

Optimal Stratification of Survey Experiments*

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Abstract

This paper studies treatment effect estimation in a two-stage model of experimentation. In the first stage, using baseline covariate information, the researcher draws a representative sample of experimental participants from a pool of eligible units. In the second stage, they assign each sampled participant to treatment or control. To implement representative sampling and balanced assignment, we introduce a new family of finely stratified randomization procedures, generalizing matched pairs designs to propensities $p(x) \neq 1/2$. When used for representative sampling, our method nonparametrically dampens the estimator variance due to treatment effect heterogeneity. When used for balanced assignment, our designs make simple difference-of-means estimation asymptotically equivalent to well-specified nonparametric regression adjustment, effectively doing nonparametric regression “by design.” Building on these results, we formalize and solve the two-stage optimal stratification problem with heterogeneous costs and fixed budget constraint, providing simple heuristics for the optimal design. We apply these results to the question of design using a pilot study, showing that implementing a pilot estimate of the optimal design is asymptotically efficient, achieving the minimal variance subject to the budget constraint. We also provide novel asymptotically exact inference methods over this entire class of designs, fully exploiting the efficiency gains from both finely stratified sampling and assignment.

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1 Introduction

Randomized controlled trials (RCTs) are increasingly common in economics research. For example, the AEA RCT registry currently lists over 7000 active experiments, spanning a wide range of subfields. Much of the literature in experimental design attempts to reduce the variance of treatment effect estimators by balancing observed covariates between the treatment and control units at design-time. This paper contributes to theory of covariate-adaptive assignment, but also models a new dimension of experimental design: the selection of the experimental participants.

In practice, the selection of participants (sampling frame) is an important part of designing an experiment. For example, [Abaluck et al. \(2021\)](#) run an experiment to estimate the effect of mask distribution on village-level covid infection rates in Bangladesh. From a pool of 1000 eligible villages, they first randomly sample 600 to be included in the experiment, then assign the sampled villages to various interventions that promote mask usage. Similarly, [Breza et al. \(2021\)](#) estimate the effect of Facebook ads discouraging holiday travel on county-level covid infection rates. Since their budget for running ads is finite, they first sample a small set of counties in which to run ads and collect outcome data, then randomly assign these sampled counties to low or high intensity of treatment. We show how to increase the efficiency of treatment effect estimation by sampling experimental units that are representative of the broader target population, and provide new inference methods that take full advantage of these efficiency gains.

To do so, this paper introduces a new family of finely stratified randomization procedures that can be used for both representative sampling of the experimental units as well finely balanced assignment of treatments. When used for treatment assignment, our method generalizes the principle of matched pairs randomization to arbitrary assignment propensities $p(x) \neq 1/2$, allowing discrete or continuous covariates in general dimension. The basic building block of our method is a new algorithm that matches the experimental units into homogeneous k -tuples, using an objective function directly linked to estimation efficiency. In addition to generalizing matched pairs, this also allows for representative sampling of the experimental participants using data-adaptive fine stratification. For example, suppose 1000 people respond to an advertisement to participate in an experiment, but our budget only allows for 400 participants. Using observed covariate information, we match the units into homogeneous 5-tuples, sampling $q = 2/5$ of the units in each tuple to participate, uniformly at random. This procedure reduces estimator variance by finely representing the distribution of treatment effect heterogeneity in the population into our smaller experiment. More generally, we provide finely stratified designs implementing heterogeneous sampling propensities $q(x)$.

We study a two-stage procedure that first samples participants then assigns treatments to the sampled units, using finely stratified randomization at both stages. We show that under finely stratified treatment assignment alone, the difference of means estimator $Y \sim 1 + D$ achieves the [Hahn \(1998\)](#) variance bound for the average treatment effect (ATE), effectively doing nonparametric regression adjustment “by design.” Finely stratified sampling provides an additional nonparametric efficiency improvement, reducing the variance due to treatment effect heterogeneity. In particular, representative sampling makes this component of the variance scale with the larger number of units that we

sample from, rather than the smaller true experiment size. We characterize the optimal stratification variables for sampling and assignment, showing that for sampling one wants to stratify on covariates that are most predictive of treatment effect heterogeneity. On the technical side, our asymptotic analysis deals with new complications introduced by joint finely stratified sampling and assignment with non-constant propensities. We prove fast rates for k-tuples matching, improving the rates currently known for matched pairs (Bai et al. (2021)). We also remove many of the technical assumptions previously used in the literature, proving our results under a simple moment condition on the outcomes and covariates. In the appendix, we also study estimation of the sample average treatment effect (SATE) over the eligible population, extending our results using design-based asymptotics.

Building on these asymptotic results, we formalize and solve an optimal design problem with fixed budget and heterogeneous costs. In development economics, for instance, the cost of including a village in an experiment can vary widely based on observable characteristics like its distance from the urban center, village size and so on, forcing applied researchers to choose a tradeoff between sample size and sample representativeness. We formalize this tradeoff, deriving the optimal sampling intensity $q^*(x)$ and assignment propensity $p^*(x)$ under finely stratified randomization. In case of homoskedasticity, for instance, the optimal sampling propensity $q^*(x) \propto C(x)^{-1/2}$, where $C(x)$ is the cost of including a unit of type $X = x$ in the experiment. Our results provide simple heuristics for optimal sampling of experimental participants, analogous to classical results on sample allocation for coarsely stratified survey design (Cochran (1977)). We show that an oracle design that implements discretizations of $q^*(x)$ and $p^*(x)$ using fine stratification minimizes the asymptotic variance of our estimator over all stratified designs, subject to the budget constraint.

We also briefly investigate exact optimality in finite samples. For propensity $p = 1/2$, we prove that the globally optimal covariate-adaptive randomization takes the form of an “alternating design”, assigning a certain optimal allocation vector $(d_i^*)_{i=1}^n$ and its mirror image $(1 - d_i^*)_{i=1}^n$ each with probability $1/2$. The optimal allocation $(d_i^*)_{i=1}^n$ solves the well-known Max-Cut graph partitioning problem (Rendl et al. (2008)), with edge weights related to the smoothness of the outcome functions.

As an application of our results, we study the problem of design using data from a pilot experiment, making use of our analytical characterization of the budget-constrained optimal stratified design. The basic idea is to estimate the optimal sampling and assignment propensities $q^*(x)$ and $p^*(x)$ using the pilot, then implement these estimates in the main experiment using fine stratification. Under large pilot asymptotics, this strategy is as efficient as the oracle design, achieving the budget-constrained minimal asymptotic variance. We also provide results under fixed pilot asymptotics, and briefly discuss potential robustifications. In the case without sampling, the problem of design using a pilot has received considerable attention in the recent literature, see for instance Hahn et al. (2011), Tabord-Meehan (2022), and Bai (2022). We derive the first fully efficient solution to this problem, asymptotically minimizing the semiparametric variance bound over all propensity scores.

Finally, we provide novel asymptotically exact inference methods over our entire family of

designs, allowing applied researchers to fully exploit the efficiency gains from finely stratified sampling and assignment with heterogeneous propensities. To do so, we generalize the method of collapsed strata (Hansen et al. (1953)) to two-stage designs with data-adaptive finely stratified sampling and assignment. See also Abadie and Imbens (2008) and Bai et al. (2021) for related results in the context of inference under matched pairs randomization. At a theoretical level, the use of non-constant sampling proportions $q(x)$ produces discontinuities in the propensity-weighted outcome functions, introducing new technical complications that we overcome with our analysis. The basic idea is to prove variance estimator consistency assuming the relevant functions are Lipschitz continuous, then appeal to density of Lipschitz functions in $L_2(X)$.

1.1 Related Literature

Our sampling model is related to the classical literature on survey sampling, e.g. as in Cochran (1977) and Lohr (2021). In contemporaneous work, Yang et al. (2021) propose a two-stage design using rerandomization for both sampling and assignment. Under rerandomization, difference of means estimation is asymptotically (almost) as efficient than ex-post linear adjustment. By contrast, we show that data-adaptive fine stratification is asymptotically equivalent to nonparametric adjustment for the imbalances in both sampling and assignment variables, see Proposition 3.11 for a formal statement.

For an overview of experimental design theory, see Rosenberger and Lachin (2016) or Athey and Imbens (2017). A representative sample of recent work on stratified treatment assignment includes Imai et al. (2009), Bugni et al. (2018), Fogarty (2018), Wang et al. (2021), Bai et al. (2021), de Chaisemartin and Ramirez-Cuellar (2021), Bai (2022), and Tabord-Meehan (2022). For treatment assignment, our work is most related to Bai (2022), who introduces finely stratified designs for constant propensity $p = a/k$ and univariate stratification variables. Aside from stratification, other recent proposals for balanced treatment assignment include Kasy (2016), Kallus (2017), Li et al. (2018), Krieger et al. (2019), and Harshaw et al. (2021). We explicitly compare with some of these methods in Remark 3.4.

Our results on design using a pilot study is related to previous results in Hahn et al. (2011), Bai (2022), and Tabord-Meehan (2022), as well as Kasy and Sautmann (2021). We provide detailed comparisons in Section 5 below. Our inference results are related to the method of collapsed strata in Hansen et al. (1953) and its modern variants studied in Abadie and Imbens (2008) and Bai et al. (2021). The rest of the paper is organized as follows. Section 2 introduces notation and discusses our matching algorithm. Section 3 states our main asymptotic results, including the equivalence with nonparametric adjustment. Section 4 formalizes and solves the optimal design problem. Section 5 discusses design using a pilot study. Section 6 provides our inference methods. Our empirical results are presented in Section 7.

2 Motivation and Description of Method

Consider running an experiment to estimate the average treatment effect (ATE). There are n eligible units, with observed baseline covariates $(X_i)_{i=1}^n$. We wish to sample proportion $q \in (0, 1]$ of these units to participate in the experiment. Denote $T_i = 1$ if a

unit is sampled and $T_i = 0$ otherwise. Sampled units are then assigned to treatment or control $D_i \in \{0, 1\}$. Let $Y_i(d)$ denote the potential outcome of unit i for $d \in \{0, 1\}$. Since outcomes are only observed for participating units¹ we write

$$Y_i = T_i[D_i Y_i(1) + (1 - D_i)Y_i(0)]$$

We focus on estimation and inference for the population ATE $= E[Y(1) - Y(0)]$, modeling the n eligible units as an independent sample from a superpopulation of interest with $(X_i, Y_i(0), Y_i(1))_{i=1}^n \sim F$ iid. For example, the eligible units could be $n = 1000$ respondents to an online advertisement to participate in an experiment with a budget constraint of 100 participants. In this context, the variable $T_i \in \{0, 1\}$ models which of the $n = 1000$ units we choose to include in the experiment. In some applications, the eligible units may comprise the entire population of interest. For example, the units $i = 1, \dots, n$ may be the entire population of villages in a country, which we sample to obtain the participating villages. To accommodate such applications, Section 9.1 in the appendix presents design-based versions of our results, targeting the sample average treatment effect SATE $= E_n[Y_i(1) - Y_i(0)]$ in the eligible population.²

Our goal is to sample a representative subset of the eligible units and assign them to treatment and control in a way that finely balances the baseline covariates $(X_i)_{i=1}^n$. To do so, we introduce a large new family of finely stratified designs, generalizing the idea of matched pairs randomization to arbitrary propensity scores $p(x) = P(D = 1|X = x)$, with continuous or discrete covariates in general dimension. We also use these new designs for finely stratified sampling, allowing us to implement arbitrary sampling propensities $q(x) = P(T = 1|X = x)$ in a finely balanced way. The basic building block of our method is a matched k -tuples design, which uses the baseline covariates to match units into homogeneous groups of k , randomly assigning a out of k units in each group to $T = 1$ during sampling or $D = 1$ during assignment. We introduce the method in the context of finely stratified sampling in the next definition.

Definition 2.1 (Local Randomization). Let $q = a/k$ with $\gcd(a, k) = 1$. Partition the eligible units into groups with $|g| = k$, so that $\{1, \dots, n\} = \bigcup_g \{i \in g\}$ disjointly. In general, there may be one remainder group with $|g| < k$. Let $\psi(X) \in \mathbb{R}^d$ be a vector of stratification variables, and suppose that the groups are homogeneous in the sense that

$$n^{-1} \sum_g \sum_{i, j \in g} |\psi(X_i) - \psi(X_j)|_2^2 = o_p(1) \quad (2.1)$$

Require that the groups only depend on the stratification variable values $\psi_{1:n} = (\psi(X_i))_{i=1}^n$, and data-independent randomness π_n , so that $g = g(\psi_{1:n}, \pi_n)$. Independently over all groups with $|g| = k$, draw sampling variables $(T_i)_{i \in g}$ by setting $T_i = 1$ for exactly a out of k units, completely at random. For units in the remainder group with $|g| < k$, draw T_i iid with $P(T_i = 1) = a/k$. We say that such a design implements q locally with respect to $\psi(x)$, denoting $T_{1:n} \sim \text{Loc}(\psi, q)$.

Equation 2.1 generalizes a similar condition in Bai et al. (2021) for the case of matched

¹If control outcomes are costlessly observed for all units, the sampling problem becomes trivial. We still contribute novel assignment designs in this case.

²Note that here the SATE is defined over the entire eligible population $i \in [n]$, not just the smaller set of experiment participants $\{i : T_i = 1\}$.

pairs $k = 2$. We discuss matching algorithms and their associated homogeneity rates in Section 2.1 below. Consider sampling and assignment propensities $q = a/k$ and $p = a'/k'$. In the rest of the paper, we study treatment effect estimation under a two-stage procedure:

- (1) Sample eligible units $T_{1:n} \sim \text{Loc}(\psi, q)$.
- (2) Assign treatments $D_{1:n} \sim \text{Loc}(\psi, p)$ to the sampled units $\{i : T_i = 1\}$.

Section 3.2 presents the most general version of the method, allowing different stratification variables ψ_1 and ψ_2 to be used for sampling and assignment, as well as varying sampling and assignment propensities $q(x)$, $p(x)$.

Example 2.2 (Matched Tuples). Suppose $n = 1000$ individuals from a target population sign up to participate in an experiment, providing basic demographic information $(X_i)_{i=1}^n$. There are only resources for 300 units to be enrolled, so $q = 3/10$. Among these 300 units, $p = 1/4$ will be assigned to the more costly treatment and $3/4$ to control. We sample using the design $T_{1:n} \sim \text{Loc}(\psi, 3/10)$, which matches the 1000 eligible units into homogeneous groups of 10 and randomly sets $T_i = 1$ for 3 out of 10 units in each group. We assign treatments $D_{1:n} \sim \text{Loc}(\psi, 1/4)$ to the $\sum_i T_i = 300$ sampled units, matching them into homogeneous tuples of four and assigning 1 out of 4 in each tuple to $D_i = 1$.

Remark 2.3 (Coarse Stratification). The procedure in Definition 2.1 produces n/k groups of k units that are tightly matched in $\psi(X)$ space, suggesting fine stratification. However, “coarse stratification” designs with a fixed number of large strata are also included in this framework. For example, we may set $\psi = 1$ and form groups $|g| = 3$ at random, automatically satisfying Equation 2.1. Assigning 2 out of 3 units in each group to treatment gives a “random matched triples” representation of complete randomization with $p = 2/3$. Similarly, coarse stratification³ with fixed strata $S(X) \in \{1, \dots, m\}$ can be obtained by setting $\psi(X) = S(X)$ and matching units with the same $S(X)$ value into groups at random. Because of this, our framework allows for unified asymptotics and inference procedures for a wide range of stratifications.

2.1 Algorithms

One possibility is to treat the left hand side of Equation 2.1 as an objective function and minimize it over all partitions of the units into groups of k . Denote $d(\psi) = \dim(\psi)$. Theorem 10.1 shows that if $E[|\psi(X)|_2^\alpha] < \infty$ for $\alpha > d(\psi) + 1$ then the optimal groups

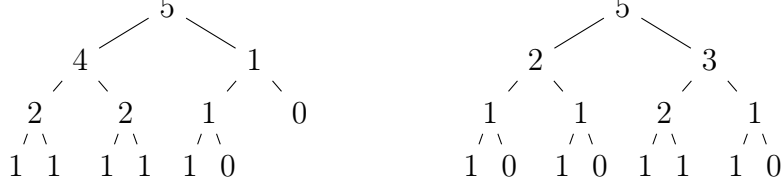
$$\min_{(g)} n^{-1} \sum_g \sum_{i,j \in g} |\psi_i - \psi_j|_2^2 = O_p(n^{2/\alpha - 2/(d(\psi)+1)}) = o_p(1) \quad (2.2)$$

If $\psi(X)$ is bounded this becomes $O_p(n^{-2/(d(\psi)+1)})$, sharpening the rate achieved under a boundedness assumption in previous work on matched pairs. For $k = 2$, the optimal groups are computable in $O(n^3)$ time using an algorithm due to Derigs (1988). Efficient algorithms for computing the optimal groups for general k are not available, and we expect the problem to be NP-hard.⁴

³Coarse stratification was previously studied using different methods in Bugni et al. (2018). We extend their results, e.g. allowing coarse stratification at both sampling and assignment stages.

⁴See Karmakar (2022) for hardness results in a related problem.

Instead of calculating the optimal partition, for $k > 2$ we iteratively apply Derigs' algorithm to match units into larger groups. For fixed k , this procedure can be shown to satisfy the same rate in Equation 2.2 above. There are many ways to implement iterative pairing for each k . For example, $k = 5$ can be obtained by pairwise matching of 4-tuples to 1-tuples, or 3-tuples to 2-tuples and so on. This is shown in the figure below, where each level of the tree represents a call to the optimal pairing algorithm.



Before the j th algorithm call, we add a certain number of “empty centroids”, represented by the 0’s in the figure. We also prohibit certain types of matches in order to guarantee the desired sequence of group sizes. For example, at the second level of the tree on the left, we set the distance to $+\infty$ between groups of size $|g| = 2$ and $|g'| = 1$, size $|g| = 1$ and $|g'| = 1$, and size $|g| = 0$ and $|g'| = 0$. There are many choices of cardinality trees for each k , not all of which can be feasibly implemented using such constraints. We provide two canonical ways of generating such sequences, their associated constraints, and the number of empty groups (zeros) to add at each level.

3 Asymptotics and Optimal Designs

This section contains our main asymptotic results, showing the nonparametric efficiency gains from finely stratified sampling and assignment. First, we state our main assumption.

Assumption 3.1. *The moments $E[Y(d)^2] < \infty$ for $d = 0, 1$ and $E[|\psi(X)|_2^\alpha] < \infty$ hold for some $\alpha > \dim(\psi) + 1$.*

Previous work on fine stratification has imposed Lipschitz continuity on the outcome function $E[Y(d)|\psi(X) = \psi]$ and variance $\text{Var}(Y(d)|\psi(X) = \psi)$, as well as boundedness of the stratification variables $\psi(X)$. Our analysis shows that these assumptions can be removed.

Estimation. Let $\hat{\theta}$ be the regression coefficient on D_i in $Y_i \sim 1 + D_i$, estimated in the sampled units $\{i : T_i = 1\}$. This is just the usual difference-of-means estimator. Before continuing to our asymptotic results, we state a variance decomposition that will be used extensively in what follows. Let $c(\psi) = E[Y(1) - Y(0)|\psi(X)]$ denote the conditional average treatment effect (CATE) and $\sigma_d^2(\psi) = \text{Var}(Y(d)|\psi(X))$ the heteroskedasticity function. Define the *balance function*

$$b(\psi; p) = E[Y(1)|\psi(X)] \left(\frac{1-p}{p} \right)^{1/2} + E[Y(0)|\psi(X)] \left(\frac{p}{1-p} \right)^{1/2}. \quad (3.1)$$

Suppose that sampling and assignment are both completely randomized, $T_{1:n} \sim \text{CR}(q)$ and $D_{1:n} \sim \text{CR}(p)$. Let $n_T = \sum_i T_i$ denote the experiment size. Our work shows that $\sqrt{n_T}(\hat{\theta} - \text{ATE}) \Rightarrow \mathcal{N}(0, V)$ with variance

$$V = \text{Var}(c(\psi)) + \text{Var}(b(\psi; p)) + E \left[\frac{\sigma_1^2(\psi)}{p} + \frac{\sigma_0^2(\psi)}{1-p} \right]. \quad (3.2)$$

We think of the first term as the variance due to treatment effect heterogeneity.⁵ We view the second term as the variance due to random assignment. It arises from the chance covariate imbalances between treatment and control units created by complete randomization $D_{1:n} \sim \text{CR}(p)$. The results in the next section show how stratified sampling and assignment nonparametrically dampen the each component of this variance expansion.

3.1 Constant Sampling and Assignment Propensities

In this section we state a central limit theorem for ATE estimation for the simplest case where the sampling and assignment propensities $q = a/k$ and $p = a'/k'$ are constant. Section 3.2 below provides the most general version of this result. Remarks 3.6-3.8 draw connections between our findings and classical results on semiparametric efficiency in the analysis of observational data.

Theorem 3.2. *Require Assumption 3.1. If sampling and assignment designs are locally randomized $T_{1:n} \sim \text{Loc}(\psi, q)$ and $D_{1:n} \sim \text{Loc}(\psi, p)$, then $\sqrt{n_T}(\hat{\theta} - \text{ATE}) \Rightarrow \mathcal{N}(0, V)$*

$$V = q \text{Var}(c(\psi)) + E \left[\frac{\sigma_1^2(\psi)}{p} + \frac{\sigma_0^2(\psi)}{1-p} \right].$$

Comparing to Equation 3.2 above, the variance component $\text{Var}(b(\psi; p))$ due to covariate imbalance between the treatment arms is now asymptotically negligible. The variance due to treatment effect heterogeneity $\text{Var}(c(\psi))$ is now dampened by the sampling proportion $q \in (0, 1]$. Observe that the normalization $\sqrt{n_T}(\hat{\theta} - \text{ATE})$ effectively holds the experiment size n_T constant as we vary q . The number of eligible units $n \approx n_T/q$ grows as $q \rightarrow 0$. For small q , there are many eligible units to sample from, allowing us to build a very representative sample to include in our experiment. Because of this, the first variance component scales with the larger sample size $n \approx n_T/q$ of eligible units, effectively “boosting” the experiment size for this component of the variance.

Example 3.3 (Matched Tuples). In Example 2.2 above, we sampled $n_T = 300$ of $n = 1000$ eligible units using the stratified design $T_{1:n} \sim \text{Loc}(\psi, 3/10)$ with $q = 3/10$. Next, we assigned $1/4$ of the sampled units to treatment by $D_{1:n} \sim \text{Loc}(\psi, 1/4)$. Theorem 3.2 shows that $\sqrt{n_T}(\hat{\theta} - \text{ATE}) \Rightarrow \mathcal{N}(0, V)$ with asymptotic variance

$$V = (3/10) \text{Var}(c(\psi)) + E \left[\frac{\sigma_1^2(\psi)}{1/4} + \frac{\sigma_0^2(\psi)}{3/4} \right].$$

Nonparametric Regression by Design. If $q = 1$ or sampling is completely randomized $T_{1:n} \sim \text{CR}(q)$ ⁶ then the asymptotic variance in Theorem 3.2 is

$$V = \text{Var}(c(\psi)) + E \left[\frac{\sigma_1^2(\psi)}{p} + \frac{\sigma_0^2(\psi)}{1-p} \right].$$

This is exactly the Hahn (1998) semiparametric variance bound for ATE with iid observations $(Y, D, \psi(X))$.⁷ We achieve the semiparametric variance bound with a simple

⁵In particular, the treatment effect heterogeneity predictable by the baseline covariates.

⁶The latter statement follows from the more general results in Section 3.2.

⁷Recent work in Armstrong (2022) shows that this efficiency bound also holds in settings with covariate-adaptive randomization, including the designs considered here.

difference-of-means estimator, without the need for nonparametric re-estimation of the known propensity (Hirano et al. (2003)) or covariate adjustment with well-specified outcome models (Robins and Rotnitzky (1995)). We can interpret this result as saying that finely stratified treatment assignment $D_{1:n} \sim \text{Loc}(\psi, p)$ does nonparametric covariate adjustment “by design.” See Proposition 3.11 below for a more formal equivalence result.

Representative Sampling. If sampling is locally randomized $T_{1:n} \sim \text{Loc}(\psi, q)$, then the variance due to treatment effect heterogeneity decreases from $\text{Var}(c(\psi))$ to $q \text{Var}(c(\psi))$. In this case, V is strictly smaller than the semiparametric variance bound when $q < 1$ and $\text{Var}(c(\psi)) > 0$. Intuitively, by using the additional covariates $(\psi(X_i))_{i=1}^n$ to select a representative sample, we finely represent the distribution of treatment effect levels $(c(\psi_i))_{i=1}^n$ in the larger eligible population into our experiment. More formally, consider an oracle setting where each we observe the treatment effect level $c(\psi_i)$ for each sampled unit $T_i = 1$, estimating the ATE by the sampled average $\hat{\theta} = (1/n_T) \sum_i T_i c(\psi_i)$. Our analysis shows that if $T_{1:n} \sim \text{Loc}(\psi, q)$ then

$$(1/n_T) \sum_{i=1}^n T_i c(\psi_i) = E_n[c(\psi_i)] + o_p(n^{-1/2}).$$

Because of this, the sampled average $(1/n_T) \sum_i T_i c(\psi_i)$ behaves like the infeasible average $E_n[c(\psi_i)]$ over all eligible units, including those not sampled into the experiment. This nonparametrically dampens the variance due to treatment effect heterogeneity from $\text{Var}(c(\psi))$ to $q \text{Var}(c(\psi))$ for $q \in (0, 1]$.

Remark 3.4 (Comparison with Other Designs). For the case without sampling, we can compare our results on fine stratification to other covariate-adaptive assignment designs. Li et al. (2018) shows that under rerandomized treatment assignment, $\hat{\theta}$ is asymptotically (almost) as efficient as interacted linear regression adjustment, effectively doing linear regression “by design.” Bugni et al. (2019) show that coarsely stratified assignment $\psi(X) \in \{1, \dots, m\}$ is asymptotically equivalent to an interacted linear regression adjustment that includes all strata indicators as covariates. Harshaw et al. (2021) suggest a novel Gram-Schmidt walk design with MSE bounded by a quantity related to linear ridge regression. By contrast, we show that the fine stratification $D_{1:n} \sim \text{Loc}(\psi, p)$ does nonparametric regression adjustment by design.

Extending Bugni et al. (2019), the next example proves an equivalence between coarse stratification and linear regression adjustment in the case where both sampling and assignment are both coarsely stratified.

Example 3.5 (Coarse Stratification). If $T_{1:n} \sim \text{Loc}(S, q)$ and $D_{1:n} \sim \text{Loc}(S, p)$ for a fixed stratification $S(X) \in \{1, \dots, m\}$, Theorem 3.2 shows that $\sqrt{n_T}(\hat{\theta} - \text{ATE}) \Rightarrow \mathcal{N}(0, V_S)$

$$V_S = q \text{Var}(c(S)) + E \left[\frac{\sigma_1^2(S)}{p} + \frac{\sigma_0^2(S)}{1-p} \right]. \quad (3.3)$$

Alternatively, suppose sampling $T_{1:n} \sim \text{CR}(q)$ and assignment $D_{1:n} \sim \text{CR}(p)$ are completely randomized, with ex-post linear adjustment for the covariate imbalances due to both sampling and assignment. To define the adjustment, let $z_i = (\mathbf{1}(S_i = k))_{k=1}^{m-1}$ denote leave-one-out strata indicators and their de-means versions $\tilde{z}_i = z_i - E_n[z_i | T_i = 1]$. Consider the linear regression $Y \sim 1 + D + \tilde{z} + D\tilde{z}$ and let $\hat{\tau}$ denote the coefficient on

D and $\hat{\beta}$ the coefficient on $D\tilde{z}$. Define the sampled covariate mean $\bar{z}_{T=1} = E_n[z_i|T_i = 1]$ and eligible covariate mean $\bar{z} = E_n[z_i]$ and consider the doubly-adjusted estimator

$$\hat{\theta}_{adj} = \hat{\tau} - \hat{\beta}'(\bar{z}_{T=1} - \bar{z})$$

Proposition 10.9 in the appendix shows that $\sqrt{n_T}(\hat{\theta}_{adj} - \text{ATE}) \Rightarrow \mathcal{N}(0, V_S)$, with the same variance V_S as in Equation 3.3. Intuitively, this result shows that under coarsely stratified sampling and assignment, the difference-of-means estimator is as precise as the doubly-adjusted estimator $\hat{\theta}_{adj}$.

We generalize Example 3.5 to an equivalence between fine stratification and nonparametric double adjustment in Proposition 3.11 in the next section. The remainder of this subsection provides more intuition for Theorem 3.2, connecting our results on fine stratification to the previous literature on semiparametric efficiency.

Remark 3.6 (Efficient Influence Function). Consider expanding the difference-of-means estimator $\hat{\theta}$ about the efficient influence function for the ATE. Denote the nonparametric regression residuals $\epsilon_i^d = Y_i(d) - E[Y_i(d)|\psi_i]$. Under locally randomized sampling and assignment

$$\begin{aligned} \hat{\theta} = & E_n[c(\psi_i)] + E_n \left[\frac{D_i \epsilon_i^1}{p} + \frac{(1 - D_i) \epsilon_i^0}{1 - p} \middle| T_i = 1 \right] \\ & + \underbrace{\text{Cov}_n(T_i, c(\psi_i)) / q}_{\text{Sampling}} + \underbrace{\text{Cov}_n(D_i, b(\psi_i) | T_i = 1) / c_p}_{\text{Assignment}} + O_p(n^{-1}). \end{aligned}$$

If $T_i = 1$ for all $i \in [n]$ then the first two terms are exactly the efficient influence function for the ATE. The third term is the estimator error due to correlation between the sampling variables T_i and treatment effect heterogeneity $c(\psi_i)$ among the eligible units. The fourth term is the estimator error due to correlation between treatment assignments D_i and outcome heterogeneity among the sampled units. Without stratification, the chance covariate imbalances produce by randomization contribute non-negligible asymptotic variance, and the errors $\sqrt{n} \text{Cov}_n(T_i, c(\psi_i)) \Rightarrow \mathcal{N}(0, v)$ with $v > 0$. By contrast, we show that if $T_{1:n} \sim \text{Loc}(\psi, q)$ then the sampling errors $\sqrt{n} \text{Cov}_n(T_i, f(\psi_i)) = o_p(1)$ for any function $E[f(\psi)^2] < \infty$, and similarly for the assignment term. Because of this, the unadjusted estimator $\hat{\theta}$ is first-order equivalent to the efficient influence function for the ATE under local randomization if $q = 1$. If $q < 1$, the situation is even better, and the first term behaves like the infeasible average $E_n[c(\psi_i)]$ over all eligible units.

Remark 3.7 (Table One). Applied researchers often report tests of covariate balance in “table one.” One common approach is to report a p-value for the test that $\beta = 0$ in the regression $f(\psi_i) = \hat{a} + \hat{\beta}D_i + e_i$, using the normal approximation $\sqrt{n}\hat{\beta} \Rightarrow \mathcal{N}(0, v)$. By contrast, if $D_{1:n} \sim \text{Loc}(\psi, q)$ then $\sqrt{n}\hat{\beta} = o_p(1)$ for any covariate $E[f(\psi)^2] < \infty$, showing that the level of such a test converges to zero. Intuitively, this shows that fine stratification with respect to $\psi(X)$ balances any square-integrable covariate $f(\psi)$ to order $o_p(n^{-1/2})$.

Remark 3.8 (Realized Propensity Score). Suppose sampling and assignment are completely randomized. For any set A with $P(\psi \in A) > 0$ define the *realized* sampling proportions in the set A by $\hat{q}_A = E_n[T_i|\psi_i \in A]$. The discrepancy between expected and realized propensities $q - \hat{q}_A = O_p(n^{-1/2})$, so q is implemented with errors of order $1/\sqrt{n}$,

and similarly for the treatment proportions p . This can be seen in Figure , where the expected and realized propensities widely diverge in certain regions of the covariate space. Such fluctuations increase estimator variance. As in [Hirano et al. \(2003\)](#), one way to fix this problem is to nonparametrically re-estimate the realized sampling and assignment proportions $\widehat{q}(\psi)$ and $\widehat{p}(\psi)$ everywhere in the space, using the propensity weighting

$$\widehat{\theta}_{ipw} = E_n \left[\frac{T_i(D_i - \widehat{p}(\psi_i))Y_i}{\widehat{q}(\psi_i)(\widehat{p} - \widehat{p}^2)(\psi_i)} \right]$$

We provide a simpler solution, showing that fine stratification gets the realized propensities right at design-time. In particular, our analysis shows that if $T_{1:n} \sim \text{Loc}(\psi, q)$ then the error $q - \widehat{q}_{A_n} = o_p(n^{-1/2})$, even for a slowly shrinking sequence of sets $P(\psi \in A_n) \rightarrow 0$. Because of this, we think of the design $T_{1:n} \sim \text{Loc}(\psi, q)$ as a “local” implementation of the propensity q with respect to $\psi(X)$.

3.2 Varying Sampling and Assignment Propensities

This section describes the most general version of our method, providing finely stratified designs for non-constant sampling and assignment proportions $q(x)$, $p(x)$. The asymptotic theory developed in this section allows us to formalize and solve the problem of optimal stratification with heterogeneous costs in Section 4 below.

First, we formally define the procedure. Suppose that $q(x) \in \{a_l/k_l : l \in L\}$ for some finite index set L . Similarly, suppose $p(x) \in \{a'_l/k'_l : l \in L'\}$ with $|L'| < \infty$. Extending our definition, let $T_{1:n} \sim \text{Loc}(\psi, q(x))$ denote the following double stratification procedure:

- (1) Partition $\{1, \dots, n\}$ into propensity strata $S_l \equiv \{i : q(X_i) = a_l/k_l\}$.
- (2) In each propensity stratum S_l , draw samples $(T_i)_{i \in S_l} \sim \text{Loc}(\psi, a_l/k_l)$.

Equivalently, we partition each propensity stratum S_l into groups $g \subseteq S_l$ of size k_l such that $S_l = \bigsqcup_{g \in \mathcal{G}_l} g$ and the homogeneity condition

$$n^{-1} \sum_g \sum_{i,j \in g} |\psi_i - \psi_j|_2^2 = n^{-1} \sum_l \sum_{g \in \mathcal{G}_l} \sum_{i,j \in g} |\psi_i - \psi_j|_2^2 = o_p(1)$$

In practice, we simply run our matching algorithm separately in each propensity stratum S_l , and draw $(T_i)_{i \in g} \sim \text{CR}(a_l/k_l)$ independently for each $g \in \mathcal{G}_l$. Treatment assignment $D_{1:n} \sim \text{Loc}(\psi, p(x))$ is defined identically, partitioning only the units $\{i : T_i = 1\}$ sampled into the experiment.

Example 3.9 (Budget and Welfare Constraints). Consider a village level experiment, where collecting outcome data is either high cost H or low cost L , depending on village proximity and labor costs. Due to budget constraints, we decide to sample $q(L) = 1/2$ of the low cost villages but only $q(H) = 1/10$ of the high cost villages. To do so, we match the high cost villages into 10-tuples and the low cost villages into pairs using publicly available covariates ψ_1 . We randomly sample one village from each 10-tuple and one from each pair. Before assigning treatments, we collect additional survey covariates ψ_2 in each sampled village $T_i = 1$. We label the sampled villages as those likely to benefit most M and least L from the intervention according to our prior. Local policymakers insist on the targeted assignment propensity $p(M) = 2/3$ and $p(L) = 1/3$. We implement this

assignment propensity using matched triples on ψ_2 , assigning $D = 1$ to 2/3 of the villages in each M -type triple and 1/3 in each L -type triple.

Before stating our main result, we extend the definition of our estimator to accommodate varying propensities. Define the double IPW estimator

$$\hat{\theta}_2 = E_n \left[\frac{T_i D_i Y_i}{q(\psi_i) p(\psi_i)} \right] - E_n \left[\frac{T_i (1 - D_i) Y_i}{q(\psi_i) (1 - p(\psi_i))} \right] \quad (3.4)$$

If $p(x) = p$ and $q(x) = q$ are constant, then $\hat{\theta}_2 = \hat{\theta} + O_p(n^{-1})$, where $\hat{\theta}$ is the difference-of-means estimator studied in the previous section.⁸ Then abusing notation we denote both estimators by $\hat{\theta}$. The following theorem gives our asymptotic results for fine stratification with varying propensities, extending the fixed propensity results in Theorem 3.2 above. We begin with the special case $\psi_1 = \psi_2 = \psi$ and $q = q(\psi)$, $p = p(\psi)$, all non-random.

Theorem 3.10 (CLT). *Suppose Assumption 3.1 holds. Assume sampling and assignment $T_{1:n} \sim \text{Loc}(\psi, q(\psi))$ and $D_{1:n} \sim \text{Loc}(\psi, p(\psi))$. Then $\sqrt{n_T}(\hat{\theta} - \text{ATE}) \Rightarrow \mathcal{N}(0, V)$*

$$V = E[q(\psi)] \left(\text{Var}(c(\psi)) + E \left[\frac{1}{q(\psi)} \left(\frac{\sigma_1^2(\psi)}{p(\psi)} + \frac{\sigma_0^2(\psi)}{1 - p(\psi)} \right) \right] \right)$$

If $q = 1$ then this is exactly the [Hahn \(1998\)](#) semiparametric variance bound for ATE with iid observations $(Y, D, \psi(X))$ and propensity $p(\psi)$. Note that the population IPW estimator is already semiparametrically efficient, and don't need to nonparametrically re-estimate the known propensities $q(\psi)$ and $p(\psi)$ as in [Hirano et al. \(2003\)](#). Next consider the efficiency gain from stratified sampling. The overall sampling proportion n_T/n has $n_T/n = E[q(\psi)] + o_p(1)$. Then defining $\bar{q} = E[q(\psi)]$, local randomization reduces the variance due to treatment effect heterogeneity from $\text{Var}(c(\psi))$ to $\bar{q} \text{Var}(c(\psi))$ for $\bar{q} \in (0, 1]$.

Regression Equivalence. Theorem 3.10 shows that if $q = 1$ then difference-of-means is semiparametrically efficient. To the best of our knowledge, no such efficiency bound is available for jointly finely stratified sampling and assignment $T_{1:n} \sim \text{Loc}(\psi, q(\psi))$ and $D_{1:n} \sim \text{Loc}(\psi, p(\psi))$ with $q \neq 1$. Instead, we show a direct equivalence between unadjusted estimation under fine stratification and nonparametric regression adjustment. In particular, the asymptotic variance V above is the same as that achieved by a doubly-robust estimator that adjusts for covariate imbalances during both sampling and assignment. To state the result, consider regression estimators \hat{m}_d for $m_d(\psi) = E[Y(d)|\psi]$. Define the doubly-augmented IPW (2-AIPW) estimator

$$\hat{\theta}_{adj} = E_n[\hat{m}_1(\psi_i) - \hat{m}_0(\psi_i)] + E_n \left[\frac{T_i D_i (Y_i - \hat{m}_1(\psi_i))}{q(\psi_i) p(\psi_i)} - \frac{T_i (1 - D_i) (Y_i - \hat{m}_0(\psi_i))}{q(\psi_i) (1 - p(\psi_i))} \right]$$

$\hat{\theta}_{adj}$ adjusts for covariate imbalances due to both sampling and assignment. We implement the estimator using cross-fitting, similar to [Chernozhukov et al. \(2017\)](#). See the proof for details. If $q = 1$ this reduces to the familiar AIPW estimator for the ATE. The next result provides an equivalence between doubly-robust nonparametric adjustment under an iid design and unadjusted estimation under a finely stratified design.

⁸This is because $E_n[D_i] = p + O(n^{-1})$ for stratified designs. It would be false for $D_i \stackrel{\text{iid}}{\sim} \text{Bernoulli}(p)$.

Proposition 3.11 (Regression Equivalence). *Require Assumption 3.1. Suppose the estimators $|\widehat{m}_d - m_d|_{2,\psi} = o_p(1)$ are well-specified and consistent. If the design is iid $T_i \overset{\text{iid}}{\sim} \text{Bernoulli}(q(\psi_i))$ and $D_i \overset{\text{iid}}{\sim} \text{Bernoulli}(p(\psi_i))$ then $\sqrt{n_T}(\widehat{\theta}_{\text{adj}} - \text{ATE}) \Rightarrow \mathcal{N}(0, V)$ with the variance V as in Theorem 3.10.*

Intuitively, Proposition 3.11 shows that local randomization makes the unadjusted estimator $\widehat{\theta}$ as efficient as a much more involved estimator, with well-specified nonparametric adjustment for covariate imbalances in both sampling and assignment.

Estimated Design Variables. We wish to formally accommodate the case where design variables $\psi, q(\psi), p(\psi)$ are estimated using previously collected data. To do so, define the random element $\xi \perp\!\!\!\perp (X_i, Y_i(1), Y_i(0))_{i=1}^n$ and define $\psi = \psi(X, \xi)$, $p = p(\psi, \xi)$, and $q = q(\psi, \xi)$. For example, we could let $\xi = (\widehat{m}_d)_{d=0,1}$ be regression estimates of $E[Y(d)|X = x]$ from a previous experiment and set $\psi(X, \xi) = (\widehat{m}_0(X), \widehat{m}_1(X))$. Our asymptotic theory will condition on ξ , allowing the study of stratification under “small pilot asymptotics” in the previous example. Define $c(\psi, \xi) = E[Y(1) - Y(0)|\psi, \xi]$ and $\sigma_d^2(\psi, \xi) = \text{Var}(Y(d)|\psi, \xi)$. The proof of Theorem 3.10 shows that if $T_{1:n} \sim \text{Loc}(\psi, q(\psi, \xi))$ and $D_{1:n} \sim \text{Loc}(\psi, p(\psi, \xi))$ then $\sqrt{n_T}(\widehat{\theta} - \text{ATE})|\xi \Rightarrow \mathcal{N}(0, V(\xi))$

$$V(\xi) = E[q(\psi, \xi)] \left(\text{Var}(c(\psi, \xi)) + E \left[\frac{1}{q(\psi, \xi)} \left(\frac{\sigma_1^2(\psi, \xi)}{p(\psi, \xi)} + \frac{\sigma_0^2(\psi, \xi)}{1 - p(\psi, \xi)} \right) \right] \right) \quad (3.5)$$

All expectations and variances are conditional on ξ . In this setting, the inference methods provided in Section 6 are asymptotically exact conditional on ξ . Note that marginally over both our experiment and the previous data ξ , the estimator is asymptotically mixed normal $\sqrt{n_T}(\widehat{\theta} - \text{ATE}) \Rightarrow Z$, where Z has characteristic function $E[\exp(-\frac{1}{2}t^2 V(\xi))]$.

Collecting Data After Sampling. In practice, experimenters may want to use different stratification variables ψ_1 for sampling and ψ_2 for assignment with $\psi_1 \neq \psi_2$. For example, ψ_1 may include publicly available administrative data, while ψ_2 includes additional survey covariates collected after sampling units into the experiment. To accommodate this, next we state the most general version of Theorem 3.10. Suppose Assumption 3.1 holds. Assume sampling and assignment $T_{1:n} \sim \text{Loc}(\psi_1, q(x))$ and $D_{1:n} \sim \text{Loc}(\psi_2, p(x))$. If $q(x) = q(\psi_1)$, require $\psi_1 \subseteq \psi_2$. Otherwise, require $(\psi_1, q) \subseteq \psi_2$. Then $\sqrt{n_T}(\widehat{\theta} - \text{ATE}) \Rightarrow \mathcal{N}(0, \bar{q}V)$ with $\bar{q} = E[q(X)]$ and asymptotic variance

$$\begin{aligned} V = & \text{Var}(c(X)) + E \left[\frac{1}{q(X)} \left(\frac{\sigma_1^2(X)}{p(X)} + \frac{\sigma_0^2(X)}{1 - p(X)} \right) \right] \\ & + E \left[\frac{1 - q(X)}{q(X)} \text{Var}(c(X)|\psi_1, q) \right] + E \left[\frac{1}{q(X)} \text{Var}(b(X)|\psi_2, p) \right] \end{aligned} \quad (3.6)$$

If $\psi_1 \neq \psi_2$, the variance V does not have a simple form like in the special cases above. Instead, we expand V relative to the efficient variance for the full vector of baseline covariates X . For example, if $q = 1$ the first line is the semiparametric variance bound for iid observations (Y, D, X) . If $\psi_1 = \psi_2 = \psi$, and $q = q(\psi)$, $p = p(\psi)$ then V can be rearranged into the form in Theorem 3.10. The requirement that the stratification is increasing⁹ $\psi_1 \subseteq \psi_2$ is subtle. We defer this technical discussion to Remark 9.2 in the

⁹In fact, we just require that $\psi_1 = g(\psi_2)$ for a measurable function g .

appendix.

Optimal Stratification Variables. Equation 3.6 has the advantage of clearly exposing the minimal efficient stratification variables. Setting $\psi_1^* = c(X)$ and $\psi_2^* = (c(X), q(X), b(X; p(X)))$, the third and fourth variance terms are identically zero, achieving the minimal variance

$$V^* = E[q(X)] \left(\text{Var}(c(X)) + E \left[\frac{1}{q(X)} \left(\frac{\sigma_1^2(X)}{p(X)} + \frac{\sigma_0^2(X)}{1-p(X)} \right) \right] \right)$$

We include (c, q) in ψ_2^* to satisfy the subvector condition $\psi_1 \subseteq \psi_2$. Since $b(X; p(X))$ and $c(X)$ are both propensity-weighted linear combinations of $m_d(X) = E[Y(d)|X]$, setting $\psi_2^* = (q, m_0, m_1)(X)$ is also optimal. If $q(x) = q(\psi_1)$, we may exclude $q(\psi_1)$ from ψ_2^* . In practice, ψ_1^* and ψ_2^* are not known. The preceding discussion suggests letting $\psi_1(X)$ be a small subvector of the baseline covariates expected to be most predictive of treatment effect heterogeneity and $\psi_2(X)$ to be a superset of $\psi_1(X)$ expected to be predictive of outcomes.¹⁰ If $c(X) = g(\psi_1(X))$ for some function g , so that $c(X)$ only varies through the subvector $\psi_1(X)$, then this choice is asymptotically optimal. If none of the baseline covariates predict treatment effect heterogeneity, so $c(X)$ is constant, then there is no efficiency loss from completely randomized sampling and $\psi_1^* = 1$.

Remark 3.12 (Curse of Dimensionality). From Equation 3.6, setting $\psi_1(X) = \psi_2(X) = X$ also minimizes the asymptotic variance V in Theorem 3.10. However, our analysis shows that if $E[Y(d)|X = x]$ is Lipschitz continuous then the finite sample variance converges to the asymptotic variance at rate $n \text{Var}(\hat{\theta}) = V + O_p(n^{-2/(\dim(\psi)+1)})$, which may be slow even in moderate dimensions if $\psi(X) = X$. In finite samples, the variance $\text{Var}(\hat{\theta})$ may be U-shaped in the dimension of the stratification variables, ordered by covariate relevance, since matching on many irrelevant variables reduces match quality on the relevant variables. This motivates the choice of stratification variables $\psi_1(x)$ and $\psi_2(x)$ of small dimension that minimize V .

4 Optimal Stratified Designs

In this section we formulate and solve the problem of optimal stratification in survey experiments, characterizing the optimal sampling and assignment propensities for budget-constrained experimentation with heterogeneous costs. For simplicity, we restrict to the case with stratification variables $\psi_1 = \psi_2 = \psi$ and sampling and assignment propensities $q = q(\psi)$ and $p = p(\psi)$.

4.1 Budget-Constrained Sampling Problem

In Section 3, we presented asymptotic results of the form $\sqrt{n_T}(\hat{\theta} - \text{ATE}) \Rightarrow \mathcal{N}(0, V)$, normalizing by the experiment size $n_T = \sum_i T_i$. This allowed for easy comparison with the previous literature on covariate-adaptive treatment assignment. However, the experiment size n_T varies with the sampling propensity $q(\psi)$, making this normalization unsuitable for studying the optimal sampling propensity. In this section, we instead normalize instead

¹⁰Since $b(X; p) \propto E[Y(1)|X]/p + E[Y(0)|X]/(1-p)$, we should prioritize predicting outcomes in the arm assigned with lowest probability.

by the number of eligible units n . Since $n_T/n \xrightarrow{p} E[q(\psi)]$, we have $\sqrt{n}(\hat{\theta} - \text{ATE}) \Rightarrow \mathcal{N}(0, V(q, p))$ with

$$V(q, p) = \text{Var}(c(\psi)) + E \left[\frac{1}{q(\psi)} \left(\frac{\sigma_1^2(\psi)}{p(\psi)} + \frac{\sigma_0^2(\psi)}{1 - p(\psi)} \right) \right] \quad (4.1)$$

Consider minimizing Equation 4.1 over all sampling propensities $q(\psi)$. Clearly the unconstrained solution is $q^*(\psi) = 1$, making the experiment as large as possible. More generally, we can formalize the problem of experimentation subject to a budget constraint. To set up the problem, define the ex-ante heteroskedasticity function $\bar{\sigma}^2(\psi) = \sigma_1^2(\psi)/p(\psi) + \sigma_0^2(\psi)/(1 - p(\psi))$. We interpret $\bar{\sigma}^2(\psi)$ as the expected residual variance from sampling a unit with $\psi_i = \psi$, prior to the realization of its treatment assignment $D_i \in \{0, 1\}$. Fixing the propensity $p(\psi)$, the asymptotic variance can be written $V(q) = \text{Var}(c(\psi)) + E[\bar{\sigma}^2(\psi)/q(\psi)]$. Let $C(\psi)$ denote the known, potentially heterogeneous, cost of including a unit of type $\psi(X) = \psi$ in the experiment. More general costs $C(x)$ can be handled easily, and are discussed in Remark 4.2 below. Then the budget-constrained variance minimization problem with budget \bar{B} can be written

$$\min_{0 < q \leq 1} E \left[\frac{\bar{\sigma}^2(\psi)}{q(\psi)} \right] \quad \text{s.t.} \quad E[C(\psi)q(\psi)] = \bar{B}. \quad (4.2)$$

The next proposition characterizes the interior solutions to this problem.

Proposition 4.1. *Suppose $\inf_x \sigma_d^2(\psi) \geq c > 0$. Define the candidate solution*

$$q^*(\psi) = \bar{B} \cdot \frac{\bar{\sigma}(\psi)C(\psi)^{-1/2}}{E[\bar{\sigma}(\psi)C(\psi)^{1/2}]} \quad (4.3)$$

If $\sup_\psi q^(\psi) \leq 1$, then q^* is optimal in Equation 4.2.*

If the feasibility constraint $\sup_\psi q^*(\psi) \leq 1$ is violated, the optimal sampling propensity may not have a simple analytical form. Remark 5.5 below provides a rounding procedure that can be used to restore feasibility in this case. To build intuition for the form of the solution, consider the following special cases.

- (a) **Homoskedasticity.** Suppose $\sigma_d^2(\psi) = \sigma_d^2$ constant for $d \in \{0, 1\}$ and $p(\psi) = p$. Then the optimal propensity $q^*(\psi) = \bar{B} \cdot C(\psi)^{-1/2}/E[C(\psi)^{1/2}]$ has $q^*(\psi) \propto 1/\sqrt{C(\psi)}$. This case provides a simple heuristic for sample allocation with known, heterogeneous costs.
- (b) **Homogeneous costs.** If $C(\psi) = 1$, then $E[q(\psi)] \leq \bar{B}$ constrains the total proportion of sampled units. In this case write $\bar{B} = \bar{q}$. The optimal solution has form $q^*(\psi) = \bar{q}\bar{\sigma}(\psi)/E[\bar{\sigma}(\psi)]$, with sampling propensity proportional to the ex-ante standard deviation. In particular, we would like to oversample ($q^*(\psi) > \bar{q}$) units of type $\psi(X) = \psi$ that have larger ex-ante residual variance than the population average $E[\bar{\sigma}(\psi)]$, and undersample in the opposite case.

Optimal Spending. Under optimal sampling, the total amount spent on units of type ψ is $C(\psi)q^*(\psi)dP(\psi) \propto \sqrt{\bar{\sigma}^2(\psi)C(\psi)}dP(\psi)$. We spend more on units with larger ex-ante variance and larger per unit cost. However, since the optimal propensity undersamples high cost units, spending grows as $\sqrt{C(\psi)}$, instead of linearly as it would if $q(\psi) = q$ were constant.

Remark 4.2 (General Costs). More generally, let $C(x)$ denote the cost of including a unit of type $X = x$. Abusing notation, define the average cost $C(\psi) = E[C(X)|\psi]$. The conclusions of Proposition 4.1 and Theorem 4.4 below remain true as stated, with budget constraint $E[C(X)q(\psi)] = E[C(\psi)q(\psi)] \leq \bar{B}$.

4.2 Globally Optimal Stratification

The main result of this section studies implementation of the globally optimal stratified design, subject to the budget constraint. First, we characterize the optimal assignment propensity. For any fixed sampling propensity $q(\psi)$, the global minimizer of Equation 4.1 is the conditional Neyman allocation $p^*(\psi) = \sigma_1(\psi)/(\sigma_1(\psi) + \sigma_0(\psi))$. In some cases, we may only be interested in implementing a constant assignment propensity $p^* \in (0, 1)$. If q is also constant, then

$$p^* = \sqrt{E[\sigma_1^2(\psi)]} \left(\sqrt{E[\sigma_1^2(\psi)]} + \sqrt{E[\sigma_0^2(\psi)]} \right)^{-1} \quad (4.4)$$

Compare this to the classical Neyman allocation $\sigma_1/(\sigma_1 + \sigma_0)$ with $\sigma_d = \text{SD}(Y(d))$. Here, only the residual variances $\sigma_d^2(\psi) = \text{Var}(Y(d)|\psi)$ enter p^* , since the fluctuations of $Y(d)$ predictable by $\psi(X)$ do not contribute to first-order asymptotic variance under fine stratification.

The jointly optimal sampling and assignment propensities are obtained by plugging the conditional Neyman allocation $p^*(\psi) = \sigma_1(\psi)/(\sigma_1(\psi) + \sigma_0(\psi))$ into the formula for $q^*(\psi; p)$ above. This gives ex-ante variance $\bar{\sigma}^2(\psi) = (\sigma_1(\psi) + \sigma_0(\psi))^2$ and jointly optimal sampling and assignment propensities

$$p^*(\psi) = \frac{\sigma_1(\psi)}{\sigma_1(\psi) + \sigma_0(\psi)} \quad q^*(\psi; p^*) = \bar{B} \frac{(\sigma_1(\psi) + \sigma_0(\psi))C(\psi)^{-1/2}}{E[(\sigma_1(\psi) + \sigma_0(\psi))C(\psi)^{1/2}]}. \quad (4.5)$$

The propensities $p^*(\psi)$ and $q^*(\psi)$ will generally need to be discretized in order to implement them using fine stratification.

Definition 4.3 (Discretization). Let $q_n^*(\psi)$ and $p_n^*(\psi)$ take values in the finite approximating propensity set $\{a_l/k_l : l \in L_n\}$ such that $|q_n^* - q^*|_\infty = o(1)$ and $|p_n^* - p^*|_\infty = o(1)$. Define $\bar{k}_n = \max_{l \in L_n} k_l$ and require $\bar{k}_n |L_n| = o(n^{1-\frac{d+1}{\alpha}})$ if $E[|\psi(X)|^\alpha] < \infty$.

If $\psi(X)$ is bounded, the final condition simplifies to $\bar{k}_n |L_n| = o(n)$. For example, one way to satisfy Definition 4.3 is to round the optimal propensities to the nearest a/k_n for some sequence $k_n \rightarrow \infty$, setting $q_n^*(\psi) = \text{argmin}\{|q^*(\psi) - a/k_n| : 1 \leq a \leq k_n - 1, k_n = \lfloor n^{1/2-\epsilon} \rfloor\}$ and similarly for $p_n^*(\psi)$. The main theorem of this section shows that such discretizations are asymptotically efficient.

Theorem 4.4 (Optimal Stratification). Suppose Assumption 3.1. If $T_{1:n} \sim \text{Loc}(\psi, q_n^*(\psi))$ and $D_{1:n} \sim \text{Loc}(\psi, p_n^*(\psi))$. Then $\sqrt{n}(\hat{\theta} - \text{ATE}) \Rightarrow \mathcal{N}(0, V^*)$

$$V^* = \text{Var}(c(\psi)) + \min_{\substack{0 < q, p \leq 1 \\ E[C(\psi)q(\psi)] = \bar{B}}} E \left[\frac{1}{q(\psi)} \left(\frac{\sigma_1^2(\psi)}{p(\psi)} + \frac{\sigma_0^2(\psi)}{1 - p(\psi)} \right) \right]$$

The design in Theorem 4.4 minimizes the asymptotic variance over all sampling and assignment propensities, subject to the budget constraint. If we set $q = 1$, then $V^* = \min_{0 \leq p \leq 1} V_H(p)$, minimizing the Hahn (1998) semiparametric efficiency bound for the ATE over all propensity scores. As noted above, Armstrong (2022) shows that this bound also applies to the designs in this paper for the case $q = 1$.

4.3 Finite Sample Optimality

In this final subsection, we briefly discuss exact optimality in finite samples. Consider the case $q = 1$ and $p(X) = a/k$ constant. In this setting, Bai (2022) shows that if the balance function $b(x)$ were known, then matching units into strata of size k according to their sorted $b(X_i)$ values minimizes $\text{MSE}(\hat{\theta}|X_{1:n})$ over all stratified designs. By contrast, here we show that if $b(x)$ were known, the class of stratified designs itself would generally be suboptimal. To see this, consider the case $p = 1/2$. Define the complete graph K_n with vertices $\{1, \dots, n\}$ and edge weights $w_{ij} = b(X_i)b(X_j)$. The Max-Cut optimization problem asks for a partition of the vertices into disjoint sets $E_1 \cup E_0 = \{1, \dots, n\}$ such that the weight of cut edges between E_1 and E_0 is maximized

$$\max_{E_0, E_1} \sum_{i,j} b(X_i)b(X_j)\mathbb{1}(i \in E_1, j \in E_0) \quad \text{s.t.} \quad E_0 \sqcup E_1 = \{1, \dots, n\} \quad (4.6)$$

For example, see Rendl et al. (2008) for an overview. Let E_0^*, E_1^* solve the Max-Cut problem in Equation 4.6. Define the optimal treatment allocation $d_{1:n}^* = d_{1:n}^*(X_{1:n})$ by $d_i^* = \mathbb{1}(i \in E_1^*)$ for $1 \leq i \leq n$. Define the *alternating design* $P^*(D_{1:n}|X_{1:n})$ by

$$P^*(D_{1:n} = d_{1:n}^*|X_{1:n}) = P^*(D_{1:n} = 1 - d_{1:n}^*|X_{1:n}) = 1/2$$

with $D_{1:n} \perp\!\!\!\perp W_{1:n}|X_{1:n}$. Our next theorem shows that the alternating design P^* is globally optimal over the set of all covariate-adaptive designs with fixed treatment probability $P(D_i = 1) = 1/2$. We denote this set of designs by $\mathcal{P}_{1/2} = \{P : P(D_i = 1) = 1/2, D_{1:n} \perp\!\!\!\perp W_{1:n}|X_{1:n}\}$. For simplicity, suppose $n = 2m$ for an integer m .

Theorem 4.5 (Optimal Design). *The design P^* has $\text{MSE}_{P^*}(\hat{\theta}|X_{1:n}) \leq \text{MSE}_P(\hat{\theta}|X_{1:n})$ for all $P \in \mathcal{P}_{1/2}$.*

The inequality is strict if Problem 4.6 has a unique solution up to permutation of set labels. In particular, note that P^* is not a matched pairs design. Nevertheless, $b(x)$ is not known, so neither the globally optimal design, nor the optimal stratified design from Bai (2022) are feasible. We also caution against plug-in approaches that use a pilot estimate of $b(X_i)$. Section 5.2 below shows that such approaches are equivalent to regression adjustment with regressions estimated in the pilot instead of the main sample, which may be much noisier if the pilot is small. For these reasons, we do not further pursue finite sample optimal designs in this paper.

5 Design with a Pilot Experiment

Suppose that data from a pilot study is available when designing the main experiment. In this section, we study a procedure that uses this data to estimate the solution to the optimal stratification problem derived in Section 4. We show that such a procedure is

asymptotically efficient, achieving the budget-constrained optimal variance of Theorem 4.4. For the case $q = 1$ (no sampling), our procedure minimizes the semiparametric variance bound over all propensity scores, providing the first asymptotically efficient solution to the question of design using a pilot study (Hahn et al. (2011)). Small pilot considerations and potential robustifications are discussed in Remark 5.3 below. Section 5.2 discusses some issues related to pilot estimation of the optimal stratification variables ψ^* derived in Section 3.

5.1 Feasible Optimal Stratification

Fix stratification variables $\psi_1 = \psi_2 = \psi$ and consider estimating the optimal design for the budget-constrained optimal stratification problem in Equation 4.2. This amounts to using pilot or proxy data to estimate the efficient sampling proportions $q^*(\psi)$ and treatment propensity $p^*(\psi)$. As a proof of concept, we first state our result under large pilot asymptotics, allowing consistent estimation of the heteroskedasticity functions $\sigma_d(\psi)$. We show that the estimated optimal design is asymptotically optimal in the sense of Theorem 4.4. Fixed pilot asymptotics and small sample considerations are discussed in Remark 5.3 below. In Section 4 we derived the optimal propensities for the budget-constrained design problem

$$q^*(\psi) = \bar{B} \frac{(\sigma_1(\psi) + \sigma_0(\psi))C(\psi)^{-1/2}}{E[(\sigma_1(\psi) + \sigma_0(\psi))C(\psi)^{1/2}]} \quad p^*(\psi) = \frac{\sigma_1(\psi)}{\sigma_1(\psi) + \sigma_0(\psi)}.$$

The sampling propensity $q^*(\psi)$ is optimal provided the feasibility condition $\sup_{\psi} q^*(\psi) \leq 1$ is satisfied. Consider pilot heteroskedasticity estimates¹¹ $\hat{\sigma}_d^2(\psi)$ for $d = 0, 1$. Define the propensity estimates¹²

$$\hat{q}(\psi) = \bar{B} \frac{(\hat{\sigma}_1(\psi) + \hat{\sigma}_0(\psi))C(\psi)^{-1/2}}{E_n[(\hat{\sigma}_1(\psi_i) + \hat{\sigma}_0(\psi_i))C(\psi_i)^{1/2}]} \quad \hat{p}(\psi) = \frac{\hat{\sigma}_1(\psi)}{\hat{\sigma}_1(\psi) + \hat{\sigma}_0(\psi)}.$$

In practice, we may find $\hat{q}(\psi_j) > 1$ for some j . This could be because the condition $\sup_{\psi} q^*(\psi) \leq 1$ is violated and the optimal sampling problem does not have an interior solution, or just due to statistical error. However, we can transform $\hat{q}(\psi)$ into an admissible sampling propensity by an iterative rounding procedure, described in Remark 5.5 below. Suppose we have done so and let $\hat{q}_n(\psi)$ and $\hat{p}_n(\psi)$ be a sequence of discretizations of $\hat{q}(\psi)$ and $\hat{p}(\psi)$, satisfying the conditions in Definition 4.3. We require the following technical conditions, including consistency of the pilot heteroskedasticity estimates.

Assumption 5.1. *Assume the interior solution condition $\sup_{\psi} q^*(\psi) \leq 1$. Require the variance regularity condition $\inf_{\psi} \sigma_d(\psi) > 0$ and $(\sigma_1/\sigma_0)(\psi) \in [c_l, c_u]$ with $0 < c_l < c_u < \infty$. Require pilot estimation rate $|\sigma_d^2 - \hat{\sigma}_d^2|_{2,\psi} = O_p(n^{-r})$ for some $r > 0$. Require discretization rate $\bar{k}_n |L_n| = o(n^{1-(\dim(\psi)+1)/\alpha_1})$ and $\bar{k}_n = \omega(n^{1/\alpha_2})$. Assume the moments $E[Y(d)^4] < \infty$, $E|\psi(X)|_2^{\alpha_1} < \infty$ for $\alpha_1 \geq \dim(\psi) + 1$, and $E[|m_d(\psi)|^{\alpha_2}] < \infty$ for $\alpha_2 \geq 1/r$. Assume costs $0 < C(\psi) < \infty$ for all ψ .*

Our main result shows that finely stratified implementation of the optimal propensity estimates is asymptotically fully efficient.

¹¹For example see Fan and Yao (1998).

¹²Note in $\hat{q}(\psi)$ the average is taken over the main experiment covariates, allowing for covariate shift between pilot and main experiment.

Theorem 5.2 (Pilot Design). *Impose Assumption 5.1. Suppose $T_{1:n} \sim \text{Loc}(\psi, \hat{q}_n(\psi))$ and $D_{1:n} \sim \text{Loc}(\psi, \hat{p}_n(\psi))$. Then $\sqrt{n}(\hat{\theta} - \text{ATE}) \Rightarrow \mathcal{N}(0, V^*)$*

$$V^* = \text{Var}(c(\psi)) + \min_{\substack{0 < q, p \leq 1 \\ E[C(\psi)q(\psi)] = \bar{B}}} E \left[\frac{1}{q(\psi)} \left(\frac{\sigma_1^2(\psi)}{p(\psi)} + \frac{\sigma_0^2(\psi)}{1 - p(\psi)} \right) \right]$$

For intuition, it is also helpful to consider certain special cases of Theorem 5.2. If we fix $q = 1$, the design $D_{1:n} \sim \text{Loc}(\psi, \hat{p}_n(\psi))$ asymptotically minimizes the Hahn (1998) variance bound over all propensity scores

$$V^* = \min_{0 \leq p \leq 1} V_H(p) = \text{Var}(c(\psi)) + \min_{0 \leq p \leq 1} E \left[\frac{\sigma_1^2(\psi)}{p(\psi)} + \frac{\sigma_0^2(\psi)}{1 - p(\psi)} \right]. \quad (5.1)$$

If \hat{p}^* is a consistent pilot estimate of the optimal constant propensity p^* in Equation 4.4, then the design $D_{1:n} \sim \text{Loc}(\psi, \hat{p}_n)$ has $\sqrt{n}(\hat{\theta} - \text{ATE}) \Rightarrow \mathcal{N}(0, V)$ with

$$V^* = \text{Var}(c(\psi)) + \min_{p \in (0,1)} E \left[\frac{\sigma_1^2(\psi)}{p} + \frac{\sigma_0^2(\psi)}{1 - p} \right]$$

Remark 5.3 (Small Pilots). For small pilots, the asymptotic results in Theorem 5.2 imposing consistent estimation of the variance function $\sigma_d^2(\psi)$ may not reflect finite sample performance. Instead consider an inconsistent variance approximation using the squared residuals from a linear regression. Let $\hat{\epsilon}_i^d$ be residuals from regressions $Y(d) \sim 1 + \psi$ in $\{D_i = d\}$. For pilot data $(W_j)_{j=1}^m$ compute $\hat{s}_d = E_m[(\hat{\epsilon}_i^d)^2]^{1/2}$ and form the pilot estimate $\hat{p}^* = \hat{s}_1/(\hat{s}_1 + \hat{s}_0)$, rounding it to a close rational number $\hat{p} = a/k$. By the estimated design variable version of Theorem 3.10, if $D_{1:n} \sim \text{Loc}(\psi, \hat{p})$ then $\sqrt{n}(\hat{\theta} - \text{ATE})|W_{1:m} \Rightarrow \mathcal{N}(0, V)$

$$V = \text{Var}(c(\psi)) + E \left[\frac{\sigma_1^2(\psi)}{\hat{p}} + \frac{\sigma_0^2(\psi)}{1 - \hat{p}} \middle| W_{1:m} \right]$$

The rounding of \hat{p}^* to $\hat{p} = a/k$ adds robustness. For example, if $p^* = 1/2$ and our pilot estimate $\hat{p}^* = .57$, we would round to $\hat{p} = 1/2$ except in very large experiments. See the discussion of discretization in Remark 5.4 below. In practice, this procedure could be further robustified by constructing a confidence interval for p^* and checking that it excludes a baseline choice such as $p = 1/2$.

Remark 5.4 (Discretization). Consider a pilot estimate $\hat{p}(\psi) = .637$. This could be rounded to any of $\hat{p} = 2/3, 3/5, 13/20, 63/100$ and so on. If ψ is bounded, Assumption 5.1 requires that $k_n = o(\sqrt{n})$ for the rounding scheme a/k_n . This condition gives some quantitative guidance about discretization fineness. For example, if $n = 400$ we would rule out $\hat{p} = 13/20$.

Remark 5.5 (Feasible Sampling). In practice, we may find that $\hat{q}(\psi_j) > 1$ for some j , violating the sampling constraint. To fix this, define an index set $J = \emptyset$ and implement the following iterative rounding procedure. (1) Find the largest $\hat{q}(\psi_j) > 1$. Set $\hat{q}(\psi_j) = 1$ and add j to J . (2) Recompute the sampling propensity according to

$$\hat{q}(\psi_i) = \frac{\bar{B}n - \sum_{i \in J} C(\psi_i)}{n - |J|} \frac{(\hat{\sigma}_1(\psi_i) + \hat{\sigma}_0(\psi_i))C(\psi_i)^{-1/2}}{E_n[(\hat{\sigma}_1(\psi_l) + \hat{\sigma}_0(\psi_l))C(\psi_l)^{1/2} | l \notin J]} \quad \forall i \notin J$$

If $\max_{i=1}^n \hat{q}(\psi_i) \leq 1$, stop. Otherwise, return to (1). This procedure satisfies the in-sample budget constraint $E_n[\hat{q}(\psi_i)C(\psi_i)] = \bar{B}$ after each iteration and terminates with $\max_{i=1}^n \hat{q}(\psi_i) \leq 1$.

Remark 5.6 (Optimal Stratification Trees). [Tabord-Meehan \(2022\)](#) suggests using pilot data to estimate a stratification \hat{S} and assignment propensity $\hat{p}(S)$ over a set of tree partitions \mathcal{T} of the covariate space. In our notation, their Theorem 3.1 shows that if $D_{1:n} \sim \text{Loc}(\hat{S}, \hat{p}(S))$ then $\sqrt{n}(\hat{\theta} - \text{ATE}) \Rightarrow \mathcal{N}(0, V)$ with asymptotic variance

$$V = \text{Var}(c(\psi)) + \min_{S \in \mathcal{T}} \left(E[\text{Var}(b(\psi; p^*(S))|S)] + E \left[\frac{\sigma_1^2(\psi)}{p^*(S)} + \frac{\sigma_0^2(\psi)}{1 - p^*(S)} \right] \right)$$

The optimal propensity $p^*(S) = \sigma_1(S)/(\sigma_1(S) + \sigma_0(S))$ with $\sigma_d^2(S) = \text{Var}(Y(d)|S)$. The display shows that the optimal stratification tree chooses a compromise between two different forces. In the first term, it tries to minimize the variance due to covariate imbalance by choosing strata S that predict propensity-weighted outcomes well, minimizing $E[\text{Var}(b(\psi; p^*(S))|S)]$. In the second term, it tries to minimize the residual variance by choosing strata S such that $p^*(S)$ is close to the globally optimal propensity $p^*(\psi) = \sigma_1(\psi)/(\sigma_1(\psi) + \sigma_0(\psi))$. By contrast, we implement a discretized consistent estimate $\hat{p}_n(\psi)$ of the optimal propensity $p^*(\psi)$ using fine stratification. This makes the middle term above asymptotically lower order and globally minimizes the residual term, without any first-order tradeoff (Equation 5.1).

5.2 Estimating Stratification Variables

Section 3 showed that the variables $\psi_1^* = c(X)$ and $\psi^* = (c, b)(X)$ were asymptotically efficient for both sampling and assignment. This suggests setting $\psi_1(X) = \hat{c}(X)$ and $\psi_2(X) = (\hat{c}, \hat{b})(X)$, using pilot estimates of the various regression functions. In our notation, the design $D_{1:n} \sim \text{Loc}(\hat{b}, p)$ was proposed in [Bai \(2022\)](#).

Comparison with Regression Adjustment. Our first result is negative, suggesting that such an approach would be dominated by throwing away the pilot data and using an iid design with regression adjustment in the main sample. For simplicity, consider $q = 1$ and let $D_{1:n} \sim \text{Loc}(\hat{b}, p)$ be the design with estimated stratification variables. Let iid treatments $\check{D}_i \sim \text{Bernoulli}(p)$ and $\hat{\theta}_{adj}$ be the cross-fit AIPW estimator of Proposition 3.11, estimated using the main experiment data $\check{W}_{1:n} = (\check{D}_i, X_i, Y_i(\check{D}_i))_{i=1}^n$, with regression estimators $\check{m}_d(\psi)$. The estimators $\hat{\theta}$ and $\hat{\theta}_{adj}$ have identical expansions

$$\begin{aligned} \hat{\theta} &= E_n[c(X_i)] + E_n[(D_i - p)(b - \hat{b})(X_i)]/c_p + R_n + O_p(n^{-1}) \\ \hat{\theta}_{adj} &= E_n[c(X_i)] + E_n[(\check{D}_i - p)(b - \check{b})(X_i)]/c_p + \check{R}_n \end{aligned}$$

for a constant c_p . The residual terms $\sqrt{n}R_n, \sqrt{n}\check{R}_n \Rightarrow \mathcal{N}(0, v)$ for the same variance $v > 0$. Denote the middle terms by $B_n = E_n[(D_i - p)(b - \hat{b})(X_i)]$ and $\check{B}_n = E_n[(\check{D}_i - p)(b - \check{b})(X_i)]$. These terms parameterize the error due to covariate imbalance between treatment and control units. Denote the pilot regression error $r_n^{pilot} = \max_d \|\hat{m}_d - m_d\|_{2,\psi}$ and main sample regression error $r_n^{main} = \max_d \|\check{m}_d - m_d\|_{2,\psi}$. It's easy to show that $\sqrt{n}B_n = O_p(r_n^{pilot})$ while $\sqrt{n}\check{B}_n = O_p(r_n^{main})$. We expect the pilot estimation error is larger $r_n^{pilot} \gg r_n^{main}$ if the pilot is much smaller than the main sample. This shows that

the design $D_{1:n} \sim \text{Loc}(\widehat{b}, p)$ behaves like a noisy version of the usual AIPW estimator $\widehat{\theta}_A$, with regressions estimated in the pilot instead of the main experiment.

Remark 5.7 (Model-Based Adjustment). The preceding discussion shows that the design $D_{1:n} \sim \text{Loc}(\widehat{b}, p)$ makes the unadjusted estimator $\widehat{\theta}$ behave like an AIPW estimator, up to lower order factors. However, the AIPW regression adjustments are estimated using data from the pilot sample instead of the main experiment, leading to worse control over the covariate imbalance term B_n . Instead, we suggest using a model-free strategy at design time, drawing $T_{1:n} \sim \text{Loc}(\psi, q(\psi))$ and $D_{1:n} \sim \text{Loc}(\psi, p(\psi))$ for a small set of covariates expected to be most predictive of outcomes and treatment effect heterogeneity. Model-based covariate adjustments such as AIPW or other regression adjustments can always be applied ex-post, using the main experiment data (or pooled data) to estimate the relevant regressions more precisely.

Residual Variance. Setting $\psi(X) = X$ in Equation 3.2 shows that under complete randomization $T_{1:n} \sim \text{CR}(q)$ and $D_{1:n} \sim \text{CR}(p)$ then $\sqrt{n_T}(\widehat{\theta} - \text{ATE}) \Rightarrow \mathcal{N}(0, V)$ with

$$V = \text{Var}(c(X)) + \text{Var}(b(X; p)) + E \left[\frac{\sigma_1^2(X)}{p} + \frac{\sigma_0^2(X)}{1-p} \right].$$

The middle term is the variance due to covariate imbalance, and the third term is the residual variance. We can think of $\text{Var}(b(X; p))$ as the “easier” term. We can make this term zero by any one of the following: (1) fine stratification on $\psi(X) = X$ under weak assumptions, (2) ex-post propensity re-weighting, or (3) ex-post regression adjustment under well-specification, or any combination of these methods. By contrast, after treatments have been assigned, neither regression adjustment nor propensity reweighting can help us further minimize the semiparametric variance bound

$$V_H(p) = \text{Var}(c(X)) + E \left[\frac{\sigma_1^2(X)}{p(X)} + \frac{\sigma_0^2(X)}{1-p(X)} \right]$$

The residual variance in this expression is the “harder” quantity. To affect it, we need to change the law of the data-generating process by changing the treatment propensity at design-time. Motivated by this, the next section studies pilot estimation of the optimal sampling and assignment propensities.

6 Inference Methods

This section provides new methods for asymptotically exact inference on the ATE under locally randomized designs, allowing researchers to report smaller confidence intervals that fully reflect the efficiency gains from finely stratified sampling and assignment. To do so, we generalize pairs-of-pairs type methods¹³ to accommodate designs with both finely stratified sampling and assignment, as well as varying propensities $q(\psi), p(\psi)$.

For each assignment group $g \in \mathcal{G}_n$, define the centroid $\bar{\psi}_g = |g|^{-1} \sum_{i \in g} \psi_i$. Let $\nu : \mathcal{G}_n \rightarrow \mathcal{G}_n$ be a bijective matching between groups satisfying $\nu(g) \neq g$, $\nu^2 = \text{Id}$, and the

¹³See [Abadie and Imbens \(2008\)](#) and [Bai et al. \(2021\)](#) for a recent analysis.

homogeneity condition

$$\frac{1}{n} \sum_{g \in \mathcal{G}_n} |\bar{\psi}_g - \bar{\psi}_{\nu(g)}|_2^2 = o_p(1) \quad (6.1)$$

In practice, ν is obtained by matching the group centroids $\bar{\psi}_g$ into pairs using the algorithm in Section 2. Let $\mathcal{G}_n^\nu = \{g \cup \nu(g) : g \in \mathcal{G}_n\}$ be the unions of paired groups formed by this matching. Define $a(g) = \sum_{i \in g} D_i$ and $k(g) = |g|$. Define the propensity weights $w_i^1 = (1 - p_i q_i)/(p_i q_i)^2$ and $w_i^0 = (1 - q_i(1 - p_i))/(q_i(1 - p_i))^2$. Finally, define the variance estimator components

$$\begin{aligned} \hat{v}_1 &= n^{-1} \sum_{g \in \mathcal{G}_n^\nu} \frac{1}{a(g) - 1} \sum_{i \neq j \in g} Y_i Y_j D_i D_j (w_i^1 w_j^1)^{1/2} \\ \hat{v}_0 &= n^{-1} \sum_{g \in \mathcal{G}_n^\nu} \frac{1}{(k - a)(g) - 1} \sum_{i \neq j \in g} Y_i Y_j (1 - D_i)(1 - D_j) (w_i^0 w_j^0)^{1/2} \\ \hat{v}_{10} &= n^{-1} \sum_{g \in \mathcal{G}_n} \frac{k}{a(k - a)}(g) \sum_{i, j \in g} Y_i Y_j D_i (1 - D_j) (q_i q_j)^{-1/2} \end{aligned}$$

Our inference strategy begins with the sample variance of the double-IPW estimator (Equation 3.4), which is consistent for the true asymptotic variance under an iid design, but too large under stratified designs. We correct this sample variance using the estimators above, which measure how well the stratification variables predict observed outcomes in local regions of the covariate space. Define the variance estimator

$$\hat{V} = \text{Var}_n \left(\frac{T_i(D_i - p(\psi_i))Y_i}{q(\psi_i)(p - p^2)(\psi_i)} \right) - \hat{v}_1 - \hat{v}_0 - 2\hat{v}_{10}. \quad (6.2)$$

Our main result shows that \hat{V} is consistent for the limiting variance of Theorem 3.10, enabling asymptotically exact inference.

Theorem 6.1 (Inference). *Assume the conditions of Theorem 3.10. If sampling $T_{1:n} \sim \text{Loc}(\psi, q(\psi))$ and assignment $D_{1:n} \sim \text{Loc}(\psi, p(\psi))$, then $\hat{V} = V + o_p(1)$.*

By Theorem 6.1, the confidence interval $\hat{C} = [\hat{\theta} \pm \hat{V}^{1/2} c_{1-\alpha/2}/\sqrt{n}]$ with $c_\alpha = \Phi^{-1}(\alpha)$ is asymptotically exact in the sense that $P(\text{ATE} \in \hat{C}) = 1 - \alpha + o(1)$. Importantly, note that the scaling is by number of eligible units n , not the experiment size $n_T = \sum_i T_i \leq n$.

7 Empirical Results

This section examines the finite sample properties of the estimation and inference methods studied above through a series of monte carlo simulation studies.

7.1 Simulations

Our first set of simulations focuses on the efficiency gain and inference properties of locally randomized *selection* into the experiment. Let $X_i \in \mathbb{R}^2$ and consider quadratic potential outcome models of the form

$$Y_i(0) = X_i' \beta(0) + \sigma_1^2(X_i) \epsilon_i(0) \quad Y_i(1) = X_i' \beta(1) + X_i' A X_i + \sigma_0^2(X_i) \epsilon_i(1)$$

We sample from the following DGP's

Model 1: $\beta(0) = (1, 1)$, $\beta(1) = (2, 2)$, $A_{12} = A_{21} = 1/2$, $A_{11} = A_{00} = 0$, with residual variance $\sigma_1^2 = \sigma_0^2 = 0.1$

Model 2: As in (1) but with $\beta(0) = (1, 1)$, $\beta(1) = (2, 2)$, $A = 0$

Model 3: As in (1) but with $\beta(0) = \beta(1) = (1, 1)$, $A = 0$

We let $X_{ij} \stackrel{\text{iid}}{\sim} 2(\text{Beta}(2, 2) - 1/2)$ and $\epsilon_i(d) \sim \mathcal{N}(0, 1)$.

Designs - In Table 1 we vary n , the number of units eligible for selection, fixing the experiment size $qn = \sum_i T_i = 100$. To highlight the marginal efficiency gains from representative selection, while also using locally randomized treatment assignment, we *fix* the assignment procedure to be “matched triples” on $\psi(x) = (x_1, x_2)$ with $p = 2/3$, as in Example 3.3. In our notation, we let selections $T_{1:n}$ and treatment assignments $D_{1:n}$ be given by

- (1) $T_{1:n} \sim \text{Loc}_n(\psi, q)$ (selection) with $n = 100/q$
- (2) $D_{1:n} \sim \text{Loc}_n(\psi, 2/3)$ (treatment assignment)

Evaluation Criteria - Table 1 presents metrics evaluating both the efficiency gains from locally-randomized selection, as well as the power and validity of our inference procedure. Note that $n = 100$ is the usual case of random selection. We evaluate each metric relative to this least efficient benchmark design using 1000 monte carlo replications drawn from the DGP's above. % Δ SD is the change in estimator standard deviation relative to random selection. ESS is the *effective sample size*. This is the size of an experiment with units selected completely at random needed to achieve the same variance as representative selection. Algebraically, $\text{ESS} = \frac{100}{1 + \% \Delta \text{Var}}$. For inference, % Δ Length is the reduction in confidence interval (CI) length, relative to the CI length under random selection.

Results - The variance reduction increases as the number of eligible units $n = 100/q$ is larger (q is smaller). This is expected from Theorem 3.10, which gives a reduction asymptotic variance of $-(1 - q) \text{Var}(c(X))$ from local randomization. The boost in effective sample size (ESS) is significant relative to completely random selection. Note that this comes essentially “for free” by being slightly more careful about choosing which units to enroll in the experiment. In Model 3, $c(x) = 0$ identically, so the theory predicts no efficiency gain due to selection. The small variance reduction seen in practice is likely a finite sample effect where balanced first-stage selection facilitates better matches during the second-stage assignment process.

Table 1 suggests our inference procedure is generally slightly conservative in finite samples. This is likely due to the between-group matching used for inference in Section 6 giving slightly worse matches than the groups produced by the design itself. For CI length, finite-sample exact inference would have % Δ Length = % Δ SD. In finite samples, our inference procedure gives a significant reduction in CI length, reflecting the increase in estimator precision due to selection.

Our second set of simulations focuses on the choice of $\psi(x)$. Asymptotically, including more baseline covariates always weakly improves estimator efficiency. However, as

| Model | n | Efficiency | | Inference | |
|-------|-----|---------------|-----|-----------|-------------------|
| | | % Δ SD | ESS | Coverage | % Δ Length |
| 1 | 100 | 0.0 | 100 | 95 | 0.0 |
| | 200 | -17.4 | 147 | 97 | -10.3 |
| | 300 | -22.8 | 168 | 97 | -13.6 |
| | 400 | -26.4 | 185 | 97 | -15.9 |
| 2 | 100 | 0.0 | 100 | 94 | 0.0 |
| | 200 | -17.7 | 148 | 96 | -9.5 |
| | 300 | -25.1 | 178 | 98 | -12.4 |
| | 400 | -28.6 | 196 | 97 | -14.5 |
| 3 | 100 | 0.0 | 100 | 96 | 0.0 |
| | 200 | -0.4 | 101 | 95 | -4.2 |
| | 300 | -5.5 | 112 | 96 | -4.9 |
| | 400 | -10.6 | 125 | 97 | -5.6 |

Table 1: Effect of Representative Selection

discussed in Section 3, the within-group matching discrepancies converge slowly in high dimensions, giving balancing rate $r_n^\psi = O(n^{-2/(d+1)})$. This suggests preferentially including covariates that are predictive of outcomes and treatment effect heterogeneity (predict functions $b_n(x)$ and $c(x)$), excluding “noise covariates” that contain little additional information. Consider the quadratic model and variable distributions above, but with $X_i \in \mathbb{R}^{10}$ and

$$\beta(0) = (1, 1, 1, .4, .4, .4, 0, 0, 0, 0) \quad \beta(1) = 2\beta(0) \quad A = (1/20)(\mathbf{1}\mathbf{1}' - \text{diag}(\mathbf{1}))$$

The first three covariates are important. The next three are less important. The last four provide no information about outcomes. We also include nonlinear interaction terms. For each $\psi(x)$ considered below, we select $q = 1/2$ of the eligible units n by matched pairs, then use matched triples for treatment assignment. Formally, we have

- (1) $T_{1:n} \sim \text{Loc}_n(\psi, 1/2)$ (selection)
- (2) $D_{1:n} \sim \text{Loc}_n(\psi, 2/3)$ (treatment assignment)

In Table 2 below, $\psi^*(x) = (b_n(x), c(x))$, the optimal design for joint selection and assignment given in Section 4. The evaluation criteria are the same as above, comparing to the least efficient design with selection and assignment done by complete randomization (CR).

Results - There are significant precision gains from locally randomized selection and assignment, as predicted by Theorem 3.10. Theorem 6.1 shows that our inference procedures are asymptotically exact, but we get slight conservativeness in finite samples, likely due to between-group match quality during inference being worse than the within-group matches. Note that, contrary to the asymptotic theory, including $x_1 \dots x_6$ is slightly *worse* than $x_1 \dots x_4$ in finite samples. Recall have the largest weight in the outcome

| (n, qn) | $\psi(x)$ | Efficiency | | Inference | |
|------------|-----------------|---------------|-----|-----------|-------------------|
| | | % Δ SD | ESS | Coverage | % Δ Length |
| (200, 100) | CR | 0.0 | 100 | 95 | 0.0 |
| | x_1, x_2 | -28.7 | 197 | 96 | -23.5 |
| | $x_1 \dots x_4$ | -37.4 | 256 | 97 | -28.2 |
| | $x_1 \dots x_6$ | -36.0 | 244 | 97 | -23.5 |
| | $x_1 \dots x_8$ | -33.9 | 229 | 98 | -19.6 |
| | $\psi^*(x)$ | -57.1 | 543 | 95 | -53.0 |
| (400, 200) | CR | 0.0 | 100 | 95 | 0.0 |
| | x_1, x_2 | -27.2 | 188 | 96 | -24.6 |
| | $x_1 \dots x_4$ | -40.4 | 282 | 98 | -32.0 |
| | $x_1 \dots x_6$ | -36.5 | 248 | 98 | -28.0 |
| | $x_1 \dots x_8$ | -33.2 | 224 | 98 | -22.8 |
| | $\psi^*(x)$ | -56.7 | 534 | 97 | -54.1 |

Table 2: Varying $\psi(x)$

model. Additionally including $x_4 \dots x_6$, which are less predictive of outcomes, degrades match quality for the important covariates $x_1 \dots x_3$, reducing efficiency in finite samples. Including noise covariates x_7, x_8 performs still worse. As in Theorem 4.4, the optimal design $\psi^*(x) = (b_n(x), c(x))$ gives the largest efficiency gains.

8 Conclusion

This paper proposes a flexible new family of designs that randomize within maximally homogeneous local groups. We show that experimenting on a representative sample of units increases estimator precision, providing a practical implementation by locally randomized sampling. We give novel asymptotically exact inference methods for covariate-adaptive sampling and assignment, allowing researchers to shrink their confidence intervals if they use our methods to design a representative experiment. We also applied our methods to the setting of design with a pilot experiment. By using pilot data to estimate the optimal treatment proportions, then locally randomizing within estimated propensity strata, we construct the first asymptotically fully efficient design in this setting.

This line of research can be extended in several interesting directions. It is clear how to extend our methods to $k > 2$ treatments, though this case is not included in our analysis. Accommodating continuous treatments is more difficult, given the discrete nature of the assignment process. Doing so would enable more efficient estimation of marginal effects and dose-response curves. We implicitly allow cluster-randomized trials by treating each cluster as a separate experimental unit. It could be helpful to practitioners to study this example more formally, explicitly accounting for heterogeneous cluster size. The design ordering we propose may not be optimal. For instance, we could reverse the process, first matching units into assignment pairs, then selecting a representative sample of pairs into

the experiment. The current ordering emphasizes representativeness at the cost of match quality during treatment assignment. We suspect these variations to be asymptotically equivalent, though they may differ in finite samples.

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9 Appendix

9.1 Finite Population Estimands

If the eligible units $\{1, \dots, n\}$ comprise the entire population of interest, then we may be interested in estimation and inference on the sample average treatment effect $\text{SATE} = E_n[Y_i(1) - Y_i(0)]$, or the average conditional treatment effect $E_n[c(X_i)]$, as in [Armstrong and Kolesár \(2021\)](#). Note that both estimands are defined over the full population of eligible units, not just the smaller set of experiment participants $\{i : T_i = 1\}$. For the SATE, suppose $T_{1:n} \sim \text{Loc}(X, q)$ and $D_{1:n} \sim \text{Loc}(X, p)$ and define the residual treatment effect variance $\sigma_\tau^2(X) = \text{Var}(Y(1) - Y(0)|X)$. Under the same conditions as Theorem 3.2, we have $\sqrt{n_T}(\hat{\theta} - \text{SATE}) \Rightarrow \mathcal{N}(0, V_{\text{SATE}})$ with

$$V_{\text{SATE}} = E \left[\frac{\sigma_1^2(X)}{p} + \frac{\sigma_0^2(X)}{1-p} - q\sigma_\tau^2(X) \right]. \quad (9.1)$$

The final variance component $E[\sigma_\tau^2(X)]$ is not identified, as in the case of SATE estimation under complete randomization. Setting $X = 1$ and $q = 1$ recovers the classical results in that setting. Observe that V_{SATE} decreases as the residual treatment effect heterogeneity $\sigma_\tau^2(X)$ increases. The negative sign reflects a competition between two opposing forces. To see this, let $\bar{Y}_i = Y_i(1)/p + Y_i(0)/(1-p)$ and consider the error decomposition

$$\hat{\theta} - \text{SATE} = \text{Cov}_n(T_i, Y_i(1) - Y_i(0))/q + \text{Cov}_n(D_i, \bar{Y}_i | T_i = 1) \quad (9.2)$$

The first term parameterizes the errors due to correlation between the sampling variables and treatment effects, and naturally increases with $\sigma_\tau^2(X) = \text{Var}(Y(1) - Y(0)|X)$. The second term is increasing in $\text{Cov}(Y(1), Y(0)|X)$ (decreasing in $\sigma_\tau^2(X)$) and larger in mean square, resulting in a net negative dependence on $\sigma_\tau^2(X)$. As $q \rightarrow 0$, the relative variance¹⁴ due to sampling increases, exactly cancelling the $E[\sigma_\tau^2(X)]$ factor due to assignment in the limit. Conservative inference for the SATE under stratified sampling and assignment can be based on either of the lower bounds $\sigma_\tau^2(X) \geq \sigma_1^2(X) + \sigma_0^2(X) - 2\sigma_1(X)\sigma_0(X) \geq 0$. We implement such a strategy in Section 6.

The average conditional treatment effect $\text{ACTE} = E_n[c(X_i)]$ is more difficult to motivate from a policy perspective. One potential application is a structural model of the form $Y_{it}(1) - Y_{it}(0) = c(X_i) + \epsilon_{it}$, with systematic component $c(X_i)$ mediated solely through observables X and transitory shock component ϵ_{it} .¹⁵ Intuitively, in this model the ACTE acts like a “denoised” version of the SATE. If the ϵ_{it} are uncorrelated, the ACTE estimated at time t is the best predictor of the SATE for a policy implemented at time $t + 1$. The proof of Theorem 3.2 shows that $\sqrt{n_T}(\hat{\theta} - E_n[c(X_i)]) \Rightarrow \mathcal{N}(0, V_c)$ with identified variance $V_c = E[\sigma_1^2(X)/p + \sigma_0^2(X)/(1-p)] \geq V_{\text{SATE}}$.

Remark 9.1 (Design-based Theory). It is interesting to compare our results with those obtained under a finite population “design-based” framework. The design based approach in statistics takes the sequence of finite populations $W_{1:n} = (X_i, Y_i(0), Y_i(1))_{i=1}^n$ as non-random, with all randomness due to treatment assignment. For example, see [Li and Ding](#)

¹⁴The normalization $\sqrt{n_T}(\hat{\theta} - \text{SATE})$ by experiment size $n_T = \sum_i T_i$ holds the number of sampled units fixed.

¹⁵Inference on a more general denoised SATE parameter in a model with transitory shocks is studied in [Deeb and de Chaisemartin \(2022\)](#).

(2018) or Wang et al. (2021). By contrast, here we impose a nonparametric generative model $W_{1:n} \sim F$, requiring F to have certain finite moments in Assumption 3.1. However, our analysis shows that $\sqrt{n_T}(\hat{\theta} - \text{SATE})|W_{1:n} \Rightarrow \mathcal{N}(0, V_{\text{SATE}})$ conditional on the data $W_{1:n}$. Since we condition on $W_{1:n}$, the generative model $W_{1:n} \sim F$ does not contribute to estimator variance. There is no hypothetical superpopulation, and all randomness is due to the sampling and assignment variables $T_{1:n}$ and $D_{1:n}$, just as in design-based theory. The distribution F imposes homogeneity on the sequence of finite populations $W_{1:n}$ as $n \rightarrow \infty$, ensuring that quantities such as $E_n[W_i]$ converge to well-defined limits. By contrast, in the finite population framework the convergence $E_n[W_i] \rightarrow W_\infty$ is assumed ex-post as a high-level condition on the sequence of populations $W_{1:n}$.

The finite population framework appears to be formally more general than our model, but also potentially more cumbersome. For example, the outcome function $E[Y(d)|X]$, which plays an important conceptual role in many parts of causal inference, does not appear to have a nonparametric design-based analogue. It would be interesting to characterize when this extra generality has “bite” and leads to substantively different prescriptions for estimation and inference than the SATE theory presented here.

9.2 Remarks

Remark 9.2 (Increasing Stratification Condition). At a high level, the condition $\psi_1 \subseteq \psi_2$ allows us to ignore the complicated effect of first-stage sampling on the joint distribution of sampled stratification variables $(\psi_{1,i})_{i:T_i=1}$. To see the problem, observe that if $q = 1/k$ then $T_i = T_j = 1$ implies that i, j cannot have been matched together during sampling. Then, for instance, we expect

$$P(|\psi_{1,i} - \psi_{1,j}| < \epsilon) > P(|\psi_{1,i} - \psi_{1,j}| < \epsilon \mid T_i = T_j = 1). \quad (9.3)$$

If $E[Y_i(d)|\psi_{1,i}] \neq 0$, such changes to the joint distribution will show up in the conditional variance $\text{Var}(\hat{\theta}|X_{1:n}, T_{1:n})$ in complicated ways that depend on the details of the matching algorithm. However, we can use the fact that the assignment stratification “partials out” $Y_i(d)$ so that, to first order, only the residuals $u_i = Y_i - E[Y_i(d)|\psi_{2,i}]$ enter this conditional variance. If $\psi_1 \subseteq \psi_2$ the residuals u_i are less affected by selection on $\psi_{1,i}$. For example, we can show $E[u_i u_j | T_i = T_j = 1] = E[u_i u_j] = 0$. We conjecture the theorem may be true without this condition if the effects in Equation 9.3 are lower order, but leave the detailed study of this issue for specific matching procedures to future work.

10 Proofs

10.1 Matching

Theorem 10.1 (Matching). *Consider triangular arrays $(\psi_{i,n})_{i=1}^n \subseteq \mathbb{R}^d$ and $(p_{i,n})_{i=1}^n \subseteq \mathbb{Q}$ with levels $p_{i,n} \in L_n = \{a_l/k_l\}$ and $\bar{k}_n = \max\{k_l : a_l/k_l \in L_n\}$. For any sequence of subsets $S_n \subseteq [n]$, there exists a sequence of partitions $(\mathcal{G}_n)_{n \geq 1}$ of S_n such that $\mathcal{G}_n = \cup_l \mathcal{G}_{nl}$ and \mathcal{G}_{nl} partitions $S_n \cap \{i : p_{i,n} = a_l/k_l\}$ with $|g| = k_l$ for all but one non-empty group*

$g \in \mathcal{G}_n$. The partition \mathcal{G}_n satisfies the homogeneity rate

$$n^{-1} \sum_{g \in \mathcal{G}_n} \frac{1}{|g|} \sum_{i,j \in g} |\psi_{i,n} - \psi_{j,n}|_2^2 \leq (1 \vee \max_{i=1}^n |\psi_{i,n}|_2^2) \cdot O((n/\bar{k}_n |L_n|)^{-2/(d+1)}). \quad (10.1)$$

Let \mathcal{G}_{nl}^* be the optimal partition of $S_n \cap \{i : p_{i,n} = a_l/k_l\}$, solving the minimization

$$\mathcal{G}_{nl}^* = \operatorname{argmin}_{\mathcal{G}_{nl}} \left[n^{-1} \sum_{g \in \mathcal{G}_{nl}} \sum_{i,j \in g} |\psi_{i,n} - \psi_{j,n}|_2^2 \right]$$

subject to the group size constraint $|g| = k_l$. Then $\mathcal{G}_n^* = \cup_l \mathcal{G}_{nl}^*$ satisfies the homogeneity rate in Equation 10.1.

Proof. First, note that for any partition \mathcal{G}_n

$$\frac{1}{n} \sum_{g \in \mathcal{G}_n} \frac{1}{|g|} \sum_{i,j \in g} |\psi_{i,n} - \psi_{j,n}|_2^2 \leq (1 \vee \max_{i=1}^n |\psi_{i,n}|_2^2) \cdot \frac{1}{n} \sum_{g \in \mathcal{G}_n} \frac{1}{|g|} \sum_{i,j \in g} \left| \frac{\psi_{i,n} - \psi_{j,n}}{1 \vee \max_{i=1}^n |\psi_{i,n}|_2} \right|_2^2$$

Clearly, $|\psi_{i,n}|_2 / (1 \vee \max_{i=1}^n |\psi_{i,n}|_2) \leq 1$ for all $1 \leq i \leq n$. Then by recentering, it suffices to show the claim for $\psi_{i,n} \in [0, 1]^d$ for all $1 \leq i \leq n$ as $n \rightarrow \infty$. Consider blocks of the form $B = \{\sum_{k=1}^d x_k e_k : x_k \in [s_k/m, (s_k+1)/m]\}$ for indices $\{s_k\}_{k=1}^d \subseteq \{0, \dots, m-1\}$. Fix an ordering of these blocks $(B_l)_{l=1}^{m^d}$ such that B_l and B_{l+1} are adjacent for all l . Intuitively, this forms a “block path” through $[0, 1]^d$, see [Bai et al. \(2021\)](#) for a picture. Define $l(i) = \min_{l=1}^{m^d} \{l : \psi_{i,n} \in B_l\}$.

Algorithm - Fix $p_a \in L_n$ and form groups $(g_{a,s})_{s=1}^{n_a}$ of units $g_{a,s} \subseteq \{i : p_{i,n} = p_a = f_a/k_a\}$ by induction as follows. (1) Form groups of size k_a arbitrarily among $\{i : l(i) = 1\}$, and $p_n(X_i, \xi_n) = p_a$, possibly using external randomness π_n . This process results in at most one partially completed group with $|g_{a,s'}| < k_a$. While $|g_{a,s'}| < k_a$, do the following: (2) increment $l \rightarrow l+1$ and (3) add an unmatched unit i with $l(i) = l+1$ to this group. If $|g_{a,s'}| = k_a$, stop. If $|g_{a,s'}| < k_a$ and there are unmatched units in B_{l+1} , goto (3). If $|g_{a,s'}| < k_a$ and there are no unmatched units in B_{l+1} , increment l and go to (1). Suppose that $g_{a,s'}$ is completed with a unit from $B_{l'}$. Then repeat the process above starting with the next group $g_{a,s'+1}$ and the units in block $B_{l'}$. Since there are $n < \infty$ units, this process terminates. Repeat this for each $a = 1, \dots, |L_n|$. By construction, this creates groups $\mathcal{G}_n = \{g_{a,s} : 1 \leq a \leq |L_n|, 1 \leq s \leq n\}$ with the ordering property

$$l(i) \leq l(j) \quad \forall i \in g_{a,s}, j \in g_{a,s'} \quad s < s' \quad a = 1, \dots, |L_n| \quad (10.2)$$

Fix an indexing of all within-group pairs $(\mathbf{p}_{a,s,t})_{t=1}^{k_a^2 - k_a} \equiv \{(i, j) : i \neq j; i, j \in g_{a,s}\}$, and denote $\mathbf{p}_{a,s,t} = (i_{a,s,t}, j_{a,s,t})$. Define $E_{a,s,t} = \{l(i_{a,s,t}) = l(j_{a,s,t})\}$, the event that a pair is in the same element of the block partition. With this notation, we have

$$\begin{aligned} n^{-1} \sum_{g \in \mathcal{G}_n} \frac{1}{|g|} \sum_{i,j \in g} |\psi_{i,n} - \psi_{j,n}|_2^2 &= n^{-1} \sum_{a=1}^{|L_n|} \sum_{s=1}^n k_a^{-1} \sum_{i,j \in g_{a,s}} |\psi_{i,n} - \psi_{j,n}|_2^2 \\ &\leq n^{-1} \sum_{a=1}^{|L_n|} \sum_{s=1}^n k_a^{-1} \sum_{t=1}^{k_a^2 - k_a} |\psi_{i_{a,s,t},n} - \psi_{j_{a,s,t},n}|_2^2 \mathbf{1}(E_{a,s,t} \neq \emptyset) \end{aligned}$$

(1) Suppose $E_{a,s,t}$ occurs. Then $d_{a,s,t} \equiv |\psi_{i_{a,s,t},n} - \psi_{j_{a,s,t},n}|_2 \leq \max_{l=1}^{m^d} \text{diam}(B_l, |\cdot|_2) \leq \sqrt{d}/m$ on this event. Then we may bound

$$\begin{aligned} n^{-1} \sum_{a=1}^{|L_n|} \sum_{s=1}^n k_a^{-1} \sum_{t=1}^{k_a^2 - k_a} d_{a,s,t}^2 \mathbb{1}(E_{a,s,t}) \mathbb{1}(g_{a,s} \neq \emptyset) &\leq \frac{d}{nm^2} \sum_{a=1}^{|L_n|} \sum_{s=1}^n k_a^{-1} \sum_{t=1}^{k_a^2 - k_a} \mathbb{1}(g_{a,s} \neq \emptyset) \\ &\leq \frac{d}{nm^2} \sum_{a=1}^{|L_n|} \sum_{s=1}^n k_a \mathbb{1}(g_{a,s} \neq \emptyset) \leq \frac{d}{m^2} \end{aligned}$$

The final inequality since the double sum exactly counts the number of units in the sample (by group) $\sum_{a=1}^{|L_n|} \sum_{s=1}^n k_a \mathbb{1}(g_{a,s} \neq \emptyset) = |S_n| \leq n$.

(2) Now consider the terms where $E_{a,s,t}$ does not occur. Fix any such pair $(i_{a,s,t}, j_{a,s,t})$. Without loss, suppose the block membership $l(i_{a,s,t}) < l(j_{a,s,t})$. For $l(i_{a,s,t}) \leq l \leq l(j_{a,s,t})$, define a sequence z_l as follows. $z_{l(i_{a,s,t})} = \psi_{i_{a,s,t},n}$, $z_{l(j_{a,s,t})} = \psi_{j_{a,s,t},n}$ and $z_l \in B_l$ chosen arbitrarily otherwise. Note that for $x \in B_l$ and $y \in B_{l+1}$, by construction of the contiguous blocks $|x - y|_2 \leq 2\sqrt{d}/m$. Then by telescoping and triangle inequality, on the event $E_{a,s,t}^c$

$$d_{a,s,t} = |\psi_{i_{a,s,t},n} - \psi_{j_{a,s,t},n}|_2 \leq \sum_{l=l(i_{a,s,t})}^{l(j_{a,s,t})-1} |z_{l+1} - z_l|_2 \leq \frac{2\sqrt{d}}{m} \cdot [l(j_{a,s,t}) - l(i_{a,s,t})]$$

Note also that if $x, y \in [0, 1]^d$ then $|x - y|_2^2 = d(|x - y|_2 / \sqrt{d})^2 \leq \sqrt{d}|x - y|_2$, using $c^2 \leq c$ for $0 \leq c \leq 1$. In particular, we have $d_{a,s,t}^2 \leq \sqrt{d} \cdot d_{a,s,t}$.

$$\begin{aligned} n^{-1} \sum_{a=1}^{|L_n|} \sum_{s=1}^n k_a^{-1} \sum_{t=1}^{k_a^2 - k_a} d_{a,s,t}^2 \mathbb{1}(E_{a,s,t}^c) &\leq \sqrt{d} n^{-1} \sum_{a=1}^{|L_n|} \sum_{s=1}^n k_a^{-1} \sum_{t=1}^{k_a^2 - k_a} d_{a,s,t} \mathbb{1}(E_{a,s,t}^c) \\ &\leq \sum_{a=1}^{|L_n|} \frac{2d}{mnk_a} \sum_{t=1}^{k_a^2 - k_a} \sum_{s=1}^n [l(i_{a,s,t}) - l(j_{a,s,t})] \mathbb{1}(E_{a,s,t}^c) \leq \sum_{a=1}^{|L_n|} \frac{2\sqrt{d}}{mnk_a} \sum_{t=1}^{k_a^2 - k_a} m^d \\ &\leq \sum_{a=1}^{|L_n|} \frac{2\sqrt{d}m^d}{mn} k_a \leq 2\sqrt{d}|L_n|\bar{k}_n n^{-1} m^{d-1} \end{aligned}$$

The second inequality follows by the ordering property in equation 10.2 above, since for each $t = 1, \dots, k_a^2 - k_a$, the intervals $([l(i_{a,s,t}), l(j_{a,s,t})])_{s=1}^n$ are non-overlapping, and there are at most m^d blocks. Summarizing the above work, we have shown that

$$\begin{aligned} n^{-1} \sum_{a=1}^{|L_n|} \sum_{s=1}^n k_a^{-1} \sum_{i,j \in g_{a,s}} |\psi_{i,n} - \psi_{j,n}|_2^2 &\leq n^{-1} \sum_{a=1}^{|L_n|} \sum_{s=1}^n k_a^{-1} \sum_{t=1}^{k_a^2 - k_a} d_{a,s,t}^2 \mathbb{1}(E_{a,s,t}) + d_{a,s,t}^2 \mathbb{1}(E_{a,s,t}^c) \\ &\leq \frac{d}{m^2} + \frac{2\sqrt{d}|L_n|\bar{k}_n m^{d-1}}{n} \end{aligned}$$

Setting $m \asymp (n/(|L_n|\bar{k}_n))^{1/(d+1)}$ gives the rate

$$n^{-1} \sum_{a=1}^{|L_n|} \sum_{s=1}^n k_a^{-1} \sum_{i,j \in g_{a,s}} |\psi_{i,n} - \psi_{j,n}|_2^2 = O\left((n/(|L_n|\bar{k}_n))^{-2/(d+1)}\right)$$

This finishes the proof of the first statement. The second statement follows by optimality of \mathcal{G}_{nl}^* and comparison with the groups just constructed. \square

10.2 CLT

Lemma 10.2 (Balancing). *Suppose $T_{1:n} \sim \text{Loc}(\psi_1, q(x))$ and $D_{1:n} \sim \text{Loc}(\psi_2, p(x))$ with $(\psi_1(X), q(X)) \in \sigma(\psi_2(X))$, and ψ_1, ψ_2, q , and p possibly dependent on $\xi \perp\!\!\!\perp W_{1:n}$. Suppose $E[|\psi_1(X)|^{\alpha_1}] < \infty$ for $\alpha_1 > d(\psi_2) + 1$ and $E[|\psi_2(X)|^{\alpha_1}] < \infty$ for $\alpha_2 > d(\psi_2) + 1$. If $E[f(\psi_2(X), p(X), \xi)^2 | \xi] < \infty$ ξ -a.s. then*

$$E_n[T_i(D_i - p(X_i))f(\psi_2(X_i), p(X_i), \xi)] = o_p(n^{-1/2})$$

If $E[g(\psi_1(X), q(X), \xi)^2 | \xi] < \infty$ ξ -a.s. then $E_n[(T_i - q(X_i))g(\psi_1(X_i), q(X_i), \xi)] = o_p(n^{-1/2})$. Suppose further that $\bar{k}_n |L_n| = o(n^{1-(d+1)/(\alpha_1 \wedge \alpha_2)})$. Then $E_n[T_i(D_i - p_n(X_i))f(\psi_2(X_i), \xi)] = o_p(n^{-1/2})$ and $E_n[(T_i - q_n(X_i))g(\psi_1(X_i), \xi)] = o_p(n^{-1/2})$ under the same conditions.

Proof. We begin with the first claim. Fix a sequence c_n with $c_n \rightarrow \infty$ and $c_n^2 = o((\bar{k}_n |L_n|)^{-2/(d+1)} n^{2/(d+1)-2/\alpha_2})$. Work on the assumed almost sure event and fix ξ . Denote $Z = (\psi_2(X), p(X))$. Define $L_2(Z) = \{g(Z) : E[g(Z)^2 | \xi] < \infty\}$. By assumption, $f(Z, \xi) \in L_2(Z)$. Our strategy is to approximate $f(Z, \xi)$ by Lipschitz functions. Define the spaces $\mathcal{L}_n = \{g(Z) \in L_2(Z) : |g|_{\text{lip}} \vee |g|_\infty \leq c_n\}$ and let $g_n \in \mathcal{L}_n$ be such that $E[(g_n(Z) - f(Z))^2] \leq 2 \inf_{g \in \mathcal{L}_n} E[(g(Z) - f(Z))^2]$. We claim that $|g_n - f|_{2,Z}^2 \rightarrow 0$. Let $\epsilon > 0$ and note that by Lemma 10.11 there exists a function $|h|_{\text{lip}} \vee |h|_\infty < \infty$ with $|h - f|_{2,Z}^2 < \epsilon$. Since $c_n \rightarrow \infty$, there exists N such that $c_n \geq |h|_{\text{lip}} \vee |h|_\infty$ for all $n \geq N$. Then for $n \geq N$ we have $|g_n - f|_{2,Z}^2 \leq 2 \inf_{g \in \mathcal{L}_n} |g - f|_{2,Z}^2 \leq 2|h - f|_{2,Z}^2 < 2\epsilon$ since $h \in \mathcal{L}_n$. Since ϵ was arbitrary, this shows the claim. Summarizing our work so far, we found $g_n(Z, \xi)$ such that $E[(f(Z, \xi) - g_n(Z, \xi))^2 | \xi] = o(1)$, $|g_n(Z, \xi) - g_n(Z', \xi)| \leq c_n |Z - Z'|_2$, and $|g_n(\cdot, \xi)|_\infty \leq c_n$ ξ -a.s. Now we expand

$$E_n[T_i(D_i - p_i)f(Z_i, \xi)] = E_n[T_i(D_i - p_i)(f(Z_i, \xi) - g_n(Z_i, \xi))] + E_n[T_i(D_i - p_i)g_n(Z_i, \xi)]$$

Denote these terms A_n, B_n . Let $\mathcal{F}_n = \sigma((\psi_2)_{1:n}, (p_i)_{1:n}, \pi_n, \tau^t, \xi)$. Since $(\psi_1, q) \in \sigma(\psi_2)$, we have $\mathcal{G}_n, f(Z_i)_{1:n}, g_n(Z_i)_{1:n} \in \mathcal{F}_n$. Then by Lemma 10.17 $E[A_n | \mathcal{F}_n] = 0$ and

$$\text{Var}(\sqrt{n} E_n[T_i(D_i - p_i)(f(Z_i, \xi) - g_n(Z_i, \xi))] | \mathcal{F}_n) \leq 2 E_n[(f(Z_i, \xi) - g_n(Z_i, \xi))^2] = o_p(1)$$

The equality by conditional Markov (Lemma 10.10) since $E[E_n[(f(Z_i, \xi) - g_n(Z_i, \xi))^2] | \xi] = |f - g_n|_{2,Z}^2 = o(1)$ ξ -a.s. as shown above. Then $A_n = o_p(n^{-1/2})$, again by conditional Markov. Next consider B_n . By Lemma 10.17, we have $E[B_n | \mathcal{F}_n] = 0$. Using the claim

and Lemma 10.17, we calculate

$$\begin{aligned}
\text{Var}(\sqrt{n}B_n|\mathcal{F}_n) &\leq n^{-1} \sum_g \frac{1}{|g|} \sum_{i,j \in g} (g_n(Z_i, \xi) - g_n(Z_j, \xi))^2 + n^{-1} \bar{k}_n |L_n| \max_{i=1}^n g_n(Z_i, \xi)^2 \\
&\leq c_n^2 n^{-1} \sum_g \frac{1}{|g|} \sum_{i,j \in g} |Z_i - Z_j|_2^2 + n^{-1} \bar{k}_n |L_n| c_n^2 \\
&= c_n^2 n^{-1} \sum_g \frac{1}{|g|} \sum_{i,j \in g} (|\psi_{2,i} - \psi_{2,j}|_2^2 + |p_i - p_j|^2) + n^{-1} \bar{k}_n |L_n| c_n^2 \\
&= c_n^2 o_p(n^{2/\alpha_2} (n/(\bar{k}_n |L_n|))^{-2/(d+1)}) + n^{-1} \bar{k}_n |L_n| c_n^2 = o_p(1)
\end{aligned}$$

The second to last equality since $p_i = p_j$ for any $i, j \in g$, $\forall g$ and by Theorem 10.1, using the fact that $\max_{i=1}^n |\psi_{2,i}|^2 = o_p(n^{2/\alpha_2})$ by Lemma 10.14. The final equality follows by our choice of c_n . Then $B_n = o_p(n^{-1/2})$, again by conditional Markov. This finishes the proof. The conclusion for sampling variables follows by the same proof, setting $T_i, q_i \rightarrow 1$, $D_i \rightarrow T_i$ and $p_i \rightarrow q_i$ and $\psi_2 \rightarrow \psi_1$. The second set of claims follows by setting $Z = \psi_2(X)$ and $p_i \rightarrow p_{i,n}$ in the above proof. \square

Assumption 10.3. *Consider the following assumptions*

- (A1) $E[Y(d)^2] < \infty$ and ψ_1, ψ_2, q, p as in Theorem 3.10 with $q_i, p_i \in [\delta, 1 - \delta] \subseteq (0, 1)$.
(A2) $E[Y(d)^4] < \infty$ and propensities $\hat{p}_{i,n} = \hat{p}_{i,n}(X_i, \xi_n)$, $\hat{q}_{i,n} = \hat{q}_{i,n}(X_i, \xi_n)$, $p_i = p(X_i)$, and $q_i = q(X_i)$ for $\xi_n \perp\!\!\!\perp \sigma(W_{1:n}, \tau^t, \tau^d, \pi_n)$. $\hat{p}_{i,n}, p_i \in [\delta, 1 - \delta] \subseteq (0, 1)$ and $\hat{q}_{i,n}, q_i \geq \delta$. Also $E_n[(\hat{p}_{i,n} - p_i)^2] = o_p(1)$ and $E_n[(\hat{q}_{i,n} - q_i)^2] = o_p(1)$.

Theorem 10.4 (CLT). (1) Suppose Assumption 10.3 (A1) holds. Let $T_{1:n} \sim \text{Loc}(\psi_1, q(x))$ and $D_{1:n} \sim \text{Loc}(\psi_2, p(x))$. Define $c_{1,i} = E[Y_i(1) - Y_i(0)|\psi_{1,i}, q_i, \xi]$. Then $\sqrt{n}(E_n[c_{1,i}] - \text{ATE})|\xi \Rightarrow \mathcal{N}(0, V_1)$ with $V_1 = \text{Var}(c_{1,i}|\xi)$. Define

$$u_i = q_i^{-1}(E[Y_i(1) - Y_i(0)|\psi_{1,i}, q_i, \xi] - E[Y_i(1) - Y_i(0)|\psi_{2,i}, p_i, \xi])$$

and $\mathcal{F}_{0,n} = \sigma(\xi, \pi_n, (\psi_{1,i})_{1:n}, (q_i)_{1:n}, \tau^t)$. Then $\sqrt{n}E_n[T_i u_i]|\mathcal{F}_{0,n} \Rightarrow \mathcal{N}(0, V_1)$ with asymptotic variance

$$V_2 = E[q_i^{-1} \text{Var}(E[Y_i(1) - Y_i(0)|\psi_{2,i}, p_i, \xi]|\psi_{1,i}, q_i, \xi)|\xi].$$

Define the residuals $\epsilon_i^d = Y_i(d) - E[Y_i(d)|\psi_{2,i}, p_i, \xi]$ and $\mathcal{G}_n = \sigma(\xi, \pi_n, \tau^t, \tau^d, (\psi_{2,i})_{1:n}, (p_i)_{1:n})$. Then $\sqrt{n}E_n[T_i D_i \epsilon_i^1 / (p_i q_i) + T_i(1 - D_i) \epsilon_i^0 / ((1 - p_i) q_i)]|\mathcal{G}_{0,n} \Rightarrow \mathcal{N}(0, V_3)$ with variance

$$V_3 = E \left[\frac{\text{Var}(Y_i(1)|\psi_{2,i}, p_i, \xi)}{q_i p_i} + \frac{\text{Var}(Y_i(0)|\psi_{2,i}, p_i, \xi)}{q_i(1 - p_i)} \middle| \xi \right].$$

(2) Alternatively, suppose Assumption 10.3 (A2) holds. Define the residuals $\epsilon_i^d = Y_i(d) - E[Y_i(d)|X]$. Let $\psi_{k,n} = \psi_{k,n}(X)$ for $k = 1, 2$. Let $T_{1:n} \sim \text{Loc}(\psi_{1,n}, \hat{q}_n(x))$ and $D_{1:n} \sim \text{Loc}(\psi_{2,n}, \hat{p}_n(x))$ and define $\mathcal{G}_{0,n} = \sigma(\pi_n, \xi_n, X_{1:n}, \tau^t, \tau^d)$. Then $\sqrt{n}E_n[T_i D_i \epsilon_i^1 / (\hat{p}_{i,n} \hat{q}_{i,n}) + T_i(1 - D_i) \epsilon_i^0 / ((1 - \hat{p}_{i,n}) \hat{q}_{i,n})]|\mathcal{G}_{0,n} \Rightarrow \mathcal{N}(0, V_3)$ with variance

$$V_3 = E \left[\frac{1}{q(X_i)} \left(\frac{\sigma_1^2(X_i)}{p(X_i)} + \frac{\sigma_0^2(X_i)}{1 - p(X_i)} \right) \right].$$

Proof. First consider (1). Define $\mathcal{H}_{0,n} = \sigma(\xi)$ and $\mathcal{H}_{k,n} = \sigma(\xi, (c_{1,i})_{i=1:k})$ for $k \geq 1$. Claim that $(\mathcal{H}_{k,n}, c_{1,k} - \text{ATE})_{k \geq 1}$ is an MDS. Adaptation is clear. For the MDS property, $E[c_{1,i} | \mathcal{F}_{i-1,n}] = E[c_{1,i} | \xi, (c_{1,k})_{k=1:i-1}] = E[c_{1,i} | \xi] = E[Y_i(1) - Y_i(0) | \xi] = \text{ATE}$ by tower law and independence. Similarly, $E[(c_{1,i} - \theta)^2 | \mathcal{F}_{i-1,n}] = E[(c_{1,i} - \theta)^2 | \xi] = \text{Var}(c_{1,i} | \xi)$. To finish, we show the Lindeberg condition. Note that by conditional Jensen inequality $\text{Var}(c_{1,i} | \xi) \leq E[c_{1,i}^2 | \xi] \leq E[(Y_1 - Y_0)^2] < \infty$. Then we have $(c_{1,i} - \theta)^2 \mathbb{1}((c_{1,i} - \theta)^2 > n\epsilon) \rightarrow 0$ ξ -a.s., so by dominated convergence as $n \rightarrow \infty$

$$E_n[E[(c_{1,i} - \theta)^2 \mathbb{1}((c_{1,i} - \theta)^2 > n\epsilon^2) | \xi]] = E[(c_{1,i} - \theta)^2 \mathbb{1}((c_{1,i} - \theta)^2 > n\epsilon^2) | \xi] \rightarrow 0$$

Then the Lindberg condition is satisfied, so by Theorem 10.6 $\sqrt{n}(E_n[c_{1,i}] - \text{ATE}) | \xi \Rightarrow \mathcal{N}(0, V_1)$ with the claimed limiting variance. Consider the second claim. Define $\mathcal{F}_{k,n} = \sigma(\mathcal{F}_{0,n}, (\psi_{2,i})_{i=1:k}, (p_i)_{i=1:k})$ for $k \geq 1$. Claim that $(T_k u_k, \mathcal{F}_{k,n})_{k \geq 1}$ is an MDS. Adaptation is clear from the definitions. Next we show the MDS property. Note that $T_{1:n} \in \mathcal{F}_{0,n}$ since $T_{1:n} = G(\tau^t, (\psi_{1,i})_{1:n}, (q_i)_{1:n}, \pi_n)$. Note the crucial fact that $(A, B) \perp\!\!\!\perp C \implies A \perp\!\!\!\perp C | B$. By randomization we have $(u_{k+1}, (\psi_{1,i})_{1:n}, (q_i)_{1:n}, \xi, (\psi_{2,i})_{i=1:k}, (p_i)_{i=1:k}) \perp\!\!\!\perp (\tau^t, \pi_n)$. Combining this with the crucial fact, gives conditional independence from (π_n, τ^t)

$$\begin{aligned} E[T_{k+1} u_{k+1} | \mathcal{F}_{k,n}] &= T_{k+1} E[u_{k+1} | \tau^t, (\psi_{1,i})_{1:n}, (q_i)_{1:n}, \pi_n, (\psi_{2,i})_{i=1:k}, (p_i)_{i=1:k}, \xi] \\ &= T_{k+1} E[u_{k+1} | (\psi_{1,i})_{1:n}, (q_i)_{1:n}, (\psi_{2,i})_{i=1:k}, (p_i)_{i=1:k}, \xi] \end{aligned}$$

Next, note that since $(X_i)_{i=1}^n$ are iid and $\xi \perp\!\!\!\perp X_{1:n}$, we have $f(X_{k+1}, \xi) \perp\!\!\!\perp g(X_{-(k+1)}, \xi) | \xi$ for any functions f, g . Let $g(X_{-(k+1)}, \xi) = ((\psi_{1,i})_{i \neq k+1}, (q_i)_{i \neq k+1}, \xi, (\psi_{2,i})_{i=1:k}, (p_i)_{i=1:k})$ and $f(X_{k+1}, \xi) = u_{k+1}$, giving conditional independence

$$\begin{aligned} E[u_{k+1} | (\psi_{1,i})_{1:n}, (q_i)_{1:n}, (\psi_{2,i})_{i=1:k}, (p_i)_{i=1:k}, \xi] &= E[u_{k+1} | \psi_{1,k+1}, q_{k+1}, \xi] \\ &= q_{k+1}^{-1} E[Y_{k+1}(1) - Y_{k+1}(0) | \psi_{1,k+1}, q_{k+1}, \xi] \\ &\quad - q_{k+1}^{-1} E[E[Y_{k+1}(1) - Y_{k+1}(0) | \psi_{2,k+1}, p_{k+1}, \xi] | \xi, \psi_{1,k+1}, q_{k+1}] = 0 \end{aligned}$$

The final equality by tower law since $\sigma(\xi, \psi_{1,k+1}, q_{k+1}) \subseteq \sigma(\psi_{2,k+1}, p_{k+1}, \xi)$ by the increasing stratification assumption. This shows the MDS property. Next, we compute the variance process $\Sigma_{k,n} = \sum_{i=1}^k E[(T_i u_i)^2 | \mathcal{F}_{i-1,n}]$. By the exact argument above $E[T_i u_i^2 | \mathcal{F}_{i-1,n}] = T_i E[u_i^2 | \psi_{1,i}, q_i, \xi]$, so that $E[T_i u_i^2 | \mathcal{F}_{i-1,n}] = T_i E[E[u_i^2 | \mathcal{F}_{i-1,n}] | \mathcal{F}_{i-1,n}] = T_i E[u_i^2 | \psi_{1,i}, q_i, \xi]$, since $(\psi_{1,i}, q_i, \xi) \in \mathcal{F}_{0,n} \subseteq \mathcal{F}_{i-1,n}$. In particular, we have shown that $\Sigma_{k,n} \in \mathcal{F}_{0,n}$ for all k, n , satisfying the variance condition of Proposition 10.6. We have

$$\Sigma_n = E_n[T_i E[u_i^2 | \psi_{1,i}, q_i, \xi]] = E_n[(T_i - q_i) E[u_i^2 | \psi_{1,i}, q_i, \xi]] + E_n[q_i E[u_i^2 | \psi_{1,i}, q_i, \xi]]$$

Call these terms A_n, B_n . By conditional Jensen and Young's, we have $E[|E[u_i^2 | \psi_{1,i}, q_i, \xi]|] \leq E[u_i^2] \lesssim E[(Y_i(1) - Y_i(0))^2] \lesssim \sum_{d=0,1} E[Y_i(d)^2] < \infty$. Then $A_n = o_p(1)$ by Lemma 10.17. For the second term B_n , we have

$$\begin{aligned} E[B_n | \xi] &= E[E_n[q_i E[u_i^2 | \psi_{1,i}, q_i, \xi]] | \xi] = E[q_i E[u_i^2 | \psi_{1,i}, q_i, \xi] | \xi] \\ &= E[q_i^{-1} \text{Var}(E[Y_i(1) - Y_i(0) | \psi_{2,i}, p_i, \xi] | \psi_{1,i}, q_i, \xi) | \xi] \end{aligned}$$

Moreover, similar to the calculation for A_n , conditional Jensen and Young's show that $E[|E[u_i^2 | \psi_{1,i}, q_i, \xi]| | \xi] \lesssim \sum_{d=0,1} E[Y_i(d)^2] < \infty$, so $B_n - E[B_n | \xi] = o_p(1)$ by conditional WLLN. Then we have shown $\Sigma_n - E[B_n | \xi] = o_p(1)$, with the claimed limit. Finally, we show the Lindeberg condition in Equation 10.3. Let $\Delta_i = Y_i(1) - Y_i(0)$. By the propensity

lower bound and Young's inequality $T_i^2 u_i^2 \leq 2\delta^{-2}(E[\Delta_i|\psi_{1,i}, q_i, \xi]^2 + E[\Delta_i|\psi_{2,i}, p_i, \xi]^2) \equiv v_i$. Note also that $E[v_i] \lesssim E[E[\Delta_i|\psi_{1,i}, q_i, \xi]^2 + E[\Delta_i|\psi_{2,i}, p_i, \xi]^2] \leq 2E[\Delta_i^2] \lesssim E[Y(1)^2 + Y(0)^2] < \infty$. The second inequality is conditional Jensen. Then for $\epsilon > 0$

$$n^{-1} \sum_{i=1}^n E[T_i^2 u_i^2 \mathbf{1}(T_i^2 u_i^2 > n\epsilon^2) | \mathcal{F}_{0,n}] \leq n^{-1} \sum_{i=1}^n E[v_i \mathbf{1}(v_i > n\epsilon^2) | \mathcal{F}_{0,n}] \equiv C_n$$

As in the conditional independence arguments above, we have

$$\begin{aligned} n^{-1} \sum_{i=1}^n E[v_i \mathbf{1}(v_i > n\epsilon^2) | \mathcal{F}_{0,n}] &= n^{-1} \sum_{i=1}^n E[v_i \mathbf{1}(v_i > n\epsilon^2) | \tau^t, \pi_n, (\psi_{1,i})_{1:n}, (q_i)_{1:n}, \xi] \\ &= n^{-1} \sum_{i=1}^n E[v_i \mathbf{1}(v_i > n\epsilon^2) | (\psi_{1,i})_{1:n}, (q_i)_{1:n}, \xi] = n^{-1} \sum_{i=1}^n E[v_i \mathbf{1}(v_i > n\epsilon^2) | \psi_{1,i}, q_i, \xi] \end{aligned}$$

Then $E[C_n] = E[E[C_n | \xi]] = E[v_i \mathbf{1}(v_i > n\epsilon^2)] \rightarrow 0$ as $n \rightarrow \infty$ since $v_i \geq 0$ and $E[v_i] < \infty$. Then $C_n \xrightarrow{P} 0$ by Markov inequality. This finishes the proof of the Lindberg condition. We skip the proof of the final claim, since it is similar to the proof of claim (2).

Next we show claim (2). Define $\mathcal{G}_{k,n} = \sigma(\mathcal{G}_{0,n}, (\epsilon_i^0, \epsilon_i^1)_{i=1:k})$. Then claim that $(z_{i,n}, \mathcal{G}_{i,n})_{i \geq 1}$ is an MDS with $z_{i,n} = T_i D_i \epsilon_i^1 / (\widehat{p}_{i,n} \widehat{q}_{i,n}) + T_i (1 - D_i) \epsilon_i^0 / ((1 - \widehat{p}_{i,n}) \widehat{q}_{i,n})$. Adaptation is clear. Next we show the MDS property. It suffices to show $E[\epsilon_i^d | \mathcal{G}_{i-1,n}] = 0$. Again we use the fact $(A, B) \perp\!\!\!\perp C \implies A \perp\!\!\!\perp C | B$. Then note that $E[\epsilon_i^d | \mathcal{G}_{i-1,n}]$ is equal to

$$E[\epsilon_i^d | \pi_n, \xi_n, X_{1:n}, \tau^t, \tau^d, (\epsilon_i^0, \epsilon_i^1)_{i=1:i-1}] = E[\epsilon_i^d | X_{1:n}, (\epsilon_i^0, \epsilon_i^1)_{i=1:i-1}] = E[\epsilon_i^d | X_i] = 0.$$

The second equality by the fact with $A = \epsilon_i^d$, $B = (X_{1:n}, (\epsilon_i^0, \epsilon_i^1)_{i=1:i-1})$, and $C = (\pi_n, \xi_n, \tau^t, \tau^d)$. The third equality by setting $A = \epsilon_i^d$, $B = X_i$, and $C = (X_{-i}, (\epsilon_i^0, \epsilon_i^1)_{i=1:i-1})$. This proves the MDS property. Next we analyze $\Sigma_{k,n} = n^{-1} \sum_{i=1}^k E[z_{i,n}^2 | \mathcal{G}_{i-1,n}]$, the variance process. By the exact reasoning above, $E[(\epsilon_i^d)^2 | \mathcal{G}_{i-1,n}] = E[(\epsilon_i^d)^2 | X_i] = \sigma_d^2(X_i)$.

$$\Sigma_{k,n} = n^{-1} \sum_{i=1}^k T_i D_i \sigma_1^2(X_i) / (\widehat{p}_{i,n} \widehat{q}_{i,n})^2 + T_i (1 - D_i) \sigma_0^2(X_i) / ((1 - \widehat{p}_{i,n}) \widehat{q}_{i,n})^2$$

so that $\Sigma_{k,n} \in \mathcal{G}_{0,n}$ for all k, n , satisfying the variance condition of Proposition 10.6.

$$E_n[T_i D_i \sigma_1^2(X_i) / (\widehat{p}_{i,n} \widehat{q}_{i,n})^2] = E_n[((T_i - \widehat{q}_{i,n}) + \widehat{q}_{i,n}) D_i \sigma_1^2(X_i) / (\widehat{p}_{i,n} \widehat{q}_{i,n})^2] = R_n^1 + R_n^2$$

By our propensity lower bounds we have $D_i \sigma_1^2(X_i) / (\widehat{p}_{i,n} \widehat{q}_{i,n})^2 \leq \delta^{-4} \sigma_1^2(X_i)$ with $E[\sigma_1^2(X)] < \infty$, so $D_i \sigma_1^2(X_i) / (\widehat{p}_{i,n} \widehat{q}_{i,n})^2$ is uniformly integrable. Then $R_n^1 = o_p(1)$ by Lemma 10.17. Similarly, $R_n^2 = E_n[((D_i - \widehat{p}_{i,n}) + \widehat{p}_{i,n}) \sigma_1^2(X_i) / \widehat{p}_{i,n}^2 \widehat{q}_{i,n}] = R_n^3 + R_n^4$ and $R_n^3 = o_p(1)$ by the same argument. Then $\Sigma_n = E_n[\sigma_1^2(X_i) / \widehat{p}_{i,n} \widehat{q}_{i,n}] + o_p(1)$.

$$\begin{aligned} |E_n[\sigma_1^2(X_i) / \widehat{p}_{i,n} \widehat{q}_{i,n} - \sigma_1^2(X_i) / p_i q_i]| &\leq \delta^{-4} |E_n[\sigma_1^2(X_i) ((q_i - \widehat{q}_{i,n}) p_i + \widehat{q}_{i,n} (p_i - \widehat{p}_{i,n}))]| \\ &\leq E_n[\sigma_1^2(X_i)^2]^{1/2} (E_n[(q_i - \widehat{q}_{i,n})^2]^{1/2} + E_n[(p_i - \widehat{p}_{i,n})^2]^{1/2}) = O_p(1) o_p(1) = o_p(1) \end{aligned}$$

Then $E_n[T_i D_i \sigma_1^2(X_i) / (\widehat{p}_{i,n} \widehat{q}_{i,n})^2] = E_n[\sigma_1^2(X_i) / p_i q_i] + o_p(1) = E[\sigma_1^2(X_i) / p_i q_i] + O_p(n^{-1/2})$ by Chebyshev inequality. By symmetry, the same holds for the $D = 0$ term. Putting this together, we have shown that $\Sigma_n \xrightarrow{P} E[\sigma_1^2(X_i) / (p_i q_i) + \sigma_0^2(X_i) / (1 - p_i) q_i]$ as claimed.

Finally, we show the Lindberg condition. Note the bound $|z_{i,n}|^2 \leq 2\delta^{-4}((\epsilon_i^1)^2 + (\epsilon_i^0)^2) \equiv v_i$, with $E[v_i] \lesssim \text{Var}(Y(1)) + \text{Var}(Y(0)) < \infty$. Then for $\epsilon > 0$ we have

$$\begin{aligned} L_n &= n^{-1} \sum_{i=1}^n E[z_{i,n}^2 \mathbf{1}(z_{i,n}^2 > n\epsilon^2) | \mathcal{G}_{0,n}] \leq n^{-1} \sum_{i=1}^n E[v_i \mathbf{1}(v_i > n\epsilon^2) | \mathcal{F}_{0,n}] \\ &= n^{-1} \sum_{i=1}^n E[v_i \mathbf{1}(v_i > n\epsilon^2) | \pi_n, \xi_n, X_{1:n}, \tau^t, \tau^d] = n^{-1} \sum_{i=1}^n E[v_i \mathbf{1}(v_i > n\epsilon^2) | X_i] \end{aligned}$$

The inequality from the upper bound just stated. The second equality from the conditional independence reasoning above. Then $E[L_n] = E[v_i \mathbf{1}(v_i > n\epsilon^2)] \rightarrow 0$ as $n \rightarrow \infty$ since $E[v_i] < \infty$. Then by Markov inequality $L_n \xrightarrow{P} 0$, which finishes the proof. \square

Proof of Theorem 3.10. Define $c_{1,i} = E[Y(1) - Y(0) | \psi_{1,i}, q_i, \xi]$ and $c_{2,i} = E[Y(1) - Y(0) | \psi_{2,i}, p_i, \xi]$. Define $b_{2,i} = E[Y(1) | \psi_{2,i}, p_i, \xi]((1-p_i)/p_i)^{1/2} + E[Y(0) | \psi_{2,i}, p_i, \xi](p_i/(1-p_i))^{1/2}$ and $\epsilon_i^d = Y_i(d) - E[Y_i(d) | \psi_{2,i}, p_i, \xi]$. Define the σ -algebras $\mathcal{F}_{k,n}$ for $1 \leq k \leq 4$

$$\mathcal{F}_{1,n} = \sigma(\psi_{1,1:n}, q_{1:n}, \xi, \pi_n) \quad \mathcal{F}_{2,n} = \sigma(\mathcal{F}_{1,n}, T_{1:n}) \quad \mathcal{F}_{3,n} = \sigma(\mathcal{F}_{2,n}, \psi_{2,1:n}, p_{1:n})$$

and $\mathcal{F}_{4,n} = \sigma(\mathcal{F}_{3,n}, D_{1:n})$. We expand our estimator by projection

$$\begin{aligned} \theta - \hat{\theta} &= (\theta - E[\hat{\theta} | \mathcal{F}_{1,n}]) + (E[\hat{\theta} | \mathcal{F}_{1,n}] - E[\hat{\theta} | \mathcal{F}_{2,n}]) + (E[\hat{\theta} | \mathcal{F}_{2,n}] - E[\hat{\theta} | \mathcal{F}_{3,n}]) \\ &\quad + (E[\hat{\theta} | \mathcal{F}_{3,n}] - E[\hat{\theta} | \mathcal{F}_{4,n}]) + (E[\hat{\theta} | \mathcal{F}_{4,n}] - \hat{\theta}) \\ &= E_n[c_{1,i} - \theta] + E_n \left[\frac{T_i - q_i}{q_i} c_{1,i} \right] + E_n \left[\frac{T_i}{q_i} (c_{2,i} - c_{1,i}) \right] + E_n \left[\frac{(D_i - p_i) T_i}{q_i \sqrt{p_i - p_i^2}} b_{2,i} \right] \\ &\quad + E_n \left[\frac{D_i T_i \epsilon_i^1}{p_i q_i} + \frac{(1 - D_i) T_i \epsilon_i^0}{(1 - p_i) q_i} \right] \equiv A_n + B_n + C_n + S_n + R_n \end{aligned}$$

We wish to apply Lemma 10.2 to show that $\sqrt{n}B_n, \sqrt{n}S_n = o_p(1)$. It suffices to check integrability. In the notation of the lemma, define $g(\psi_{1,i}, q_i, \xi) = c_{1,i}/q_i$, and note that by our propensity bound, Young's inequality, and contraction

$$E[E[c_{1,i}^2/q_i^2 | \xi]] \leq \delta^{-2} E[c_{1,i}^2] \lesssim E \left[\sum_{d=0,1} E[Y(d) | \psi_{1,i}, q_i, \xi]^2 \right] \leq \sum_{d=0,1} E[Y(d)^2]$$

This shows $E[c_{1,i}^2/q_i^2 | \xi] < \infty$ ξ -a.s. Similarly, $E[b_{2,i}^2 q_i^{-2} (p_i(1-p_i))^{-1} | \xi] < \infty$ ξ -a.s. Then $\sqrt{n}B_n, \sqrt{n}S_n = o_p(1)$ by the lemma.

The conditions of Theorem 10.4 are satisfied by assumption. Then $\sqrt{n}A_n | \xi \Rightarrow \mathcal{N}(0, V_1)$ with $V_1 = \text{Var}(c_{1,i} | \xi)$, $\sqrt{n}C_n | \mathcal{F}_{2,n} \Rightarrow \mathcal{N}(0, V_2)$ with $V_2 = E[q_i^{-1} \text{Var}(c_{2,i} | \psi_{1,i}, q_i, \xi) | \xi]$, and $\sqrt{n}R_n | \mathcal{F}_{4,n} \Rightarrow \mathcal{N}(0, V_3)$ with $V_3 = E[q_i^{-1} (\sigma_{1,i}^2 p_i^{-1} + \sigma_{0,i}^2 (1-p_i)^{-1}) | \xi]$ for $\sigma_{d,i}^2 = \text{Var}(Y(d) | \psi_{2,i}, p_i, \xi)$. Then by Slutsky and Lemma 10.12, we have $\sqrt{n}(\hat{\theta} - \text{ATE}) | \xi \Rightarrow \mathcal{N}(0, V_1 + V_2 + V_3)$.

To finish the proof, we use algebra to reformulate the variance $V = V_1 + V_2 + V_3$. By the law of total variance $\text{Var}(c(X) | \xi) = \text{Var}(E[c(X) | \psi_1, q, \xi] | \xi) + E[\text{Var}(c(X) | \psi_1, q, \xi) | \xi]$,

with $E[c(X_i)|\psi_{1,i}, q_i, \xi] = c_{1,i}$. Then since $\xi \perp\!\!\!\perp W_{1:n}$

$$V_1 = \text{Var}(c(X)) - E[\text{Var}(c(X)|\psi_1, q, \xi)|\xi].$$

Next, by the law of total variance

$$\text{Var}(c(X_i)|\psi_{1,i}, q_i, \xi) = E[\text{Var}(c(X)|\psi_{2,i}, p_i, \xi)|\psi_{1,i}, q_i, \xi] + \text{Var}(E[c(X)|\psi_{2,i}, p_i, \xi]|\psi_{1,i}, q_i, \xi)$$

Using this equality and applying tower law gives

$$\begin{aligned} V_2 &= E[q_i^{-1} \text{Var}(c_{2,i}|\psi_{1,i}, q_i, \xi)|\xi] = E[q_i^{-1} \text{Var}(E[c(X)|\psi_{2,i}, p_i, \xi]|\psi_{1,i}, q_i, \xi)|\xi] \\ &= E[q_i^{-1} \text{Var}(c(X_i)|\psi_{1,i}, q_i, \xi)|\xi] - E[q_i^{-1} E[\text{Var}(c(X)|\psi_{2,i}, p_i, \xi)|\psi_{1,i}, q_i, \xi]|\xi] \\ &= E[q_i^{-1} \text{Var}(c(X_i)|\psi_{1,i}, q_i, \xi)|\xi] - E[q_i^{-1} \text{Var}(c(X)|\psi_{2,i}, p_i, \xi)|\xi] \end{aligned}$$

Also, $\text{Var}(Y_i(d)|\psi_{2,i}, p_i, \xi) = \text{Var}(E[Y_i(d)|X_i, \xi]|\psi_{2,i}, p_i, \xi) + E[\text{Var}(Y_i(d)|X_i, \xi)|\psi_{2,i}, p_i, \xi]$. Using this fact gives

$$\begin{aligned} V_3 &= E[(q_i p_i)^{-1} \sigma_{1,i}^2|\xi] + E[(q_i(1-p_i))^{-1} \sigma_{0,i}^2|\xi] \\ &= E[(q_i p_i)^{-1} \text{Var}(m_1(X_i)|\psi_{2,i}, p_i, \xi)|\xi] + E[(q_i p_i)^{-1} E[\sigma_1^2(X_i)|\psi_{2,i}, p_i, \xi]|\xi] \\ &\quad + E[(q_i(1-p_i))^{-1} \text{Var}(m_0(X_i)|\psi_{2,i}, p_i, \xi)|\xi] + E[(q_i(1-p_i))^{-1} E[\sigma_0^2(X_i)|\psi_{2,i}, p_i, \xi]|\xi] \\ &= E[(q_i p_i)^{-1} \text{Var}(m_1(X_i)|\psi_{2,i}, p_i, \xi)|\xi] + E[(q_i(1-p_i))^{-1} \text{Var}(m_0(X_i)|\psi_{2,i}, p_i, \xi)|\xi] \\ &\quad + E[(q_i p_i)^{-1} \sigma_1^2(X_i)|\xi] + E[(q_i(1-p_i))^{-1} \sigma_0^2(X_i)|\xi] \end{aligned}$$

The final equality by tower law and since $q_i \in \sigma(\psi_{2,i}, p_i, \xi)$ by construction. Putting this all together, we get total variance

$$\begin{aligned} V_1 + V_2 + V_3 &= \text{Var}(c(X_i)) - E[\text{Var}(c(X_i)|\psi_{1,i}, q_i, \xi)|\xi] \\ &\quad + E[q_i^{-1} \text{Var}(c(X_i)|\psi_{1,i}, q_i, \xi)|\xi] - E[q_i^{-1} \text{Var}(c(X)|\psi_{2,i}, p_i, \xi)|\xi] \\ &\quad + E[(q_i p_i)^{-1} \text{Var}(m_1(X_i)|\psi_{2,i}, p_i, \xi)|\xi] + E[(q_i(1-p_i))^{-1} \text{Var}(m_0(X_i)|\psi_{2,i}, p_i, \xi)|\xi] \\ &\quad + E[(q_i p_i)^{-1} \sigma_1^2(X_i)|\xi] + E[(q_i(1-p_i))^{-1} \sigma_0^2(X_i)|\xi] \\ &= T_1 + T_2 + T_3 + T_4 + T_5 + T_6 + T_7 + T_8 \end{aligned}$$

Note that $T_2 + T_3 = E[q_i^{-1}(1-q_i) \text{Var}(c(X_i)|\psi_{1,i}, q_i, \xi)|\xi]$. Also note that

$$\begin{aligned} & - \text{Var}(c(X_i)|\psi_{2,i}, p_i, \xi) + p_i^{-1} \text{Var}(m_1(X_i)|\psi_{2,i}, p_i, \xi) + (1-p_i)^{-1} \text{Var}(m_0(X_i)|\psi_{2,i}, p_i, \xi) \\ &= p_i^{-1}(1-p_i) \text{Var}(m_1(X_i)|\psi_{2,i}, p_i, \xi) + p_i(1-p_i)^{-1} \text{Var}(m_0(X_i)|\psi_{2,i}, p_i, \xi) \\ &\quad + 2 \text{Cov}(m_1(X_i), m_0(X_i)|\psi_{2,i}, p_i, \xi) = \text{Var}(b(X_i; p_i)|\psi_{2,i}, p_i, \xi) \end{aligned}$$

Then we have $T_4 + T_5 + T_6 = E[q_i^{-1} \text{Var}(b(X_i; p_i)|\psi_{2,i}, p_i, \xi)|\xi]$. Then as claimed

$$\begin{aligned} V &= \sum_{k=1}^8 T_k = \text{Var}(c(X)) + E[q_i^{-1}(1-q_i) \text{Var}(c(X_i)|\psi_{1,i}, q_i, \xi)|\xi] \\ &\quad + E[q_i^{-1} \text{Var}(b(X_i; p_i)|\psi_{2,i}, p_i, \xi)|\xi] + E \left[q_i^{-1} \left(\frac{\sigma_1^2(X_i)}{p_i} + \frac{\sigma_0^2(X_i)}{1-p_i} \right) \middle| \xi \right]. \end{aligned}$$

Finally, note $E_n[T_i] = E_n[T_i - q_i] + E_n[q_i] = o_p(n^{-1/2}) + E_n[q_i] = E[q_i|\xi] + O_p(n^{-1/2})$. The second equality by Lemma 10.2, and the third by conditional Chebyshev (Lemma

10.10). Then $\sqrt{n_T}(\hat{\theta} - \text{ATE}) = \sqrt{n}(E_n[T_i])^{1/2}(\hat{\theta} - \text{ATE}) = E[q_i|\xi]^{1/2}\sqrt{n}(\hat{\theta} - \text{ATE}) + o_p(1) \Rightarrow N(0, E[q(X, \xi)] \cdot V)$ by continuous mapping theorem and Slutsky. This finishes the proof \square

Definition 10.5 (Conditional Weak Convergence). For random variables $A_n, A \in \mathbb{R}^d$ and σ -algebras $(\mathcal{F}_n)_n, \mathcal{G}$ define conditional weak convergence

$$A_n|\mathcal{F}_n \Rightarrow A|\mathcal{G} \iff E[e^{it'A_n}|\mathcal{F}_n] = E[e^{it'A}|\mathcal{G}] + o_p(1) \quad \forall t \in \mathbb{R}^d$$

We require a slight modification of the martingale difference CLT in Billingsley, allowing the weak limit to be a mixture of normals.

Proposition 10.6 (MDS-CLT). Consider probability spaces $(\Omega_n, \mathcal{G}_n, P_n)$ each equipped with filtration $(\mathcal{F}_{k,n})_{k \geq 0}$. Suppose $(Y_{k,n})_{k=1}^n$ is adapted to $(\mathcal{F}_{k,n})_{k \geq 0}$ and has $E[Y_{k,n}|\mathcal{F}_{k-1,n}] = 0$ for all $k \geq 1$ with $n \rightarrow \infty$. Make the following definitions

$$S_{k,n} = \sum_{j=1}^k Y_{j,n} \quad \sigma_{k,n}^2 = E[Y_{k,n}^2|\mathcal{F}_{k-1,n}] \quad \Sigma_{k,n} = \sum_{j=1}^k \sigma_{j,n}^2$$

Denote $S_n \equiv S_{n,n}$ and $\Sigma_n \equiv \Sigma_{n,n}$. Suppose that $\sigma_{k,n}^2 \in \mathcal{F}_{0,n}$ for all k, n and $\Sigma_n = \sigma^2 + o_p(1)$ with $\sigma^2 \in \mathcal{F}_{0,n}$. Also, suppose for each $\epsilon > 0$

$$L_n^\epsilon = \sum_{k=1}^n E[Y_{k,n}^2 \mathbf{1}(|Y_{k,n}| \geq \epsilon) | \mathcal{F}_{0,n}] = o_p(1) \quad (10.3)$$

Then $E[e^{itS_n}|\mathcal{F}_{0,n}] = e^{-\frac{1}{2}t^2\sigma^2} + o_p(1)$.

Proof. We modify the argument in Theorem 35.12 of Billingsley (1995).

$$\begin{aligned} E \left[e^{itS_n} - e^{-\frac{1}{2}t^2\sigma^2} | \mathcal{F}_{0,n} \right] &= E[e^{itS_n}(1 - e^{\frac{1}{2}t^2\Sigma_n}e^{-\frac{1}{2}t^2\sigma^2}) | \mathcal{F}_{0,n}] \\ &\quad + E[e^{-\frac{1}{2}t^2\sigma^2}(e^{\frac{1}{2}t^2\Sigma_n}e^{itS_n} - 1) | \mathcal{F}_{0,n}] \end{aligned}$$

For the first term, by conditional Jensen inequality

$$\begin{aligned} |E[e^{itS_n}(1 - e^{\frac{1}{2}t^2\Sigma_n}e^{-\frac{1}{2}t^2\sigma^2}) | \mathcal{F}_{0,n}]| &\leq E[|(1 - e^{\frac{1}{2}t^2\Sigma_n}e^{-\frac{1}{2}t^2\sigma^2})| | \mathcal{F}_{0,n}] \\ &= |(1 - e^{\frac{1}{2}t^2\Sigma_n}e^{-\frac{1}{2}t^2\sigma^2})| = o_p(1) \end{aligned}$$

The first equality since $\Sigma_n, \sigma^2 \in \mathcal{F}_{0,n}$. Since $\Sigma_n = \sigma^2 + o_p(1)$, the second equality follows by continuous mapping. The second term has

$$\begin{aligned} |E[e^{-\frac{1}{2}t^2\sigma^2}(e^{\frac{1}{2}t^2\Sigma_n}e^{itS_n} - 1) | \mathcal{F}_{0,n}]| &= e^{-\frac{1}{2}t^2\sigma^2} |E[(e^{\frac{1}{2}t^2\Sigma_n}e^{itS_n} - 1) | \mathcal{F}_{0,n}]| \\ &= e^{-\frac{1}{2}t^2\sigma^2} \left| \sum_{k=1}^n E[e^{itS_{k-1,n}} e^{\frac{1}{2}t^2\Sigma_{k,n}} (e^{itY_{k,n}} - e^{-\frac{1}{2}t^2\sigma_{k,n}^2}) | \mathcal{F}_{0,n}] \right| \\ &\leq e^{-\frac{1}{2}t^2\sigma^2} e^{\frac{1}{2}t^2\Sigma_n} \sum_{k=1}^n E[|e^{itS_{k-1,n}} E[e^{itY_{k,n}} - e^{-\frac{1}{2}t^2\sigma_{k,n}^2} | \mathcal{F}_{k-1,n}]| | \mathcal{F}_{0,n}] \\ &= e^{-\frac{1}{2}t^2\sigma^2} e^{\frac{1}{2}t^2\Sigma_n} \sum_{k=1}^n E[|E[e^{itY_{k,n}} - e^{-\frac{1}{2}t^2\sigma_{k,n}^2} | \mathcal{F}_{k-1,n}]| | \mathcal{F}_{0,n}] \equiv o_p(1) Z_n \end{aligned}$$

The first equality since $\sigma^2 \in \mathcal{F}_{0,n}$. The second equality by telescoping. The first inequality by triangle inequality and since $\Sigma_{k,n} \in \mathcal{F}_{0,n}$, $\Sigma_{k,n} \leq \Sigma_n$, and $S_{k-1,n} \in \mathcal{F}_{k-1,n}$ for $1 \leq k \leq n$. The final equality by continuous mapping since $\Sigma_n \xrightarrow{p} \sigma^2$. We want to show that $Z_n = O_p(1)$. Fix $\epsilon > 0$ and let $I_{k,n} = \mathbb{1}(|Y_{k,n}| > \epsilon)$. Note the facts $|e^{ix} - (1 + ix - (1/2)x^2)| \leq (1/6)|x|^3 \wedge |x|^2$ and $|e^z - (1 + z)| \leq |z|^2 e^{|z|}$ for real x , complex z . By the MDS property and $E[Y_{k,n}^2 | \mathcal{F}_{k-1,n}] = \sigma_{k,n}^2$, combined with these facts

$$\begin{aligned} |E[e^{itY_{k,n}} - e^{-\frac{1}{2}t^2\sigma_{k,n}^2} | \mathcal{F}_{k-1,n}]| &\leq E[|tY_{k,n}|^3 \wedge |tY_{k,n}|^2 + (1/4)t^4\sigma_{k,n}^4 e^{\frac{1}{2}t^2\sigma_{k,n}^2} | \mathcal{F}_{k-1,n}] \\ &\leq (t^2 + t^4 + |t|^3 + e^{\frac{1}{2}t^2\Sigma_n})E[(\epsilon|Y_{k,n}|^2 + |Y_{k,n}|^2 I_{k,n} + \sigma_{k,n}^4) | \mathcal{F}_{k-1,n}] \\ &\equiv A_{n,t}(\epsilon\sigma_{k,n}^2 + E[|Y_{k,n}|^2 I_{k,n} | \mathcal{F}_{k-1,n}] + E[\sigma_{k,n}^4 | \mathcal{F}_{k-1,n}]) \end{aligned}$$

Then we have

$$\begin{aligned} Z_n &\leq A_{n,t} \sum_{k=1}^n E[\epsilon\sigma_{k,n}^2 + E[|Y_{k,n}|^2 I_{k,n} | \mathcal{F}_{k-1,n}] + E[\sigma_{k,n}^4 | \mathcal{F}_{k-1,n}] | \mathcal{F}_{0,n}] \\ &= A_{n,t}(\epsilon\Sigma_n + L_n^\epsilon) + A_{n,t} \sum_{k=1}^n E[\sigma_{k,n}^4 | \mathcal{F}_{0,n}] \leq A_{n,t}(\epsilon\Sigma_n + L_n^\epsilon + \Sigma_n(\epsilon^2 + L_n^\epsilon)) \end{aligned}$$

To see the final inequality, note that $\sigma_{k,n}^4 \leq \sigma_{k,n}^2 \max_{k=1}^n \sigma_{k,n}^2$ and We have $\sigma_{k,n}^2 = E[Y_{k,n}^2 | \mathcal{F}_{k-1,n}] \leq \epsilon^2 + E[Y_{k,n}^2 I_{k,n} | \mathcal{F}_{k-1,n}] \leq \epsilon^2 + \sum_{j=1}^n E[Y_{j,n}^2 I_{j,n} | \mathcal{F}_{j-1,n}]$. Taking $\max_{k=1}^n$ on both sides gives $\max_{k=1}^n \sigma_{k,n}^2 \leq \epsilon^2 + \sum_{j=1}^n E[Y_{j,n}^2 I_{j,n} | \mathcal{F}_{j-1,n}]$. Then $\sum_{k=1}^n E[\sigma_{k,n}^4 | \mathcal{F}_{0,n}] \leq \sum_{k=1}^n E[\sigma_{k,n}^2(\epsilon^2 + \sum_{j=1}^n E[Y_{j,n}^2 I_{j,n} | \mathcal{F}_{j-1,n}]) | \mathcal{F}_{0,n}] = \Sigma_n(\epsilon^2 + L_n^\epsilon)$. Note that since $\Sigma_n \xrightarrow{p} \sigma^2$, we have $A_{n,t}, \Sigma_n = O_p(1)$ and $L_n^\epsilon = o_p(1)$ by assumption. Since ϵ was arbitrary, this shows $Z_n = o_p(1)$. \square

10.3 Optimal Stratification and Pilot Design

Proof of Proposition 4.1. Define $V(q) = E[\bar{\sigma}^2(\psi)/q(\psi)]$. Define the sets

$$\mathcal{Q}' = \{q \in \mathbb{R}^\psi : |q|_\infty < \infty, q > 0, E[C(\psi)q(\psi)] \leq \bar{B}, V(q) < \infty\} \quad \mathcal{Q} = \mathcal{Q}' \cap \{0 < q \leq 1\}$$

and recall the candidate optimal solution $q^*(\psi) = \bar{B} \cdot \bar{\sigma}(\psi)C(\psi)^{-1/2} / E[\bar{\sigma}(\psi)C(\psi)^{1/2}]$. Let $t \in [0, 1]$ and $q_1, q_2 \in \mathcal{Q}'$. By convexity of $y \rightarrow 1/y$ on $(0, \infty)$, for each $\psi \in \psi$ we have

$$\frac{\bar{\sigma}^2(\psi)}{tq_1(\psi) + (1-t)q_2(\psi)} \leq t \frac{\bar{\sigma}^2(\psi)}{q_1(\psi)} + (1-t) \frac{\bar{\sigma}^2(\psi)}{q_2(\psi)}$$

Taking expectations of both sides gives $V(tq_1 + (1-t)q_2) \leq tV(q_1) + (1-t)V(q_2)$, so V is convex on $\{q > 0\}$ and \mathcal{Q}' is convex. We claim that $q^* \in \mathcal{Q}'$. First, $q^* \in (0, 1]$ by assumption. Since $\sup_{\psi \in \psi} q^* \leq 1$ we have $E[\bar{\sigma}(\psi)C(\psi)^{1/2}] > 0$. By Holder $E[\bar{\sigma}(\psi)C(\psi)^{1/2}] \leq E[\bar{\sigma}^2(\psi)]^{1/2} E[C(\psi)]^{1/2} < \infty$. Then we have $V(q^*) = \bar{B}^{-1} E[\bar{\sigma}(\psi)C(\psi)^{1/2}]^2 < \infty$. Also clearly $E[C(\psi)q^*(\psi)] = \bar{B}$. This shows the claim. Next, suppose that $q^* + t\Delta \in \mathcal{Q}'$ for some $t \in [0, 1]$ and $|\Delta|_\infty < M$. Since $q^* \in \mathcal{Q}'$, by convexity of \mathcal{Q}' , $q^* + t'\Delta \in \mathcal{Q}'$ for all $0 \leq t' \leq t$. Then $dV_{q^*}[\Delta] \equiv \limsup_{t \rightarrow 0+} t^{-1}(V(q^* + t\Delta) - V(q^*))$ is well-defined. We claim that this limit exists. Under our assumptions $q^* > q_l \geq 0$ for some q_l , so that for any ψ

and all $t \leq q_t^2/2M < \infty$

$$\left| t^{-1} \left(\frac{\bar{\sigma}^2(\psi)}{q^*(\psi) + t\Delta(\psi)} - \frac{\bar{\sigma}^2(\psi)}{q^*(\psi)} \right) \right| = \frac{|\bar{\sigma}^2(\psi)\Delta(\psi)|}{((q^*)^2 + q^*t\Delta)(\psi)} \leq \frac{M|\bar{\sigma}^2(\psi)|}{q_t^2 - tM} \leq \frac{M|\bar{\sigma}^2(\psi)|}{2q_t^2}$$

Since $\|\bar{\sigma}^2\|_1 < \infty$, dominated convergence implies

$$\begin{aligned} dV_{q^*}[\Delta] &= \lim_{t \rightarrow 0^+} t^{-1}(V(q^* + t\Delta) - V(q^*)) = E \left[\lim_{t \rightarrow 0^+} \frac{-\bar{\sigma}^2(\psi)\Delta(\psi)}{((q^*)^2 + q^*t\Delta)(\psi)} \right] \\ &= -E \left[\frac{\bar{\sigma}^2(\psi)\Delta(\psi)}{(q^*)^2(\psi)} \right] = (1/\bar{q})^2 E[C(\psi)\Delta(\psi)] E[\bar{\sigma}(\psi)C(\psi)^{1/2}]^2 = 0 \end{aligned}$$

The last line since $q^* + t\Delta \in \mathcal{Q}'$ implies $\bar{B} = E[q^*(\psi)C(\psi) + t\Delta(\psi)C(\psi)] = \bar{B} + tE[\Delta(\psi)C(\psi)]$, so that $E[\Delta(\psi)C(\psi)] = 0$. Let $q \in \mathcal{Q}'$, so that $\Delta = q - q^*$ has $|\Delta|_\infty < \infty$. Then by convexity $V(q) - V(q^*) \geq dV_{q^*}[q - q^*] = 0$, showing that $q^* = \operatorname{argmin}_{q \in \mathcal{Q}'} V(q)$. Since $q^* \in \mathcal{Q}$ by assumption, it is optimal over $\mathcal{Q} \subseteq \mathcal{Q}'$ as well. \square

Proof of Theorem 5.2. Suppose that $T_{1:n} \sim \operatorname{Loc}(\psi, \hat{q}_n(\psi))$ and $D_{1:n} \sim \operatorname{Loc}(\psi, \hat{p}_n(\psi))$. Similar to the proof of Theorem 3.10, define σ -algebras $\mathcal{F}_{1,n} = \sigma(\psi_{1:n}, \pi_n, \xi_n)$, $\mathcal{F}_{2,n} = \sigma(\mathcal{F}_{1,n}, T_{1:n})$, and $\mathcal{F}_{3,n} = \sigma(\mathcal{F}_{2,n}, D_{1:n})$. Next, expand the estimator

$$\begin{aligned} \hat{\theta} - \theta &= E_n[c(\psi_i) - \theta] + E_n \left[\frac{T_i - \hat{q}_{i,n}}{\hat{q}_{i,n}} c(\psi_i) \right] + E_n \left[\frac{T_i(D_i - \hat{p}_{i,n})}{\hat{q}_{i,n} \sqrt{\hat{p}_{i,n}(1 - \hat{p}_{i,n})}} b(\psi_i; \hat{p}_n) \right] \\ &\quad + E_n \left[\frac{D_i T_i \epsilon_i^1}{\hat{p}_{i,n} \hat{q}_{i,n}} + \frac{(1 - D_i) T_i \epsilon_i^0}{(1 - \hat{p}_{i,n}) \hat{q}_{i,n}} \right] \equiv A_n + B_n + C_n + R_n \end{aligned}$$

Our main task is to show that $\sqrt{n}B_n, \sqrt{n}C_n = o_p(1)$. We do this by modifying the proof of Lemma 10.2.

(1) *Balancing Argument* - Note that the term C_n has

$$C_n = E_n \left[\frac{T_i(D_i - \hat{p}_n(\psi_i))}{\hat{q}_n(\psi_i)} \left(\frac{m_1(\psi_i)}{\hat{p}_n(\psi_i)} + \frac{m_0(\psi_i)}{1 - \hat{p}_n(\psi_i)} \right) \right] = C_{n1} + C_{n2}.$$

Recall $q^*(\psi) = \bar{B}(\sigma_1(\psi) + \sigma_0(\psi))C(\psi)^{-1/2}/E[(\sigma_1(\psi) + \sigma_0(\psi))C(\psi)^{1/2}]$. Under our assumptions, it's clear that $\inf_\psi q^*(\psi) > 0$. Consider the term above involving $m_1(\psi)$. We have $E[m_1(\psi)^2/q^*(\psi)^2] \lesssim E[m_0(\psi)^2] < \infty$. Then the construction in Lemma 10.2 provides functions $h_n \in L_2(\psi)$ with $|h_n|_{\text{lip}} \vee |h_n|_\infty \leq c_n$ and $|h_n - (m_1/q^*)|_{2,\psi} \rightarrow 0$. Define $\mathcal{F}_n = \sigma(\psi_{1:n}, \xi_n, \pi_n, \tau^t)$. Note that $\mathcal{G}_n, (q^*(\psi_i))_{1:n}, (\hat{q}_n(\psi_i))_{1:n}, (\hat{p}_n(\psi_i))_{1:n} \in \mathcal{F}_n$. Then by Lemma 10.17 $E[C_{n1}|\mathcal{F}_n] = 0$. We can expand C_{n1} as

$$E_n \left[\frac{T_i(D_i - \hat{p}_n(\psi_i))}{\hat{p}_n(\psi_i)} \left(\left(\frac{m_1(\psi_i)}{\hat{q}_n(\psi_i)} - \frac{m_1(\psi_i)}{\hat{q}(\psi_i)} \right) + \left(\frac{m_1(\psi_i)}{\hat{q}(\psi_i)} - \frac{m_1(\psi_i)}{q^*(\psi_i)} \right) + \frac{m_1(\psi_i)}{q^*(\psi_i)} \right) \right]$$

and write $C_{n1} = T_{n1} + T_{n2} + T_{n3}$. Consider the first term. By construction, $|\hat{q}_n - \hat{q}|_\infty \leq 1/\bar{k}_n$. Then by Lemma 10.17 we have $E[B_{n1}|\mathcal{F}_n] = 0$ and

$$\operatorname{Var}(\sqrt{n}T_{n1}|\mathcal{F}_n) \lesssim E_n[m_1(\psi_i)^2(\hat{q}_n - \hat{q})^2(\psi_i)/\hat{p}_n(\psi_i)^2] \lesssim o(1)E_n[m_1(\psi_i)^2] = o_p(1)$$

Again by Lemma 10.17 we have

$$\text{Var}(\sqrt{n}T_{n2}|\mathcal{F}_n) \lesssim E_n[m_1(\psi_i)^2(\hat{q} - q^*)^2(\psi_i)/\hat{p}_n(\psi_i)^2] \lesssim \max_{i=1}^n m_1(\psi_i)^2 E_n[(\hat{q} - q^*)^2(\psi_i)]$$

The final expression is $o_p(n^{2/\alpha_2-2r}) + O(n^{2/\alpha_2}/\bar{k}_n^2)$, using Lemma 10.14 and the bound from Lemma 10.7. This is $o_p(1)$ under our assumptions. Consider the final term T_{n3}

$$T_{n3} = E_n \left[\frac{T_i(D_i - \hat{p}_n(\psi_i))}{\hat{p}_n(\psi_i)} \left(\left(\frac{m_1(\psi_i)}{q^*(\psi_i)} - h_n(\psi_i) \right) + h_n(\psi_i) \right) \right] = T_{n4} + T_{n5}$$

As above $E[T_{n4}|\mathcal{F}_n] = 0$ and $\text{Var}(\sqrt{n}T_{n4}|\mathcal{F}_n) \lesssim E_n[(m_1(\psi_i)/q^*(\psi_i) - h_n(\psi_i))^2] = o_p(1)$ by Markov inequality since $|h_n - (m_1/q^*)|_{2,\mathcal{X}} \rightarrow 0$. Finally, by the last part of Lemma 10.17 we have $E[T_{n5}|\mathcal{F}_n] = 0$ and

$$\begin{aligned} \text{Var}(\sqrt{n}T_{n5}|\mathcal{F}_n) &= n^{-1} \sum_{g \in \mathcal{G}_n} \frac{1}{|g|} \sum_{i,j \in g} (h_n(\psi_i) - h_n(\psi_j))^2 + n^{-1} \bar{k}_n |L_n| \max_{i=1}^n h_n(\psi_i)^2 \\ &\lesssim c_n^2 n^{-1} \sum_{g \in \mathcal{G}_n} \frac{1}{|g|} \sum_{i,j \in g} |\psi_i - \psi_j|_2^2 + n^{-1} \bar{k}_n |L_n| \max_{i=1}^n h_n(\psi_i)^2 \\ &\lesssim c_n^2 \max_{i=1}^n |\psi_i|_2^2 \cdot O((n/\bar{k}_n |L_n|)^{-2/(d+1)}) + n^{-1} \bar{k}_n |L_n| O_p(c_n^2) \\ &= o_p(c_n^2 n^{2/\alpha_1-2/(d+1)} (\bar{k}_n |L_n|)^{2/(d+1)}) = o_p(1) \end{aligned}$$

The first inequality by Lipschitz continuity, the second inequality by Theorem 10.1. The final equality holds if $\bar{k}_n |L_n| = o(n^{1-\frac{d+1}{\alpha_1}})$ and $c_n = O(n^{2/\alpha_1-2/(d+1)} (\bar{k}_n |L_n|)^{2/(d+1)})$ with $c_n \rightarrow \infty$. This finishes the proof that $\sqrt{n}B_n = o_p(1)$. An identical proof shows that $\sqrt{n}C_n = o_p(1)$.

(2) *CLT Argument* - By the previous argument, we have $\sqrt{n}(\hat{\theta} - \theta) = \sqrt{n}A_n + \sqrt{n}R_n + o_p(1)$. $\sqrt{n}A_n \Rightarrow \mathcal{N}(0, V_1)$ with $V_1 = \text{Var}(c(\psi))$ by vanilla CLT. We want to apply Theorem 10.4 to the term $\sqrt{n}R_n$. The moment bound is satisfied by assumption. For condition (a2), set $\psi_{k,n}(\psi) = \psi$, $p_i = p_i^*$, and $q_i = q_i^*$. The propensity lower and upper bounds follow from our assumptions on $\sigma_d(\psi)$. Condition (c) is shown in Lemma 10.7. Then by Theorem 10.4 $\sqrt{n}R_n|\mathcal{F}_{3,n} \Rightarrow \mathcal{N}(0, V_2)$ with $V_2 = E[(q_i^* p_i^*)^{-1} \sigma_1^2(\psi_i) + (q_i^* (1 - p_i^*))^{-1} \sigma_0^2(\psi_i)]$ by Theorem 10.4. Then by Lemma 10.12, $\sqrt{n}(A_n + R_n) \Rightarrow \mathcal{N}(0, V_1 + V_2)$. Applying Lemma 10.17 with $h_n = 1$ shows that $\text{Var}(E_n[(T_i - \hat{q}_{i,n})]|\mathcal{F}_{1,n}) \leq n^{-1} \bar{k}_n |L_n|$. Then $E_n[T_i] = E_n[\hat{q}_{i,n}] + O_p((\bar{k}_n |L_n|/n)^{1/2})$ and $E_n[\hat{q}_{i,n}] = E_n[\hat{q}_i] + O(1/\bar{k}_n)$, since $|\hat{q} - \hat{q}_n|_\infty = O(1/\bar{k}_n)$ by discretization. Finally, $E_n[\hat{q}_i] = E_n[q_i^*] + o_p(1)$ by Lemma 10.7. Putting this all together, we have $E_n[T_i] = E_n[q_i^*] + o_p(1) = E[q_i^*] + o_p(1)$, so that $\sqrt{n_T}(\hat{\theta} - \text{ATE}) = \sqrt{n}(E_n[T_i])^{1/2}(\hat{\theta} - \text{ATE}) = E[q^*(\psi_i)]^{1/2} \sqrt{n}(\hat{\theta} - \text{ATE}) + o_p(1) \Rightarrow N(0, E[q^*(\psi_i)] \cdot V)$ by continuous mapping theorem and Slutsky. This finishes the proof. The conclusion for Theorem 4.4 follows by setting $\hat{q} = q^*$ and $\hat{q}_n = q_n^*$, so that the error T_{n2} above is identically zero. \square

Proof of Theorem 4.5. Fix $P \in \mathcal{P}_{1/2}$. By the fundamental expansion of the IPW estimator, $\hat{\theta} = E_n[c(X_i)] + 2E_n[(D_i - 1/2)b(X_i)] + 2E_n[D_i \epsilon_i^1 + (1 - D_i) \epsilon_i^0] \equiv A_n + B_n + C_n$. Then

by the law of total variance

$$\begin{aligned}\text{Var}(\hat{\theta}|X_{1:n}) &= \text{Var}(B_n + C_n|X_{1:n}) = E[\text{Var}(B_n + C_n|X_{1:n}, D_{1:n})|X_{1:n}] \\ &+ \text{Var}(E[B_n + C_n|X_{1:n}, D_{1:n}]|X_{1:n}) = E[\text{Var}(C_n|X_{1:n}, D_{1:n})|X_{1:n}] + \text{Var}(B_n|X_{1:n})\end{aligned}$$

The last line follows since $E[C_n|X_{1:n}, D_{1:n}] = 0$. One can show $E[\text{Var}(C_n|X_{1:n}, D_{1:n})|X_{1:n}] = (2/n)E_n[\sigma_1^2(X_i) + \sigma_0^2(X_i)]$, which does not depend on P . Then $\text{argmin}_{P \in \mathcal{P}_{1/2}} \text{Var}(\hat{\theta}|X_{1:n}) = \text{argmin}_{P \in \mathcal{P}_{1/2}} \text{Var}(B_n|X_{1:n})$. Denote $Z_i = D_i - 1/2$. Then $n^2 \text{Var}(B_n|X_{1:n})$ is equal to

$$E[(Z'_{1:n} b_{1:n})^2] = \sum_{d_{1:n} \in \{0,1\}^n} P(d_{1:n}|X_{1:n}) (Z'_{1:n} b_{1:n})^2 \geq \min_{d_{1:n} \in \{0,1\}^n} ((D_{1:n} - 1/2)' b_{1:n})^2$$

Let $d_{1:n}^* = d_{1:n}^*(X_{1:n})$ be a vector achieving the RHS minimum. Define $P^* \in \mathcal{P}_{1/2}$ by $P^*(d_{1:n}^*|X_{1:n}) = P^*(1 - d_{1:n}^*|X_{1:n}) = 1/2$. Then

$$\begin{aligned}n^2 \min_{P \in \mathcal{P}_{1/2}} \text{Var}(B_n|X_{1:n}) &\geq ((d_{1:n}^* - (1/2)\mathbf{1}_n)' b_{1:n})^2 = E_{P^*}[(D_{1:n} - (1/2)\mathbf{1}_n)' b_{1:n}]^2 \\ &= n^2 \text{Var}_{P^*}(B_n|X_{1:n}) \geq n^2 \min_{P \in \mathcal{P}_{1/2}} \text{Var}(B_n|X_{1:n})\end{aligned}$$

The final equality since $P^* \in \mathcal{P}_{1/2}$ by construction. Then equality holds throughout, and $P^* \in \text{argmin}_{P \in \mathcal{P}_{1/2}} \text{Var}(B_n|X_{1:n}) = \text{argmin}_{P \in \mathcal{P}_{1/2}} \text{Var}(\hat{\theta}|X_{1:n})$. Next, we characterize the optimal vector $d_{1:n}^*$. Observe that $(d_i - 1/2)(d_j - 1/2) = (d_i - 1/2)(d_j - 1/2) - 1/4 + 1/4 = -(1/2)\mathbf{1}(d_i \neq d_j) + 1/4$. Then $\text{argmin}_{d_{1:n} \in \{0,1\}^n} ((d_{1:n} - (1/2)\mathbf{1}_n)' b_{1:n})^2$ is equal to

$$\text{argmin}_{d_{1:n} \in \{0,1\}^n} \sum_{i,j=1}^n (d_i - 1/2)(d_j - 1/2) b_i b_j = \text{argmax}_{d_{1:n} \in \{0,1\}^n} \sum_{i \neq j}^n \mathbf{1}(d_i \neq d_j) b_i b_j$$

The latter problem is equivalent to Max-Cut with edge weights $w_{ij} = b_i b_j$, as claimed. \square

Lemma 10.7 (Propensity Convergence). *Suppose that $|\hat{\sigma}_d^2 - \sigma_d^2|_{2,\psi}^2 = O_p(n^{-r})$ for $d = 0, 1$. Then $|\hat{p}_{i,n} - p_i^*|_{2,n}^2 \vee |\hat{q}_{i,n} - q_i^*|_{2,n}^2 = O(1/\bar{k}_n^2) + O_p(n^{-r})$.*

Proof. We have $|\hat{p}_{i,n} - p_i^*|_{2,n}^2 \leq 2|\hat{p}_{i,n} - \hat{p}_i|_{2,n}^2 + 2|\hat{p}_i - p_i^*|_{2,n}^2 \leq O(1/\bar{k}_n^2) + |\hat{p}_i - p_i^*|_{2,n}^2$ by discretization, and similarly for $|\hat{q}_{i,n} - q_i^*|_{2,n}^2$. Then it suffices to bound $|\hat{p}_i - p_i^*|_{2,n}^2$. Note

$$|\hat{\sigma}_d - \sigma_d|_{2,\psi}^2 = \int_{\psi} \frac{(\hat{\sigma}_d^2 - \sigma_d^2)^2(\psi)}{(\hat{\sigma}_d + \sigma_d)^2(\psi)} dP(\psi) \leq c_l^{-1} |\hat{\sigma}_d^2 - \sigma_d^2|_{2,\psi}^2 = O_p(n^{-r}).$$

Then $|\hat{p}_i - p_i^*|_{2,n}^2$ is equal to

$$\begin{aligned}E_n[(\hat{\sigma}_{1i}/(\hat{\sigma}_{1i} + \hat{\sigma}_{0i}) - \sigma_{1i}/(\sigma_{1i} + \sigma_{0i}))^2] &\leq 4c_l^{-2} E_n[(\sigma_{1i} - \hat{\sigma}_{1i})\hat{\sigma}_{0i} + \hat{\sigma}_{1i}(\hat{\sigma}_{0i} - \sigma_{0i})]^2 \\ &\lesssim E_n[(\sigma_{1i} - \hat{\sigma}_{1i})^2 \hat{\sigma}_{0i}^2] + E_n[\hat{\sigma}_{1i}^2 (\hat{\sigma}_{0i} - \sigma_{0i})^2] \lesssim |\hat{\sigma}_0 - \sigma_0|_{2,n}^2 + |\hat{\sigma}_1 - \sigma_1|_{2,n}^2\end{aligned}$$

The last expression is equal to $\sum_{d=0,1} |\hat{\sigma}_d - \sigma_d|_{2,\psi}^2 + O_p(n^{-1/2}) = O_p(n^{-r}) + O_p(n^{-1/2})$ by conditional Markov inequality. Next consider $|\hat{q} - q^*|_{2,n}^2$. Denote $C_i = C(\psi_i)$. Define $\mu_n = E_n[C_i^{1/2}(\hat{\sigma}_{1i} + \hat{\sigma}_{0i})]$ and $\mu = E[C_i^{1/2}(\sigma_{1i} + \sigma_{0i})]$. Using $\xi_n \perp\!\!\!\perp W_{1:n}$, by conditional

Chebyshev $E_n[C_i^{1/2}\hat{\sigma}_{di}] = E[C_i^{1/2}\hat{\sigma}_{di}|\xi_n] + O_p(n^{-1/2})$. Then we have

$$\begin{aligned} (\mu_n - \mu)^2 &\lesssim \sum_{d=0,1} (E_n[C_i^{1/2}\hat{\sigma}_{di}] - E[C_i^{1/2}\hat{\sigma}_{di}|\xi_n] + E[C_i^{1/2}\hat{\sigma}_{di}|\xi_n] - E[C_i^{1/2}\sigma_{di}])^2 \\ &\lesssim \sum_{d=0,1} (E_n[C_i^{1/2}\hat{\sigma}_{di}] - E[C_i^{1/2}\hat{\sigma}_{di}|\xi_n])^2 + |C|_\infty E_\psi[(\hat{\sigma}_{di} - \sigma_{di})^2] = O_p(n^{-1}) + O_p(n^{-r}) \end{aligned}$$

The first inequality is Young's, the second by Young's and Jensen. Define $\hat{v}_i = C_i^{-1/2}(\hat{\sigma}_{1i} + \hat{\sigma}_{0i})$ and $v_i = C_i^{-1/2}(\sigma_{1i} + \sigma_{0i})$.

$$\begin{aligned} |\hat{q} - q^*|_{2,n}^2 &= \bar{q}^2 E_n[(\hat{v}_i/\mu_n - v_i/\mu)^2] \lesssim (\mu_n \mu)^{-2} (E_n[\hat{v}_i^2](\mu - \mu_n) + \mu_n E_n[(\hat{v}_i - v_i)^2]) \\ &\lesssim (\mu - \mu_n)^2 + E_n[(\hat{v}_i - v_i)^2] = O_p(n^{-r}) + O_p(n^{-1/2}) \end{aligned}$$

The first inequality is Young's, the second using our bounded variance assumption. The final equality follows from work above. This finishes the proof. \square

10.4 Inference

Proof of Theorem 6.1. First, consider the sample variance. The second moment is

$$E_n \left[\left(\frac{T_i(D_i - p(\psi_i))Y_i}{q(\psi_i)(p - p^2)(\psi_i)} \right)^2 \right] = E_n \left[\frac{T_i D_i Y_i(1)^2}{p_i^2 q_i^2} \right] + E_n \left[\frac{T_i(1 - D_i)Y_i(0)^2}{(1 - p_i)^2 q_i^2} \right]$$

Consider the first term

$$\begin{aligned} E_n[T_i D_i Y_i(1)^2/(p_i^2 q_i^2)] &= E_n[T_i(D_i - p_i)Y_i(1)^2/(p_i^2 q_i^2)] + E_n[(T_i - q_i)Y_i(1)^2/(p_i q_i^2)] \\ &+ E_n[Y_i(1)^2/(p_i q_i)] = E_n[Y_i(1)^2/(p_i q_i)] + o_p(1) = E[Y_i(1)^2/(p_i q_i)] + o_p(1) \end{aligned}$$

The first equality is by Lemma 10.17, and the second by WLLN. Then by symmetry $E_n[T_i(1 - D_i)Y_i(0)^2/(1 - p_i)^2 q_i^2] = E[Y_i(0)^2/(1 - p_i)q_i] + o_p(1)$. Then the sample variance in the theorem converges in probability to $v = E[Y_i(1)^2/(p_i q_i)] + E[Y_i(0)^2/(1 - p_i)q_i] - \text{ATE}^2$. With V the limiting variance of Theorem 3.10, simple algebra shows that

$$v - V = E[m_{1i}^2(1 - p_i q_i)/p_i q_i] + E[m_{0i}^2(1 - q_i(1 - p_i))/(q_i(1 - p_i))] + 2E[m_{1i}m_{0i}].$$

Then the conclusion follows from the consistency results in Lemma 10.8. \square

Lemma 10.8 (Inference). *Suppose that $E[Y(d)^2] < \infty$ for $d = 0, 1$. Replace w_i^d in the variance estimators in Section 6 with a bounded non-negative function $a_i = a(\psi_i)$. Then we have $\hat{v}_1 \xrightarrow{p} E[q_i p_i m_{1i}^2 a_i]$, $\hat{v}_0 \xrightarrow{p} E[q_i(1 - p_i) m_{0i}^2 a_i]$, and $\hat{v}_{10} \xrightarrow{p} E[q_i m_{1i} m_{0i} a_i]$.*

Proof. First, we show the homogeneity condition $n^{-1} \sum_{g \in \mathcal{G}_n^\nu} \sum_{i,j \in g} |\psi_i - \psi_j|_2^2 = o(1)$ also holds for the matched groups $\mathcal{G}_n^\nu = \{g \cup \nu(g) : g \in \mathcal{G}_n\}$. Note that we may write

$$\begin{aligned} n^{-1} \sum_{g \in \mathcal{G}_n^\nu} \sum_{i,j \in g} |\psi_i - \psi_j|_2^2 &= (2n)^{-1} \sum_{g \in \mathcal{G}_n} \sum_{i,j \in g \cup \nu(g)} |\psi_i - \psi_j|_2^2 \\ &= (2n)^{-1} \sum_{g \in \mathcal{G}_n} \sum_{i \in g, j \in \nu(g)} |\psi_i - \psi_j|_2^2 + (2n)^{-1} \sum_{g \in \mathcal{G}_n} \sum_{i \neq j} |\psi_i - \psi_j|_2^2 \mathbf{1}(ij \in g \text{ or } ij \in \nu(g)) \end{aligned}$$

The second term is $o(1)$ by Theorem 10.1. By Young's inequality, the first term is

$$\begin{aligned} &\lesssim n^{-1} \sum_{g \in \mathcal{G}_n} \sum_{i \in g, j \in \nu(g)} |\psi_i - \bar{\psi}_g|_2^2 + |\bar{\psi}_g - \bar{\psi}_{\nu(g)}|_2^2 + |\bar{\psi}_{\nu(g)} - \psi_j|_2^2 \\ &\leq 2n^{-1} \sum_{g \in \mathcal{G}_n} k(g) \sum_{i \in g} |\psi_i - \bar{\psi}_g|_2^2 + \bar{k}^2 n^{-1} \sum_{g \in \mathcal{G}_n} |\bar{\psi}_g - \bar{\psi}_{\nu(g)}|_2^2 \end{aligned}$$

The second term is $o(1)$ again by Theorem 10.1. The first term is

$$2n^{-1} \sum_{g \in \mathcal{G}_n} k(g) \sum_{i \in g} |k(g)^{-1} \sum_{j \in g} (\psi_i - \psi_j)|_2^2 \leq 2n^{-1} \sum_{g \in \mathcal{G}_n} \sum_{i, j \in g} |\psi_i - \psi_j|_2^2 = o(1)$$

again by triangle inequality, Jensen, and Theorem 10.1. This finishes our proof of the claim.

Denote $v_i = m_1(\psi_i) \sqrt{a(\psi_i)}$. By Lemma 10.11, there exists a sequence $(z_n)_{n \geq 1}$ of functions with $|z_n|_{lip} \leq c_n$ and $|z_n - m_1(\psi)^2 a(\psi)|_{2, \psi} = o(1)$. Observe that $|z_n|_{2, \psi} = O(1)$ since $m_1(\psi)^2 a(\psi) \in L_2(\psi)$. For matched groups $g = g_1 \sqcup g_2$ with assignment propensities a_1/k_1 and a_2/k_2 in each group, define the group weight $w_g = 1/(a_1 + a_2 - 1)$ and variance estimator

$$\begin{aligned} \hat{v}_1 &= n^{-1} \sum_{g \in \mathcal{G}_n^\nu} w_g \sum_{i, j \in g} Y_i(1) Y_j(1) (a_i a_j)^{1/2} D_i D_j \mathbb{1}(i \neq j) \\ &= n^{-1} \sum_{g \in \mathcal{G}_n^\nu} w_g \sum_{\substack{i, j \in g \\ i \neq j}} (m_1(\psi_i) + \epsilon_i^1)(m_1(\psi_j) + \epsilon_j^1) (a_i a_j)^{1/2} D_i D_j = A_n + B_n + C_n + R_n. \end{aligned}$$

Consider the first term A_n . Noting that $ab = -(1/2)[(a-b)^2 - a^2 - b^2]$, we have

$$\begin{aligned} A_n &= n^{-1} \sum_{g \in \mathcal{G}_n^\nu} w_g \sum_{i, j \in g} v_i v_j D_i D_j \mathbb{1}(i \neq j) = -(2n)^{-1} \sum_{g \in \mathcal{G}_n^\nu} w_g \sum_{i, j \in g} (v_i - v_j)^2 D_i D_j \\ &\quad + (2n)^{-1} \sum_{g \in \mathcal{G}_n^\nu} w_g \sum_{i, j \in g} (v_i^2 + v_j^2) D_i D_j \mathbb{1}(i \neq j) \equiv A_n^1 + A_n^2 \end{aligned}$$

Denote $z_{in} = z_n(\psi_i)$ and let $\bar{k} = \max_{g \in \mathcal{G}} |g|$. Observe that by Jensen's inequality

$$\begin{aligned} |A_n^1| &\leq (2n)^{-1} \sum_{g \in \mathcal{G}_n^\nu} \sum_{\substack{i, j \in g \\ i \neq j}} (v_i - v_j)^2 = (2n)^{-1} \sum_{g \in \mathcal{G}_n^\nu} \sum_{\substack{i, j \in g \\ i \neq j}} (v_i - z_{in} + z_{in} - z_{jn} + z_{jn} - v_j)^2 \\ &\leq 3(2n)^{-1} \sum_{g \in \mathcal{G}_n^\nu} \sum_{\substack{i, j \in g \\ i \neq j}} |v_i - z_{in}|^2 + |z_{in} - z_{jn}|^2 + |z_{jn} - v_j|^2 \\ &\lesssim n^{-1} \sum_{g \in \mathcal{G}_n^\nu} (|g| - 1) \sum_{i \in g} |v_i - z_{in}|^2 + c_n^2 \cdot n^{-1} \sum_{g \in \mathcal{G}_n^\nu} \sum_{i, j \in g} |\psi_i - \psi_j|_2^2 \\ &\leq (\bar{k} - 1) n^{-1} E_n[T_i |v_i - z_{in}|^2] + o_p(1) \lesssim E_n[|v_i - z_{in}|^2] + o_p(1) = o_p(1) \end{aligned}$$

For the second to last inequality, set $c_n = O(r_n^{-2})$ for $n^{-1} \sum_{g \in \mathcal{G}_n^\nu} \sum_{i, j \in g} |\psi_i - \psi_j|_2^2 = o(r_n)$, as guaranteed by the fact above. For the final equality, note that $E[E_n[|v_i - z_{in}|^2]] =$

$|v_i - z_{in}|_{2,\psi}^2 = o(1)$, so that $E_n[|v_i - z_{in}|^2] = o_p(1)$ by Markov inequality. Next,

$$\begin{aligned} A_n^2 &= (2n)^{-1} \sum_{g \in \mathcal{G}_n^\nu} w_g \sum_{\substack{i,j \in g \\ i \neq j}} (v_i^2 + v_j^2) D_i D_j = 2 \cdot (2n)^{-1} \sum_{g \in \mathcal{G}_n^\nu} w_g \sum_{i \in g} v_i^2 D_i \sum_{\substack{j \in g \\ j \neq i}} D_j \\ &= 2 \cdot (2n)^{-1} \sum_{g \in \mathcal{G}_n^\nu} \sum_{i \in g} v_i^2 D_i w_g (a_1 + a_2 - 1) = n^{-1} \sum_{g \in \mathcal{G}_n^\nu} \sum_{i \in g} v_i^2 D_i = E_n[v_i^2 D_i T_i] \end{aligned}$$

Finally, note that $E_n[v_i^2 D_i T_i] = E[v_i^2 p_i q_i] + o_p(1)$ by Lemma 10.17. Next we claim that $R_n = o_p(1)$. It suffices to verify the conditions of Lemma 10.13 for

$$u_g = w_g \sum_{\substack{i,j \in g \\ i \neq j}} \epsilon_i^1 \epsilon_j^1 (a_i a_j)^{1/2} D_i D_j \mathbf{1}(i \neq j) = w_g \sum_{i \neq j} \epsilon_i^1 \epsilon_j^1 (a_i a_j)^{1/2} D_i D_j \mathbf{1}(ij \in g)$$

Define $\mathcal{F}_n = \sigma(\psi_{1:n}, \tau^t, \pi_n)$, so that $\mathcal{G}_n^\nu, D_{1:n} \in \mathcal{F}_n$. Note that $u_g \perp\!\!\!\perp u_{g'} | \mathcal{F}_n$ for $g \neq g'$ by Lemma 10.15. We also need to show $E[u_g | \mathcal{F}_n] = 0$ for each $g \in \mathcal{G}_n^\nu$. Note that $E[u_g | \mathcal{F}_n] = E[E[u_g | \mathcal{F}_n'] | \mathcal{F}_n]$ with $\mathcal{F}_n' = \sigma(\mathcal{F}_n, \tau^d)$ and $D_{1:n} \in \mathcal{F}_n'$. Repeatedly using the fact that $(A, B) \perp\!\!\!\perp C \implies A \perp\!\!\!\perp C | B$, we have $E[\epsilon_i^1 \epsilon_j^1 | \mathcal{F}_n'] = E[\epsilon_i^1 \epsilon_j^1 | \psi_i, \psi_j] = E[\epsilon_j^1 E[\epsilon_i^1 | \psi_i] | \psi_i, \psi_j] = E[\epsilon_j^1 E[\epsilon_i^1 | \psi_i] | \psi_i, \psi_j] = 0$. Then apparently $E[u_g | \mathcal{F}_n] = 0$ for all $g \in \mathcal{G}_n^\nu$. Finally, Lemma 10.13 requires the condition $\frac{1}{n} \sum_{g \in \mathcal{G}_n^\nu} E[|u_g| \mathbf{1}(|u_g| > c_n) | \mathcal{F}_n] = o_p(1)$. Observe that for any positive constants $(a_k)_{k=1}^m$ we have $\sum_k a_k \mathbf{1}(\sum_k a_k > c) \leq m \sum_k a_k \mathbf{1}(a_k > c/m)$ and $ab \mathbf{1}(ab > c) \leq a^2 \mathbf{1}(a^2 > c) + b^2 \mathbf{1}(b^2 > c)$. Let $c_n \rightarrow \infty$ and let $\bar{k} = \max_{g \in \mathcal{G}_n^\nu} |g|$. Using the indicator function facts above, for any group $g \in \mathcal{G}_n^\nu$, $E[|u_g| \mathbf{1}(|u_g| > c_n) | \mathcal{F}_n]$ is bounded above by

$$\begin{aligned} &E \left[\sum_{i,j \in g} |\epsilon_i^1| |\epsilon_j^1| (a_i a_j)^{1/2} \mathbf{1}(i \neq j) \mathbf{1} \left(\sum_{i,j \in g} |\epsilon_i^1| |\epsilon_j^1| (a_i a_j)^{1/2} \mathbf{1}(i \neq j) > c_n \right) \middle| \mathcal{F}_n \right] \\ &\leq \bar{k}^2 \sum_{\substack{i,j \in g \\ i \neq j}} E \left[|\epsilon_i^1| |\epsilon_j^1| (a_i a_j)^{1/2} \mathbf{1} (|\epsilon_i^1| |\epsilon_j^1| (a_i a_j)^{1/2} > c_n / \bar{k}^2) \middle| \mathcal{F}_n \right] \\ &\leq 2\bar{k}^3 \sum_{i \in g} E \left[a_i (\epsilon_i^1)^2 \mathbf{1} (a_i (\epsilon_i^1)^2 > c_n / \bar{k}^2) \middle| \mathcal{F}_n \right] = 2\bar{k}^3 \sum_{i \in g} E \left[a_i (\epsilon_i^1)^2 \mathbf{1} (a_i (\epsilon_i^1)^2 > c_n / \bar{k}^2) \middle| \psi_i \right] \end{aligned}$$

The first line by triangle inequality and since $w_g \leq 1$. The second and third inequalities use the fact about indicator functions above. The last line again uses the fact that $(A, B) \perp\!\!\!\perp C \implies A \perp\!\!\!\perp C | B$, combined with iid sampling and independence of (τ^t, π_n) from the data $W_{1:n}$. Then since \mathcal{G}_n^ν is a partition of $\{T_i = 1\} \subseteq [n]$, we have

$$\frac{1}{n} \sum_{g \in \mathcal{G}_n^\nu} E[|u_g| \mathbf{1}(|u_g| > c_n) | \mathcal{F}_n] \lesssim \frac{2\bar{k}^3}{n} \sum_{i=1}^n E \left[a_i (\epsilon_i^1)^2 \mathbf{1} (a_i (\epsilon_i^1)^2 > c_n / \bar{k}^2) \middle| \psi_i \right]$$

Taking an expectation of the RHS gives $2\bar{k}^3 E[a_i (\epsilon_i^1)^2 \mathbf{1} (a_i (\epsilon_i^1)^2 > c_n / \bar{k}^2)] = o(1)$ as $n \rightarrow \infty$ since $E[a_i (\epsilon_i^1)^2] \leq |a|_\infty E[\text{Var}(Y(1) | \psi)] \lesssim \text{Var}(Y(1)) < \infty$. Then by Markov inequality $\frac{1}{n} \sum_{g \in \mathcal{G}_n^\nu} E[|u_g| \mathbf{1}(|u_g| > c_n) | \mathcal{F}_n] = o_p(1)$, so $R_n = o_p(1)$ by Lemma 10.13. An identical argument shows that $B_n, C_n = o_p(1)$. Then we have shown that $\hat{v}_1 = E[v_i^2 p_i q_i] + o_p(1)$. By symmetry, $\hat{v}_0 = E[v_i^2 (1 - p_i) q_i] + o_p(1)$.

Next, consider the cross-term estimator.

$$\begin{aligned}\widehat{v}_{10} &= n^{-1} \sum_{g \in \mathcal{G}_n} w_g \sum_{i,j \in g} Y_i(1)Y_j(0)(a_i a_j)^{1/2} D_i(1 - D_j) \\ &= n^{-1} \sum_{g \in \mathcal{G}_n} w_g \sum_{i,j \in g} (m_{1i} + \epsilon_i^1)(m_{0j} + \epsilon_j^0)(a_i a_j)^{1/2} D_i(1 - D_j) = A_n + B_n + C_n + R_n\end{aligned}$$

Denote $v_i^d = m_{di} a_i^{1/2}$ and let $\mathcal{F}_n = \sigma(\psi_{1:n}, \tau^t, \pi_n)$ as before. Note the equality $a_i b_j + b_i a_j = -(a_i - a_j)(b_i - b_j) + a_i b_i + a_j b_j$. Then the conditional expectation $E[A_n | \mathcal{F}_n]$ is equal to

$$n^{-1} \sum_{g \in \mathcal{G}_n} w_g \sum_{i,j \in g} E[v_i^1 v_j^0 D_i(1 - D_j) | \mathcal{F}_n] = n^{-1} \sum_{g \in \mathcal{G}_n} w_g \sum_{i,j \in g} v_i^1 v_j^0 \frac{a(k-a)}{k(k-1)} \mathbf{1}(i \neq j)$$

Expanding using the fact above gives

$$\begin{aligned}&= n^{-1} \sum_{g \in \mathcal{G}_n} w_g \frac{a(k-a)}{k(k-1)} \sum_{i < j \in g} (v_i^1 v_j^0 + v_j^1 v_i^0) = -n^{-1} \sum_{g \in \mathcal{G}_n} w_g \frac{a(k-a)}{k(k-1)} \sum_{i < j \in g} (v_i^1 - v_j^1)(v_i^0 - v_j^0) \\ &+ n^{-1} \sum_{g \in \mathcal{G}_n} w_g \frac{a(k-a)}{k(k-1)} \sum_{i < j \in g} (v_i^1 v_i^0 + v_j^1 v_j^0) \equiv T_{n1} + T_{n2}\end{aligned}$$

First consider T_{n1} . By Young's inequality we have

$$|T_{n1}| \lesssim n^{-1} \sum_{g \in \mathcal{G}_n} \sum_{i < j \in g} |v_i^1 - v_j^1| |v_i^0 - v_j^0| \leq (2n)^{-1} \sum_{g \in \mathcal{G}_n} \sum_{i,j \in g} |v_i^1 - v_j^1|^2 + |v_i^0 - v_j^0|^2$$

The form of each expression is identical to our analysis of A_n^1 above. Then $T_{n1} = o_p(1)$ by the same argument. Next consider T_{n2} . By counting, it's easy to see that $\sum_{i < j \in g} (v_i^1 v_i^0 + v_j^1 v_j^0) = (k-1) \sum_{i \in g} v_i^1 v_i^0$. Then setting $w_g = k/(a(k-a))$ gives

$$\begin{aligned}T_{n2} &= n^{-1} \sum_{g \in \mathcal{G}_n} w_g \frac{a(k-a)}{k(k-1)} (k-1) \sum_{i \in g} v_i^1 v_i^0 = n^{-1} \sum_{g \in \mathcal{G}_n} \sum_{i \in g} v_i^1 v_i^0 \\ &= E_n[T_i v_i^1 v_i^0] = E_n[q_i v_i^1 v_i^0] + E_n[(T_i - q_i) v_i^1 v_i^0] = E[q_i m_{1i} m_{0i} a_i] + o_p(1)\end{aligned}$$

The third equality since \mathcal{G}_n partitions the units $\{T_i = 1\}$. The final equality follows from Lemma 10.17. Next consider

$$A_n - E[A_n | \mathcal{F}_n] = n^{-1} \sum_{g \in \mathcal{G}_n} \sum_{i,j \in g} w_g v_i^1 v_j^0 \left(D_i(1 - D_j) - \frac{a(k-a)}{k(k-1)} \right) \mathbf{1}(i \neq j) \equiv n^{-1} \sum_{g \in \mathcal{G}_n} u_g$$

We have $E[u_g | \mathcal{F}_n] = 0$ for all $g \in \mathcal{G}_n$ by the calculation of $E[A_n | \mathcal{F}_n]$ above. Also, note $u_g \perp\!\!\!\perp u_{g'} | \mathcal{F}_n$ for all $g \neq g'$ by Lemma 10.15. Then by Lemma 10.13, it suffices to show that $\frac{1}{n} \sum_{g \in \mathcal{G}_n} E[|u_g| \mathbf{1}(|u_g| > c_n) | \mathcal{F}_n] = o_p(1)$ for some $c_n \rightarrow \infty$ with $c_n = o(n^{1/2})$. It's easy to see that $w_g \leq 2$, so we have

$$|u_g| \leq \sum_{i,j \in g} w_g |v_i^1| |v_j^0| \left| D_i(1 - D_j) - \frac{a(k-a)}{k(k-1)} \right| \mathbf{1}(i \neq j) \leq \sum_{i,j \in g} |v_i^1| |v_j^0|$$

Then using the indicator function bounds in our analysis above, $E[|u_g| \mathbf{1}(|u_g| > c_n) | \mathcal{F}_n]$

is bounded above by

$$\begin{aligned}
& E \left[\sum_{i,j \in g} |v_i^1| |v_j^0| \mathbf{1} \left(\sum_{i,j \in g} |v_i^1| |v_j^0| > c_n \right) | \mathcal{F}_n \right] \leq \bar{k}^2 \sum_{i,j \in g} E [|v_i^1| |v_j^0| \mathbf{1} (|v_i^1| |v_j^0| > c_n / (\bar{k})^2) | \mathcal{F}_n] \\
& \leq \bar{k}^3 \sum_{i \in g} (E [|v_i^1|^2 \mathbf{1} (|v_i^1|^2 > c_n / (\bar{k})^2) | \mathcal{F}_n] + E [|v_i^0|^2 \mathbf{1} (|v_i^0|^2 > c_n / (\bar{k})^2) | \mathcal{F}_n]) \\
& \leq \bar{k}^3 \sum_{i \in g} (E [|v_i^1|^2 \mathbf{1} (|v_i^1|^2 > c_n / (\bar{k})^2) | \psi_i] + E [|v_i^0|^2 \mathbf{1} (|v_i^0|^2 > c_n / (\bar{k})^2) | \psi_i])
\end{aligned}$$

Then $\frac{1}{n} \sum_{g \in \mathcal{G}_n^v} E[|u_g| \mathbf{1}(|u_g| > c_n) | \mathcal{F}_n]$ is bounded above by

$$\begin{aligned}
& \bar{k}^3 \frac{1}{n} \sum_{g \in \mathcal{G}_n^v} \sum_{i \in g} (E [|v_i^1|^2 \mathbf{1} (|v_i^1|^2 > c_n / (\bar{k})^2) | \psi_i] + E [|v_i^0|^2 \mathbf{1} (|v_i^0|^2 > c_n / (\bar{k})^2) | \psi_i]) \\
& = \bar{k}^3 E_n[T_i E[|v_i^1|^2 \mathbf{1} (|v_i^1|^2 > c_n / (\bar{k})^2) | \psi_i]] + E_n[T_i E[|v_i^0|^2 \mathbf{1} (|v_i^0|^2 > c_n / (\bar{k})^2) | \psi_i]] \\
& \leq \bar{k}^3 E_n[E[|v_i^1|^2 \mathbf{1} (|v_i^1|^2 > c_n / (\bar{k})^2) | \psi_i]] + E_n[E[|v_i^0|^2 \mathbf{1} (|v_i^0|^2 > c_n / (\bar{k})^2) | \psi_i]]
\end{aligned}$$

Taking expectation of the final line gives $\sum_{d=0,1} E[|v_i^d|^2 \mathbf{1} (|v_i^d|^2 > c_n / (\bar{k})^2)] \rightarrow 0$ as $n \rightarrow \infty$ by dominated convergence. Then by Markov inequality, we have shown that $\frac{1}{n} \sum_{g \in \mathcal{G}_n} E[|u_g| \mathbf{1}(|u_g| > c_n) | \mathcal{F}_n]$, so $A_n - E[A_n | \mathcal{F}_n] = o_p(1)$ by Lemma 10.13. Next, note that $B_n, C_n, R_n = o_p(1)$ by the exact argument used in our analysis of \hat{v}_1 . Then we have shown that $\hat{v}_{10} \xrightarrow{P} E[q_i m_{1i} m_{0i} a_i]$ as claimed. \square

10.5 Regression Equivalence Propositions

Proposition 10.9 (Coarse Stratification). *Assume $E[Y(d)^2] < \infty$ and $P(S = k) > 0$ for all $k \in [m]$. Suppose that $T_{1:n} \sim \text{CR}(q)$ and $D_{1:n} \sim \text{CR}(p)$. The estimator $\hat{\theta} = \hat{\tau} - \hat{\beta}'(\bar{z}_{T=1} - \bar{z})$ defined in Example 3.5 has $\sqrt{n_T}(\hat{\theta} - \text{ATE}) \Rightarrow \mathcal{N}(0, V)$ with the claimed variance.*

Proof. First, consider the regression $Y \sim 1 + D + \tilde{z} + D\tilde{z}$ defining $\hat{\tau}$ and $\hat{\beta}$. Define the within-arm regression coefficients $\hat{\alpha}_d = \text{Var}_n(\tilde{z}_i | D_i = d, T_i = 1)^{-1} \text{Cov}_n(\tilde{z}_i, Y_i(d) | D_i = d, T_i = 1)$ for $d \in \{0, 1\}$. E.g. by the calculations in Lin (2013), we have $\hat{\beta} = \hat{\alpha}_1 - \hat{\alpha}_0$ and

$$\hat{\tau} = \hat{\theta} - (\hat{\alpha}_1/p + \hat{\alpha}_0/(1-p)) E_n[(D_i - p)\tilde{z}_i | T_i = 1]$$

Consider $\hat{\alpha}_1$, for instance.

$$\text{Var}_n(\tilde{z}_i | D_i = 1, T_i = 1) = \frac{E_n[\tilde{z}_i \tilde{z}_i' D_i T_i]}{E_n[D_i T_i]} - \frac{E_n[\tilde{z}_i D_i T_i] E_n[\tilde{z}_i' D_i T_i]}{E_n[D_i T_i]^2}$$

By Definition 2.1, we have $E_n[D_i T_i] = E_n[(D_i - p)T_i] + pE_n[T_i - q] + pq = pq + O(1/n)$. Similarly, we expand $E_n[\tilde{z}_i \tilde{z}_i' D_i T_i] = E_n[\tilde{z}_i \tilde{z}_i' (D_i - p)T_i] + pE_n[\tilde{z}_i \tilde{z}_i' (T_i - q)] + pqE_n[\tilde{z}_i \tilde{z}_i']$ and $E_n[\tilde{z}_i' D_i T_i] = E_n[\tilde{z}_i' (D_i - p)T_i] + pE_n[\tilde{z}_i' (T_i - q)] + pqE_n[\tilde{z}_i]$. It's easy to check that the conditions of Lemma 10.17 are satisfied, so that $E_n[\tilde{z}_i \tilde{z}_i' D_i T_i] = pqE_n[\tilde{z}_i \tilde{z}_i'] + O_p(n^{-1/2})$ and $E_n[\tilde{z}_i' D_i T_i] = pqE_n[\tilde{z}_i] + O_p(n^{-1/2})$. Putting these facts together gives $\text{Var}_n(\tilde{z}_i | D_i = 1, T_i =$

1) = $E_n[\tilde{z}_i \tilde{z}'_i] - E_n[\tilde{z}_i]E_n[\tilde{z}'_i] + O_p(n^{-1/2})$. Similarly, $E_n[z_i|T_i = 1] = E_n[z_i] + O_p(n^{-1/2})$ so

$$\begin{aligned} E_n[\tilde{z}_i \tilde{z}'_i] - E_n[\tilde{z}_i]E_n[\tilde{z}'_i] &= E_n[z_i z'_i] - E_n[z_i]E_n[z'_i|T_i = 1] - E_n[z_i|T_i = 1]E_n[z'_i] \\ &\quad + E_n[z_i|T_i = 1]E_n[z'_i|T_i = 1] = E_n[z_i z'_i] - E_n[z_i]E_n[z'_i] + O_p(n^{-1/2}) \\ &= \text{Var}(z) + O_p(n^{-1/2}) \end{aligned}$$

Similar reasoning shows that $\text{Cov}_n(\tilde{z}_i, Y_i(d)|D_i = d, T_i = 1) = \text{Cov}(z, Y(d)) + O_p(n^{-1/2})$ if $E[Y(d)^2] < \infty$. Note that $\text{Var}(z) \succ 0$ by the assumption that $P(S = k) > 0 \forall k \in [m]$. Then we have $\hat{\alpha}_d = \text{Var}(z)^{-1} \text{Cov}(z, Y(d)) + O_p(n^{-1/2})$. Recall $z = (\mathbf{1}(S = k))_{k=1}^{m-1}$. For a function $f(X) \in L_1(X)$ define $\gamma(f) = \text{Var}(z)^{-1} \text{Cov}(z, f(X))$. Then it's easy to see that $\gamma(f)'z = E[f(X)|S] - E[f(X)|S = m]$. Now by above work $\hat{\beta} = \hat{\alpha}_1 - \hat{\alpha}_0 = \text{Var}(z)^{-1} \text{Cov}(z, Y(1) - Y(0)) + O_p(n^{-1/2}) = \gamma(c) + O_p(n^{-1/2})$. Similarly, $\hat{\alpha}_1/p + \hat{\alpha}_0/(1-p) = \text{Var}(z)^{-1} \text{Cov}(z, Y(1)/p + Y(0)/(1-p)) + O_p(n^{-1/2}) = \gamma(b)(p-p)^{-1/2} + O_p(n^{-1/2})$. Then by the fundamental expansion of the IPW estimator, we have

$$\begin{aligned} \check{\theta} &= \hat{\theta} - (\hat{\alpha}_1/p + \hat{\alpha}_0/(1-p)) E_n[(D_i - p)\tilde{z}_i|T_i = 1] - \hat{\beta}' E_n \left[\frac{z_i(T_i - q)}{q} \right] + O_p(n^{-1}) \\ &= \hat{\theta} - \gamma(b)' E_n \left[\frac{(D_i - p)z_i}{q\sqrt{p-p^2}} \right] - \gamma(c)' E_n \left[\frac{z_i(T_i - q)}{q} \right] + O_p(n^{-1}) = E_n[c(X_i)] \\ &\quad + E_n \left[\frac{T_i - q}{q} (c(X_i) - \gamma(c)'z_i) \right] + E_n \left[\frac{T_i(D_i - p)}{q\sqrt{p-p^2}} (b(X_i) - \gamma(b)'z_i) \right] + R_n + O_p(n^{-1}) \\ &= E_n[c(X_i)] + E_n \left[\frac{T_i - q}{q} (c(X_i) - E[c(X_i)|S_i]) \right] + E_n \left[\frac{T_i(D_i - p)}{q\sqrt{p-p^2}} (b(X_i) - E[b|S_i]) \right] \\ &\quad + R_n + O_p(n^{-1}) \end{aligned}$$

The first line uses $E_n[z_i|T_i = 1] = E_n[z_i]/q + O(n^{-1})$ and $E_n[(D_i - p)E_n[z_i|T_i = 1]] = O(n^{-1})$ by stratification. The second line uses the probability limits established above and that $E_n[(D_i - p)z_i] = O_p(n^{-1/2})$ and $E_n[(T_i - q)z_i] = O_p(n^{-1/2})$. The final equality follows from the characterization of $\text{Var}(z)^{-1} \text{Cov}(z, f(X))$ above and since $E_n[(D_i - p)E[b(X_i)|S_i = m]] = O(n^{-1})$ and $E[(T_i - q)E[c(X_i)|S_i = m]] = O(n^{-1})$ by stratification. The final expansion is identical to that in the proof of Theorem 3.10. The conclusion then follows by the same argument. \square

Proof of Proposition 3.11. First we discuss cross-fitting. Denote $Z_i = (W_i, D_i, T_i)$. Let $I_1 \sqcup I_2 = [n]$ be a random partition with $|I_1| \asymp n$ and $(I_1, I_2) \in \sigma(\pi_n)$ for randomness $\pi_n \perp\!\!\!\perp Z_{1:n}$. The estimator has form $E_n[F(Z_i, \hat{m}_{I(i)})]$. Use the cross-fitting pattern $I(i) = I_1$ if $i \in I_2$ and similarly for $i \in I_1$. We may rearrange the summand of $\hat{\theta}_{adj}$ as

$$\begin{aligned} &\frac{T_i D_i Y_i}{q_i p_i} - \frac{Y_i(1 - D_i)T_i}{q_i(1 - p_i)} + \hat{m}_1(\psi_i) - \hat{m}_0(\psi_i) - \frac{T_i}{q_i} \left(\frac{D_i \hat{m}_1(\psi_i)}{p_i} - \frac{(1 - D_i) \hat{m}_0(\psi_i)}{(1 - p_i)} \right) \\ &= \frac{T_i D_i Y_i}{q_i p_i} - \frac{Y_i(1 - D_i)T_i}{q_i(1 - p_i)} + \hat{m}_1(\psi_i) - \hat{m}_0(\psi_i) - \frac{T_i}{q_i} (\hat{m}_1(\psi_i) - \hat{m}_0(\psi_i)) \\ &\quad - \frac{T_i(D_i - p_i)}{q_i} \left(\frac{\hat{m}_1(\psi_i)}{p_i} - \frac{\hat{m}_0(\psi_i)}{(1 - p_i)} \right) \end{aligned}$$

Continuing the calculation, this is

$$\begin{aligned}
&= \frac{T_i D_i Y_i}{q_i p_i} - \frac{Y_i(1 - D_i)T_i}{q_i(1 - p_i)} - \frac{(T_i - q_i)}{q_i} \widehat{c}(\psi_i) - \frac{T_i(D_i - p_i)}{q_i(p_i - p_i^2)^{1/2}} \widehat{b}(\psi_i) \\
&= c(\psi_i) + \frac{(T_i - q_i)}{q_i} (c - \widehat{c})(\psi_i) - \frac{T_i(D_i - p_i)}{q_i(p_i - p_i^2)^{1/2}} (b - \widehat{b})(\psi_i) + \frac{D_i T_i \epsilon_i^1}{p_i q_i} + \frac{(1 - D_i) T_i \epsilon_i^0}{(1 - p_i) q_i}
\end{aligned}$$

Consider the term $T_{1,n} = n^{-1} \sum_{i \in I_1} q_i^{-1} (T_i - q_i) (c - \widehat{c})(\psi_i, Z_{I_2})$. By Lemma 10.15 with $m = 2$, $g = I$ and $h, \tau_s = 1$ we have $Z_{I_1} \perp\!\!\!\perp Z_{I_2} | \pi_n$. Then applying Lemma 10.15

$$\begin{aligned}
E[nT_{1,n}^2 | \pi_n, Z_{I_2}] &= n^{-1} \sum_{i \in I_1} E[q_i^{-2} (T_i - q_i)^2 (c - \widehat{c})(\psi_i, Z_{I_2})^2 | \pi_n, Z_{I_2}] \\
&= n^{-1} \sum_{i \in I_1} E[q_i^{-1} (1 - q_i) (c - \widehat{c})(\psi_i, Z_{I_2})^2 | \pi_n, Z_{I_2}] \\
&= n^{-1} |I_1| \int (1 - q)(\psi) / q(\psi) (c - \widehat{c})(\psi, Z_{I_2})^2 dP(\psi) \lesssim |c - \widehat{c}(Z_{I_2})|_{2,\psi} = o_p(1)
\end{aligned}$$

Then $\sqrt{n}T_{1,n} = o_p(1)$. Then by symmetry $\sqrt{n}E_n[q_i^{-1} (T_i - q_i) (c - \widehat{c})(\psi_i)] = o_p(1)$. Similarly, $\sqrt{n}E_n[\frac{T_i(D_i - p_i)}{q_i(p_i - p_i^2)^{1/2}} (b - \widehat{b})(\psi_i)] = o_p(1)$. The claim now follows from vanilla CLT, noting that $n_T/n \xrightarrow{P} E[q_i]$. \square

10.6 Lemmas

Lemma 10.10 (Conditional Convergence). *Let $(\mathcal{G}_n)_{n \geq 1}$ and $(A_n)_{n \geq 1}$ a sequence of σ -algebras and RV's. Define conditional convergence*

$$\begin{aligned}
A_n = o_{p,\mathcal{G}_n}(1) &\iff P(|A_n| > \epsilon | \mathcal{G}_n) = o_p(1) \quad \forall \epsilon > 0 \\
A_n = O_{p,\mathcal{G}_n}(1) &\iff P(|A_n| > s_n | \mathcal{G}_n) = o_p(1) \quad \forall s_n \rightarrow \infty
\end{aligned}$$

Then the following results hold

- (i) $A_n = o_p(1) \iff A_n = o_{p,\mathcal{G}_n}(1)$ and $A_n = O_p(1) \iff A_n = O_{p,\mathcal{G}_n}(1)$
- (ii) $E[|A_n| | \mathcal{G}_n] = o_p(1) / O_p(1) \implies A_n = o_p(1) / O_p(1)$
- (iii) $\text{Var}(A_n | \mathcal{G}_n) = o_p(c_n^2) / O_p(c_n^2) \implies A_n - E[A_n | \mathcal{G}_n] = o_p(c_n) / O_p(c_n)$.
- (iv) If $(A_n)_{n \geq 1}$ has $A_n \leq \bar{A} < \infty$ \mathcal{G}_n -a.s. $\forall n$ and $A_n = o_p(1) \implies E[|A_n| | \mathcal{G}_n] = o_p(1)$

Proof. (i) Consider that for any $\epsilon > 0$

$$P(|A_n| > \epsilon) = E[\mathbf{1}(|A_n| > \epsilon)] = E[E[\mathbf{1}(|A_n| > \epsilon) | \mathcal{G}_n]] = E[P(|A_n| > \epsilon | \mathcal{G}_n)]$$

If $A_n = o_p(1)$, then $E[P(|A_n| > \epsilon | \mathcal{G}_n)] = o(1)$, so $P(|A_n| > \epsilon | \mathcal{G}_n) = o_p(1)$ by Markov inequality. Conversely, if $P(|A_n| > \epsilon | \mathcal{G}_n) = o_p(1)$, then $E[P(|A_n| > \epsilon | \mathcal{G}_n)] = o(1)$ since $(P(|A_n| > \epsilon | \mathcal{G}_n))_{n \geq 1}$ is uniformly bounded, hence UI. Then $P(|A_n| > \epsilon) = o(1)$. The second equivalence follows directly from the first. (ii) follows from (i) and conditional Markov inequality. (iii) is an application of (ii). For (iv), note that for any $\epsilon > 0$

$$E[|A_n| | \mathcal{G}_n] \leq \epsilon + E[|A_n| \mathbf{1}(|A_n| > \epsilon) | \mathcal{G}_n] \leq \epsilon + \bar{A} P(|A_n| > \epsilon | \mathcal{G}_n) = \epsilon + o_p(1)$$

The equality is by (i) and our assumption. Since $\epsilon > 0$ was arbitrary $E[|A_n||\mathcal{G}_n] = o_p(1)$. \square

Lemma 10.11 (Lipschitz Approximation). *Let $Z \in \mathbb{R}^d$ be a random variable. Define $\mathcal{L} = \{g(Z) \in L_2(Z) : |g|_{lip} \vee |g|_\infty < \infty\}$. Then \mathcal{L} is dense in $L_2(Z)$.*

Proof. Let $\mathbb{1}(Z \in A)$ P -measurable for non-empty A . Define $f_n(z) = (1 + nd(z, A))^{-1}$ with $d(z, A) = \inf_{y \in A} |z - y|_2$. The function $z \rightarrow d(z, A)$ is 1-Lipschitz (e.g. reverse triangle inequality). Then $|f_n(z) - f_n(y)| \leq n|z - y|_2$ for any $z, y \in \mathbb{R}^d$ and $|f_n|_\infty \leq 1$ so $f_n(Z) \in \mathcal{L}$. Observe that $f_n(z) \rightarrow \mathbb{1}(z \in A)$ pointwise as $n \rightarrow \infty$. Then by dominated convergence $\int (f_n(z) - \mathbb{1}(z \in A))^2 dP(z) \rightarrow 0$, since the integrand converges pointwise and is dominated by $2 \in L_1(Z)$. Then bounded Lipschitz functions are dense in the set of indicator functions of measurable sets. Next, consider a simple function $\sum_k a_k \mathbb{1}(Z \in A_k)$ with $|a_k| < \infty$ for all k . If $g_{nk}(z) = (1 + nd(z, A_k))^{-1}$ the same argument shows that $\sum_k a_k g_{nk}(Z) - \sum_k a_k \mathbb{1}(Z \in A_k) \rightarrow 0$ in $L_2(Z)$. The left hand sum is bounded Lipschitz, showing that the Lipschitz functions are dense in the set of simple functions in $L_2(Z)$. The bounded simple functions are dense in $L_2(Z)$ (e.g. [Folland \(1999\)](#)), so bounded Lipschitz functions are dense in $L_2(Z)$ by transitivity. \square

Lemma 10.12 (Asymptotic Independence). *Consider a probability space (Ω, \mathcal{G}, P) with σ -algebras $\mathcal{F} = \mathcal{F}_{n,0}$ and $\mathcal{F}_{n,k} \subseteq \mathcal{F}_{n,k+1} \subseteq \mathcal{G}$ for all $n \geq 1$ and $0 \leq k \leq m-1$. Let $X_{n,k}$ be $\mathcal{F}_{n,k}$ -measurable random variables for all n and $1 \leq k \leq m$. Suppose that $X_{n,k}|\mathcal{F}_{n,k-1} \Rightarrow L_k|\mathcal{F}$, as in Definition 10.5. Then $(X_{n,1}, \dots, X_{n,m})|\mathcal{F} \Rightarrow (L_1, \dots, L_m)|\mathcal{F}$, with jointly independent limit.*

Proof. By Levy continuity, it suffices to show $E[e^{it'(X_{1,n}, \dots, X_{m,n})}|\mathcal{F}] \rightarrow \prod_{k=1}^m E[e^{it_k L_k}|\mathcal{F}]$ for all $t \in \mathbb{R}^m$. We work by induction on k . By assumption, $E[e^{itX_{n,1}}|\mathcal{F}] \xrightarrow{p} E[e^{itL_1}|\mathcal{F}]$ for all $t \in \mathbb{R}$. Assume by induction that the conclusion holds for $1 \leq k \leq k' \leq m$. Then

$$\begin{aligned} E[e^{i \sum_{k=1}^{k'+1} t_k X_{n,k}}|\mathcal{F}] &= E[e^{i \sum_{k=1}^{k'} t_k X_{n,k}} E[e^{it_{k'+1} X_{n,k'+1}}|\mathcal{F}_{n,k'}]|\mathcal{F}] \\ &= E[e^{i \sum_{k=1}^{k'} t_k X_{n,k}} (E[e^{it_{k'+1} X_{n,k'+1}}|\mathcal{F}_{n,k'}] - E[e^{it_{k'+1} L_{k'+1}}|\mathcal{F}])|\mathcal{F}] \\ &\quad + E[e^{it_{k'+1} L_{k'+1}}|\mathcal{F}] E[e^{i \sum_{k=1}^{k'} t_k X_{n,k}}|\mathcal{F}] = \prod_{k=1}^{k'+1} E[e^{it_{k'+1} L_{k'+1}}|\mathcal{F}] + o_p(1) \end{aligned}$$

The first equality is by tower law and our measurability and increasing σ -algebra assumption. For the final equality, note that $|e^{i \sum_{k=1}^{k'} t_k X_{n,k}} (E[e^{it_{k'+1} X_{n,k'+1}}|\mathcal{F}_{n,k'}] - E[e^{it_{k'+1} L_{k'+1}}|\mathcal{F}])| \leq |E[e^{it_{k'+1} X_{n,k'+1}}|\mathcal{F}_{n,k'}] - E[e^{it_{k'+1} L_{k'+1}}|\mathcal{F}]| \xrightarrow{p} 0$. Then the first term in the sum above is $o_p(1)$ since the integrand converges in probability and is bounded, hence UI. The final equality also uses our inductive hypothesis. This finishes the proof. \square

Lemma 10.13 (LLN). *Consider $A_n = n^{-1} \sum_{g \in \mathcal{G}_n} u_g$, with \mathcal{G}_n a collection of disjoint subsets of $[n]$. Let $(\mathcal{F}_n)_{n \geq 1}$ be σ -algebras such that \mathcal{G}_n is \mathcal{F}_n -measurable, $E[u_g|\mathcal{F}_n] = 0$, and for all $g \neq g' \in \mathcal{G}_n$ $u_g \perp\!\!\!\perp u_{g'}|\mathcal{F}_n$. If $n^{-1} \sum_{g \in \mathcal{G}_n} E[|u_g| \mathbb{1}(|u_g| > c_n)|\mathcal{F}_n] \xrightarrow{p} 0$ for $c_n = \omega(1)$, $c_n = o(n^{1/2})$, then $A_n \xrightarrow{p} 0$.*

Proof. By disjointness $|\mathcal{G}_n| \leq n$. Fix an indexing $\mathcal{G}_n = \{g_s : 1 \leq s \leq n\}$, possibly with $g_s = \emptyset$ for some s . Define $\bar{u}_{sn} = u_s \mathbb{1}(|u_s| \leq c_n)$ and $\bar{\mu}_{sn} = E[u_s \mathbb{1}(|u_s| \leq c_n)|\mathcal{F}_n]$. Expand

$$\frac{1}{n} \sum_{g \in \mathcal{G}_n} u_g = \frac{1}{n} \sum_{s=1}^n u_s = \frac{1}{n} \sum_{s=1}^n [(u_s - \bar{u}_{sn}) + (\bar{u}_{sn} - \bar{\mu}_{sn}) + \bar{\mu}_{sn}] = T_{n1} + T_{n2} + T_{n3}$$

Observe that $E[|T_{n1}||\mathcal{F}_n] \leq (1/n) \sum_{s=1}^n E[|u_s| \mathbf{1}(|u_s| > c_n)|\mathcal{F}_n] = o_p(1)$ by assumption. Then $T_{n1} = o_p(1)$ by conditional Markov (Lemma 10.10). Next consider T_{n2} . Note that by definition $E[\bar{u}_{sn} - \bar{\mu}_{sn}|\mathcal{F}_n] = 0$ for each $1 \leq s \leq n$. Note that for $s \neq s'$ $\text{Cov}(\bar{u}_{sn}, \bar{u}_{s'n}|\mathcal{F}_n) = 0$ by the conditional independence assumption. Then $\text{Var}(T_{n2}|\mathcal{F}_n) = n^{-2} \sum_{s=1}^n \text{Var}(\bar{u}_{sn}|\mathcal{F}_n) \leq n^{-2} \sum_{s=1}^n E[\bar{u}_{sn}^2|\mathcal{F}_n] \leq n^{-1} c_n^2 = o(1)$, so that $T_{n2} = o_p(1)$ by conditional Chebyshev. Finally, since $E[u_s|\mathcal{F}_n] = 0$, we have $\bar{\mu}_{sn} = -E[u_s \mathbf{1}(|u_s| > c_n)|\mathcal{F}_n]$. Then $E[|T_{n3}||\mathcal{F}_n] \leq (1/n) \sum_{s=1}^n E[|u_s| \mathbf{1}(|u_s| > c_n)|\mathcal{F}_n] = o_p(1)$, so that $T_{n3} = o_p(1)$ by conditional Markov as before. This finishes the proof. \square

Lemma 10.14. Suppose $E[|X|^p] < \infty$ for $p > 0$. Then $\max_{i=1}^n |X_i| = o_p(n^{1/p})$.

Proof. For $\epsilon > 0$ we have $P(\max_{i=1}^n |X_i| > \epsilon n^{1/p}) \leq nP(|X_i| > \epsilon n^{1/p}) = nP(|X_i|^p > \epsilon^p n) \leq n(\epsilon^p n)^{-1} E[|X_i|^p \mathbf{1}(|X_i|^p > \epsilon^p n)] \lesssim E[|X_i|^p \mathbf{1}(|X_i|^p > \epsilon^p n)] \rightarrow 0$. The first inequality by union bound, the equality by monotonicity of $x \rightarrow x^p$. The second inequality is Markov's, and the final statement by dominated convergence, since $E[|X_i|^p] < \infty$. \square

Lemma 10.15 (Random Partitions). Let $\mathcal{G}_n = \{g_s\}_{s=1}^m$ a random collection of disjoint subsets of $[n]$. Let variables $((\tau_s)_s, W_{1:n}, \pi)$ jointly independent for some random elements π and $(\tau_s)_{s=1}^m$. Let $h_i = h(W_i)$ for a fixed measurable function h and suppose that the partition \mathcal{G}_n is \mathcal{F}_n -measurable for $\mathcal{F}_n = \sigma(h_{1:n}, \pi)$. Then for $s \neq r$ and measurable F_s, F_r , we have $F_s((W_i)_{i \in g_s}, \tau_s) \perp\!\!\!\perp F_r((W_i)_{i \in g_r}, \tau_r)|\mathcal{F}_n$.

Proof. Since F_s, F_r are arbitrary, it suffices to show $E[F_s F_r|\mathcal{F}_n] = E[F_s|\mathcal{F}_n]E[F_r|\mathcal{F}_n]$. Denote $W_I = (W_i)_{i \in I}$ and similarly for h_I . Then $F_s = \sum_{I \in 2^{[n]}} \mathbf{1}(g_s = I) F_s(W_I, \tau_s)$. We claim that it suffices to show

$$E[F_s(W_I, \tau_s) F_r(W_J, \tau_r)|\mathcal{F}_n] = E[F_s(W_I, \tau_s)|\mathcal{F}_n] E[F_r(W_J, \tau_r)|\mathcal{F}_n]$$

for all disjoint $I \cap J = \emptyset$. To see this, note that in this case by measurability of \mathcal{G}_n with respect to \mathcal{F}_n and disjointness of the groups we would have

$$\begin{aligned} E[F_s F_r|\mathcal{F}_n] &= E \left[\sum_{I \in 2^{[n]}} \sum_{J \in 2^{[n]}} \mathbf{1}(g_s = I) \mathbf{1}(g_r = J) F_s(W_I, \tau_s) F_r(W_J, \tau_r) | \mathcal{F}_n \right] \\ &= \sum_{\substack{I, J \in 2^{[n]} \\ I \cap J = \emptyset}} \mathbf{1}(g_s = I) \mathbf{1}(g_r = J) E[F_s(W_I, \tau_s)|\mathcal{F}_n] E[F_r(W_J, \tau_r)|\mathcal{F}_n] = E[F_s|\mathcal{F}_n] E[F_r|\mathcal{F}_n] \end{aligned}$$

Then consider such $I \cap J = \emptyset$. Note the fact (1) if $(A, B) \perp\!\!\!\perp C$ then $A \perp\!\!\!\perp C|B$. By applying (1) with $A = (W_I, \tau_s, W_J, \tau_r)$, $B = h_{1:n}$ and $C = \pi$, we have $E[F_s(W_I, \tau_s) F_r(W_J, \tau_r)|\mathcal{F}_n] = E[F_s(W_I, \tau_s) F_r(W_J, \tau_r)|h_{1:n}]$. Then it suffices to show that (a) $(W_I, \tau_s) \perp\!\!\!\perp (W_J, \tau_r)|h_{1:n}$. By fact (1) to show (a) it suffices to prove (b) $(W_I, \tau_s, h_I) \perp\!\!\!\perp (W_J, \tau_r)|h_{I^c}$. By fact (1) again, it suffices to show (c) $(W_I, \tau_s, h_I) \perp\!\!\!\perp (W_J, \tau_r, h_{I^c})$. This is true by disjointness and joint independence of the $(\tau_s)_s$, which finishes the proof. \square

Lemma 10.16 (Design Properties). Let $D_{1:n} \sim \text{Loc}(\psi_n, p_n)$. Denote group randomization variables $\tau^d = (\tau_{a,s}^d)_{a,s}$, jointly independent with $(\tau_{a,s,\ell}^d)_{\ell=1}^{k_a} \sim \text{CR}(q_a/k_a)$ for $1 \leq s \leq n-1$ and remainder group $(\tau_{a,n,\ell}^d)_{\ell=1}^{k_a} \sim \text{SRS}(q_a/k_a)$. Let $(\mathcal{F}_n)_{n \geq 0}$ a sequence of σ -algebras with $\mathcal{G}_n \subseteq \mathcal{F}_n$ and $\mathcal{F}_n \perp\!\!\!\perp \tau^d$. Then the following hold

- (i) For each $i \in [n]$ we have $E[D_i|\mathcal{F}_n] = \sum_{a=1}^{|L_n|} \sum_{s=1}^n \mathbf{1}(i \in g_{a,s}) \cdot p_a$. In particular, $E[D_i \mathbf{1}(i \in g_{a,s})|\mathcal{F}_n] = \mathbf{1}(i \in g_{a,s}) \cdot p_a$.

(ii) For $1 \leq i \leq n$ and $1 \leq s \leq n$ we have $\text{Var}(D_i|\mathcal{F}_n)\mathbb{1}(i \in g_{a,s}) = p_a(1-p_a)\mathbb{1}(i \in g_{a,s})$.
For $1 \leq i, j \leq n$ distinct indices and $1 \leq s \leq n-1$

$$\text{Cov}(D_i, D_j|\mathcal{F}_n)\mathbb{1}(i, j \in g_{a,s}) = -\frac{q_a(k_a - q_a)}{k_a^2(k_a - 1)}\mathbb{1}(i, j \in g_{a,s})$$

In particular, $|\text{Cov}(D_i, D_j|\mathcal{F}_n)\mathbb{1}(i, j \in g_{a,s})| \leq k_a^{-1}\mathbb{1}(i, j \in g_{a,s})\mathbb{1}(s \neq n)$. Moreover, $\text{Cov}(D_i, D_j|\mathcal{F}_n)\mathbb{1}(g(i) \neq g(j)) = 0$.

Proof. For the first statement, note that

$$D_i = \sum_{a=1}^{|L_n|} \sum_{s=1}^n \sum_{\ell=1}^{k_a} D_i \mathbb{1}(i = g_{a,s,\ell}) = \sum_{a=1}^{|L_n|} \sum_{s=1}^n \sum_{\ell=1}^{k_a} \tau_{a,s,\ell}^d \mathbb{1}(i = g_{a,s,\ell})$$

Then since $\mathcal{G}_n \in \mathcal{F}_n$ and $\tau_{a,s,\ell}^d \perp\!\!\!\perp \mathcal{F}_n$ we have

$$\begin{aligned} E[D_i|\mathcal{F}_n] &= \sum_{a=1}^{|L_n|} \sum_{s=1}^n \sum_{\ell=1}^{k_a} E[\mathbb{1}(i = g_{a,s,\ell})\tau_{a,s,\ell}^d|\mathcal{F}_n] = \sum_{a=1}^{|L_n|} \sum_{s=1}^n \sum_{\ell=1}^{k_a} \mathbb{1}(i = g_{a,s,\ell}) E[\tau_{a,s,\ell}^d|\mathcal{F}_n] \\ &= \sum_{a=1}^{|L_n|} \sum_{s=1}^n \sum_{\ell=1}^{k_a} \mathbb{1}(i = g_{a,s,\ell}) E[\tau_{a,s,\ell}^d] = \sum_{a=1}^{|L_n|} \sum_{s=1}^n \sum_{\ell=1}^{k_a} \mathbb{1}(i = g_{a,s,\ell}) p_a \end{aligned}$$

To finish, note that $\sum_{\ell=1}^{k_a} \mathbb{1}(i = g_{a,s,\ell}) = \mathbb{1}(i \in g_{a,s})$ by definition. For (ii), by the decomposition above and measurability assumption, for $1 \leq j \neq i \leq n$

$$\text{Cov}(D_i, D_j|\mathcal{F}_n) = \sum_{a,a'=1}^{|L_n|} \sum_{s,s'=1}^n \sum_{\ell=1}^{k_a} \sum_{\ell'=1}^{k_{a'}} \mathbb{1}(i = g_{a,s,\ell}) \mathbb{1}(j = g_{a',s',\ell'}) \text{Cov}(\tau_{a,s,\ell}^d, \tau_{a',s',\ell'}^d|\mathcal{F}_n)$$

By σ -algebra independence and joint independence of groupwise randomizations

$$\begin{aligned} \text{Cov}(\tau_{a,s,\ell}^d, \tau_{a',s',\ell'}^d|\mathcal{F}_n) &= \text{Cov}(\tau_{a,s,\ell}^d, \tau_{a',s',\ell'}^d) \\ &= \begin{cases} 0 & (a, s) \neq (a', s') \\ p_a - p_a^2 & (a, s, \ell) = (a', s', \ell') \\ -\frac{q_a(k_a - q_a)}{k_a^2(k_a - 1)} & (a, s) = (a', s'); \quad \ell \neq \ell' \quad 1 \leq s < n \\ 0 & (a, s) = (a', s'); \quad \ell \neq \ell' \quad s = n \end{cases} \end{aligned}$$

The third line follows since by definition of $\text{CR}(q_a/k_a)$, for $(a, s) = (a', s')$ we have

$$\begin{aligned} \text{Cov}(\tau_{a,s,\ell}^d, \tau_{a',s',\ell'}^d) &= P(\tau_{a,s,\ell}^d = \tau_{a',s',\ell'}^d = 1) - (q_a/k_a)^2 = \binom{k_a}{q_a}^{-1} \binom{k_a - 2}{q_a - 2} - (q_a/k_a)^2 \\ &= \frac{q_a(q_a - 1)}{k_a(k_a - 1)} - (q_a/k_a)^2 = -\frac{q_a(k_a - q_a)}{k_a^2(k_a - 1)} \end{aligned}$$

The bounds follow by inspection. \square

Lemma 10.17 (Stochastic Balance). *Let $(\mathcal{F}_n)_{n \geq 1}$ such that treatment groups $\mathcal{G}_n, (h_n(W_i))_{i=1}^n \in \mathcal{F}_n$ for a sequence of functions $(h_n)_{n \geq 1}$ and $\mathcal{F}_n \perp\!\!\!\perp \tau^d$. Let $D_{1:n} \sim \text{Loc}(\psi_n, p_n)$ and $T_{1:n} \in \{0, 1\}^n$ such that $\{i : T_i = 1\} = \sqcup_{g \in \mathcal{G}_n} g$. The following hold*

- (1) $E[E_n[T_i(D_i - p_n(X_i))h_n(W_i)]|\mathcal{F}_n] = 0$
- (2) $\text{Var}(E_n[T_i(D_i - p_n(X_i))h_n(W_i)]|\mathcal{F}_n) \leq 2n^{-1}E_n[T_i h_n(W_i)^2] \leq 2n^{-1}E_n[h_n(W_i)^2]$
- (3) Suppose $\exists (\mathcal{F}_n)_{n \geq 1}$ satisfying the conditions above. If $\sup_{n \geq 1} E[h_n(W)^2] < \infty$ then $E_n[T_i(D_i - p_n(X_i))h_n(W_i)] = O_p(n^{-1/2})$. If $(h_n(W))_{n \geq 1}$ is uniformly integrable then $E_n[T_i(D_i - p_n(X_i))h_n(W_i)] = o_p(1)$.
- (4) The variance $\text{Var}(\sqrt{n}E_n[T_i(D_i - p_n(X_i))h_n(W_i)]|\mathcal{F}_n)$ is bounded above by

$$n^{-1} \sum_{g \in \mathcal{G}_n} |g|^{-1} \sum_{i,j \in g} (h_n(W_i) - h_n(W_j))^2 + n^{-1} \bar{k}_n |L_n| \cdot \max_{i=1}^n h_n(W_i)^2$$

Proof. First, by assumption $\{i : T_i = 1\} = \bigsqcup_{g \in \mathcal{G}} g$, so $E_n[T_i(D_i - p_{i,n})h_n(W_i)]$ is equal to

$$n^{-1} \sum_{g \in \mathcal{G}_n} \sum_{i \in g} (D_i - p_{i,n})h_n(W_i) = n^{-1} \sum_{a=1}^{|L_n|} \sum_{s=1}^n \sum_{i=1}^n (D_i - p_a)h_n(W_i) \mathbf{1}(i \in g_{a,s})$$

By Lemma 10.16 and our measurability assumptions

$$E[(D_i - p_a)h_n(W_i) \mathbf{1}(i \in g_{a,s})|\mathcal{F}_n] = h_n(W_i)E[(D_i - p_a) \mathbf{1}(i \in g_{a,s})|\mathcal{F}_n] = 0$$

By linearity, this shows the claim. By Lemma 10.16, $\text{Var}(E_n[(D_i - p_{i,n})h_n(W_i)]|\mathcal{F}_n)$ is

$$\begin{aligned} & n^{-2} \sum_{a,a'}^{|L_n|} \sum_{s,s'=1}^n \sum_{i,j}^n \text{Cov}((D_i - p_a)h_n(W_i) \mathbf{1}(i \in g_{a,s}), (D_j - p_{a'})h_n(W_j) \mathbf{1}(j \in g_{a',s'})|\mathcal{F}_n) \\ &= n^{-2} \sum_{a,a'}^{|L_n|} \sum_{s,s'=1}^n \sum_{i,j}^n h_n(W_i)h_n(W_j) \mathbf{1}(i \in g_{a,s}) \mathbf{1}(j \in g_{a',s'}) \text{Cov}(D_i, D_j|\mathcal{F}_n) \\ &= n^{-2} \sum_{a=1}^{|L_n|} \sum_{s=1}^n \sum_{i,j}^n h_n(W_i)h_n(W_j) \mathbf{1}(i, j \in g_{a,s}) \text{Cov}(D_i, D_j|\mathcal{F}_n) \end{aligned}$$

The final equality follows from Lemma 10.16. By triangle inequality and the covariance bound in Lemma 10.16, this is bounded above by

$$\begin{aligned} & n^{-2} \sum_{a=1}^{|L_n|} \sum_{s=1}^n \left[\sum_{i=1}^n h_n(W_i)^2 \mathbf{1}(i \in g_{a,s}) + \sum_{i \neq j}^n |h_n(W_i)| |h_n(W_j)| \mathbf{1}(i, j \in g_{a,s}) k_a^{-1} \mathbf{1}(s \neq n) \right] \\ & \leq n^{-1} E_n[T_i h_n(W_i)^2] + \sum_{a=1}^{|L_n|} k_a^{-1} \sum_{s=1}^n \sum_{i \neq j}^n |h_n(W_i)| |h_n(W_j)| \mathbf{1}(i, j \in g_{a,s}) \\ & \leq n^{-1} E_n[T_i h_n(W_i)^2] + n^{-2} \sum_{a=1}^{|L_n|} k_a^{-1} \sum_{s=1}^n \left(\sum_{i=1}^n |h_n(W_i)| \mathbf{1}(i \in g_{a,s}) \right)^2 \end{aligned}$$

Continuing the calculation, by Jensen's inequality the second term is equal to

$$n^{-2} \sum_{a=1}^{|L_n|} k_a \sum_{s=1}^n \left(k_a^{-1} \sum_{i \in g_{a,s}} |h_n(W_i)| \right)^2 \leq n^{-2} \sum_{a=1}^{|L_n|} \sum_{s=1}^n k_a k_a^{-1} \sum_{i \in g_{a,s}} |h_n(W_i)|^2 = n^{-1} E_n[T_i h_n(W_i)^2]$$

This completes the proof of (2). Claim (3) follows by applying (1), (2) with $(\mathcal{F}_n)_{n \geq 1}$ any sequence satisfying the conditions, followed by Markov inequality (Lemma 10.10). For the final part of (3), let $c_n \rightarrow \infty$ with $c_n = o(\sqrt{n})$ and define $\bar{h}_{in} = h_n(W_i) \mathbf{1}(|h_n(W_i)| \leq c_n)$. Consider the expansion $E_n[T_i(D_i - p_{i,n})h_n(W_i)] = E_n[T_i(D_i - p_{i,n})(h_n(W_i) - \bar{h}_{in}(W_i))] + E_n[T_i(D_i - p_{i,n})\bar{h}_{in}(W_i)] \equiv A_n + B_n$. We may write $|A_n| \leq E_n[|h_n(W_i) - \bar{h}_{in}(W_i)|] = E_n[|h_n(W_i)| \mathbf{1}(|h_n(W_i)| > c_n)]$. Then we have $E[|A_n|] \leq E[|h_n(W_i)| \mathbf{1}(|h_n(W_i)| > c_n)] \rightarrow 0$ as $n \rightarrow \infty$ by uniform integrability. Then $A_n = o_p(1)$ by Markov inequality. By the first part of (3), $\text{Var}(B_n | \mathcal{F}_n) \leq 2n^{-1} E_n[\bar{h}_{in}(W_i)^2] \leq 2n^{-1} c_n^2 = o(1)$. Then $B_n = o_p(1)$ by conditional Chebyshev.

For the final identity (5), note that from Lemma 10.16 for $s \neq n$ and $p_a = q_a/k_a$

$$\begin{aligned} \sum_{i,j \in g_{a,s}} h_n(W_i) h_n(W_j) \text{Cov}(D_i, D_j | \mathcal{F}_n) &= \frac{q_a(k_a - q_a)}{k_a^2} \sum_{i \in g_{a,s}} h_n(W_i)^2 \\ &\quad + \frac{q_a(k_a - q_a)}{k_a^2(k_a - 1)} \sum_{i,j \in g_{a,s}} (-2) h_n(W_i) h_n(W_j) \end{aligned}$$

Note that $-2ab = (a - b)^2 - a^2 - b^2$. Then the second sum is

$$\begin{aligned} \sum_{i,j \in g_{a,s}} (-2) h_n(W_i) h_n(W_j) &= \sum_{i,j \in g_{a,s}} [(h_n(W_i) - h_n(W_j))^2 - h_n(W_i)^2 - h_n(W_j)^2] \\ &= \sum_{i,j \in g_{a,s}} (h_n(W_i) - h_n(W_j))^2 - (k_a - 1) \sum_{i \in g_{a,s}} h_n(W_i)^2 \end{aligned}$$

Substituting in the first display above, the diagonal terms cancel. For the claimed constant, note that $\max_{p \in (0,1)} p(1-p) \leq 1/4$ and $\max_{k \geq 2} \frac{k}{k-1} \leq 2$, so $\frac{q_a(k_a - q_a)}{k_a^2(k_a - 1)} \leq k_a^{-1}$. Aggregating over (a, s) gives

$$n^{-1} \sum_{a=1}^{|L_n|} \left[\sum_{s=1}^{n-1} \sum_{i,j \in g_{a,s}} (h_n(W_i) - h_n(W_j))^2 + \sum_{i=1}^n h_n(W_i)^2 \mathbf{1}(i \in g_{a,n}) \right]$$

The second term is

$$\begin{aligned} n^{-1} \sum_{a=1}^{|L_n|} \sum_{i=1}^n h_n(W_i)^2 \mathbf{1}(i \in g_{a,n}) &\leq n^{-1} \max_{i=1}^n h_n(W_i)^2 \sum_{a=1}^{|L_n|} \sum_{i=1}^n \mathbf{1}(i \in g_{a,n}) \\ &\leq n^{-1} \bar{k}_n |L_n| \cdot \max_{i=1}^n h_n(W_i)^2 \end{aligned}$$

This finishes the proof. \square