \) for any sufficiently small \( \delta > 0 \). Moreover, since \( p^A(x) = \lim_{\delta \to 0} p^A(x; \delta) \in (0, 1) \), \( p^A(x; \delta) \in (0, 1) \) for any sufficiently small \( \delta > 0 \). This implies that we can find points \( x_{0,\delta}, x_{1,\delta} \in B(x, \delta) (\subset \mathcal{X}) \) such that \( A(x_{0,\delta}) < 1 \) and \( A(x_{1,\delta}) > 0 \) for any sufficiently small \( \delta > 0 \), for otherwise \( p^A(x; \delta) \in \{0, 1\} \). Noting that \( x_{0,\delta} \to x \) and \( x_{1,\delta} \to x \) as \( \delta \to 0 \),

\[
\lim_{\delta \to 0} (E[Y_i | X_i = x_{1,\delta}, Z_i = 1] - E[Y_i | X_i = x_{0,\delta}, Z_i = 0]) = \lim_{\delta \to 0} (E[Y_{1i} | X_i = x_{1,\delta}] - E[Y_{0i} | X_i = x_{0,\delta}]) \\
= E[Y_{1i} - Y_{0i} | X_i = x],
\]

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where the first equality follows from Property 1, and the second from Assumption 2. □

Proof of Part (b).
Suppose to the contrary that \( \mathcal{L}^p(\{x \in S : p^A(x) \in \{0, 1\}\}) > 0 \). Without loss of generality, assume \( \mathcal{L}^p(\{x \in S : p^A(x) = 1\}) > 0 \). The proof proceeds in four steps.

Step C.1.1. \( \mathcal{L}^p(S \cap \mathcal{X}_1) > 0 \).

Proof. By Assumption 1, \( A \) is continuous almost everywhere. Part 1 of Corollary A.1 then implies that \( p^A(x) = A(x) \) for almost every \( x \in \{x^* \in S : p^A(x^*) = 1\} \). Since \( \mathcal{L}^p(\{x \in S : p^A(x) = 1\}) > 0 \), \( \mathcal{L}^p(\{x \in S : p^A(x) = 1\}) > 0 \), and hence \( \mathcal{L}^p(S \cap \mathcal{X}_1) > 0 \). □

Step C.1.2. \( S \cap \text{int}(\mathcal{X}_1) \neq \emptyset \).

Proof. Suppose that \( S \cap \text{int}(\mathcal{X}_1) = \emptyset \). Then, we must have that \( S \cap \mathcal{X}_1 \subset \mathcal{X}_1 \setminus \text{int}(\mathcal{X}_1) \). It then follows that \( \mathcal{L}^p(S \cap \mathcal{X}_1) \leq \mathcal{L}^p(\mathcal{X}_1 \setminus \text{int}(\mathcal{X}_1)) = \mathcal{L}^p(\mathcal{X}_1) - \mathcal{L}^p(\text{int}(\mathcal{X}_1)) = 0 \), where the last equality holds by Assumption 1. But this is a contradiction to the result from Step C.1.1. □

Step C.1.3. \( p^A(x) = 1 \) for any \( x \in \text{int}(\mathcal{X}_1) \).

Proof. Pick any \( x \in \text{int}(\mathcal{X}_1) \). By the definition of interior, \( B(x, \delta) \subset \mathcal{X}_1 \) for any sufficiently small \( \delta > 0 \). Therefore, \( p^A(x; \delta) = 1 \) for any sufficiently small \( \delta > 0 \). □

Step C.1.4. \( E[Y_{1i} - Y_{0i}|X_i \in S] \) is not identified.

Proof. We first introduce some notation. Let \( Q \) be the set of all distributions of \((Y_{1i}, Y_{0i}, X_i, Z_i)\) satisfying Property 1 and Assumptions 1 and 2. Let \( P \) be the set of all distributions of \((Y_i, X_i, Z_i)\). Let \( T : Q \to P \) be a function such that, for \( Q \in Q \), \( T(Q) \) is the distribution of \((Z_i Y_{1i} + (1 - Z_i) Y_{0i}, X_i, Z_i)\), where the distribution of \((Y_{1i}, Y_{0i}, X_i, Z_i)\) is \( Q \). Let \( Q_0 \) and \( P_0 \) denote the true distributions of \((Y_{1i}, Y_{0i}, X_i, Z_i)\) and \((Y_i, X_i, Z_i)\), respectively. Given \( P_0 \), the identified set of \( E[Y_{1i} - Y_{0i}|X_i \in S] \) is given by \{ \( E_Q[Y_{1i} - Y_{0i}|X_i \in S] \) : \( P_0 = T(Q), Q \in Q \) \}, where \( E_Q[\cdot] \) is the expectation operator under distribution \( Q \). We show that this set contains two distinct values.

In what follows, \( \text{Pr}(\cdot) \) and \( E[\cdot] \) without a subscript denote the probability and expectation under the true distributions \( Q_0 \) and \( P_0 \) as updated until now.

Now pick any \( x^* \in S \cap \text{int}(\mathcal{X}_1) \). Since \( A \) and \( \text{int}(\mathcal{X}_1) \) are open, there is some \( \delta > 0 \) such that \( B(x^*, \delta) \subset S \cap \text{int}(\mathcal{X}_1) \). Let \( \epsilon = \frac{\delta}{2} \), and consider a function \( f : \mathcal{X} \to \mathbb{R} \) such that \( f(x) = E[Y_{0i}|X = x] \) for all \( x \in \mathcal{X} \setminus B(x^*, \epsilon) \) and \( f(x) = E[Y_{0i}|X = x] - 1 \) for all \( x \in B(x^*, \epsilon) \). Below, we show that \( f \) is continuous at any point \( x \in \mathcal{X} \) such that \( p^A(x) \in (0, 1) \) and \( A(x) \in \{0, 1\} \).

Pick any \( x \in \mathcal{X} \) such that \( p^A(x) \in (0, 1) \) and \( A(x) \in \{0, 1\} \). Since \( B(x^*, \delta) \subset \text{int}(\mathcal{X}_1) \) and \( \text{int}(\mathcal{X}_1) \subset \{x' \in \mathcal{X} : p^A(x') = 1\} \) by Step C.1.3, \( x \notin B(x^*, \delta) \). Hence, \( B(x, \epsilon) \subset \mathcal{X} \setminus B(x^*, \epsilon) \). By Assumption 2 and the definition of \( f \), \( f \) is continuous at \( x \).

Now take any random vector \((Y_{1i}^*, Y_{0i}^*, X_i^*, Z_i^*)\) that is distributed according to the true distribution \( Q_0 \). Let \( Q \) be the distribution of \((Y_{1i}^Q, Y_{0i}^Q, X_i^Q, Z_i^Q)\), where \((Y_{1i}^Q, X_i^Q, Z_i^Q) = (Y_{1i}^*, X_i^*, Z_i^*)\), and

\[
Y_{0i}^* = \begin{cases} 
Y_{0i}^Q & \text{if } X_i^* \in \mathcal{X} \setminus B(x^*, \epsilon) \\
Y_{0i}^Q - 1 & \text{if } X_i^* \in B(x^*, \epsilon)
\end{cases}
\]
Note first that if \( E[Y_{i}^*|X_i^* = x] = E[Y_{i1}^*|X_i^* = x] \) and \( E_{Q}[Y_{i}^Q|X_i^Q = x] = f(x) \),
where \( E[Y_{i}^*|X_i^*] \) and \( f \) are both continuous at any point \( x \in \mathcal{X} \) such that \( p^A(x) \in (0,1) \) and \( A(x) \in \{0,1\} \). Also, \( Z_{i}^Q = Z_{i}^* = 1 \) if \( X_i^* \in B(x^*, \epsilon) \). It then follows that

\[
Y_i^Q = Z_i^QY_{i1}^Q + (1 - Z_i^Q)Y_{0i}^Q
\]

\[
= \begin{cases} 
Z_i^*Y_{i1}^* + (1 - Z_i^*)Y_{0i}^* & \text{if } X_i^* \in \mathcal{X} \setminus B(x^*, \epsilon) \\
Z_i^*Y_{i1}^* & \text{if } X_i^* \in B(x^*, \epsilon)
\end{cases}
\]

and

\[
Y_i^* = Z_i^*Y_{i1}^* + (1 - Z_i^*)Y_{0i}^*
\]

\[
= \begin{cases} 
Z_i^*Y_{i1}^* + (1 - Z_i^*)Y_{0i}^* & \text{if } X_i^* \in \mathcal{X} \setminus B(x^*, \epsilon) \\
Z_i^*Y_{i1}^* & \text{if } X_i^* \in B(x^*, \epsilon)
\end{cases}
\]

Thus, \( Y_i^Q = Y_i^* \), and hence \( T(Q) = T(Q_0) = P_0 \).

Using \( E_{Q}[Y_{i1}^Q|X_i^Q = x] = E[Y_{i1}^*|X_i^* = x] \) and \( E_{Q}[Y_{0i}^Q|X_i^Q = x] = f(x) \), we have

\[
E_{Q}[Y_{i1}^Q - Y_{0i}^Q|X_i^Q \in S] = E_{Q}[E_{Q}[Y_{i1}^Q|X_i^Q]|X_i^Q \in S] - E_{Q}[E_{Q}[Y_{0i}^Q|X_i^Q]|X_i^Q \in S, X_i^Q \notin B(x^*, \epsilon)]\Pr_{Q}(X_i^Q \notin B(x^*, \epsilon)|X_i^Q \in S)
\]

\[
- E_{Q}[E_{Q}[Y_{0i}^Q|X_i^Q]|X_i^Q \in B(x^*, \epsilon)]\Pr_{Q}(X_i^Q \in B(x^*, \epsilon)|X_i^Q \in S)
\]

\[
= E[Y_{i1}^*|X_i^* \in S] - E[f(X_i^*)|X_i^* \in S, X_i^* \notin B(x^*, \epsilon)]\Pr(X_i^* \notin B(x^*, \epsilon)|X_i^* \in S) - E[f(X_i^*)|X_i^* \in B(x^*, \epsilon)]\Pr(X_i^* \in B(x^*, \epsilon)|X_i^* \in S)
\]

\[
= E[Y_{0i}^*|X_i^* \in S] - E[Y_{0i}^*|X_i^* \in S, X_i^* \notin B(x^*, \epsilon)]\Pr(X_i^* \notin B(x^*, \epsilon)|X_i^* \in S) - E[Y_{0i}^* - 1|X_i^* \in B(x^*, \epsilon)]\Pr(X_i^* \in B(x^*, \epsilon)|X_i^* \in S)
\]

\[
= E[Y_{i1}^* - Y_{0i}^*|X_i^* \in S] + \Pr(X_i^* \in B(x^*, \epsilon)|X_i^* \in S).
\]

By the definition of support, \( \Pr(X_i^* \in B(x^*, \epsilon)) > 0 \). Since \( T(Q) = T(Q_0) = P_0 \) but \( E_{Q}[Y_{i1}^Q - Y_{0i}^Q|X_i^Q \in S] \neq E[Y_{i1}^* - Y_{0i}^*|X_i^* \in S] \), \( E[Y_{i1} - Y_{0i}|X_i \in S] \) is not identified.

\[\square\]

C.2 Proof of Corollary 1

If \( \Pr(D_i(1) - D_i(0)) = 1|X_i = x = 1 \), \( \Pr(Y_{i1} - Y_{0i} = Y_i(1) - Y_i(0)|X_i = x = 1 \), and hence \( E[Y_{i1} - Y_{0i}|X_i = x] = E[Y_i(1) - Y_i(0)|X_i = x] \). Then, Part (a) follows from Proposition 1 (a). If \( \Pr(D_i(1) \geq D_i(0)|X_i = x = 1 \), we have

\[
E[Y_{i1} - Y_{0i}|X_i = x] = E[(D_i(1) - D_i(0))(Y_i(1) - Y_i(0))|X_i = x]
\]

\[
= \Pr(D_i(1) \neq D_i(0)|X_i = x)E[Y_i(1) - Y_i(0)|D_i(1) \neq D_i(0), X_i = x].
\]
If in addition \( \Pr(D_i(1) \neq D_i(0)|X_i = x) > 0 \), we obtain
\[
E[Y_i(1) - Y_i(0)|D_i(1) \neq D_i(0), X_i = x] = \frac{E[Y_{i1} - Y_{0i}|X_i = x]}{\Pr(D_i(1) \neq D_i(0)|X_i = x)} \frac{E[Y_{i1} - Y_{0i}|X_i = x]}{E[D_i(1) - D_i(0)|X_i = x]}.
\]

Then, Part (b) follows from Proposition 1 (a).

C.3 Proof of Theorem 1

We prove consistency and asymptotic normality of the following estimators without imposing Assumption 3 (c). These estimators are asymptotically equivalent to the estimators defined in Section 4.1 if Assumption 3 (c) holds.

First, consider the following 2SLS regression using the observations with \( p^A(X_i; \delta_n) \in (0, 1) \):
\[
D_i = \gamma_0 (1 - I_n) + \gamma_1 Z_i + \gamma_2 p^A(X_i; \delta_n) + \nu_i \tag{15}
\]
\[
Y_i = \beta_0 (1 - I_n) + \beta_1 D_i + \beta_2 p^A(X_i; \delta_n) + \epsilon_i. \tag{16}
\]

Here \( I_n \) is a dummy random variable which equals one if there exists a constant \( q \in (0, 1) \) such that \( A(X_i) \in \{0, q, 1\} \) for all \( i \in \{1, \ldots, n\} \). \( I_n \) is the indicator that \( A(X_i) \) takes on only one nondegenerate value in the sample. If the support of \( A(X_i) \) (in the population) contains only one value in \( (0, 1) \), \( p^A(X_i; \delta_n) \) is asymptotically constant conditional on \( p^A(X_i; \delta_n) \in (0, 1) \). To avoid the multicollinearity between asymptotically constant \( p^A(X_i; \delta_n) \) and a constant, we do not include the constant term if \( I_n = 1 \). Let \( I_{i,n} = 1 \{p^A(X_i; \delta_n) \in (0, 1)\} \), \( D_{i,n} = (1, D_i, p^A(X_i; \delta_n))' \), \( Z_{i,n} = (1, Z_i, p^A(X_i; \delta_n))' \), \( D_{i,n}^{nc} = (D_i, p^A(X_i; \delta_n))' \), and \( Z_{i,n}^{nc} = (Z_i, p^A(X_i; \delta_n))' \). The 2SLS estimator \( \hat{\beta} \) from this regression is then given by
\[
\hat{\beta} = \begin{cases} 
(\sum_{i=1}^n Z_{i,n} D_{i,n}' I_{i,n})^{-1} \sum_{i=1}^n Z_{i,n} Y_{i,n} I_{i,n} & \text{if } I_n = 0 \\
(\sum_{i=1}^n Z_{i,n}^{nc} D_{i,n}^{nc}' I_{i,n})^{-1} \sum_{i=1}^n Z_{i,n}^{nc} Y_{i,n} I_{i,n} & \text{if } I_n = 1.
\end{cases}
\]

Let \( \hat{\beta}_1 \) denote the 2SLS estimator of \( \beta_1 \) in the above regression.

Similarly, consider the following simulation version of the 2SLS regression using the observations with \( p^s(X_i; \delta_n) \in (0, 1) \):
\[
D_i = \gamma_0 (1 - I_n) + \gamma_1 Z_i + \gamma_2 p^s(X_i; \delta_n) + \nu_i \tag{17}
\]
\[
Y_i = \beta_0 (1 - I_n) + \beta_1 D_i + \beta_2 p^s(X_i; \delta_n) + \epsilon_i. \tag{18}
\]

Let \( \hat{\beta}_1^s \) denote the 2SLS estimator of \( \beta_1 \) in the simulation-based regression.

Below, we prove the following result.

**Theorem C.1.** Suppose that Assumptions 1 and 3 hold except Assumption 3 (c), and that \( \delta_n \to 0, n\delta_n \to \infty \) and \( S_n \to \infty \) as \( n \to \infty \). Then the 2SLS estimators \( \hat{\beta}_1 \) and \( \hat{\beta}_1^s \) converge in probability to
\[
\beta_1 \equiv \lim_{\delta \to 0} E[\omega(\delta)(Y_i(1) - Y_i(0))],
\]

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where
\[ \omega_i(\delta) = \frac{p^A(X_i; \delta)(1 - p^A(X_i; \delta))(D_i(1) - D_i(0))}{E[p^A(X_i; \delta)(1 - p^A(X_i; \delta))(D_i(1) - D_i(0))]}. \]

Suppose, in addition, that Assumptions 4 and 5 hold and that \( n\delta^2_n \to 0 \) as \( n \to \infty \). Then
\[ \hat{s}_{1n}^{-1}(\hat{\beta}_1 - \beta_1) \xrightarrow{d} \mathcal{N}(0, 1), \]
\[ (\hat{\sigma}_{n}^n)^{-1}(\hat{\sigma}_n^n - \beta_1) \xrightarrow{d} \mathcal{N}(0, 1). \]

where we define \( \hat{s}_{1n}^{-1} \) and \( (\hat{\sigma}_n^n)^{-1} \) as follows: let
\[ \hat{\Sigma}_n = \begin{cases} \sum_{i=1}^n \hat{\epsilon}_{i,n}^2 & \text{if } I_n = 0 \\ (\sum_{i=1}^n \hat{\epsilon}_{i,n} \hat{\epsilon}_{i,n}' \hat{\Sigma}_n) - \sum_{i=1}^n \hat{\epsilon}_{i,n} \hat{\epsilon}_{i,n}' & \text{if } I_n = 1, \end{cases} \]
where
\[ \hat{\epsilon}_{i,n} = \begin{cases} Y_i - \hat{D}_{i,n} \hat{\beta} & \text{if } I_n = 0 \\ Y_i - \hat{D}_{i,n} \hat{\beta} & \text{if } I_n = 1. \end{cases} \]

Let \( \hat{\sigma}_n^n \) denote the estimator for the variance of \( \hat{\beta}_1 \). That is, \( \hat{\sigma}_n^n \) is the second diagonal element of \( \hat{\Sigma}_n \) when \( I_n = 0 \) and is the first diagonal element of \( \hat{\Sigma}_n \) when \( I_n = 1 \). \( (\hat{\sigma}_n^n)^2 \) is the analogously-defined estimator for the variance of \( \hat{\beta}_1^n \) from the simulation-based regression.

Throughout the proof, we omit the subscript \( n \) from \( I_{i,n}, D_{i,n}, Z_{i,n}, \hat{\epsilon}_{i,n}, \hat{\Sigma}_n, \hat{\beta}_n, \) etc. for notational brevity. We provide proofs separately for the two cases, the case in which \( \Pr(A(X_i) \in (0, 1)) > 0 \) and the case in which \( \Pr(A(X_i) \in (0, 1)) = 0 \). For each case, we first prove consistency and asymptotic normality of \( \hat{\beta}_1 \), and then prove consistency and asymptotic normality of \( \hat{\beta}_1^n \).

**C.3.1 Consistency and Asymptotic Normality of \( \hat{\beta}_1 \) When \( \Pr(A(X_i) \in (0, 1)) > 0 \)**

By Lemma B.6,
\[ \lim_{\delta \to 0} E[p^A(X_i; \delta)(1 - p^A(X_i; \delta))(D_i(1) - D_i(0))] = E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))]. \]

When \( \Pr(A(X_i) \in (0, 1)) > 0 \), \( E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))] = E[p^A(X_i)(1 - p^A(X_i))(D_i(1) - D_i(0))] \), since \( p^A(x) = A(x) \) for almost every \( x \in \mathcal{X} \) by Proposition A.1. Under Assumption 3 (b), \( E[p^A(X_i)(1 - p^A(X_i))(D_i(1) - D_i(0))] > 0 \). Again by Lemma B.6,
\[ \lim_{\delta \to 0} E[\omega_i(\delta)(Y_i(1) - Y_i(0))] = \frac{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))Y_i(1) - Y_i(0)]}{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))]]. \]
Let \( \beta_1 = \frac{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))Y_i(1) - Y_i(0)]}{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))]} \). Let
\[ \hat{\beta}_c = \sum_{i=1}^n \hat{Z}_i \hat{D}_i' \hat{I}_i \]
\[ \hat{\beta}_nc = (\sum_{i=1}^n \hat{Z}_i^{nc} (\hat{D}_i^{nc})' \hat{I}_i \sum_{i=1}^n \hat{Z}_i^{nc} \hat{Y}_i \hat{I}_i, \]

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and let $\hat{\beta}_1 = (0,1,0)\hat{\beta}^c$ and $\hat{\beta}^{nc}_1 = (1,0)\hat{\beta}^{nc}$. $\hat{\beta}_1$ is given by

$$
\hat{\beta}_1 = \hat{\beta}^c_1(1 - I_n) + \hat{\beta}^{nc}_1 I_n.
$$

Also, let $\tilde{D}_i = (1, D_i, A(X_i))'$, $\tilde{Z}_i = (1, Z_i, A(X_i))'$, $\tilde{D}_i^{nc} = (D_i, A(X_i))'$, $\tilde{Z}_i^{nc} = (Z_i, A(X_i))'$, and $I^A_i = 1\{A(X_i) \in (0,1)\}$.

We claim that $\Pr(I_n = 1) \to 0$ when $\text{Var}(A(X_i)|I^A_i = 1) > 0$, and that $\Pr(I_n = 1) \to 1$ when $\text{Var}(A(X_i)|I^A_i = 1) = 0$. To show the first claim, observe that $I_n = 1$ if and only if $\hat{V}_n = 0$, where

$$
\hat{V}_n = \frac{\sum_{i=1}^n (A(X_i) - \frac{\sum_{i=1}^n A(X_i)I^A_i}{\sum_{i=1}^n I^A_i})^2 I^A_i}{\sum_{i=1}^n I^A_i}
$$

is the sample variance of $A(X_i)$ conditional on $I^A_i = 1$. When $\text{Var}(A(X_i)|I^A_i = 1) > 0$,

$$
\Pr(I_n = 1) = \Pr(\hat{V}_n = 0) \\
\leq \Pr(|\hat{V}_n - \text{Var}(A(X_i)|I^A_i = 1)| \geq \text{Var}(A(X_i)|I^A_i = 1)) \\
\to 0,
$$

where the convergence follows since $\hat{V}_n \xrightarrow{p} \text{Var}(A(X_i)|I^A_i = 1) > 0$.

To show the second claim, note that, when $\text{Var}(A(X_i)|I^A_i = 1) = 0$, there exists $q \in (0,1)$ such that $\Pr(A(X_i) = q|I^A_i = 1) = 1$. It follows that

$$
\Pr(I_n = 0) = \Pr(A(X_i) \in \{0,1\} \text{ for all } i = 1, \ldots, n) \\
\quad + \Pr(A(X_i) = q' \text{ and } A(X_j) = q'' \text{ for some } q', q'' \in (0,1) \text{ with } q' \neq q'' \text{ for some } i, j \in \{1, \ldots, n\}) \\
\quad = \Pr(A(X_i) \in \{0,1\} \text{ for all } i = 1, \ldots, n) \\
\quad = (1 - \Pr(A(X_i) \in (0,1)))^n,
$$

which converges to zero as $n \to \infty$, since $\Pr(A(X_i) \in (0,1)) > 0$.

The above claims imply that $\hat{\beta}_1 = \hat{\beta}^c_1$ with probability approaching one when $\text{Var}(A(X_i)|I^A_i = 1) > 0$, and that $\hat{\beta}_1 = \hat{\beta}^{nc}_1$ with probability approaching one when $\text{Var}(A(X_i)|I^A_i = 1) = 0$. Therefore, to prove consistency and asymptotic normality of $\hat{\beta}_1$, it suffices to show those of $\hat{\beta}^c_1$ when $\text{Var}(A(X_i)|I^A_i = 1) > 0$ and those of $\hat{\beta}^{nc}_1$ when $\text{Var}(A(X_i)|I^A_i = 1) = 0$.

Below we first show that, if Assumptions 1 and 3 hold and $\delta_n \to 0$ as $n \to \infty$, then $\hat{\beta}_1 \xrightarrow{p} \beta_1$. We then show that, if, in addition, Assumption 4 holds and $n\delta^2_n \to 0$ as $n \to \infty$, then $\hat{\sigma}^{-1}(\hat{\theta}_1 - \theta_1) \xrightarrow{d} N(0,1)$.

**Proof of Consistency.** To prove consistency of $\hat{\beta}_1$, we first show that $\hat{\beta}^c_1 \xrightarrow{p} \beta_1$ when $\text{Var}(A(X_i)|I^A_i = 1) > 0$. We then show that $\hat{\beta}^{nc}_1 \xrightarrow{p} \beta_1$ whether or not $\text{Var}(A(X_i)|I^A_i = 1) > 0$. By Lemma B.6,

$$
\hat{\beta}^c = (\sum_{i=1}^n Z_i D_i' I_i)^{-1} \sum_{i=1}^n Z_i Y_i I_i \xrightarrow{p} (E[\tilde{Z}_i \tilde{D}_i(I^A_i)]^{-1} E[\tilde{Z}_i Y_i I^A_i])
$$
provided that $E[	ilde{Z}_i\tilde{D}_i'I_i^A]$ is invertible. After a few lines of algebra, we have
\[
\det(E[\tilde{Z}_i\tilde{D}_i'I_i^A]) = \Pr(I_i^A = 1)^2 \Var(A(X_i)|I_i^A = 1) E[D_i(Z_i - A(X_i))I_i^A]
\]
\[
= \Pr(I_i^A = 1)^2 \Var(A(X_i)|I_i^A = 1) E[(Z_iD_i(1) + (1 - Z_i)D_i(0))(Z_i - A(X_i))I_i^A]
\]
\[
= \Pr(I_i^A = 1)^2 \Var(A(X_i)|I_i^A = 1) E[((Z_i - Z_iA(X_i))D_i(1) - (1 - Z_i)A(X_i)D_i(0))I_i^A]
\]
\[
= \Pr(I_i^A = 1)^2 \Var(A(X_i)|I_i^A = 1) E[(A(X_i) - A(X_i)^2)D_i(1) - (1 - A(X_i))A(X_i)D_i(0))I_i^A]
\]
\[
= \Pr(I_i^A = 1)^2 \Var(A(X_i)|I_i^A = 1) E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))I_i^A]
\]
\[
= \Pr(I_i^A = 1)^2 \Var(A(X_i)|I_i^A = 1) E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))],
\]
where the fourth equality follows from Property 1. Therefore, $E[\tilde{Z}_i\tilde{D}_i'I_i^A]$ is invertible when $\Var(A(X_i)|I_i^A = 1) > 0$. Another few lines of algebra gives
\[
(E[\tilde{Z}_i\tilde{D}_i'I_i^A])^{-1} = \frac{1}{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))]} \begin{bmatrix}
* & * & *
0 & 1 & -1
* & * & *
\end{bmatrix}
\]
when $\Var(A(X_i)|I_i^A = 1) > 0$. Therefore, when $\Var(A(X_i)|I_i^A = 1) > 0$,
\[
\hat{\beta}_i \xrightarrow{p} \frac{E[Z_iY_iI_i^A] - E[A(X_i)Y_iI_i^A]}{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))]} = \frac{E[Z_iY_iI_i^A] - E[A(X_i)(Z_iY_i + (1 - Z_i)Y_0)I_i^A]}{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))]} = \frac{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))]}{E[A(X_i)(1 - A(X_i))]} = \beta_1,
\]
where the third line follows from Property 1, and the second last follows from the definitions of $Y_i$ and $Y_0$.

We next consider $\hat{\beta}_i^{nc}$. By Lemma B.6,
\[
\hat{\beta}_i^{nc} = \left( \sum_{i=1}^{n} Z_i^{nc}(D_i^{nc})'I_i \right)^{-1} \sum_{i=1}^{n} Z_i^{nc}Y_i I_i \xrightarrow{p} (E[Z_i^{nc}(D_i^{nc})'I_i^A])^{-1} E[Z_i^{nc}Y_i I_i^A]
\]
provided that $E[Z_i^{nc}(D_i^{nc})'I_i^A]$ is invertible. After a few lines of algebra, we have
\[
\det(E[Z_i^{nc}(D_i^{nc})'I_i^A]) = E[A(X_i)^2I_i^A]E[D_i(Z_i - A(X_i))I_i^A]
\]
\[
= E[A(X_i)^2I_i^A]E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))]
\]
\[
> 0.
\]
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By the law of total conditional variance,
\[
(E[\tilde{Z}_i^{nc}(\tilde{D}_i^{nc})'|I^A])^{-1} = \frac{1}{E[A(X_i)(1 - A(x_i))(D_i(1) - D_i(0))]} \begin{pmatrix} 1 & -1 \\ * & * \end{pmatrix}.
\]
Therefore,
\[
\hat{\beta}_i^{nc} \overset{p}{\rightarrow} \frac{E[Z_i Y_i I^A] - E[A(X_i)Y_i I^A]}{E[A(X_i)(1 - A(x_i))(D_i(1) - D_i(0))]} = \beta_1.
\]

**Proof of Asymptotic Normality.** Let \( (\hat{\sigma}^c)^2 \) be the second diagonal element of
\[
\Sigma^c = (\sum_{i=1}^n Z_i D'_i I_i)^{-1} (\sum_{i=1}^n \epsilon_i^2 Z_i Z'_i I_i)(\sum_{i=1}^n D_i Z'_i I_i)^{-1}
\]
and \( (\hat{\sigma}^{nc})^2 \) be the first diagonal element of
\[
\Sigma^{nc} = (\sum_{i=1}^n Z_i^{nc} D'_i^{nc} I_i)^{-1} (\sum_{i=1}^n \epsilon_i^2 Z_i^{nc} Z'_i^{nc} I_i)(\sum_{i=1}^n D_i^{nc} (Z_i^{nc})' I_i)^{-1}.
\]
We only show that \((\hat{\sigma}^c)^{-1}(\hat{\beta}^c - \beta_1) \overset{d}{\rightarrow} \mathcal{N}(0, 1)\) when \(\text{Var}(A(X_i)|I^A_i) = 1 > 0\). We can show that \((\hat{\sigma}^{nc})^{-1}(\hat{\beta}^{nc} - \beta_1) \overset{d}{\rightarrow} \mathcal{N}(0, 1)\) by an analogous argument. The proof proceeds in six steps.

**Step C.3.1.1.** Let \( \hat{\beta}_n = (E[\tilde{Z}_i D'_i I_i])^{-1} E[\tilde{Z}_i Y_i I_i] \), and let \( \hat{\beta}_{1,n} \) denote the second element of \( \hat{\beta}_n \). Then \( \hat{\beta}_{1,n} = \beta_1 \) for any choice of \( \delta_n > 0 \).

**Proof.** Note first that, for every \( \delta > 0 \), \( p^A(x; \delta) \in (0, 1) \) for almost every \( x \in \{x' \in \mathcal{X} : A(x') \in (0, 1)\} \), since by almost everywhere continuity of \( A \), for almost every \( x \in \{x' \in \mathcal{X} : A(x') \in (0, 1)\} \), there exists an open ball \( B \subset B(x, \delta) \) such that \( A(x') \in (0, 1) \) for every \( x' \in B \). After a few lines of algebra, we have
\[
\det(E[\tilde{Z}_i D'_i I_i]) = \text{Pr}(I_i = 1)^2 \text{Var}(A(X_i)|I_i = 1) E[D_i(Z_i - A(X_i)) I_i] = \text{Pr}(I_i = 1)^2 \text{Var}(A(X_i)|I_i = 1) E[A(X_i)(1 - A(x_i))(D_i(1) - D_i(0)) I_i] = \text{Pr}(I_i = 1)^2 \text{Var}(A(X_i)|I_i = 1) E[A(X_i)(1 - A(x_i))(D_i(1) - D_i(0))],
\]
where the last equality holds since \( p^A(x; \delta) \in (0, 1) \) for almost every \( x \in \{x' \in \mathcal{X} : A(x') \in (0, 1)\} \).

By the law of total conditional variance,
\[
\text{Var}(A(X_i)|I_i = 1) \geq E[\text{Var}(A(X_i)|I_i = 1, I^A)|I_i = 1] + \text{Var}(E[A(X_i)|I_i = 1, I^A]|I_i = 1) \
\geq \sum_{t \in \{0, 1\}} \text{Var}(A(X_i)|I_i = 1, I^A = t) \text{Pr}(I^A = t|I_i = 1) \
\geq \text{Var}(A(X_i)|I_i = 1, I^A = 1) \text{Pr}(I^A = 1|I_i = 1) = \text{Var}(A(X_i)|I^A = 1) \text{Pr}(I^A = 1|I_i = 1) > 0.
\]
Therefore, \( E[\hat{Z}_i \hat{D}'_i I_i] \) is invertible. Another few lines of algebra gives
\[
(E[\hat{Z}_i \hat{D}'_i I_i])^{-1} = \frac{1}{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))]} \begin{bmatrix}
* & * & * \\
0 & 1 & -1 \\
* & * & *
\end{bmatrix}.
\]
It follows that
\[
\tilde{\beta}_{1,n} = \frac{E[Z_i Y_i I_i] - E[A(X_i) Y_i I_i]}{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))]} = \frac{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))(Y_i(1) - Y_i(0))I_i]}{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))]} = \frac{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))(Y_i(1) - Y_i(0))]}{E[A(X_i)(1 - A(X_i))(D_i(1) - D_i(0))]} = \beta_1.
\]

We can write
\[
\sqrt{n}(\hat{\beta} - \beta_n) = \left( \frac{1}{n} \sum_{i=1}^{n} \frac{Z_i D'_i I_i}{\sqrt{n}} \right)^{-1} \frac{1}{\sqrt{n}} \sum_{i=1}^{n} Z_i Y_i I_i - \left( \frac{1}{n} \sum_{i=1}^{n} \frac{Z_i D'_i I_i}{\sqrt{n}} \right)^{-1} \frac{1}{\sqrt{n}} \sum_{i=1}^{n} \tilde{Z}_i Y_i I_i = (A)
\]
\[
+ \left( \frac{1}{n} \sum_{i=1}^{n} \frac{Z_i D'_i I_i}{\sqrt{n}} \right)^{-1} \frac{1}{\sqrt{n}} \sum_{i=1}^{n} \tilde{Z}_i Y_i I_i - (E[\hat{Z}_i \hat{D}'_i I_i])^{-1} \sqrt{n} E[\hat{Z}_i Y_i I_i] = (B).
\]

We first consider (B). Let \( \tilde{\epsilon}_{i,n} = Y_i - \hat{D}'_i \beta_n \) so that
\[
E[\hat{Z}_i \tilde{\epsilon}_{i,n} I_i] = E[\hat{Z}_i (Y_i - \hat{D}'_i \beta_n) I_i] = E[\hat{Z}_i Y_i I_i] - E[\hat{Z}_i \hat{D}'_i I_i] \beta_n = 0.
\]

Then
\[
(B) = \left( \frac{1}{n} \sum_{i=1}^{n} \tilde{Z}_i D'_i I_i \right)^{-1} \frac{1}{\sqrt{n}} \sum_{i=1}^{n} \tilde{Z}_i (D'_i \beta_n + \tilde{\epsilon}_{i,n}) I_i - (E[\hat{Z}_i \hat{D}'_i I_i])^{-1} \sqrt{n} E[\hat{Z}_i (D'_i \beta_n + \tilde{\epsilon}_{i,n}) I_i]
\]
\[
= \sqrt{n}(\hat{\beta}_n - \beta_n) + \left( \frac{1}{n} \sum_{i=1}^{n} \tilde{Z}_i D'_i I_i \right)^{-1} \frac{1}{\sqrt{n}} \sum_{i=1}^{n} \tilde{Z}_i \tilde{\epsilon}_{i,n} I_i - (E[\hat{Z}_i \hat{D}'_i I_i])^{-1} \sqrt{n} E[\hat{Z}_i \tilde{\epsilon}_{i,n} I_i]
\]
\[
= \left( \frac{1}{n} \sum_{i=1}^{n} \tilde{Z}_i D'_i I_i \right)^{-1} \frac{1}{\sqrt{n}} \sum_{i=1}^{n} \tilde{Z}_i \tilde{\epsilon}_{i,n} I_i.
\]

Step C.3.1.2.
\[
\frac{1}{\sqrt{n}} \sum_{i=1}^{n} \tilde{Z}_i \tilde{\epsilon}_{i,n} I_i \overset{d}{\rightarrow} N(0, E[\tilde{\epsilon}_{i,n}^2 \tilde{Z}_i \tilde{D}'_i I_i])
\]
Proof. We use the triangular-array Lyapunov CLT and the Cramér-Wold device. Pick a nonzero \( \lambda \in \mathbb{R}^p \), and let \( V_{i,n} = \frac{1}{\sqrt{n}} \lambda' \tilde{Z}_i \tilde{e}_{i,n} I_i \). First, we have

\[
\sum_{i=1}^{n} E[V_{i,n}^2] = \lambda' E[\tilde{e}_{i,n}^2 \tilde{Z}_i \tilde{Z}_i' I_i] \lambda.
\]

By Lemma B.6,

\[
\tilde{\beta}_n \to (E[\tilde{Z}_i \tilde{D}_i' I_i^A])^{-1} E[\tilde{Z}_i Y_i I_i^A]
\]

as \( n \to \infty \). Let \( \beta = (E[\tilde{Z}_i \tilde{D}_i' I_i^A])^{-1} E[\tilde{Z}_i Y_i I_i^A] \) and \( \tilde{c}_i = Y_i - \tilde{D}_i' \beta \). We have

\[
E[\tilde{c}_{i,n}^2 \tilde{Z}_i \tilde{Z}_i' I_i] = E[(Y_i - \tilde{D}_i' \tilde{\beta}_n)^2 \tilde{Z}_i \tilde{Z}_i' I_i]
\]

\[
= E[(\tilde{c}_i - \tilde{D}_i' \tilde{\beta}_n - \beta_i)^2 \tilde{Z}_i \tilde{Z}_i' I_i]
\]

\[
= E[\tilde{c}_i^2 \tilde{Z}_i \tilde{Z}_i' I_i] - 2E[\tilde{c}_i((\tilde{\beta}_{0,n} - \beta_0) + D_i(\tilde{\beta}_{1,n} - \beta_1) + A(X_i)(\tilde{\beta}_{2,n} - \beta_2)) \tilde{Z}_i \tilde{Z}_i' I_i]
\]

\[
+ E[((\tilde{\beta}_{0,n} - \beta_0) + D_i(\tilde{\beta}_{1,n} - \beta_1) + A(X_i)(\tilde{\beta}_{2,n} - \beta_2))^2 \tilde{Z}_i \tilde{Z}_i' I_i]
\]

\[
\to E[\tilde{c}_i^2 \tilde{Z}_i \tilde{Z}_i' I_i^A]
\]

as \( n \to \infty \), where the convergence follows from Lemma B.6 and from the fact that \( \tilde{\beta}_n \to \beta \). Therefore,

\[
\sum_{i=1}^{n} E[V_{i,n}^2] \to \lambda' E[\tilde{c}_i^2 \tilde{Z}_i \tilde{Z}_i' I_i^A] \lambda.
\]

We next verify the Lyapunov condition: for some \( t > 0 \),

\[
\sum_{i=1}^{n} E[|V_{i,n}|^{2+t}] \to 0.
\]

We have

\[
\sum_{i=1}^{n} E[|V_{i,n}|^4] = \frac{1}{n} E[|\lambda' \tilde{Z}_i \tilde{e}_{i,n} I_i|^4].
\]

We use the \( c_r \)-inequality: \( E[|X + Y|^r] \leq 2^{r-1} E[|X|^r + |Y|^r] \) for \( r \geq 1 \). Repeating using the \( c_r \)-inequality gives

\[
E[|\lambda' \tilde{Z}_i \tilde{e}_{i,n} I_i|^4] = E[|\lambda' \tilde{Z}_i (Y_i - \tilde{\beta}_{0,n} - \tilde{\beta}_{1,n} D_i - \tilde{\beta}_{2,n} A(X_i))|^4 I_i]
\]

\[
\leq 2^{3c} E[(|\lambda' \tilde{Z}_i|^4)(|Y_i|^4 + |\tilde{\beta}_{0,n}|^4 + |\tilde{\beta}_{1,n}|^4 D_i + |\tilde{\beta}_{2,n}|^4 A(X_i)^4 I_i]
\]

\[
\leq 2^{3c}(\lambda_1 + \lambda_2 + \lambda_3)^4 (E[Y_i^4] + \tilde{\beta}_{0,n}^4 + \tilde{\beta}_{1,n}^4 + \tilde{\beta}_{2,n}^4)
\]

for some finite constant \( c \), and the right-hand side converges to

\[
2^{3c}(\lambda_1 + \lambda_2 + \lambda_3)^4 (E[Y_i^4] + \tilde{\beta}_{0,n}^4 + \tilde{\beta}_{1,n}^4 + \tilde{\beta}_{2,n}^4),
\]

which is finite under Assumption 3 (a). Therefore,

\[
\sum_{i=1}^{n} E[|V_{i,n}|^4] \to 0,
\]

and the conclusion follows from the Lyapunov CLT and the Cramér-Wold device. \( \square \)
We next consider (A). We can write

\[ (A) = \left( \frac{1}{n} \sum_{i=1}^{n} Z_i D'_i I_i \right)^{-1} \frac{1}{\sqrt{n}} \sum_{i=1}^{n} (Z_i Y_i I_i - \tilde{Z}_i Y_i I_i) \]

\[-\left( \frac{1}{n} \sum_{i=1}^{n} Z_i D'_i I_i \right)^{-1} \left[ \frac{1}{\sqrt{n}} \sum_{i=1}^{n} (Z_i D'_i I_i - \tilde{Z}_i D'_i I_i) \right] \left( \frac{1}{n} \sum_{i=1}^{n} \tilde{Z}_i D'_i I_i \right)^{-1} \frac{1}{n} \sum_{i=1}^{n} \tilde{Z}_i Y_i I_i. \]

**Step C.3.1.3.** Let \( \{V_i\}_{i=1}^{\infty} \) be i.i.d. random variables such that \( E[|V_i|] < \infty \) and that \( E[V_i|X_i] \) is bounded on \( N(D^*, \delta') \cap X \) for some \( \delta' > 0 \). Then,

\[ E[V_i p^A(X_i; \delta) - A(X_i)]1\{p^A(X_i; \delta) \in (0, 1)\}] = O(\delta) \]

for \( l = 0, 1 \).

**Proof.** For every \( x \notin N(D^*, \delta) \), \( B(x, \delta) \cap D^* = \emptyset \), so \( A \) is continuously differentiable on \( B(x, \delta) \).

By the mean value theorem, for every \( x \notin N(D^*, \delta) \) and \( a \in B(0, \delta) \),

\[ A(x + a) = A(x) + \nabla A(y(x, a))'a \]

for some point \( y(x, a) \) on the line segment connecting \( x \) and \( x + a \). For every \( x \notin N(D^*, \delta) \),

\[ p^A(x; \delta) = \frac{\int_{B(0, 1)} A(x + \delta u) du}{\int_{B(0, 1)} du} \]

\[ = \frac{\int_{B(0, 1)} (A(x) + \delta \nabla A(y(x, \delta u))'u) du}{\int_{B(0, 1)} du} \]

\[ = A(x) + \delta \frac{\int_{B(0, 1)} \nabla A(y(x, \delta u))'u du}{\int_{B(0, 1)} du}. \]

Now, we can write

\[ E[V_ip^A(X_i; \delta) - A(X_i)]1\{p^A(X_i; \delta) \in (0, 1)\}] \]

\[ = E[V_ip^A(X_i; \delta) - A(X_i)]1\{p^A(X_i; \delta) \in (0, 1)\}1\{X_i \notin N(D^*, \delta)\}] \]

\[ + E[V_ip^A(X_i; \delta) - A(X_i)]1\{p^A(X_i; \delta) \in (0, 1)\}1\{X_i \in N(D^*, \delta)\}]. \]

For the first term,

\[ |E[V_ip^A(X_i; \delta) - A(X_i)]1\{p^A(X_i; \delta) \in (0, 1)\}1\{X_i \notin N(D^*, \delta)\}]| \]

\[ = \delta |E[V_ip^A(X_i; \delta) - A(X_i)]1\{p^A(X_i; \delta) \in (0, 1)\}1\{X_i \notin N(D^*, \delta)\}]| \]

\[ \leq \delta |E[V_i p^A(X_i; \delta)] \frac{\int_{B(0, 1)} \nabla A(y(X_i(\delta u), \delta u))'u du}{\int_{B(0, 1)} du} 1\{p^A(X_i; \delta) \in (0, 1)\}1\{X_i \notin N(D^*, \delta)\}]| \]

\[ \leq \delta \sum_{k=1}^{p} \sup_{x \in \mathbb{C}^*} |\frac{\partial A(x)}{\partial x_k}| \frac{\int_{B(0, 1)} |u_k| du}{\int_{B(0, 1)} du} \]

\[ = O(\delta), \]

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where we use the assumption that the partial derivatives of $A$ is bounded on $C^*$. For the second term, for sufficiently small $\delta > 0$,

$$|E[V_i p^A(X_i; \delta) (p^A(X_i; \delta) - A(X_i)) \mathbb{1}\{p^A(X_i; \delta) \in (0, 1)\} \mathbb{1}\{X_i \in N(D, \delta)\}]|$$

$$\leq E[|E[V_i|X_i]\{X_i \in N(D, \delta)\}]$$

$$\leq CE[\mathbb{1}\{X_i \in N(D, \delta)\}]$$

$$= C \Pr(X_i \in N(D, \delta))$$

$$= O(\delta),$$

where $C$ is some constant, the second inequality follows from the assumption that $E[V_i|X_i]$ is bounded on $N(D, \delta') \cap \mathcal{X}$ for some $\delta' > 0$, and the last equality follows from Assumption 4 (a).

**Step C.3.1.4.** $\frac{1}{\sqrt{n}} \sum_{i=1}^{n} (\tilde{Z}_i Y_i I_i - \tilde{Z}_i Y_i I_i) = o_p(1)$ and $\frac{1}{\sqrt{n}} \sum_{i=1}^{n} (\tilde{Z}_i D_i I_i - \tilde{Z}_i \tilde{D}_i I_i) = o_p(1)$.

**Proof.** We only show that $\frac{1}{\sqrt{n}} \sum_{i=1}^{n} (p^A(X_i; \delta_n) - A(X_i)^2) I_i = o_p(1)$. The proofs for the other elements are similar. As for bias,

$$E[\frac{1}{\sqrt{n}} \sum_{i=1}^{n} (p^A(X_i; \delta_n) - A(X_i)^2) I_i]$$

$$= \sqrt{n} E[(p^A(X_i; \delta_n) - A(X_i)^2) I_i]$$

$$= \sqrt{n} E[(p^A(X_i; \delta_n) + A(X_i))(p^A(X_i; \delta_n) - A(X_i)) I_i]$$

$$= \sqrt{n} O(\delta_n)$$

$$= 0,$$

where the third equality follows from Step C.3.1.3 and the last from the assumption that $n \delta_n^2 \rightarrow 0$.

As for variance, by Lemma B.6,

$$\text{Var}\left(\frac{1}{\sqrt{n}} \sum_{i=1}^{n} (p^A(X_i; \delta_n) - A(X_i)^2) I_i\right)$$

$$\leq E[(p^A(X_i; \delta_n)^2 - A(X_i)^2)^2 I_i]$$

$$= E[(p^A(X_i; \delta_n)^4 - 2p^A(X_i; \delta_n)^2 A(X_i)^2 + A(X_i)^4) I_i]$$

$$\rightarrow E[(A(X_i)^4 - 2A(X_i)^2 A(X_i)^2 + A(X_i)^4) I_i^4]$$

$$= 0.$$

**Step C.3.1.5.** $n \Sigma_c \xrightarrow{p} (E[Z_i D_i^A])^{-1} E[\epsilon^2 Z_i Z_i^A](E[D_i Z_i^A])^{-1}.$
Proof. Let $\epsilon_i = Y_i - D_i'\beta$. We have
\[
\frac{1}{n} \sum_{i=1}^{n} \epsilon_i^2 Z_i Z_i' I_i = \frac{1}{n} \sum_{i=1}^{n} (Y_i - D_i' \beta_c^c)^2 Z_i Z_i' I_i
\]
\[
= \frac{1}{n} \sum_{i=1}^{n} (\epsilon_i - D_i' (\beta_c^c - \beta))^2 Z_i Z_i' I_i
\]
\[
= \frac{1}{n} \sum_{i=1}^{n} \epsilon_i^2 Z_i Z_i' I_i
\]
\[- \frac{2}{n} \sum_{i=1}^{n} (Y_i - D_i' \beta)((\beta_0^c - \beta_0) + D_i (\beta_0^c - \beta_1) + p^A (X_i; \delta_n) (\beta_2^c - \beta_2)) Z_i Z_i' I_i
\]
\[
+ \frac{1}{n} \sum_{i=1}^{n} ((\beta_0^c - \beta_0) + D_i (\beta_1^c - \beta_1) + p^A (X_i; \delta_n) (\beta_2^c - \beta_2))^2 Z_i Z_i' I_i
\]
\[
= \frac{1}{n} \sum_{i=1}^{n} \epsilon_i^2 Z_i Z_i' I_i + o_p(1) O_p(1),
\]
where the last equality follows from the result that $\hat{\beta}_c^c - \beta = o_p(1)$ and from Lemma B.6. Again by Lemma B.6,
\[
\frac{1}{n} \sum_{i=1}^{n} \epsilon_i^2 Z_i Z_i' I_i = \frac{1}{n} \sum_{i=1}^{n} (Y_i^2 - 2 Y_i D_i' \beta + \beta D_i D_i' \beta) Z_i Z_i' I_i
\]
\[
\xrightarrow{p} E[(Y_i^2 - 2 Y_i D_i' \beta + \beta D_i D_i' \beta) Z_i Z_i' I_i^A]
\]
\[
= E[\epsilon_i^2 Z_i Z_i' I_i^A],
\]
and
\[
\frac{1}{n} \sum_{i=1}^{n} Z_i D_i' I_i \xrightarrow{p} E[Z_i D_i' I_i^A].
\]
The conclusion then follows. 

\(\square\)

Step C.3.1.6. \((\hat{\beta}_c^c)^{-1}(\hat{\beta}_1^c - \beta_1) \xrightarrow{d} \mathcal{N}(0, 1)\).

Proof. By combining the results from Steps C.3.1.2–C.3.1.4 and by Lemma B.6,

\((A) \xrightarrow{p} 0,\)
\((B) \xrightarrow{d} \mathcal{N}(0, (E[Z_i D_i' I_i^A])^{-1} E[\epsilon_i^2 Z_i Z_i' I_i^A] (E[Z_i D_i' I_i^A])^{-1}),\)

and therefore,
\[
\sqrt{n}(\hat{\beta}_c^c - \beta_n) \xrightarrow{d} \mathcal{N}(0, (E[Z_i D_i' I_i^A])^{-1} E[\epsilon_i^2 Z_i Z_i' I_i^A] (E[Z_i D_i' I_i^A])^{-1}).
\]
The conclusion then follows from Steps C.3.1.1 and C.3.1.5.
C.3.2 Consistency and Asymptotic Normality of $\hat{\beta}_1^c$ When $\Pr(A(X_i) \in (0, 1)) > 0$

Let $I_i = 1\{p^s(X_i; \delta_n) \in (0, 1)\}$, $D_i^c = (1, D_i, p^s(X_i; \delta_n))^\prime$ and $Z_i^c = (1, Z_i, p^s(X_i; \delta_n))^\prime$. Let

$$\hat{\beta}^{c,s} = (\sum_{i=1}^n Z_i^c(D_i^c)'I_i^c)^{-1}\sum_{i=1}^n Z_i^cY_i^cI_i^c$$

and

$$\hat{\Sigma}^{c,s} = (\sum_{i=1}^n Z_i^c(D_i^c)'I_i^c)^{-1}(\sum_{i=1}^n (\epsilon_i^c)^2Z_i^c(Z_i^c)^\prime I_i^c)(\sum_{i=1}^n D_i^c(Z_i^c)^\prime I_i^c)^{-1},$$

where $\epsilon_i^c = Y_i - (D_i^c)'\hat{\beta}^{c,s}$. Here, we only show that $\hat{\beta}_1^{c,s} \xrightarrow{p} \beta_1$ if $S_n \to \infty$ and that $(\hat{\sigma}^s)^{-1}(\hat{\beta}_1^{c,s} - \beta_1) \xrightarrow{d} \mathcal{N}(0, 1)$ if Assumption 5 holds when $\Var(A(X_i)|I_i^A = 1) > 0$. For that, it suffices to show that

$$\hat{\beta}^{c,s} - \hat{\beta}^c = o_p(1)$$

if $S_n \to \infty$ and that

$$\sqrt{n}(\hat{\beta}^{c,s} - \hat{\beta}^c) = o_p(1),$$

$$n\hat{\Sigma}^{c,s} \xrightarrow{p} (E[\bar{Z}_i\bar{D}_i^s|I_i^A])^{-1}E[\epsilon_i^2\bar{Z}_i\bar{Z}_i^s|E[\bar{D}_i\bar{Z}_i^s|I_i^A]]^{-1}$$

if Assumption 5 holds. We have

$$\hat{\beta}^{c,s} - \hat{\beta}^c = \left\{\begin{array}{l}
\frac{1}{n}\sum_{i=1}^n Z_i^c(D_i^c)'I_i^c)^{-1}\frac{1}{n}\sum_{i=1}^n Z_i^cY_i^cI_i^c - \left(\frac{1}{n}\sum_{i=1}^n Z_i^c(D_i^c)'I_i^c\right)^{-1}\frac{1}{n}\sum_{i=1}^n Z_i^cY_i^cI_i^c \\
- \left(\frac{1}{n}\sum_{i=1}^n Z_i^c(D_i^c)'I_i^c\right)^{-1}(\frac{1}{n}\sum_{i=1}^n Z_i^cY_i^cI_i^c - \frac{1}{n}\sum_{i=1}^n Z_i^cY_i^cI_i^c) \\
- \left(\frac{1}{n}\sum_{i=1}^n Z_i^c(D_i^c)'I_i^c\right)^{-1}(\frac{1}{n}\sum_{i=1}^n Z_i^c(D_i^c)'I_i^c - \frac{1}{n}\sum_{i=1}^n Z_i^c(D_i^c)'I_i^c)(\frac{1}{n}\sum_{i=1}^n Z_i^cD_i^s(I_i^c)I_i^c)^{-1}\frac{1}{n}\sum_{i=1}^n Z_i^cY_i^cI_i^c.
\end{array}\right.$$
so that
\[
- \frac{1}{n} \sum_{i=1}^{n} (\xi_i^s)^2 Z_i^s (Z_i^s)' I_i^s \overset{p}{\to} E[\xi_i^2 \hat{Z}_i \hat{Z}_i' I_i^A].
\]
Also, \( \frac{1}{n} \sum_{i=1}^{n} Z_i^s(D_i^r)' I_i^s \overset{p}{\to} E[\hat{Z}_i D_i^r I_i^A] \) by using Lemma B.8. The conclusion then follows. \( \square \)

C.3.3 Consistency and Asymptotic Normality of \( \hat{\beta}_1 \) When \( \Pr(A(X_i) \in (0, 1)) = 0 \)

Since \( \Pr(A(X_i) \in (0, 1)) = 0, I_n = 0 \) with probability one. Hence,
\[
\hat{\beta} = \left( \sum_{i=1}^{n} Z_i D_i I_i \right)^{-1} \sum_{i=1}^{n} Z_i Y_i I_i
\]
with probability one. We use the notation and results provided in Appendix B. By Lemma B.5, under Assumption 3 (e), there exists \( \mu > 0 \) such that \( d_{\Omega^*}^s \) is twice continuously differentiable on \( N(\partial \Omega^*, \mu) \) and that
\[
\int_{N(\partial \Omega^*, \delta)} g(x)dx = \int_{-\delta}^{\delta} \int_{\partial \Omega^*} g(u + \lambda \nu_{\Omega^*}(u)) J_{\nu_{\Omega^*}}(u, \lambda) dH^{p-1}(u) d\lambda
\]
for every \( \delta \in (0, \mu) \) and every function \( g : \mathbb{R}^p \to \mathbb{R} \) that is integrable on \( N(\partial \Omega^*, \delta) \).

Below we show that \( \hat{\beta}_1 \overset{p}{\to} \beta_1 \) if \( n\delta_n \to \infty \) and \( \delta_n \to 0 \) and that \( \hat{\sigma}^{-1}(\hat{\beta}_1 - \beta_1) \overset{d}{\to} N(0, 1) \) if \( n\delta_n \to 0 \) in addition. The proof proceeds in eight steps.

**Step C.3.3.1.** There exist \( \bar{\delta} > 0 \) and a bounded function \( r : \partial \Omega^* \cap N(\mathcal{X}, \bar{\delta}) \times (-1, 1) \times (0, \bar{\delta}) \to \mathbb{R} \) such that
\[
p^A(u + \delta \nu_{\Omega^*}(u); \delta) = k(v) + \delta r(u, v, \delta)
\]
for every \( (u, v, \delta) \in \partial \Omega^* \cap N(\mathcal{X}, \bar{\delta}) \times (-1, 1) \times (0, \bar{\delta}) \), where
\[
k(v) = \begin{cases} 
1 - \frac{1}{2} I_{(1-v^2)}(\frac{p+1}{2}, \frac{1}{2}) & \text{for } v \in [0, 1) \\
\frac{1}{2} I_{(1-v^2)}(\frac{p+1}{2}, \frac{1}{2}) & \text{for } v \in (-1, 0].
\end{cases}
\]
Here \( I_*(\alpha, \beta) \) is the regularized incomplete beta function (the cumulative distribution function of the beta distribution with shape parameters \( \alpha \) and \( \beta \)).

**Proof.** By Assumption 3 (f) (ii), there exists \( \delta \in (0, \frac{\mu}{2}) \) such that \( A(x) = 0 \) for almost every \( x \in N(\mathcal{X}, 3\delta) \setminus \Omega^* \). By Taylor’s theorem, for every \( u \in \partial \Omega^* \cap N(\mathcal{X}, \bar{\delta}) \) and \( a \in B(0, 2\delta) \),
\[
d_{\Omega^*}^s(u + a) = d_{\Omega^*}^s(u) + \nabla d_{\Omega^*}^s(u)' a + a'R(u, a)a,
\]
where
\[
R(u, a) = \int_0^1 (1 - t)D^2 d_{\Omega^*}^s(u + ta) dt.
\]
Since \( D^2 d_{\Omega^*}^s \) is continuous and \( \text{cl}(N(\partial \Omega^*, 2\delta)) \) is bounded and closed, \( D^2 d_{\Omega^*}^s \) is bounded on \( \text{cl}(N(\partial \Omega^*, 2\delta)) \). Therefore, \( R(\cdot, \cdot) \) is bounded on \( \partial \Omega^* \cap N(\mathcal{X}, \bar{\delta}) \times B(0, 2\delta) \). It also follows that
\[
d_{\Omega^*}^s(u + a) = \nu_{\Omega^*}(u)' a + a'R(u, a)a,
\]
for every \( (u, \delta) \in N(\mathcal{X}, 3\delta) \setminus \Omega^* \).
since \(d_{\Omega}^s(u) = 0\) and \(\nabla d_{\Omega}^s(u) = \nu_{\Omega^*}(u)\) for every \(u \in \partial \Omega^* \cap N(\mathcal{X}, \delta)\) by Lemma B.1. For \((u, v, \delta) \in \partial \Omega^* \cap N(\mathcal{X}, \delta) \times (-1, 1) \times (0, \delta),\)

\[
p^A(u + \delta \nu_{\Omega^*}(u); \delta) = \frac{\int_{B(0, 1)} A(u + \delta \nu_{\Omega^*}(u) + \delta w) dw}{\text{Vol}_p} = \frac{\int_{B(0, 1)} 1\{u + \delta \nu_{\Omega^*}(u) + \delta w \in \Omega^*\} dw}{\text{Vol}_p} = \frac{\int_{B(0, 1)} 1\{d_{\Omega}^s(u + \delta (\nu_{\Omega^*}(u) + w)) \geq 0\} dw}{\text{Vol}_p} = \frac{\int_{B(0, 1)} 1\{\delta \nu_{\Omega^*}(u)(v_{\nu_{\Omega^*}}(u) + w) + \delta^2(v_{\nu_{\Omega^*}}(u) + w)' R(u, \delta(v_{\nu_{\Omega^*}}(u) + w))(v_{\nu_{\Omega^*}}(u) + w) \geq 0\} dw}{\text{Vol}_p},
\]

where \(\text{Vol}_p\) denotes the volume of the \(p\)-dimensional unit ball, and the second equality follows since \(u + \delta \nu_{\Omega^*}(u) + \delta w \in N(\mathcal{X}, 3\delta)\) and hence \(A(u + \delta \nu_{\Omega^*}(u) + \delta w) = 0\) for almost every \(w \in B(0, 1)\) such that \(u + \delta \nu_{\Omega^*}(u) + \delta w \notin \Omega^*.\) Observe that

\[
1\{\delta \nu_{\Omega^*}(u)(v_{\nu_{\Omega^*}}(u) + w) + \delta^2(v_{\nu_{\Omega^*}}(u) + w)' R(u, \delta(v_{\nu_{\Omega^*}}(u) + w))(v_{\nu_{\Omega^*}}(u) + w) \geq 0\}
= 1\{v + \nu_{\Omega^*}(u) \cdot w + \delta(v_{\nu_{\Omega^*}}(u) + w)' R(u, \delta(v_{\nu_{\Omega^*}}(u) + w))(v_{\nu_{\Omega^*}}(u) + w) \geq 0\}
= 1\{v + \nu_{\Omega^*}(u) \cdot w \geq 0\}
- 1\{v + \nu_{\Omega^*}(u) \cdot w \geq 0, v + \nu_{\Omega^*}(u) \cdot w + \delta(v_{\nu_{\Omega^*}}(u) + w)' R(u, \delta(v_{\nu_{\Omega^*}}(u) + w))(v_{\nu_{\Omega^*}}(u) + w) < 0\}
\]

Note that the set \(\{w \in B(0, 1) : v + \nu(u) \cdot w \geq 0\}\) is a region of the \(p\)-dimensional unit ball cut off by the plane \(\{w \in \mathbb{R}^p : v + \nu(u) \cdot w = 0\}\). The distance from the center of the unit ball to the plane is \(|v|\). Using the formula for the volume of a hyperspherical cap (see e.g. Li (2011)), we have

\[
\int_{B(0, 1)} 1\{v + \nu(u) \cdot w \geq 0\} dw = \begin{cases} 
\text{Vol}_p - \frac{1}{2} \text{Vol}_p I_{2(1-v) - (1-v)^2} (\frac{p+1}{2}, \frac{1}{2}) & \text{for } v \in [0, 1) \\
\frac{1}{2} \text{Vol}_p I_{2(1+v) - (1+v)^2} (\frac{p+1}{2}, \frac{1}{2}) & \text{for } v \in (-1, 0].
\end{cases}
\]

Therefore, for every \((u, v, \delta) \in \partial \Omega^* \cap N(\mathcal{X}, \delta) \times (-1, 1) \times (0, \delta),\)

\[
p^A(u + \delta \nu_{\Omega^*}(u); \delta) = k(v) + \frac{\int_{B(0, 1)} (-a(u, v, w, \delta) + b(u, v, w, \delta)) dw}{\text{Vol}_p}.
\]

Now let \(r(u, v, \delta) = \delta^{-1}(p^A(u + \delta \nu_{\Omega^*}(u); \delta) - k(v))\). Since \(R(\cdot, \cdot)\) is bounded on \(\partial \Omega^* \cap N(\mathcal{X}, \delta) \times B(0, 2\delta)\) and \(\|\nu_{\Omega^*}(u)\| = 1\), there exists \(\bar{r} > 0\) such that

\[
|(v_{\nu_{\Omega^*}}(u) + w)' R(u, \delta(v_{\nu_{\Omega^*}}(u) + w))(v_{\nu_{\Omega^*}}(u) + w)| \leq \bar{r}
\]
for every \((u, v, w, \delta) \in \partial \Omega^* \cap N(\mathcal{X}, \delta) \times (-1, 1) \times B(0, 1) \times (0, \delta)\). Therefore,

\[
0 \leq a(u, v, w, \delta) \leq 1\{0 \leq v + \nu_{\Omega^*}(u) \cdot w < \delta r\}
\]

and

\[
0 \leq b(u, v, w, \delta) \leq 1\{-\delta r \leq v + \nu_{\Omega^*}(u) \cdot w < 0\}.
\]

It then follows that

\[
\frac{\int_{B(0,1)} 1\{0 \leq v + \nu_{\Omega^*}(u) \cdot w < \delta r\}dw}{\text{Vol}_p} \leq \frac{\int_{B(0,1)} (-a(u, v, w, \delta) + b(u, v, w, \delta))dw}{\text{Vol}_p} \leq \frac{\int_{B(0,1)} 1\{-\delta r \leq v + \nu_{\Omega^*}(u) \cdot w < 0\}dw}{\text{Vol}_p}.
\]

The set \(\{w \in B(0,1) : 0 \leq v + \nu_{\Omega^*}(u) \cdot w < \delta r\}\) is a region of the \(p\)-dimensional unit ball cut off by the two planes \(\{w \in \mathbb{R}^p : v + \nu_{\Omega^*}(u) \cdot w = 0\}\) and \(\{w \in \mathbb{R}^p : v + \nu_{\Omega^*}(u) \cdot w = \delta r\}\). Its Lebesgue measure is at most the volume of the \((p - 1)\)-dimensional unit ball times the distance between the two planes, so

\[
-\delta \text{Vol}_{p-1}\bar{r} \leq -\int_{B(0,1)} 1\{0 \leq v + \nu_{\Omega^*}(u) \cdot w < \delta r\}dw.
\]

Likewise,

\[
\int_{B(0,1)} 1\{-\delta r \leq v + \nu_{\Omega^*}(u) \cdot w < 0\}dw \leq \delta \text{Vol}_{p-1}\bar{r}.
\]

Therefore,

\[
\frac{-\delta \text{Vol}_{p-1}\bar{r}}{\text{Vol}_p} \leq \frac{\int_{B(0,1)} (-a(u, v, w, \delta) + b(u, v, w, \delta))dw}{\text{Vol}_p} \leq \frac{\delta \text{Vol}_{p-1}\bar{r}}{\text{Vol}_p}.
\]

It follows that

\[
\frac{r(u, v, \delta)}{\text{Vol}_p} = \frac{-1}{\text{Vol}_p} \int_{B(0,1)} (-a(u, v, w, \delta) + b(u, v, w, \delta))dw \\
\in [-\frac{\text{Vol}_{p-1}\bar{r}}{\text{Vol}_p}, \frac{\text{Vol}_{p-1}\bar{r}}{\text{Vol}_p}],
\]

and hence \(r\) is bounded on \(\partial \Omega^* \cap N(\mathcal{X}, \delta) \times (-1, 1) \times (0, \delta)\). \(\square\)

**Step C.3.3.2.** For every \((u, v, \delta) \in \partial \Omega^* \cap N(\mathcal{X}, \delta) \times (-1, 1) \times (0, \delta), p^A(u + \delta v\nu_{\Omega^*}(u); \delta) \in (0, 1)\).

**Proof.** Fix \((u, v, \delta) \in \partial \Omega^* \cap N(\mathcal{X}, \delta) \times (-1, 1) \times (0, \delta).\) Suppose \(v = 0\). By Step C.3.3.1, \(p^A(u) = \lim_{\delta \to 0} p^A(u; \delta') = k(0) = \frac{1}{2}\). This implies that there exists \(\delta' \in (0, \delta)\) such that \(p^A(w; \delta') \in (0, 1)\). It then follows that \(0 < L^p(B(u, \delta') \cap \Omega^*) \leq L^p(B(x, \delta) \cap \Omega^*)\) and that \(0 < L^p(B(x, \delta') \cap \Omega^*) \leq L^p(B(x, \delta) \cap \Omega^*)\). Therefore, \(p^A(u; \delta) = \frac{L^p(B(u, \delta') \cap \Omega^*)}{L^p(B(x, \delta) \cap \Omega^*)} \in (0, 1)\).

Suppose \(v \neq 0\) and let \(\epsilon \in (0, \delta(1 - |v|))\). Note that \(B(u, \epsilon) \subset B(u + \delta v\nu_{\Omega^*}(u), \delta)\), since for any \(x \in B(u, \epsilon), \|u + \delta v\nu_{\Omega^*}(u) - x\| \leq \|\delta v\nu_{\Omega^*}(u)\| + \|u - x\| \leq \delta |v| + \epsilon < \delta\). Since \(p^A(u) = \frac{1}{2}\), there exists \(\epsilon' \in (0, \epsilon)\) such that \(p^A(u; \epsilon') \in (0, 1)\). It then follows that \(0 < L^p(B(u, \epsilon') \cap \Omega^*) \leq L^p(B(u, \epsilon) \cap \Omega^*) \leq L^p(B(x, \epsilon') \cap \Omega^*) \leq L^p(B(x, \epsilon) \cap \Omega^*) \leq L^p(B(u + \delta v\nu_{\Omega^*}(u), \delta) \cap \Omega^*)\). Therefore, \(p^A(u + \delta v\nu_{\Omega^*}(u); \delta) = \frac{L^p(B(u + \delta v\nu_{\Omega^*}(u), \delta) \cap \Omega^*)}{L^p(B(x, \epsilon) \cap \Omega^*)} \in (0, 1)\). \(\square\)
Step C.3.3.3. Let $g : \mathbb{R}^p \to \mathbb{R}$ be a function that is bounded on $N(\partial \Omega^*, \delta') \cap N(\mathcal{X}, \delta')$ for some $\delta' > 0$. Then, for $l \geq 0$, there exist $\tilde{\delta} > 0$ and constant $C > 0$ such that

$$|\delta^{-1} E[p^A(X_i; \delta) g(X_i) 1\{p^A(X_i; \delta) \in (0, 1)\}]| \leq C$$

for every $\delta \in (0, \tilde{\delta})$. If $g$ is continuous on $N(\partial \Omega^*, \delta') \cap N(\mathcal{X}, \delta')$ for some $\delta' > 0$, then

$$\delta^{-1} E[p^A(X_i; \delta) g(X_i) 1\{p^A(X_i; \delta) \in (0, 1)\}] = \int_{\Omega^*} k(v) dv \int_{\partial \Omega^*} g(x) f_X(x) d\mathcal{H}^{p-1}(x) + o(1)$$

for $l \geq 0$. Furthermore, if $g$ is continuously differentiable and $\nabla g$ is bounded on $N(\partial \Omega^*, \delta') \cap N(\mathcal{X}, \delta')$ for some $\delta' > 0$, then

$$\delta^{-1} E[p^A(X_i; \delta) g(X_i) 1\{p^A(X_i; \delta) \in (0, 1)\}] = \int_{\Omega^*} k(v) dv \int_{\partial \Omega^*} g(x) f_X(x) d\mathcal{H}^{p-1}(x) + O(\delta)$$

for $l \geq 0$.

Proof. Let $\tilde{\delta}$ be given in Step C.3.3.1. Under Assumption 3 (g), there exists $\tilde{\delta} \in (0, \tilde{\delta})$ such that $f_X$ is bounded, is continuously differentiable, and has bounded partial derivatives on $N(\partial \Omega^*, 2\tilde{\delta}) \cap N(\mathcal{X}, 2\tilde{\delta})$. Let $\tilde{\delta} \in (0, \tilde{\delta})$ be such that both $g$ and $f_X$ are bounded on $N(\partial \Omega^*, 2\tilde{\delta}) \cap N(\mathcal{X}, 2\tilde{\delta})$. We first show that $p^A(x; \delta) \in (0, 1)$ for every $x \in \mathcal{X} \setminus N(\partial \Omega^*, \delta)$ for every $\delta \in (0, \tilde{\delta})$. Pick $x \in \mathcal{X} \setminus N(\partial \Omega^*, \delta)$ and $\delta \in (0, \tilde{\delta})$. Since $B(x, \delta) \cap \partial \Omega^* = \emptyset$, either $B(x, \delta) \subset \text{int}(\Omega^*)$ or $B(x, \delta) \subset \text{int}(\mathbb{R}^p \setminus \Omega^*)$. If $B(x, \delta) \subset \text{int}(\Omega^*)$, $p^A(x; \delta) = 1$. If $B(x, \delta) \subset \text{int}(\mathbb{R}^p \setminus \Omega^*)$, $p^A(x; \delta) = 0$, since $A(x') = 0$ for almost every $x' \in B(x, \delta) \subset N(\mathcal{X}, 3\tilde{\delta}) \setminus \Omega^*$ by the choice of $\tilde{\delta}$. Therefore, \{x \in \mathcal{X} : p^A(x; \delta) \in (0, 1)\} \subset N(\partial \Omega^*, \delta)$ for every $\delta \in (0, \tilde{\delta})$. By this and Lemma B.5, for $\delta \in (0, \tilde{\delta})$,

$$\delta^{-1} E[p^A(X_i; \delta) g(X_i) 1\{p^A(X_i; \delta) \in (0, 1)\}] = \delta^{-1} \int p^A(x; \delta) g(x) 1\{p^A(x; \delta) \in (0, 1)\} f_X(x) dx$$

$$= \delta^{-1} \int_{N(\partial \Omega^*, \delta)} p^A(x; \delta) g(x) 1\{p^A(x; \delta) \in (0, 1)\} f_X(x) dx$$

$$= \delta^{-1} \int_{-\delta}^\delta \int_{\partial \Omega^*} p^A(u + \lambda \nu_{\Omega^*} (u); \delta) j g(u + \lambda \nu_{\Omega^*} (u)) 1\{p^A(u + \lambda \nu_{\Omega^*} (u); \delta) \in (0, 1)\}$$

$$\times f_X(u + \lambda \nu_{\Omega^*} (u)) J^\Omega_{\nu_{\Omega^*}} (u, \lambda) d\mathcal{H}^{p-1}(u) d\lambda.$$
For every \((u, v, \delta) \in \partial\Omega^* \setminus N(\mathcal{X}, \tilde{\delta}) \times (-1, 1) \times (0, \tilde{\delta}), u + \delta v \nu_{\Omega^*}(u) \notin \mathcal{X},\) so
\[
\delta^{-1} E[p^A(X_i; \delta) g(X_i) 1\{p^A(X_i; \delta) \in (0, 1)\}]
= \int_{-1}^{1} \int_{\partial\Omega^* \cap N(\mathcal{X}, \tilde{\delta})} p^A(u + \delta v \nu_{\Omega^*}(u); \delta) 1\{p^A(u + \delta v \nu_{\Omega^*}(u); \delta) \in (0, 1)\}
\times g(u + \delta v \nu_{\Omega^*}(u)) f_X(u + \delta v \nu_{\Omega^*}(u)) J^\partial_{\psi}^\Omega \psi_{\Omega^*}(u, \delta v) dH^{p-1}(u) dv
= \int_{-1}^{1} \int_{\partial\Omega^* \cap N(\mathcal{X}, \tilde{\delta})} (k(v) + \delta r(u, v, \delta)) g(u + \delta v \nu_{\Omega^*}(u)) f_X(u + \delta v \nu_{\Omega^*}(u)) J^\partial_{\psi}^\Omega \psi_{\Omega^*}(u, \delta v) dH^{p-1}(u) dv,
\]
where the second equality follows from Steps C.3.3.1 and C.3.3.2. By Lemma B.5, \(J^\partial_{\psi}^\Omega \psi_{\Omega^*}(\cdot, \cdot)\) is bounded on \(\partial\Omega^* \times (-\tilde{\delta}, \tilde{\delta}).\) Since \(r, g\) and \(f_X\) are also bounded, for some constant \(C > 0,
\|
\delta^{-1} E[p^A(X_i; \delta) 1\{p^A(X_i; \delta) \in (0, 1)\}]\|
\leq C \int_{-1}^{1} \int_{\partial\Omega^* \cap N(\mathcal{X}, \tilde{\delta})} dH^{p-1}(u) dv,
\]
which is finite by Assumption 3 (f) (i). Moreover, if \(g\) and \(f_X\) are continuous on \(N(\partial\Omega^*, 2\tilde{\delta}) \cap N(\mathcal{X}, 2\tilde{\delta}),\) by the Dominated Convergence Theorem,
\[
\delta^{-1} E[p^A(X_i; \delta) g(X_i) 1\{p^A(X_i; \delta) \in (0, 1)\}] \to \int_{-1}^{1} k(v) dv \int_{\partial\Omega^*} g(u) f_X(u) dH^{p-1}(u),
\]
where we use the fact from Lemma B.5 that \(J^\partial_{\psi}^\Omega \psi_{\Omega^*}(u, \delta)\) is continuous in \(\lambda\) and \(J^\partial_{\psi}^\Omega \psi_{\Omega^*}(u, 0) = 1.\)

Note that \(A(x) = 1\) for every \(x \in \Omega^*\) and \(A(x) = 0\) for almost every \(x \in N(\mathcal{X}, 2\tilde{\delta}) \setminus \Omega^*\). Also, for every \((u, v, \delta) \in \partial\Omega^* \cap N(\mathcal{X}, \tilde{\delta}) \times (-1, 1) \times (0, \tilde{\delta}), u + \delta v \nu_{\Omega^*}(u) \in \Omega^*\) if \(v \in (0, 1)\) and \(u + \delta v \nu_{\Omega^*}(u) \in \mathcal{X}, 2\tilde{\delta}) \setminus \Omega^*\) if \(v \in (-1, 0]\). Therefore,
\[
\delta^{-1} E[Z_p p^A(X_i; \delta) g(X_i) 1\{p^A(X_i; \delta) \in (0, 1)\}]
= \delta^{-1} E[A(X_i) p^A(X_i; \delta) g(X_i) 1\{p^A(X_i; \delta) \in (0, 1)\}]
= \int_{-1}^{1} \int_{\partial\Omega^* \cap N(\mathcal{X}, \tilde{\delta})} A(u + \delta v \nu_{\Omega^*}(u)) (k(v) + \delta r(u, v, \delta)) g(u + \delta v \nu_{\Omega^*}(u))
\times f_X(u + \delta v \nu_{\Omega^*}(u)) J^\partial_{\psi}^\Omega \psi_{\Omega^*}(u, \delta v) dH^{p-1}(u) dv
= \int_{0}^{1} \int_{\partial\Omega^* \cap N(\mathcal{X}, \tilde{\delta})} (k(v) + \delta r(u, v, \delta)) g(u + \delta v \nu_{\Omega^*}(u)) f_X(u + \delta v \nu_{\Omega^*}(u)) J^\partial_{\psi}^\Omega \psi_{\Omega^*}(u, \delta v) dH^{p-1}(u) dv
\to \int_{0}^{1} k(v) dv \int_{\partial\Omega^*} g(u) f_X(u) dH^{p-1}(u).
\]

Now suppose that \(g\) and \(f_X\) are continuously differentiable on \(N(\partial\Omega^*, 2\tilde{\delta}) \cap N(\mathcal{X}, 2\tilde{\delta})\) and that \(\nabla g\) and \(\nabla f\) are bounded on \(N(\partial\Omega^*, 2\tilde{\delta}) \cap N(\mathcal{X}, 2\tilde{\delta})\). Using the mean-value theorem, we obtain that, for any \((u, v, \delta) \in \partial\Omega^* \cap N(\mathcal{X}, \tilde{\delta}) \times (-1, 1) \times (0, \tilde{\delta}),\)
\[
g(u + \delta v \nu_{\Omega^*}(u)) = g(u) + \nabla g(y_g(u, \delta v \nu_{\Omega^*}(u))) \delta v \nu_{\Omega^*}(u),

f_X(u + \delta v \nu_{\Omega^*}(u)) = f_X(u) + \nabla f_X(y_f(u, \delta v \nu_{\Omega^*}(u))) \delta v \nu_{\Omega^*}(u)
\]
for some \( y_g(u, \delta v y_\Omega (u)) \) and \( y_f(u, \delta v y_\Omega (u)) \) that are on the line segment connecting \( u \) and \( u + \delta v y_\Omega (u) \). In addition,

\[
J_{p-1}^{\partial \Omega^*} \psi y_\Omega (u, \delta v) = J_{p-1}^{\partial \Omega^*} \psi y_\Omega (u, 0) + \frac{\partial J_{p-1}^{\partial \Omega^*} \psi y_\Omega (u, y_f(u, \delta v))}{\partial \lambda} \delta v
\]

\[
= 1 + \frac{\partial J_{p-1}^{\partial \Omega^*} \psi y_\Omega (u, y_f(u, \delta v))}{\partial \lambda} \delta v
\]

for some \( y_f(u, \delta v) \) that is on the line segment connecting 0 and \( \delta v \). By Lemma B.5, \( \frac{\partial J_{p-1}^{\partial \Omega^*} \psi y_\Omega (\cdot)}{\partial \lambda} \) is bounded on \( \partial \Omega^* \times (-\delta, \delta) \). We then have

\[
\delta^{-1} E[p^A(X_1; \delta)^l g(X_1) 1\{p^A(X_1; \delta) \in (0, 1)\}] = \int_{-1}^{1} \int_{\partial \Omega^* \cap N(\hat{X}, \delta)} (k(v) + \delta r(u, v, \delta))^l (g(u) + \nabla g(y_g(u, \delta v y_\Omega (u))))^l \delta v y_\Omega (u) (1 + \frac{\partial J_{p-1}^{\partial \Omega^*} \psi y_\Omega (u, y_f(u, \delta v))}{\partial \lambda} \delta v) dH^{p-1}(u) dv
\]

\[
= \int_{-1}^{1} \int_{\partial \Omega^* \cap N(\hat{X}, \delta)} (k(v))^l g(u) f_X(u) + \delta h(u, v, \delta) dH^{p-1}(u) dv
\]

\[
= \int_{-1}^{1} k(v)^l dv \int_{\partial \Omega^*} g(u) f_X(u) dH^{p-1}(u) + \delta \int_{-1}^{1} \int_{\partial \Omega^* \cap N(\hat{X}, \delta)} h(u, v, \delta) dH^{p-1}(u) dv
\]

for some function \( h \) bounded on \( \partial \Omega^* \cap N(\hat{X}, \delta) \times (-1, 1) \times (0, \delta) \). It then follows that

\[
\delta^{-1} E[p^A(X_1; \delta)^l g(X_1) 1\{p^A(X_1; \delta) \in (0, 1)\}] = \int_{-1}^{1} k(v)^l dv \int_{\partial \Omega^*} g(u) f_X(u) dH^{p-1}(u) + O(\delta).
\]

Also,

\[
\delta^{-1} E[Z_i p^A(X_1; \delta)^l g(X_1) 1\{p^A(X_1; \delta) \in (0, 1)\}] = \int_{0}^{1} \int_{\partial \Omega^* \cap N(\hat{X}, \delta)} (k(v) + \delta r(u, v, \delta))^l g(u + \delta v y_\Omega (u)) f_X(u + \delta v y_\Omega (u)) J_{p-1}^{\partial \Omega^*} \psi y_\Omega (u, \delta v) dH^{p-1}(u) dv
\]

\[
= \int_{0}^{1} k(v)^l dv \int_{\partial \Omega^*} g(u) f_X(u) dH^{p-1}(u) + O(\delta).
\]

\[\Box\]

**Step C.3.3.4.** Let \( S_D = \lim_{\delta \to 0} \delta^{-1} E[Z_i D_i 1\{p^A(X_1; \delta) \in (0, 1)\}] \) and \( S_Y = \lim_{\delta \to 0} \delta^{-1} E[Z_i Y_i 1\{p^A(X_1; \delta) \in (0, 1)\}] \). Then the second element of \( S_D^{-1} S_Y \) is \( \beta_1 \).

**Proof.** Note that \( D_i = Z_i D_i(1) + (1 - Z_i) D_i(0) \) and \( Y_i = Z_i Y_i + (1 - Z_i) Y_i \). By Step C.3.3.3,

\[
S_D = \begin{bmatrix}
2f_X & \int_{\partial \Omega^*} E[D_i(1) + D_i(0)|X_1 = x] f_X(x) dH^{p-1}(x) & \int_{-1}^{1} k(v)^l dv f_X \\
f_X & \int_{\partial \Omega^*} E[D_i(1)|X_1 = x] f_X(x) dH^{p-1}(x) & \int_{-1}^{1} k(v)^l dv f_X \\
\int_{-1}^{1} k(v) dv f_X & \int_{\partial \Omega^*} \int_{-1}^{1} k(v) dv E[D_i(1)|X_1 = x] f_X(x) dH^{p-1}(x) & \int_{-1}^{1} k(v)^2 dv f_X
\end{bmatrix},
\]

where \( f_{X} \) is the density function of \( X_1 \), and \( f_{D_i} \) is the density function of \( D_i \).
where \( \tilde{f}_X = \int_{\partial \Omega^*} f_X(x) d\mathcal{H}^{p-1}(x) \), and
\[
S_Y = \left[ \begin{array}{c}
\int_{\partial \Omega^*} E[Y_{1i} + Y_{0i} | X_i = x] f_X(x) d\mathcal{H}^{p-1}(x) \\
\int_{\partial \Omega^*} E[Y_{1i} | X_i = x] f_X(x) d\mathcal{H}^{p-1}(x) \\
\int_{\partial \Omega^*} \left( \int_0^1 k(v) dv E[Y_{1i} | X_i = x] \right) + \int_0^1 k(v) dv E[Y_{0i} | X_i = x] f_X(x) d\mathcal{H}^{p-1}(x)
\end{array} \right].
\]
After a few lines of algebra, we have
\[
\text{det}(S_D) = \tilde{f}_X \int_{\partial \Omega^*} E[D_i(1) - D_i(0)|X_i = x] f_X(x) d\mathcal{H}^{p-1}(x) \times \left( \int_{-1}^1 (k(v) - \int_{-1}^0 k(s) ds)^2 dv + \int_0^1 (k(v) - \int_0^1 k(s) ds)^2 dv \right),
\]
which is nonzero under Assumption 3 (b) and (f) (i). After another few lines of algebra, we obtain that the second element of \( S_D^{-1} S_Y \) is
\[
\int_{\partial \Omega^*} E[(D_i(1) - D_i(0))(Y_{1i}(1) - Y_{1i}(0))|X_i = x] f_X(x) d\mathcal{H}^{p-1}(x) / \int_{\partial \Omega^*} E[D_i(1) - D_i(0)|X_i = x] f_X(x) d\mathcal{H}^{p-1}(x).
\]
On the other hand, by Step C.3.3.3,
\[
\beta_1 = \lim_{\delta \to 0} E[w_i(\delta)(Y_{1i}(1) - Y_{1i}(0))]
= \lim_{\delta \to 0} \frac{\delta^{-1} E[p^A(X_i; \delta)(1 - p^A(X_i; \delta))(D_i(1) - D_i(0))(Y_{1i}(1) - Y_{1i}(0))1\{p^A(X_i; \delta) \in (0, 1)\}]}{\delta^{-1} E[p^A(X_i; \delta)(1 - p^A(X_i; \delta))(D_i(1) - D_i(0))(Y_{1i}(1) - Y_{1i}(0))1\{p^A(X_i; \delta) \in (0, 1)\}]} 
= \int_{-1}^1 k(v)(1 - k(v)) dv \int_{\partial \Omega^*} E[(D_i(1) - D_i(0))(Y_{1i}(1) - Y_{1i}(0))|X_i = x] f_X(x) d\mathcal{H}^{p-1}(x) 
= \int_{\partial \Omega^*} E[(D_i(1) - D_i(0))(Y_{1i}(1) - Y_{1i}(0))|X_i = x] f_X(x) d\mathcal{H}^{p-1}(x) / \int_{\partial \Omega^*} E[D_i(1) - D_i(0)|X_i = x] f_X(x) d\mathcal{H}^{p-1}(x).
\]
\[\square\]

**Step C.3.3.5.** If \( n\delta_n \to \infty \) as \( n \to \infty \), then \( \hat{\beta}_1 \to p \beta_1 \).

**Proof.** It suffices to verify that the variance of each element of \( \frac{1}{n\delta_n} \sum_{i=1}^n Z_i D_i I_i \) and \( \frac{1}{n\delta_n} \sum_{i=1}^n Z_i Y_i I_i \) is \( o(1) \). Here, we only verify that \( \text{Var}\left( \frac{1}{n\delta_n} \sum_{i=1}^n p^A(X_i; \delta_n) Y_i I_i \right) = o(1) \). Note that
\[
E[Y_i^2 | X_i] = E[Z_i Y_i^2 + (1 - Z_i) Y_i^2 | X_i] \leq E[Y_i^2 + Y_i^2 | X_i] .
\]

Under Assumption 3 (g), there exists \( \delta' > 0 \) such that \( E[Y_i^2 + Y_i^2 | X_i] \) is continuous on \( N(\partial \Omega^*, \delta') \). Since \( cl(N(\partial \Omega^*, \frac{1}{2}\delta')) \) is closed and bounded, \( E[Y_i^2 + Y_i^2 | X_i] \) is bounded on \( cl(N(\partial \Omega^*, \frac{1}{2}\delta')) \). We have
\[
\text{Var}\left( \frac{1}{n\delta_n} \sum_{i=1}^n p^A(X_i; \delta_n) Y_i I_i \right) \leq \frac{1}{n\delta_n} \delta_n^{-1} E[p^A(X_i; \delta_n)^2 Y_i^2 I_i] 
= \frac{1}{n\delta_n} \delta_n^{-1} E[p^A(X_i; \delta_n)^2 E[Y_i^2 | X_i] I_i] 
\leq \frac{1}{n\delta_n} C
\]
for some \( C > 0 \), where the last inequality follows from Step C.3.3.3. The conclusion follows since \( n\delta_n \to \infty \). \[\square\]
Now let \( \beta = (\beta_0, \beta_1, \beta_2)' = S^{-1}_D S_Y \) and let \( \epsilon_i = Y_i - D_i' \beta \). We can write
\[
\sqrt{n\delta_n}(\hat{\beta} - \beta) = \left( \frac{1}{n\delta_n} \sum_{i=1}^{n} Z_i D_i' I_i \right)^{-1} \frac{1}{\sqrt{n\delta_n}} \sum_{i=1}^{n} Z_i \epsilon_i I_i = \left( \frac{1}{n\delta_n} \sum_{i=1}^{n} Z_i D_i' I_i \right)^{-1} \frac{1}{\sqrt{n\delta_n}} \sum_{i=1}^{n} \{ (Z_i \epsilon_i I_i - E[Z_i \epsilon_i I_i]) + E[Z_i \epsilon_i I_i] \}.
\]

**Step C.3.3.6.**
\[
\frac{1}{\sqrt{n\delta_n}} \sum_{i=1}^{n} (Z_i \epsilon_i I_i - E[Z_i \epsilon_i I_i]) \xrightarrow{d} N(0, V),
\]
where \( V = \lim_{n \to \infty} \delta_n^{-1} E[\epsilon_i^2 Z_i Z_i'] I_i \).

**Proof.** We use the triangular-array Lyapunov CLT and the Cramér-Wold device. Pick a nonzero \( \lambda \in \mathbb{R}^p \), and let \( V_{i,n} = \frac{1}{\sqrt{n\delta_n}} \lambda'(Z_i \epsilon_i I_i - E[Z_i \epsilon_i I_i]) \). First,
\[
\sum_{i=1}^{n} E[V_{i,n}^2] = \delta_n^{-1} \lambda'(E[\epsilon_i^2 Z_i Z_i'] I_i - E[Z_i \epsilon_i I_i] E[Z_i \epsilon_i I_i]) \lambda.
\]

By Step C.3.3.3,
\[
E[Z_i \epsilon_i I_i] = E[Z_i (Y_i - D_i' \beta)] I_i = O(\delta_n),
\]
so
\[
\delta_n^{-1} E[Z_i \epsilon_i I_i] E[Z_i \epsilon_i I_i] = o(1).
\]
We have
\[
E[\epsilon_i^2 Z_i Z_i'] I_i = E[(Y_i - \beta_0 - \beta_1 D_i - \beta_2 p^A(X_i; \delta_n))^2 Z_i Z_i'] I_i = E[Z_i (Y_i' - \beta_0 - \beta_1 D_i(1) - \beta_2 p^A(X_i; \delta_n))^2 Z_i Z_i'] I_i + E[(1 - Z_i)(Y_0 - \beta_0 - \beta_1 D_i(0) - \beta_2 p^A(X_i; \delta_n))^2 Z_i Z_i'] I_i.
\]

Since \( E[Y_i X_i], E[Y_i 0 X_i], E[D_i(1) X_i], E[D_i(0) X_i], E[Y_i^2 X_i], E[Y_i 0^2 X_i], E[Y_i D_i(1) X_i] \) and \( E[Y_i D_i(0) X_i] \) are continuous on \( N(\delta \Omega^*, \delta') \) for some \( \delta' > 0 \) under Assumption 3 (g), \( \lim_{n \to \infty} \delta_n^{-1} E[\epsilon_i^2 Z_i Z_i'] I_i \) exists and finite. Therefore,
\[
\sum_{i=1}^{n} E[V_{i,n}^2] \to \lambda' V \lambda < 0.
\]
We next verify the Lyapunov condition: for some \( t > 0 \),
\[
\sum_{i=1}^{n} E[[V_{i,n}]^{2+t}] \to 0.
\]
We have
\[
\sum_{i=1}^{n} E[[V_{i,n}]^4] = \frac{1}{n\delta_n} \delta_n^{-1} E[[\lambda'(Z_i \epsilon_i I_i - E[Z_i \epsilon_i I_i])]^4] \leq \frac{1}{n\delta_n} 2^{3\epsilon} \delta_n^{-1} \{ E[|\lambda' Z_i \epsilon_i I_i|^4] + |\lambda' E[Z_i \epsilon_i I_i]|^4 \}
\]

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by the $c_r$-inequality. Repeating using the $c_r$-inequality gives

$$
\delta_n^{-1}E[|\lambda'Z_i \epsilon_i I_i|^4] = \delta_n^{-1}E[|\lambda'Z_i(Y_i - \beta_0 - \beta_1 D_i - \beta_2 p^A(X_i; \delta_n))|^4 I_i]
\leq 2^{2c} \delta_n^{-1}E[|\lambda'Z_i|^4(|Y_i|^4 + |\beta_0|^4 + |\beta_1|^4 D_i + |\beta_2|^4 p^A(X_i; \delta_n)^4)| I_i]
\leq 2^{3c} (\lambda_1 + \lambda_2 + \lambda_3)^4 \delta_n^{-1}E[|Y_i|^4 + \beta_0^4 + \beta_1^4 + \beta_2^4)| I_i]
= 2^{3c} O(1)
$$

for some finite constant $c$, where the last equality holds by Step C.3.3.3 under Assumption 3 (g). Moreover,

$$
\delta_n^{-1}|\lambda' E[Z_i \epsilon_i I_i]|^4 = \delta_n^3 |\lambda' \delta_n^{-1} E[Z_i \epsilon_i I_i]|^4
= \delta_n^3 O(1)
= o(1).
$$

Therefore, when $n \delta_n \to \infty$,

$$
\sum_{i=1}^n E[|V_i|^4] \to 0,
$$

and the conclusion follows from the Lyapunov CLT and the Cramér-Wold device. \qed

**Step C.3.3.7.** $n \delta_n \hat{\Sigma} \xrightarrow{p} S_D^{-1} \mathbf{V} (S_D^{-1})^{-1}$.

**Proof.** We have

$$
\frac{1}{n \delta_n} \sum_{i=1}^n \hat{\epsilon}_i^2 Z_i Z_i' I_i = \frac{1}{n \delta_n} \sum_{i=1}^n (Y_i - D_i' \hat{\beta})^2 Z_i Z_i' I_i
= \frac{1}{n \delta_n} \sum_{i=1}^n (\epsilon_i - D_i' (\hat{\beta} - \beta))^2 Z_i Z_i' I_i
= \frac{1}{n \delta_n} \sum_{i=1}^n \epsilon_i^2 Z_i Z_i' I_i
- \frac{2}{n \delta_n} \sum_{i=1}^n (Y_i - D_i' \beta)((\hat{\beta}_0 - \beta_0) + D_i(\hat{\beta}_1 - \beta_1) + p^A(X_i; \delta_n)(\hat{\beta}_2 - \beta_2)) Z_i Z_i' I_i
+ \frac{1}{n \delta_n} \sum_{i=1}^n ((\hat{\beta}_0 - \beta_0) + D_i(\hat{\beta}_1 - \beta_1) + p^A(X_i; \delta_n)(\hat{\beta}_2 - \beta_2))^2 Z_i Z_i' I_i
= \frac{1}{n \delta_n} \sum_{i=1}^n \epsilon_i^2 Z_i Z_i' I_i + o_p(1) O_p(1),
$$

where the last equality follows from the result that $\hat{\beta} - \beta = o_p(1)$ and from application of Step C.3.3.3 as in Steps C.3.3.5 and C.3.3.6. To show $\frac{1}{n \delta_n} \sum_{i=1}^n \epsilon_i^2 Z_i Z_i' I_i \xrightarrow{p} \mathbf{V}$, it suffices to verify that the variance of each element of $\frac{1}{n \delta_n} \sum_{i=1}^n \epsilon_i^2 Z_i Z_i' I_i$ is $o(1)$. We only verify that
Var\left( \frac{1}{n \delta_n} \sum_{i=1}^{n} \epsilon_i^2 p^A(X_i; \delta_n)^2 I_i \right) = o(1). Using the c_r-inequality, we have that for some constant c, 
\[
\begin{align*}
\text{Var}\left( \frac{1}{n \delta_n} \sum_{i=1}^{n} \epsilon_i^2 p^A(X_i; \delta_n)^2 I_i \right) &\leq \frac{1}{n \delta_n^2} \delta_n^{-1} E[\epsilon_i^4 I_i] \\
&= \frac{1}{n \delta_n^2} \delta_n^{-1} E[(Y_i - \beta_0 - \beta_1 D_i - \beta_2 p^A(X_i))^4 I_i] \\
&\leq \frac{1}{n \delta_n^2} 2^{3c} \delta_n^{-1} E[(Y_i^4 + \beta_0^4 + \beta_1^4 D_i + \beta_2^4 p^A(X_i))^4 I_i] \\
&\leq \frac{1}{n \delta_n^2} 2^{3c} \delta_n^{-1} E[(Y_i^4 + \beta_0^4 + \beta_1^4 + \beta_2^4) I_i] \\
&= \frac{1}{n \delta_n} 2^{3c} O(1) \\
&= o(1),
\end{align*}
\]
where the second last equality holds by Step C.3.3.3 under Assumption 3 (g). Therefore, 
\[
\frac{1}{n \delta_n} \sum_{i=1}^{n} \epsilon_i^2 Z_i Z_i' I_i \xrightarrow{p} V.
\]

It follows that 
\[
n \delta_n \hat{\Sigma} = \left( \frac{1}{n \delta_n} \sum_{i=1}^{n} Z_i D_i' I_i \right)^{-1} \left( \frac{1}{n \delta_n} \sum_{i=1}^{n} \epsilon_i^2 Z_i Z_i' I_i \right) \left( \frac{1}{n \delta_n} \sum_{i=1}^{n} D_i Z_i' I_i \right)^{-1} \xrightarrow{p} S_D^{-1} V(S_D')^{-1}.
\]

\hfill \Box

**Step C.3.3.8.** \( \hat{\sigma}^{-1}(\hat{\beta} - \beta_1) \xrightarrow{d} N(0, 1) \).

**Proof.** Let \( \beta_n = S_D^{-1} \delta_n^{-1} E[Z_i Y_i I_i] \). We then have 
\[
\begin{align*}
\frac{1}{\sqrt{n \delta_n}} \sum_{i=1}^{n} E[Z_i \epsilon_i I_i] &\equiv \sqrt{n \delta_n \delta_n^{-1}} E[Z_i(Y_i - D_i' \beta)] \\
&= \sqrt{n \delta_n} \delta_n^{-1} E[Z_i(Y_i - D_i' \beta_n + D_i' (\beta_n - \beta))] \\
&= \sqrt{n \delta_n} \delta_n^{-1} \{ E[Z_i Y_i I_i] - E[Z_i D_i I_i] \beta_n + E[Z_i D_i I_i] (\beta_n - \beta) \} \\
&= \sqrt{n \delta_n} \{ S_D - \delta_n^{-1} E[Z_i D_i I_i] S_D^{-1} \delta_n^{-1} E[Z_i Y_i I_i] \\
&+ \delta_n^{-1} E[Z_i D_i I_i] S_D^{-1} (\delta_n^{-1} E[Z_i Y_i I_i] - S_Y) \} \\
&= \sqrt{n \delta_n} (O(\delta_n) O(1) + O(1) O(\delta_n)) \\
&= O(\sqrt{n \delta_n \delta_n}),
\end{align*}
\]
where we use Step C.3.3.3 for the second last equality. Thus, when \( n \delta_n^3 \to 0 \), 
\[
\sqrt{n \delta_n}(\hat{\beta} - \beta) = \left( \frac{1}{n \delta_n} \sum_{i=1}^{n} Z_i D_i I_i \right)^{-1} \frac{1}{\sqrt{n \delta_n}} \sum_{i=1}^{n} \{ (Z_i \epsilon_i I_i - E[Z_i \epsilon_i I_i]) + E[Z_i \epsilon_i I_i] \} \\
\xrightarrow{d} N(0, S_D^{-1} V(S_D')^{-1}).
\]
The conclusion then follows from Step C.3.3.7. \hfill \Box
C.3.4 Consistency and Asymptotic Normality of $\hat{\beta}_s^i$ When $\Pr(A(X_i) \in (0, 1)) = 0$

Let $I_i^s = 1\{p^s(X_i; \delta_n) \in (0, 1)\}$, $D_i^s = (1, D_i, p^s(X_i; \delta_n))'$ and $Z_i^s = (1, Z_i, p^s(X_i; \delta_n))'$. $\hat{\beta}^s$ and $\hat{\Sigma}^s$ are given by

$$\hat{\beta}^s = \left(\sum_{i=1}^{n} Z_i^s(D_i^s)'I_i^s\right)^{-1}\sum_{i=1}^{n} Z_i^s Y_i I_i^s,$$

and

$$\hat{\Sigma}^s = \left(\sum_{i=1}^{n} Z_i^s(D_i^s)'I_i^s\right)^{-1}\left(\sum_{i=1}^{n} (\hat{\epsilon}_i^s)^2 Z_i^s(Z_i^s)'I_i^s\right)\left(\sum_{i=1}^{n} D_i^s(Z_i^s)'I_i^s\right)^{-1},$$

where $\hat{\epsilon}_i^s = Y_i - (D_i^s)'\hat{\beta}^s$. It is sufficient to show that

$$\hat{\beta}^s - \hat{\beta} = o_p(1),$$

if $S_n \to \infty$ and that

$$\sqrt{n\delta_n}(\hat{\beta}^s - \hat{\beta}) = o_p(1),$$

$$n\delta_n \hat{\Sigma}^s \overset{p}{\to} S_D^{-1}V(S_D')^{-1},$$

if Assumption 5 holds.

**Step C.3.4.1.** Let $\{V_i\}_{i=1}^{\infty}$ be i.i.d. random variables. If $E[V_i | X_i]$ and $E[V_i^2 | X_i]$ are bounded on $N(\delta \Omega, \delta') \cap N(\chi, \delta')$ for some $\delta' > 0$, and $S_n \to \infty$, then

$$\frac{1}{n\delta_n} \sum_{i=1}^{n} V_ip^s(X_i; \delta_n)'I_i^s - \frac{1}{n\delta_n} \sum_{i=1}^{n} V_ip^A(X_i; \delta_n)'I_i = o_p(1)$$

for $l = 0, 1, 2, 3, 4$. If, in addition, Assumption 5 holds, then

$$\frac{1}{\sqrt{n\delta_n}} \sum_{i=1}^{n} V_ip^s(X_i; \delta_n)'I_i^s - \frac{1}{\sqrt{n\delta_n}} \sum_{i=1}^{n} V_ip^A(X_i; \delta_n)'I_i = o_p(1)$$

for $l = 0, 1, 2$.

**Proof.** We have

$$\frac{1}{n\delta_n} \sum_{i=1}^{n} V_ip^s(X_i; \delta_n)'I_i^s - \frac{1}{n\delta_n} \sum_{i=1}^{n} V_ip^A(X_i; \delta_n)'I_i$$

$$= \frac{1}{n\delta_n} \sum_{i=1}^{n} V_ip^s(X_i; \delta_n)'(I_i^s - I_i) + \frac{1}{n\delta_n} \sum_{i=1}^{n} V_i(p^s(X_i; \delta_n)' - p^A(X_i; \delta_n))I_i.$$
We first consider \( \frac{1}{\eta n} \sum_{i=1}^{n} V_i(p^*(X_i; \delta_n)^3 - p^A(X_i; \delta_n)^3)I_i \). By using the argument in the proof of Step C.3.3.3 in Section C.3.3, we have

\[
E \left[ \frac{1}{\eta n} \sum_{i=1}^{n} V_i(p^*(X_i; \delta_n)^3 - p^A(X_i; \delta_n)^3)I_i \right]
\]

\( = \delta_n^{-1} E[V_i|X_i]E[p^*(X_i; \delta_n)^3 - p^A(X_i; \delta_n)^3]I_i \]

\( \leq \delta_n^{-1} E[V_i|X_i]|E[p^*(X_i; \delta_n)^3 - p^A(X_i; \delta_n)^3]|I_i \]

\( = \int_{-1}^{1} \int_{\partial \Omega \cap N(X_i, \delta)} |E[V_i|X_i = u + \delta_n \nu \nu \nu \nu (u)]|E[p^*(u + \delta_n \nu \nu \nu \nu (u); \delta_n)^3 - p^A(u + \delta_n \nu \nu \nu \nu (u); \delta_n)^3]|
\]

\( \times f_X(u + \delta_n \nu \nu \nu \nu (u)) \Omega_{p-1}(u, \delta_n \nu \nu \nu \nu) d\mathcal{H}^{p-1}(u)dv, \)

where the choice of \( \tilde{\delta} \) is as in the proof of Step C.3.3.3. By Lemma B.7, for \( l = 0, 1, 2, \)

\[
E \left[ \frac{1}{\eta n} \sum_{i=1}^{n} V_i(p^*(X_i; \delta_n)^l - p^A(X_i; \delta_n)^l)I_i \right]
\]

\( \leq \frac{1}{S_n} \int_{-1}^{1} \int_{\partial \Omega \cap N(X_i, \delta)} |E[V_i|X_i = u + \delta_n \nu \nu \nu \nu (u)]|f_X(u + \delta_n \nu \nu \nu \nu (u)) \Omega_{p-1}(u, \delta_n \nu \nu \nu \nu) d\mathcal{H}^{p-1}(u)dv
\]

\( = O(S_n^{-1}). \)

Also, by Lemma B.7,

\[
E \left[ \frac{1}{\eta n} \sum_{i=1}^{n} V_i(p^*(X_i; \delta_n)^3 - p^A(X_i; \delta_n)^3)I_i \right]
\]

\( = \delta_n^{-1} E[V_i(p^*(X_i; \delta_n)^3 - p^A(X_i; \delta_n)^3)]E[p^*(X_i; \delta_n) - p^A(X_i; \delta_n)]p^A(X_i; \delta_n) + p^A(X_i; \delta_n)^2)I_i \]

\( \leq \delta_n^{-1} E[V_i|X_i]E[p^*(X_i; \delta_n)^3 - p^A(X_i; \delta_n)]p^A(X_i; \delta_n) + p^A(X_i; \delta_n)^2)|X_i|I_i \]

\( \leq 3\delta_n^{-1} E[V_i|X_i]E[p^*(X_i; \delta_n)^3 - p^A(X_i; \delta_n)]|X_i|I_i \]

\( = \int_{-1}^{1} \int_{\partial \Omega \cap N(X_i, \delta)} |E[V_i|X_i = u + \delta_n \nu \nu \nu \nu (u)]|E[p^*(u + \delta_n \nu \nu \nu \nu (u); \delta_n)^3 - p^A(u + \delta_n \nu \nu \nu \nu (u); \delta_n)^3]|
\]

\( \times f_X(u + \delta_n \nu \nu \nu \nu (u)) \Omega_{p-1}(u, \delta_n \nu \nu \nu \nu) d\mathcal{H}^{p-1}(u)dv
\]

\( \leq \left( \frac{1}{S_n} \right)^{\epsilon} + O(1) \)

for every \( \epsilon > 0 \). We can make the right-hand side arbitrarily close to zero by taking sufficiently small \( \epsilon > 0 \) and sufficiently large \( S_n \), which implies that \( E[\frac{1}{\eta n} \sum_{i=1}^{n} V_i(p^*(X_i; \delta_n)^3 - p^A(X_i; \delta_n)^3)I_i] = o(1) \) if \( S_n \rightarrow \infty \). Likewise,

\[
E \left[ \frac{1}{\eta n} \sum_{i=1}^{n} V_i(p^*(X_i; \delta_n)^4 - p^A(X_i; \delta_n)^4)I_i \right]
\]

\( = \delta_n^{-1} E[V_i(p^*(X_i; \delta_n)^4 + p^A(X_i; \delta_n)^4)]E[p^*(X_i; \delta_n) + p^A(X_i; \delta_n)](p^*(X_i; \delta_n) - p^A(X_i; \delta_n))I_i \]

\( \leq \delta_n^{-1} E[V_i|X_i]E[p^*(X_i; \delta_n)^4 + p^A(X_i; \delta_n)^4)]E[p^*(X_i; \delta_n) + p^A(X_i; \delta_n)](p^*(X_i; \delta_n) - p^A(X_i; \delta_n))|X_i|I_i \]

\( \leq 8\delta_n^{-1} E[V_i|X_i]E[p^*(X_i; \delta_n) - p^A(X_i; \delta_n)]|X_i|I_i \]

\( = o(1). \)
As for variance, for \( l = 0, 1, 2, \)

\[
\text{Var}(\frac{1}{n\delta_n}\sum_{i=1}^{n} V_i(p^\delta(X_i; \delta_n)^l - p^A(X_i; \delta_n)^l)I_i) \leq \frac{1}{n\delta_n} \delta_n^{-1}E[V_i^2|X_i]E[(p^\delta(X_i; \delta_n)^l - p^A(X_i; \delta_n)^l)^2|X_i]I_i] \\
\leq \frac{1}{n\delta_n} \delta_n^{-1}E[V_i^2|X_i]E[(p^\delta(X_i; \delta_n)^l - p^A(X_i; \delta_n)^l)^2|X_i]I_i] \\
\leq \frac{4}{n\delta_n S_n}\delta_n^{-1}E[V_i^2|X_i]I_i] \\
= O((n\delta_n S_n)^{-1}),
\]

and for \( l = 3, 4, \)

\[
\text{Var}(\frac{1}{n\delta_n}\sum_{i=1}^{n} V_i(p^\delta(X_i; \delta_n)^l - p^A(X_i; \delta_n)^l)I_i) \leq \frac{1}{n\delta_n} \delta_n^{-1}E[V_i^2|X_i]E[(p^\delta(X_i; \delta_n)^l - p^A(X_i; \delta_n)^l)^2|X_i]I_i] \\
\leq \frac{1}{n\delta_n} \delta_n^{-1}E[V_i^2I_i] \\
= o(1).
\]

Therefore, \( \frac{1}{n\delta_n}\sum_{i=1}^{n} V_i(p^\delta(X_i; \delta_n)^l - p^A(X_i; \delta_n)^l)I_i = o_p(1) \) if \( S_n \to \infty \) for \( l = 0, 1, 2, 3, 4, \) and \( \frac{1}{\sqrt{n\delta_n}}\sum_{i=1}^{n} V_i(p^\delta(X_i; \delta_n)^l - p^A(X_i; \delta_n)^l)I_i \to o_p(1) \) if \( n^{-1/2} S_n \to \infty \) for \( l = 0, 1, 2. \)

We next show that \( \frac{1}{n\delta_n}\sum_{i=1}^{n} V_i p^\delta(X_i; \delta_n)^l (I_i^s - I_i) = o_p(1) \) if \( S_n \to \infty \) for \( l \geq 0. \) We have

\[
|E[\frac{1}{n\delta_n}\sum_{i=1}^{n} V_ip^\delta(X_i; \delta_n)^l (I_i^s - I_i)]| = \delta_n^{-1}|E[Vip^\delta(X_i; \delta_n)^l (I_i^s - I_i)]| \\
\leq \delta_n^{-1}E[|E[V_i|X_i]|E[p^\delta(X_i; \delta_n)^l (I_i^s - I_i)|X_i]|] \\
= \delta_n^{-1}E[|E[V_i|X_i]|E|I_i^s - I_i||X_i]|] .
\]

Since \( I_i^s - I_i \leq 0 \) with strict inequality only if \( I_i = 1, \)

\[
E[I_i^s - I_i||X_i|] = -E[I_i^s - I_i|X_i]I_i = (1 - E[I_i^s|X_i])I_i = \text{Pr}(p^\delta(X_i; \delta_n) \in \{0, 1\}|X_i)I_i.
\]

We then have

\[
|E[\frac{1}{n\delta_n}\sum_{i=1}^{n} V_ip^\delta(X_i; \delta_n)^l (I_i^s - I_i)]| \\
\leq \delta_n^{-1}E[|E[V_i|X_i]|\text{Pr}(p^\delta(X_i; \delta_n) \in \{0, 1\}|X_i)I_i] \\
\leq \delta_n^{-1}E[|E[V_i|X_i]|((1 - p^A(X_i; \delta_n))^{S_n} + p^A(X_i; \delta_n)^{S_n})I_i] \\
\leq \int_{-1}^{1} \int_{\partial \Omega^* \cap N(x, \hat{\delta})} |E[V_i|X_i = u + \delta_n v\psi_{\Omega^*}(u)]|\{(1 - p^A(u + \delta_n v\psi_{\Omega^*}(u); \delta_n))^{S_n} \\
+ p^A(u + \delta_n v\psi_{\Omega^*}(u); \delta_n)^{S_n}\} f_X(u + \delta_n v\psi_{\Omega^*}(u))J_{p-1}^{\text{arp}} \psi_{\Omega^*}(u, \delta_n, v) d\mathcal{H}^{p-1}(u) dv,
\]

where the second inequality follows from Lemma B.7. Note that for every \((u, v) \in \partial \Omega^* \cap N(x, \hat{\delta}) \times (-1, 1), \lim_{\delta \to 0} p^A(u + \delta_n v\psi_{\Omega^*}(u); \delta_n) = k(v) \in (0, 1)\) by Step C.3.3.1 in Section C.3.3. Since \( E[V_i|X_i], f_X \) and \( J_{p-1}^{\text{arp}} \psi_{\Omega^*} \) are bounded, by the Bounded Convergence Theorem,

\[
|E[\frac{1}{n\delta_n}\sum_{i=1}^{n} V_ip^\delta(X_i; \delta_n)^l (I_i^s - I_i)]| = o(1)
\]
if $S_n \to \infty$.

As for variance,

$$
\text{Var}\left(\frac{1}{n}\sum_{i=1}^{n} V_i p^s(X_i; \delta) (I_i^s - I_i)\right) \leq \frac{1}{n\delta_n} \delta_n^{-1} E[V_i^2 p^s(X_i; \delta)]^2 (I_i^s - I_i)^2 \\
\leq \frac{1}{n\delta_n} \delta_n^{-1} E[V_i^2 |I_i^s - I_i|] \\
= \frac{1}{n\delta_n} \delta_n^{-1} E[V_i^2 |X_i|] E[|I_i^s - I_i||X_i|] \\
= o(1).
$$

Lastly, we show that, for $l \geq 0$, $\frac{1}{\sqrt{n\delta_n}} \sum_{i=1}^{n} V_i p^s(X_i; \delta)(I_i^s - I_i) = o(1)$ if Assumption 5 holds. Let $\eta_n = \gamma \frac{\log n}{S_n}$, where $\gamma$ is the one satisfying Assumption 5. We have

$$
E\left[\frac{1}{\sqrt{n\delta_n}} \sum_{i=1}^{n} V_i p^s(X_i; \delta) (I_i^s - I_i)\right] \\
\leq \sqrt{n\delta_n} \sup_{x \in \mathcal{N}(\partial \Omega, 2\delta) \cap I} |E[V_i |X_i = x]| (\sqrt{n\delta_n} \Pr(p^A(X_i; \delta) \in (0, \eta_n) \cup (1 - \eta_n, 1)) \\
+ 2\sqrt{n\delta_n} (1 - \eta_n)^{S_n} \delta_n^{-1} E[1\{p^A(X_i; \delta) \in (\eta_n, 1 - \eta_n)\}]).
$$

By Assumption 5, $\sqrt{n\delta_n} \Pr(p^A(X_i; \delta) \in (0, \eta_n) \cup (1 - \eta_n, 1)) = o(1)$. For the second term,

$$
2\sqrt{n\delta_n} (1 - \eta_n)^{S_n} \delta_n^{-1} E[1\{p^A(X_i; \delta) \in (\eta_n, 1 - \eta_n)\}] \leq 2\sqrt{n\delta_n} (1 - \eta_n)^{S_n} \delta_n^{-1} E[I_i] \\
= 2\sqrt{n\delta_n} (1 - \eta_n)^{S_n} O(1).
$$

Observe that $\eta_n = \gamma \frac{\log n}{S_n} = \gamma \frac{\log n}{n^{1/2}S_n} \to 0$, since $n^{-1/2}S_n \to \infty$ and $\log n / n^{1/2} \to 0$. Using the fact that $e^t \geq 1 + t$ for every $t \in \mathbb{R}$, we have

$$
\sqrt{n\delta_n} (1 - \eta_n)^{S_n} \leq \sqrt{n\delta_n} (e^{\eta_n})^{S_n} \\
= \sqrt{n\delta_n} e^{-\eta_n S_n} \\
= \sqrt{n\delta_n} e^{-\gamma \log n} \\
= \sqrt{n\delta_n} e^{-\gamma} \\
= n^{1/2 - \gamma} \delta_n^{1/2} \\
\to 0,
$$

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since $\gamma > 1/2$. As for variance,
\[
\text{Var}\left( \frac{1}{\sqrt{n\delta_n}} \sum_{i=1}^{n} V_ip^s(X_i;\delta_n)^t(I_i^s - I_i) \right) \leq \delta_n^{-1} E\left[ V_i^2 p^s(X_i;\delta_n)^2(I_i^s - I_i)^2 \right] \\
\leq \delta_n^{-1} E\left[ E[V_i^2|X_i] E[|I_i^s - I_i||X_i]I_i \right] \\
= o(1).
\]

We have
\[
\hat{\beta} - \hat{\beta} = \left( \frac{1}{n\delta_n} \sum_{i=1}^{n} Z_i^s(D_i^s)'I_i^s \right)^{-1} \frac{1}{n\delta_n} \sum_{i=1}^{n} Z_i^sY_iI_i^s - \left( \frac{1}{n\delta_n} \sum_{i=1}^{n} Z_i^sD_i^s' \right)^{-1} \frac{1}{n\delta_n} \sum_{i=1}^{n} Z_iY_iI_i
\]
\[
= \left( \frac{1}{n\delta_n} \sum_{i=1}^{n} Z_i^s(D_i^s)'I_i^s \right)^{-1} \frac{1}{n\delta_n} \sum_{i=1}^{n} Z_i^sY_iI_i^s - \frac{1}{n\delta_n} \sum_{i=1}^{n} Z_iY_iI_i
\]
\[
- \left( \frac{1}{n\delta_n} \sum_{i=1}^{n} Z_i^s(D_i^s)'I_i^s \right)^{-1} \frac{1}{n\delta_n} \sum_{i=1}^{n} Z_i^s(D_i^s)'Y_i^s - \frac{1}{n\delta_n} \sum_{i=1}^{n} Z_iD_i^s'I_i^s \left( \frac{1}{n\delta_n} \sum_{i=1}^{n} Z_iD_i^s'I_i^s \right)^{-1} \frac{1}{n\delta_n} \sum_{i=1}^{n} Z_iY_iI_i.
\]

By Step C.3.4.1, $\hat{\beta} - \hat{\beta} = o_p(1)$ if $S_n \to \infty$, and $\sqrt{n}\delta_n(\hat{\beta} - \hat{\beta}) = o_p(1)$ if Assumption 5 holds.

By proceeding as in Step C.3.3.7 in Section C.3.3, we have
\[
\frac{1}{n\delta_n} \sum_{i=1}^{n} (\epsilon_i^s)^2Z_i^s(Z_i^s)'I_i^s = \frac{1}{n\delta_n} \sum_{i=1}^{n} (\epsilon_i^s)^2Z_i^s(Z_i^s)' + o_p(1),
\]
where $\epsilon_i^s = Y_i - (D_i^s)'\beta$. Then, by Step C.3.4.1,
\[
\frac{1}{n\delta_n} \sum_{i=1}^{n} (\epsilon_i^s)^2Z_i^s(Z_i^s)'I_i^s - \frac{1}{n\delta_n} \sum_{i=1}^{n} \epsilon_i^2Z_iZ_i' = \frac{1}{n\delta_n} \sum_{i=1}^{n} (Y_i^2 - 2Y_i\beta')D_i^s'\beta Z_i^s(Z_i^s)'I_i^s - \frac{1}{n\delta_n} \sum_{i=1}^{n} (Y_i^2 - 2Y_iD_i^s'\beta + \beta'D_i^sD_i^s'\beta)Z_iZ_i'I_i + o_p(1)
\]
\[
= o_p(1)
\]
so that
\[
\frac{1}{n\delta_n} \sum_{i=1}^{n} (\epsilon_i^s)^2Z_i^s(Z_i^s)'I_i^s \xrightarrow{p} V.
\]

Also, $\frac{1}{n\delta_n} \sum_{i=1}^{n} Z_i^s(D_i^s)'I_i^s \xrightarrow{p} S_D$ by using Step C.3.4.1. The conclusion then follows. \qed
C.4 Proof of Proposition A.2

With change of variables $u = \frac{x^* - x}{\delta}$, we have

$$p^A(x; \delta) = \frac{\int_{B(x, \delta)} A(x^*) dx^*}{\int_{B(x, \delta)} dx^*}$$

$$= \frac{\delta^p \int_{B(0, 1)} A(x + \delta u) du}{\delta^p \int_{B(0, 1)} du}$$

$$= \frac{\int_{\cup_{q \in Q} U_{x,q}} A(x + \delta u) du + \int_{B(0, 1)} \setminus \cup_{q \in Q} U_{x,q} A(x + \delta u) du}{\int_{B(0, 1)} du}$$

$$= \frac{\sum_{q \in Q} \int_{U_{x,q}} A(x + \delta u) du}{\int_{B(0, 1)} du},$$

where the last equality follows from the assumption that $\mathcal{L}^p(\cup_{q \in Q} U_{x,q}) = \mathcal{L}^p(B(0, 1))$. By the definition of $U_{x,q}$, for each $q \in Q$, $\lim_{\delta \to 0} A(x + \delta u) = q$ for any $u \in U_{x,q}$. By the Dominated Convergence Theorem,

$$p^A(x) = \lim_{\delta \to 0} p^A(x; \delta) = \frac{\sum_{q \in Q} q \mathcal{L}^p(U_{x,q})}{\mathcal{L}^p(B(0, 1))}.$$

The numerator exists, since $q \leq 1$ for all $q \in Q$ and $\sum_{q \in Q} \mathcal{L}^p(U_{x,q}) = \mathcal{L}^p(B(0, 1))$. \hfill \Box

C.5 Proof of Corollary A.1

1. Suppose that $A$ is continuous at $x \in \mathcal{X}$, and let $q = A(x)$. Then, by definition, $U_{x,q} = B(0, 1)$. By Proposition A.2, $p^A(x)$ exists, and $p^A(x) = q$. \hfill \Box

2. Pick any $x \in \text{int}(\mathcal{X}_q)$. $A$ is continuous at $x$, since there exists $\delta > 0$ such that $B(x, \delta) \subset \mathcal{X}_q$ by the definition of interior. By the previous result, $p^A(x)$ exists, and $p^A(x) = q$. \hfill \Box

3. Let $\mathcal{N}$ be the neighborhood of $x$ on which $f$ is continuously differentiable. By the mean value theorem, for any sufficiently small $\delta > 0$,

$$f(x + \delta u) = f(x) + \nabla f(\tilde{x}_\delta) \cdot \delta u$$

for some $\tilde{x}_\delta$ which is on the line segment connecting $x$ and $x + \delta u$. Since $\tilde{x}_\delta \to x$ as $\delta \to 0$ and $\nabla f$ is continuous on $\mathcal{N}$, $\nabla f(\tilde{x}_\delta) \cdot u \to \nabla f(x) \cdot u$ as $\delta \to 0$. Therefore, if $\nabla f(x) \cdot u > 0$, then $f(x + \delta u) = \nabla f(\tilde{x}_\delta) \cdot \delta u > 0$ for any sufficiently small $\delta > 0$, and if $\nabla f(x) \cdot u < 0$, then $f(x + \delta u) = \nabla f(\tilde{x}_\delta) \cdot \delta u < 0$ for any sufficiently small $\delta > 0$. We then have

$$U^+_x \equiv \{u \in B(0, 1) : \nabla f(x) \cdot u > 0\} \subset U_{x,q_1}$$

$$U^-_x \equiv \{u \in B(0, 1) : \nabla f(x) \cdot u < 0\} \subset U_{x,q_2}.$$
Let $V$ be the Lebesgue measure of a half $p$-dimensional unit ball. Since $V = L^p(U_1^+) \leq L^p(U_x^+) \leq L^p(U_x,q_1)$, and $L^p(U_x,q_1) + L^p(U_x,q_2) \leq L^p(B(0,1)) = 2V$, it follows that $L^p(U_x,q_1) = L^p(U_x,q_2) = V$. By Proposition A.2, $p^A(x)$ exists, and $p^A(x) = \frac{1}{2}(q_1 + q_2)$.

4. We have that $U_{0,q_1} = \{(u_1, u_2) \in B(0,1) : u_1 \leq 0 \text{ or } u_2 \leq 0\}$ and $U_{0,q_2} = \{(u_1, u_2) \in B(0,1) : u_1 > 0, u_2 > 0\}$. By Proposition A.2, $p^A(x)$ exists, and $p^A(x) = \frac{\sqrt{q_1 L^2(U_{0,q_1}) + q_2 L^2(U_{0,q_2})}}{L^2(B(0,1))} = \frac{3}{4}q_1 + \frac{1}{4}q_2$.

### C.6 Proof of Proposition A.1

Since $A$ is a $L^p$-measurable and bounded function, $A$ is locally integrable with respect to the Lebesgue measure, i.e., for every ball $B \subset \mathbb{R}^p$, $\int_B A(x)dx$ exists. An application of the Lebesgue differentiation theorem (see e.g. Theorem 1.4 in Chapter 3 of Stein and Shakarchi (2005)) to the function $A$ shows that
\[
\lim_{\delta \to 0} \frac{\int_{B(x,\delta)} A(x^*)dx^*}{\int_{B(x,\delta)} dx^*} = A(x)
\]
for almost every $x \in \mathbb{R}^p$.

### C.7 Proof of Proposition A.3

We can prove Part (a) using the same argument in the proof of Proposition 1 (a). For Part (b), suppose to the contrary that there exists $x_d \in X^S_d$ such that $L^p_c(\{x_c \in X^S_c(x_d) : p^A(x_d, x_c) \in \{0,1\}\}) > 0$. Without loss of generality, assume $L^p_c(\{x_c \in X^S_c(x_d) : p^A(x_d, x_c) = 1\}) > 0$. The proof proceeds in five steps.

**Step C.7.1.** $L^p_c(X^S_c(x_d)) \cap X^S_c(x_d) > 0$.

**Step C.7.2.** $X^S_c(x_d) \cap \text{int}(X^S_c(x_d)) \neq \emptyset$.

**Step C.7.3.** $p^A(x_d, x_c) = 1$ for any $x_c \in \text{int}(X^S_c(x_d))$.

**Step C.7.4.** For every $x^*_c \in X^S_c(x_d) \cap \text{int}(X^S_c(x_d))$, there exists $\delta > 0$ such that $B(x^*_c, \delta) \subset X^S_c(x_d) \cap \text{int}(X^S_c(x_d))$.

**Step C.7.5.** $E[Y_{i} - Y_{0}|X_i \in S]$ is not identified.

Following the argument in the proof of Proposition 1 (b), we can prove Steps C.7.1–C.7.3. Once Step C.7.4 is established, we prove Step C.7.5 by following the proof of Step C.1.4 in Proposition 1 (b) with $B(x^*_c, \delta)$ and $B(x^*_c, \epsilon)$ in place of $B(x^*, \delta)$ and $B(x^*, \epsilon)$, respectively, using the fact that $\Pr(X_{ci} \in B(x^*_c, \epsilon)|X_{di} = x_d) > 0$ by the definition of support. Here, we provide the proof of Step C.7.4.

**Proof of Step C.7.4.** Pick an $x^*_c \in X^S_c(x_d) \cap \text{int}(X^S_c)$. Then, $x^* = (x_d, x^*_c) \in S$. Since $S$ is open relative to $X$, there exists an open set $U \in \mathbb{R}^p$ such that $S = U \cap X$. This implies that for any sufficiently small $\delta > 0$, $B(x^*, \delta) \cap X \subset U \cap X = S$. It then follows that $\{x_c \in \mathbb{R}^p :$
(x_d, x_c) ∈ B(x^*, δ) ∩ X \subset \{x_c \in \mathbb{R}^p : (x_d, x_c) \in S\}$, equivalently, $B(x^*_c, δ) \cap X_c(x_d) \subset X^S_c(x_d)$. By choosing a sufficiently small $δ > 0$ so that $B(x^*_c, δ) \subset \text{int}(X_c, x_d) \subset X_c(x_d)$, we have $B(x^*_c, δ) \subset X^S_c(x_d) \cap \text{int}(X_c, x_d)$.

C.8 Proof of Theorem A.1

The proof is analogous to the proof of Theorem 1. The only difference is that, when we prove the convergence of expectations, we show the convergence of the expectations conditional on $X_{di}$, and then take the expectations over $X_{di}$.

D Machine Learning Simulation: Details

Parameter Choice. For the variance-covariance matrix $\Sigma$ of $X_i$, we first create a $100 \times 100$ symmetric matrix $V$ such that the diagonal elements are one, $V_{ij}$ is nonzero and equal to $V_{ji}$ for $(i, j) \in \{2, 3, 4, 5, 6\} \times \{35, 66, 78\}$, and everything else is zero. We draw values from Unif($-0.5, 0.5$) independently for the nonzero off-diagonal elements of $V$. We then create matrix $\Sigma = V \times V$, which is a positive semidefinite matrix.

For $\alpha_0$ and $\alpha_1$, we first draw $\tilde{\alpha}_{0j}$, $j = 51, ... , 100$, from Unif($-100, 100$) independently across $j$, and draw $\tilde{\alpha}_{1j}$, $j = 1, ... , 100$, from Unif($-150, 200$) independently across $j$. We then set $\tilde{\alpha}_{0j} = \tilde{\alpha}_{1j}$ for $j = 1, ... , 50$, and calculate $\alpha_0$ and $\alpha_1$ by normalizing $\tilde{\alpha}_0$ and $\tilde{\alpha}_1$ so that $\text{Var}(X'_i\alpha_0) = \text{Var}(X'_i\alpha_1) = 1$.

Training of Prediction Model. We first randomly split the sample $\{(\tilde{Y}_i, \tilde{X}_i, \tilde{D}_i, \tilde{Z}_i)\}_{i=1}^{\tilde{n}}$ into train (80%) and test datasets (20%). We use random forests on the training sample to obtain the prediction model $\mu_z$ and validate its performance on the test sample. The trained algorithm has an accuracy of 97% on the test data.

E Empirical Policy Application: Details

E.1 Hospital Cost Data

We use publicly available Healthcare Cost Report Information System (HCRIS) data,\textsuperscript{41} to project\textsuperscript{42} safety net eligibility and funding amounts for all hospitals in the dataset. This data set contains information on various hospital characteristics including utilization, number of employees, medicare cost data and financial statement data.

The data is available from financial year 1996 to 2019. As the coverage is higher for 2018 (compared to 2019), we utilize the data corresponding to the 2018 financial year. Hospitals are uniquely identified in a financial year by their CMS (Center for Medicaid and Medicare Services)...

\textsuperscript{41}We use the RAND cleaned version of this dataset, which can be accessed https://www.hospitaldatasets.org/

\textsuperscript{42}We use the methodology detailed in the CARES ACT website to project funding based on 2018 financial year cost reports.
Certification Number. We have data for 4,705 providers for the 2018 financial year. We focus on 4,648 acute care and critical access hospitals that are either located in one of the 50 states or Washington DC.

**Disproportionate patient percentage**

Disproportionate patient percentage is equal to the percentage of Medicare inpatient days attributable to patients eligible for both Medicare Part A and Supplemental Security Income (SSI) summed with the percentage of total inpatient days attributable to patients eligible for Medicaid but not Medicare Part A.\(^{43}\) In the data, this variable is missing for 1560 hospitals. We impute the disproportionate patient percentage to 0 when it is missing.

**Uncompensated care per bed**

Cost of uncompensated care refers to the care provided by the hospital for which no compensation was received from the patient or the insurer. It is the sum of a hospital’s bad debt and the financial assistance it provides.\(^{44}\) The cost of uncompensated care is missing for 86 hospitals, which we impute to 0. We divide the cost of uncompensated care by the number of beds in the hospital to obtain the cost per bed. The data on bed count is missing for 15 hospitals, which we drop from the analysis, leaving us with 4,633 hospitals in 2,473 counties.

**Profit Margin**

Hospital profit margins are indicative of the financial health of the hospitals. We calculate profit margins as the ratio of net income to total revenue where total revenue is the sum of net patient revenue and total other income. After the calculation, profit margins are missing for 92 hospitals, which we impute to 0.

**Funding**

We calculate the projected funding using the formula on the CARES ACT website. Hospitals that do not qualify on any of the three dimensions are not given any funding. Each eligible hospital is assigned an individual facility score, which is calculated as the product of disproportionate patient percentage and number of beds in that hospital. We calculate cumulative facility score as the sum of all individual facility scores in the dataset. Each hospital receives a share of $10 billion, where the share is determined by the ratio of individual facility score of that hospital to the cumulative facility score. The amount of funding received by hospitals is bounded below at $5 million and capped above at $50 million.

\(^{43}\)For the precise definition, see https://www.cms.gov/Medicare/Medicare-Fee-for-Service-Payment/AcuteInpatientPPS/dsh.

\(^{44}\)The precise definition can be found at https://www.aha.org/fact-sheets/2020-01-06-fact-sheet-uncompensated-hospital-care-cost.
E.2 Hospital Utilization Data

We use the publicly available COVID-19 Reported Patient Impact and Hospital Capacity by Facility dataset for our outcome variables. This provides facility level data on hospital utilization aggregated on a weekly basis, from July 31st onwards. These reports are derived from two main sources – (1) HHS TeleTracking and (2) reporting provided directly to HHS Protect by state/territorial health departments on behalf of health care facilities.\(^{45}\)

The hospitals are uniquely identified for a given collection week (which goes from Friday to Thursday) by their CMS Certification number. All hospitals that are registered with CMS by June 1st 2020 are included in the population. We merge the hospital cost report data with the utilization data using the CMS certification number. According to the terms and conditions of the CARES Health Care Act, the recipients may use the relief funds only to “prevent, prepare for, and respond to coronavirus” and for “health care related expenses or lost revenues that are attributable to coronavirus”. Therefore, for our analysis we focus on 4 outcomes that were directly affected by COVID-19, for the week spanning July 31st to August 6th 2020. The outcome measures are described below.\(^{46}\)

1. Total reports of patients currently hospitalized in an adult inpatient bed who have laboratory-confirmed or suspected COVID-19, including those in observation beds reported during the 7-day period.

2. Total reports of patients currently hospitalized in an adult inpatient bed who have laboratory-confirmed COVID-19 or influenza, including those in observation beds. Including patients who have both laboratory-confirmed COVID-19 and laboratory confirmed influenza during the 7-day period.

3. Total reports of patients currently hospitalized in a designated adult ICU bed who have suspected or laboratory-confirmed COVID-19.

4. Total reports of patients currently hospitalized in a designated adult ICU bed who have laboratory-confirmed COVID-19 or influenza, including patients who have both laboratory-confirmed COVID-19 and laboratory-confirmed influenza.\(^{47}\)

---


\(^{46}\)We conduct sanity checks and impute observations to missing if they fail our checks. For example, we impute the value # Confirmed/ Suspected COVID Patients and # Confirmed COVID Patients to missing when the latter is greater than the former. # Confirmed/ Suspected COVID Patients should be greater than or equal to # Confirmed COVID Patients as the former includes the latter. Similarly, we impute # Confirmed/ Suspected COVID Patients in ICU and # Confirmed COVID Patients in ICU to be missing when the latter is greater than the former.

\(^{47}\)In the dataset, when the values of the 7 day sum are reported to be less than 4, they are replaced with -999,999. We recode these values to be missing. The results in Table 4 remain almost the same even if we impute the suppressed values (coded as -999,999) with 0s. Results are available upon request.
E.3 Computing Approximate Propensity Score

As the three determinants of safety net eligibility are continuous variables, we can think of this setting as a multi-dimensional regression discontinuity design and a suitable setting to apply our method. In this setting, the $X_i$ are disproportionate patient percentage, uncompensated care per bed and profit margin. Funding eligibility ($Z_i$) is determined algorithmically using these 3 dimensions. $D_i$ is the amount of funding received by hospital $i$, which depends on both safety net eligibility status $Z_i$, number of beds in the hospital, and disproportionate patient percentage. Before calculating APS, we normalize each characteristic of $X_i$ to have mean 0 and variance 1. For each hospital and every $\delta \in \{0.01, 0.025, 0.05, 0.075, 0.1, 0.25, 0.5\}$, we draw 10,000 times from a $\delta$-ball around the normalized covariate space and calculate APS by averaging funding eligibility $Z_i$ over these draws.

E.4 Additional Empirical Results

Figure A.2: APS Estimation with Varying Simulations $S$

Notes: The above figure plots the APS estimates for 10 randomly selected hospitals along the eligibility margin for varying numbers of simulations $S$. Each line represents a different hospital. The dotted line at $10^4$ indicates the number of simulations we use for our main analysis.
Table A.1: Differential Attrition

<table>
<thead>
<tr>
<th>Ineligible Hospitals</th>
<th>No Controls</th>
<th>$\delta = 0.01$</th>
<th>$\delta = 0.025$</th>
<th>$\delta = 0.05$</th>
<th>$\delta = 0.075$</th>
<th>$\delta = 0.1$</th>
<th>$\delta = 0.25$</th>
<th>$\delta = 0.5$</th>
</tr>
</thead>
<tbody>
<tr>
<td>#Confirmed/Suspected Covid Patients</td>
<td>.745</td>
<td>38.19***</td>
<td>-15.51</td>
<td>-24.80</td>
<td>-44.34</td>
<td>-57.95</td>
<td>-40.34</td>
<td>2.05</td>
</tr>
<tr>
<td></td>
<td>(8.55)</td>
<td>(85.67)</td>
<td>(70.81)</td>
<td>(70.09)</td>
<td>(63.06)</td>
<td>(48.58)</td>
<td>(22.20)</td>
<td>(15.67)</td>
</tr>
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<td>N=3532</td>
<td>N=73</td>
<td>N=195</td>
<td>N=392</td>
<td>N=547</td>
<td>N=719</td>
<td>N=1389</td>
<td>N=1947</td>
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<tr>
<td>#Confirmed Covid Patients</td>
<td>.754</td>
<td>33.97***</td>
<td>0.85</td>
<td>-30.81</td>
<td>21.32</td>
<td>1.96</td>
<td>-0.39</td>
<td>-1.28</td>
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<tr>
<td></td>
<td>(7.44)</td>
<td>(73.28)</td>
<td>(55.22)</td>
<td>(33.46)</td>
<td>(29.41)</td>
<td>(25.14)</td>
<td>(15.75)</td>
<td>(12.56)</td>
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<tr>
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<td>N=3558</td>
<td>N=70</td>
<td>N=191</td>
<td>N=385</td>
<td>N=539</td>
<td>N=709</td>
<td>N=1366</td>
<td>N=1923</td>
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<tr>
<td>#Confirmed/Suspected Covid Patients in ICU</td>
<td>.728</td>
<td>13.18***</td>
<td>13.68</td>
<td>9.54</td>
<td>5.71</td>
<td>-0.83</td>
<td>2.34</td>
<td>-0.46</td>
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<tr>
<td></td>
<td>(2.74)</td>
<td>(23.41)</td>
<td>(17.74)</td>
<td>(11.91)</td>
<td>(10.68)</td>
<td>(9.01)</td>
<td>(5.78)</td>
<td>(4.64)</td>
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<td>N=186</td>
<td>N=374</td>
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<td>N=1314</td>
<td>N=1846</td>
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<tr>
<td>#Confirmed Covid Patients in ICU</td>
<td>.744</td>
<td>12.16***</td>
<td>7.97</td>
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<td>0.65</td>
<td>1.87</td>
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<td>(2.58)</td>
<td>(25.63)</td>
<td>(18.89)</td>
<td>(11.25)</td>
<td>(9.97)</td>
<td>(8.52)</td>
<td>(5.57)</td>
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<td>N=1321</td>
<td>N=1868</td>
</tr>
</tbody>
</table>

Notes: This table reports differential safety net eligibility effects on the availability of outcome data at the hospital level. Column 1 presents the average of the availability indicators of the outcome variables for the ineligible hospitals. In column 2, we regress the availability indicator on dummy for safety net eligibility without any controls. In columns 3-9, we run this regression controlling for the Approximate Propensity Score with different values of bandwidth $\delta$ on the sample with nondegenerate Approximate Propensity Score. All Approximate Propensity Scores are computed by averaging 10,000 simulation draws. The outcome variables are the 7 day totals for the week spanning July 31st, 2020 to August 6th, 2020. Confirmed or Suspected COVID patients refer to the sum of patients in inpatient beds with lab-confirmed/suspected COVID-19. Confirmed COVID patients refer to the sum of patients in inpatient beds with lab-confirmed COVID-19, including those with both lab-confirmed COVID-19 and influenza. Inpatient bed totals also include observation beds. Similarly, Confirmed/Suspected COVID patients in ICU refer to the sum of patients in ICU beds with lab-confirmed or suspected COVID-19. Confirmed COVID patients in ICU refers to the sum of patients in ICU beds with lab-confirmed COVID-19, including those with both lab-confirmed COVID-19 and influenza. Robust standard errors are reported in parenthesis.