Policy design in experiments with unknown interference*

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First Version: November, 2020;
This Version: February, 2022

Abstract

This paper proposes an experimental design for estimation and inference on welfare-maximizing policies in the presence of spillover effects. I consider a setting where units are organized into a *finite* number of large clusters and interact in unknown ways within each cluster. As a first contribution, I introduce a single-wave experiment which, by carefully varying the randomization across pairs of clusters, estimates the marginal effect of a change in treatment probabilities, taking spillover effects into account. Using the marginal effect, I propose a practical test for policy optimality. The idea is that researchers should report the marginal effect and test for policy optimality: the marginal effect indicates the direction for a welfare improvement, and the test provides evidence on whether it is worth conducting additional experiments to estimate a welfare-improving treatment allocation. As a second contribution, I design a multiple-wave experiment to estimate treatment assignment rules and maximize welfare. I derive strong small-sample guarantees on the difference between the maximum attainable welfare and the welfare evaluated at the estimated policy, which converges linearly in the number of clusters and iterations. Simulations calibrated to existing experiments on information diffusion and cash-transfer programs show welfare improvements up to fifty percentage points.

Keywords: Experimental Design, Spillovers, Welfare Maximization, Causal Inference.
JEL Codes: C31, C54, C90.

*Previous versions can be found at https://arxiv.org/abs/2011.08174. I am especially grateful to my advisor Graham Elliott, James Fowler, Paul Niehaus, Yixiao Sun, and Kaspar Wüthrich for their continuous advice and support. I wish to thank Karun Adusumilli, Peter Aronow, Timothy Christensen, Aureo de Paula, Tjeerd de Vries, Brian Karrer, Matt Goldman, Amanda Kowalski, Xinwei Ma, Craig Mcintosh, Karthik Muralidharan, Ariel Pakes, Jack Porter, Cyrus Samii, Pedro Sant'Anna, Fredrik Sävje, Jesse Shapiro, Pietro Spini, Alex Tetenov, Ye Wang, Andrei Zeleneev, and participants at numerous seminars and presentations for helpful comments and discussion. All mistakes are my own.

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1 Introduction

One of the goals of a government or NGO is to estimate the welfare-maximizing policy. Network interference is often a challenge: treating an individual may also generate spillovers and affect the design of the policy. For instance, approximately 40% of experimental papers published in the "top-five" economic journals in 2020 mention spillover effects as a possible threat when estimating the effect of the program.[1] Researchers have become increasingly interested in experimental designs for choosing the treatment rule (policy) which maximizes welfare.[2] But when it comes to experiments on networks, standard approaches are geared towards the estimation of treatment effects. Estimation of treatment effects, on its own, is not sufficient for welfare-maximization.[3] For example, when assigning cash transfers, these may have the largest direct effect when given to people living in remote areas but generate the smallest spillovers. This trade-off has significant policy implications when treating each individual is costly or infeasible.

This paper studies experimental designs in the presence of network interference when the goal is welfare maximization. The main difficulty in these settings is that interactions can be challenging to measure, and collecting network information can be very costly as it may require enumerating all individuals and their connections in the population.[4] We, therefore, focus on a setting with limited information on the network. This is formalized by assuming that units are organized into a *finite* number of large clusters, such as schools, districts, or regions, and interact through an unobserved network (and in unknown ways) within each cluster. In the cash-transfer program, we may expect that treatments generate spillovers to those living in the same or nearby villages[5], but spillovers are negligible between individuals in different regions. We propose the first experimental design to analyze and estimate *welfare-maximizing* treatment rules in the presence of unobserved spillovers on networks.

We make two main contributions. As a first contribution, we introduce a design where researchers randomize treatments and collect outcomes once (*single-wave experiment*) to (i) test whether one or more treatment allocation rules, such as the one currently implemented by the policymaker, maximize welfare; (ii) estimate how we can improve welfare with a (small) change to allocation rules. The experiment is based on a simple idea. With a

[1] This is based on the author's calculation. Top-five economic journals are *American Economic Review*, *Econometrica*, *Journal of Political Economy*, *Quarterly Journal of Economics*, *Review of Economic Studies*.

[2] See Bubeck et al. (2012) and Kasy and Sautmann (2019) for a discussion.

[3] Examples of treatment effects are the direct effects of the treatment and the overall effect, i.e., the effect if we treat all individuals, compared to treating *none*. For welfare maximization, none of these estimands is sufficient. The direct effect ignores spillovers, while the optimal rule may only treat some but not individuals either because of treatment costs or because of decreasing marginal returns from neighbors' treatments.

[4] See Breza et al. (2020) for a discussion on the cost associated with collecting network information.

[5] For instance, Egger et al. (2019) document spillovers from cash-transfers between nearby villages.

small number of clusters, we do not have enough information to precisely estimate the welfare-maximizing treatment rule. However, if we take *two* clusters and assign treatments in each cluster independently with slightly different (locally perturbated) probabilities, we can estimate the marginal effect of a change in the treatment assignment rule (which we will refer to as marginal *policy* effect, MPE). For instance, in the cash-transfer example, the MPE defines the marginal effect of treating more people in remote areas, taking spillover effects into account.[6] Using the MPE, we introduce a practical test for whether there exists a welfare-improving treatment allocation rule. As this paper suggests, researchers should report estimates of the MPE and test for welfare-maximizing policies. The MPE indicates the *direction* for a welfare improvement, and the test provides evidence on whether it is worth conducting additional experiments to estimate a welfare-improving treatment allocation.

Specifically, the experiment *pairs* clusters and randomizes treatments independently within clusters, with local perturbations to treatment probabilities within each pair. The difference in treatment probabilities balances the bias and variance of a difference-in-differences estimator. We show that the estimator for each pair converges to the marginal effect and derive properties for inference with finitely many clusters. The experiment separately estimates the direct and spillover effects, which are of independent interest. These are the effect on the recipients and the marginal effect of increasing the neighbors' treatment probability.[7]

As a second contribution, we offer an adaptive (i.e., *multiple-wave*) experiment to estimate welfare-maximizing allocation rules. Our goal here is to adaptively randomize treatments to estimate the welfare-maximizing policy while also improving participants' welfare.[8] We propose an experiment which guarantees tight small-sample upper bounds for *both* the (i) out-of-sample regret, i.e., the difference between the maximum attainable welfare and the welfare evaluated at the estimated policy deployed on a new population, and the (ii) in-sample regret, i.e., the regret of the experimental participants. The experiment groups clusters into pairs, using as many pairs as the number of iterations (or more); every iteration, it randomizes treatments in a cluster and perturbs the treatment probability within each pair; finally, it updates policies sequentially, using the information on the marginal effects from a different pair via gradient descent. We illustrate the existence of a bias in adaptive experiments with repeated sampling and develop a novel algorithm with circular updates to avoid the bias.

From a theoretical perspective, a corollary of our sequential experiment's small-sample

[6]This is the derivative of welfare with respect to the policy's parameters, taking spillovers into account, different from what known in observational studies as marginal treatment effect (Carneiro et al., 2010), which instead depends on the individual selection into treatment mechanism.

[7]Also, between-clusters local perturbations accommodate settings where policymakers cannot allow much variation in how treatments are assigned between different clusters because of exogenous constraints.

[8]Improving participants' welfare is desirable for large-scale experiments, common on online platforms (Karrer et al., 2021), and of increasing interest in development studies (Muralidharan and Niehaus, 2017).

guarantees is that the out-of-sample regret converges at a faster-than-parametric rate in the number of clusters and iterations, and similarly the in-sample regret.[9] We note that there are no regret guarantees tailored to unobserved interference. Existing results for treatment choice with *i.i.d.* data, treating clusters as sampled observations, would instead imply a slower convergence in the number of clusters.[10] We achieve a faster rate by (a) exploiting *within*-cluster variation in assignments and *between* clusters' local perturbations; (b) deriving concentration within each cluster as the cluster's size increases; (c) assuming and leveraging decreasing marginal effects of increasing neighbors' treatment probability.

We illustrate the numerical properties of the method with calibrated experiments that use data from an information diffusion experiment (Cai et al., 2015) and a cash-transfer program (Alatas et al., 2012, 2016). We show that our test can, in expectation, lead to welfare improvements up to fifty percentage points if, upon rejections of the null hypothesis that increasing treatment probabilities does not improve welfare, policy-makers increase treatment probabilities by five percent. When designing an adaptive experiment, the proposed method substantially improves both out-of-sample and in-sample regret, even with few iterations.

Throughout the text, we assume that the maximum degree grows at an appropriate slower rate than the cluster size; covariates and potential outcomes are identically distributed between clusters; treatment effects do not carry over in time. In the Appendix, we relax these assumptions and study three extensions: (a) experimental design with a global interference mechanism; (b) matching clusters with covariates drawn from cluster-specific distributions, and introduce matching via distributional embeddings; (c) experimental design with *dynamic* treatment effects, and propose a novel experimental design in this setting.

Our paper adds to both the literature on single-wave and multiple-wave experiments. In the context of *single-wave* (or two-wave) experiments, existing network experiments include clustered experiments (Eckles et al., 2017; Ugander et al., 2013; Karrer et al., 2021) and saturation designs (Baird et al., 2018; Pouget-Abadie, 2018). References with observed networks include Basse and Airoldi (2018b), Jagadeesan et al. (2020), Viviano (2020) among others.[11] See also Bai (2019); Tabord-Meehan (2018) with *i.i.d.* data. These authors study

[9] The average out-of-sample regret converges at a rate $1/T$, where T is the number of iterations and proportional to the number of clusters, and $\log(T)/T$ the in-sample regret. For the out-of-sample regret, we derive an exponential rate $\exp(-c_0 T)$, for a positive constant c_0 under additional restrictions (see Sec 4.1).

[10] For treatment choice, Kitagawa and Tetenov (2018) establish distribution-free lower bounds of order $1/\sqrt{n}$. In the literature on bandit feedback Shamir (2013) provides lower bounds of order $1/\sqrt{n}$ for continuous stochastic optimization procedure with strongly-convex functions (see also Bubeck et al. (2011)); optimization connects to bandits of Flaxman et al. (2004); Agarwal et al. (2010) which, however, provide slower rates for high-probability bounds (see also Section 4.1). We note that Wager and Xu (2021) provide rates of order $1/T$ for in-sample regret but leverage an explicit model for market interactions with infinite asymptotics. Here, we do not impose assumptions on the interference mechanism and consider finitely many clusters. Kasy and Sautmann (2019) provides bounds for either (but not both) notions of regret in finite sample.

[11] For the analysis on the bias of average treatment effect estimators with interference see also Basse and

experimental designs for inference on treatment effects only, but not the problem of inference on welfare-maximizing policies. Different from all the above references, we propose a design that allows us to identify the marginal effect under interference (and treatment effects). We introduce the first test and design for inference on policy optimality under unobserved interference. This difference motivates us to introduce a novel design consisting of local perturbations between clusters. The idea of using the information on marginal effects connects to the literature on optimal taxation (Saez, 2001; Chetty, 2009), which differently consider observational studies with independent units.

With *multiple-wave* experiments, we introduce a framework for adaptive experimentation with unknown interference. We connect to the literature on adaptive exploration (Bubeck et al., 2012; Russo et al., 2017; Kasy and Sautmann, 2019, among others), and the one on derivative free stochastic optimization (Flaxman et al., 2004; Kleinberg, 2005; Shamir, 2013; Agarwal et al., 2010, among others). These references do not study the problem of network interference. Here, we leverage between-clusters perturbations and within-cluster concentration to obtain high-probability bounds on the regret with fast rates. In related work, Wager and Xu (2021) have studied prices estimation via local experimentation in the different context of a single market, with asymptotically independent agents. They assume infinitely many individuals and an explicit model for market prices. As noted by the authors, the structural assumptions imposed in the above reference do not allow for spillovers on a network (i.e., individuals may depend arbitrarily on neighbors' assignments). Our setting differs due to network spillovers and the fact that individuals are organized into finitely many independent components (clusters), where such spillovers are unobserved. These differences motivate (i) our design mechanism, which exploits two-level randomization at the cluster and individual level instead of individual-level randomization, (ii) pairing and perturbations between clusters. From a theoretical perspective, network dependence and repeated sampling induce novel challenges for an adaptive experiment which we address here. Also, none of the above references on adaptive experiments studies inference.

We relate to the literature on inference under interference and draw from Hudgens and Halloran (2008) for definitions of potential outcomes. Differently from our paper, this literature does not study experimental design and welfare-maximization. Aronow and Samii (2017); Manski (2013); Leung (2020); Ogburn et al. (2017); Goldsmith-Pinkham and Imbens (2013); Li and Wager (2020) assume an observed network, while Vazquez-Bare (2017), Hudgens and Halloran (2008), Ibragimov and Müller (2010) consider clusters among others. Sävje et al. (2021) study inference of the direct effect of treatment only. Our focus on policy

Feller (2018), Johari et al. (2020), Basse and Airoldi (2018a), and Imai et al. (2009) when matching with different-sized clusters for overall average treatment effects.

optimality and experimental design differs from the above references. We show that inference on policy-optimality requires information on the MPE, here estimated with a clusters pair.

More broadly, we connect to the literature on statistical treatment rules, on estimation Manski (2004); Kitagawa and Tetenov (2018); Athey and Wager (2021); Bhattacharya and Dupas (2012); Stoye (2009); Mbakop and Tabord-Meehan (2021); Kitagawa and Wang (2021); Sasaki and Ura (2020); Viviano (2019), among others, and inference Andrews et al. (2019); Rai (2018); Armstrong and Shen (2015); Kasy (2016);[12] see also the literature on classification (Elliott and Lieli, 2013). This literature considers an existing experiment and does not study experimental design. Here, instead, we leverage an adaptive procedure to maximize out-of-sample and participants' welfare. Also, this literature has not studied policy design with unobserved interference. We broadly relate also to the literature on targeting on networks (see for a review Bloch et al., 2019), which mainly focus on model-based approaches with a single observed network – different from here where we leverage clusters' variations; the one on peer groups composition (Graham et al., 2010), and the one on inference with externalities (e.g., Bhattacharya et al., 2013, which, however, does not study inference on policy optimality). These references also do not study experimental designs.

The rest of the paper is organized as follows. Section 2 introduces the setup, and an illustration of the method. Section 3 studies the single-wave experiment. Section 4 presents the adaptive experiment. Section 5 collects the numerical experiments. Section 6 presents dynamic treatments. Section 7 gives our conclusions. Appendices A, B present extensions.

2 Setup and method's overview

This section introduces conditions, estimands, and a brief overview of the method.

We consider a setting with K clusters, where K is an even number. We assume that each cluster has N individuals, while our framework directly extends to clusters of different but proportional sizes. Observables and unobservables are jointly independent between clusters but not necessarily within clusters. Each cluster k is associated with a vector of outcomes, covariates, treatment assignments, and an adjacency matrix which is different for each cluster. These are $Y_{i,t}^{(k)} \in \mathcal{Y}, D_{i,t}^{(k)} \in \{0,1\}, X_i^{(k)} \in \mathcal{X} \subseteq \mathbb{R}^p, A^{(k)} \in \mathcal{A}$, respectively. Here, $(Y_{i,t}^{(k)}, D_{i,t}^{(k)})$ denote the outcome and treatment assignment of individual i at time t in cluster k, respectively, $X_i^{(k)}$ are time-invariant (baseline) covariates, and $A^{(k)}$ is a cluster-specific adjacency matrix. Each period t, researchers only observe a random subsample

$$\left(Y_{i,t}^{(k)}, X_i^{(k)}, D_{i,t}^{(k)}\right)_{i=1}^{n}, \quad n = \lambda N, \quad \lambda \in (0,1],$$

[12] See also Kato and Kaneko (2020); Hadad et al. (2019); Imai and Li (2019); Hirano and Porter (2020), which do not allow for testing for policy optimality, but construct confidence bands for the welfare.

where n defines the sample size of observations from each cluster and proportional to the cluster size.[13] There are T periods in total. While units sampled each period may or may not be the same, with abuse of notation, we index sampled units $i \in \{1, \cdots, n\}$. Whenever we provide asymptotic analyses, we let N grow[14] and K be fixed.

2.1 Setup: covariates, network and potential outcomes

Next, we introduce conditions on the covariates, network, and potential outcomes. These conditions guarantee that Lemma 2.1 (in Section 2.2) holds; practitioners may skip this subsection and directly refer to Section 2.2 for their implications, keeping in mind the covariates' distribution in Equation (1). We now discuss the network and covariates. We assume that individuals can form a link with a subset of individuals in each cluster. Formally, in each cluster, nodes are spaced under some latent space (Lubold et al., 2020) and can interact with at most the $\gamma_N^{1/2}$ closest nodes under the latent space. We say $1\{i_k \leftrightarrow j_k\} = 1$ if individual i can interact with j in cluster k and zero otherwise. Conditional on the indicators $1\{i_k \leftrightarrow j_k\}$,

$$(X_i^{(k)}, U_i^{(k)}) \sim_{i.i.d.} F_{U|X} F_X, \quad A_{i,j}^{(k)} = l\Big(X_i^{(k)}, X_j^{(k)}, U_i^{(k)}, U_j^{(k)}\Big) 1\{i_k \leftrightarrow j_k\} \tag{1}$$

for an arbitrary and unknown function $l(\cdot)$ and unobservables $U_i^{(k)}$. Equation (1) states that whether two individuals interact depends on: (i) whether they are close enough within a certain latent space (captured by $1\{i_k \leftrightarrow j_k\}$); (ii) their covariates and unobserved individual heterogeneity (i.e., X_i, U_i) which capture homophily. Equation (1) also states that covariates are $i.i.d.$ unconditionally on $A^{(k)}$, but not necessarily conditionally.[15] Figure 1 provides an illustration. Here, we condition on the indicators $1\{i_k \leftrightarrow j_k\}$ (which can differ across clusters) to control the network's maximum degree but not on the network $A^{(k)}$. Equivalently, we can interpret such indicators as exogenously drawn from some arbitrary distribution.[16]

Appendix B.8 generalizes our framework as $l(X_i, X_j, U_i, U_j, \omega_{i,j})$ also depends on exogenous (and independent) unobservables $\omega_{i,j}$, here omitted for expositional convenience.

Assumption 2.1 (Network)**.** For $i \in \{1, \cdots, N\}, k \in \{1, \cdots, K\}$, let (i) Equation (1) hold given the indicators $1\{i_k \leftrightarrow j_k\}$, for some unknown $l(\cdot)$; (ii) $\sum_{j=1}^{n} 1\{i_k \leftrightarrow j_k\} = \gamma_N^{1/2}$.

Assumption 2.1 states the following: before being born, each individual may interact with

[13] Our framework directly extends when $n = g(N)$ for a generic monotonic function $g(\cdot)$, and when we sample a different but proportional number of individuals from each cluster.

[14] Here, we consider a sequence of data-generating processes.

[15] We might also augment $l(\cdot)$ with additional $i.i.d.$ exogenous η_{ij} without affecting our results.

[16] Formally, $\mathcal{I}_k \sim \mathcal{P}_k, \quad (X_i^{(k)}, U_i^{(k)}) | \mathcal{I}_k \sim_{i.i.d.} F_{U|X} F_X, \quad A_{i,j}^{(k)} | \mathcal{I}_k = l\Big(X_i^{(k)}, X_j^{(k)}, U_i^{(k)}, U_j^{(k)}\Big) 1\{i_k \leftrightarrow j_k\}$, where $\mathcal{I}_k$ is the matrix of such indicators in cluster k and $\mathcal{P}_k$ is a cluster-specific distribution left unspecified.

$\gamma_N^{1/2}$ many other individuals. After the birth, the individual's gender, income, parental status determine her type and the distribution of her and her potential connections' edges.[17] Here $\gamma_N^{1/2}$ captures the degree of dependence. Whenever $\gamma_N^{1/2}$ equals N, we impose no restriction on the individual's number of connections. In Theorem 3.1 the maximum degree can be equal to the cluster's size, while subsequent results require stricter restrictions.[18]

We now discuss potential outcomes. Under interference, outcomes depend on individuals' covariates, unobservables and neighbors' treatments. That is, $Y_{i,t}^{(k)}(\mathbf{d}_1^{(k)}, \cdots, \mathbf{d}_t^{(k)})$ denotes the potential outcome of individual i at time t. Here $\mathbf{d}_s^{(k)} \in \{0,1\}^N$ denotes the treatment assignments at time s of all individuals in cluster k. The following condition is imposed.

Assumption 2.2 (Potential outcomes)**.** Suppose that for any $i, t, k, \mathbf{d}_s^{(k)} \in \{0,1\}^N, s \leq t$

$$Y_{i,t}^{(k)}(\mathbf{d}_1^{(k)}, \cdots, \mathbf{d}_t^{(k)}) = r\left(\mathbf{d}_{i,t}^{(k)}, \mathbf{d}_{j:A_{i,j}^{(k)}>0,t}^{(k)}, X_i^{(k)}, X_{j:A_{i,j}^{(k)}>0}^{(k)}, U_i, U_{j:A_{i,j}^{(k)}>0}, A_{i,\cdot}^{(k)}, \nu_{i,t}^{(k)}\right) + \tau_k + \alpha_t$$

for some unknown function $r(\cdot)$, stationary (but possibly serially dependent) unobservables $\nu_{i,\cdot}^{(k)} | X^{(k)}, U^{(k)} \sim_{i.i.d.} P_\nu$, fixed effects τ_k, α_t.

Assumption 2.2 imposes three conditions. First, treatment effects do not carry-over in time. Second, potential outcomes are stationary up to separable fixed effects (unobservables can depend over time).[19] Third, potential outcomes can depend on neighbors' assignments, neighbors' covariates, and neighbors' identities, with arbitrary and unknown heterogeneity in spillover effects. Note that $A_{i,\cdot}^{(k)}$ denotes the set of connections of individual i in cluster k, with $\{j : A_{i,j}^{(k)} > 0\}$ denoting those individuals with some connection to i. Assumption 2.2 imposes no condition on the dependence on neighbors' assignments but assumes that cluster fixed effects do not depend on treatment assignments, which is important for identification.

Remark 1 (Extensions)**.** In Appendix A and Appendix B we present extensions. In Appendix A.1 individuals depend on general equilibrium effects, and in Appendix A.3 on the past assignments. Appendix B.3 presents non-separable fixed effects. Appendix B.5 studies staggered adoption. Appendix A.2, A.5 present effects heterogeneous on clusters' types. □

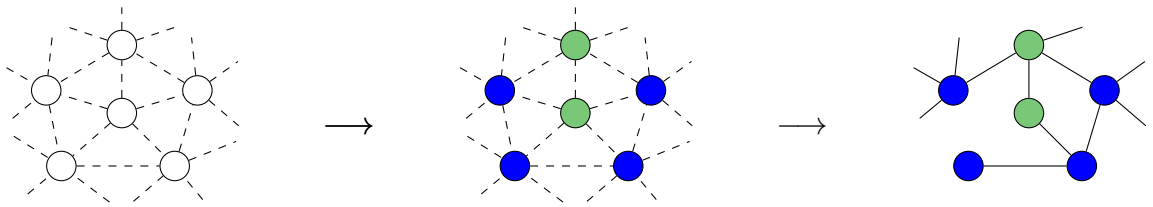
2.2 Policy and welfare maximization

The goal of this paper is to estimate a policy (treatment assignment rule) that maximizes welfare. We focus on a parametric class of policies, indexed by some parameter β. Formally,

[17] Networks formed from pairwise interactions have also been discussed in Jackson and Wolinsky (1996), Li and Wager (2020), Leung (2019). Here, we impose such restrictions to obtain easy-to-interpret conditions on the degree. Assumption 2.1 is sufficient but not necessary.

[18] Assumption 2.1 would not be required if we were to observe neighbors' assignments as in Viviano (2019).

[19] Such condition is often implicitly imposed in studies on experimental design (Kasy and Sautmann, 2019). For a discussion on the no-carry-over assumption, see Athey and Imbens (2018). We relax it in Section A.3.



Possible connections Types' assignment Network formation

Figure 1: Example of the network formation model, with $\gamma_N = 5$. Individuals' are assigned different types which may or may not be observed by the researcher (corresponding to different colors). Individuals interact based on their types and form links among the possible connections. The possible connections and the realized adjacency matrix remain unobserved.

a policy

$$\pi(\cdot;\beta) : \mathcal{X} \mapsto [0,1], \quad \beta \in \mathcal{B},$$

is a map that prescribes the individual treatment probability based on covariates. Here $\mathcal{B}$ is a compact parameter space, and $\pi(x,\beta)$ is twice differentiable in β. The experiment assigns treatments independently based on $\pi(\cdot)$, and time/cluster-specific parameters $\beta_{k,t}$.

Assumption 2.3 (Treatment assignments in the experiment)**.** For given parameters $\beta_{k,t}$

$$D_{i,t}^{(k)} | X^{(k)}, \beta_{k,t} \sim_{i.i.d.} \text{Bern}\Big(\pi(X_i^{(k)};\beta_{k,t})\Big),$$

which, for short of notation, we refer to as $D_{i,t}^{(k)} | X^{(k)}, \beta_{k,t} \sim \pi(X_i^{(k)}, \beta_{k,t})$.

Assumption 2.3 defines a treatment rule in experiments. Treatments are assigned independently based on covariates and time and cluster-specific parameters $\beta_{k,t}$ (whose choice is discussed in the next sections). The assignment in Assumption 2.3 is easy to implement: it can be implemented in an online fashion and does not require network information, which justifies its choice; also, it generalizes assignments in saturation designs studied for inference on treatment effects (Baird et al., 2018). Our goal is to estimate the welfare-maximizing β (see Remark 2).[20] Our framework extends to continuous treatments, omitted for brevity.[21]

An example of assignment rule is treating individuals with equal probability (Akbarpour et al., 2018), i.e., $\pi(\cdot;\beta) = \beta \in [0,1]$. We can also *target* treatments, i.e., $\pi(x;\beta) = \beta_x$, indicating the treatment probability for $X_i^{(k)} = x$ (with $\mathcal{X}$ discrete).

Throughout our discussion, whenever we write $\pi(\cdot;\beta)$, omitting the subscripts (k,t), we refer to a generic exogenous (i.e., not data dependent) vector of parameters β. We define $\mathbb{E}_\beta[\cdot]$ the expectation taken over the distribution of treatments assigned according to $\pi(\cdot;\beta)$.

[20] In Theorem 4.5 we show that the optimum obtained under Assumption 2.3 is asymptotically equivalent to the one with arbitrary dependent assignments under additional conditions on spillovers and costs.

[21] All our results hold for $D_{i,t}^{(k)} | X_i^{(k)}, \beta_{k,t} = \pi(X_i^{(k)}, \beta_{k,t})$, where $\pi(\cdot;\beta)$ is smooth in β.

Lemma 2.1 (Outcomes)**.** *Under Assumption 2.1, 2.2, under an assignment in Assumption 2.3 with exogenous (i.e., not data-dependent) $\beta_{k,t}$ the following holds:*

$$Y_{i,t}^{(k)} = y\left(X_i^{(k)}, \beta_{k,t}\right) + \varepsilon_{i,t}^{(k)} + \alpha_t + \tau_k, \quad \mathbb{E}_{\beta_{k,t}}\left[\varepsilon_{i,t}^{(k)} | X_i^{(k)}\right] = 0, \tag{2}$$

for some function $y(\cdot)$ unknown to the researcher. In addition, for some unknown $m(\cdot)$, $\mathbb{E}_{\beta_{k,t}}\left[Y_{i,t}^{(k)} | D_{i,t}^{(k)} = d, X_i^{(k)} = x\right] = m(d, x, \beta_{k,t}) + \alpha_t + \tau_k$.

The proof is in Appendix C.2.3. Equation (B.9) states that the outcome depends on two components. The first is the conditional expectation given the individual covariates, and the parameter $\beta_{k,t}$, *unconditional* on covariates, adjacency matrix, individual, and neighbors' assignments. We can interpret the functions $y(\cdot)$ and $m(\cdot)$ as functions which depend on observables only. The dependence with $\beta_{k,t}$ captures spillover effects, since treatments' distribution depends on $\beta_{k,t}$, while we average over neighbors' treatments and covariates. The second component $\varepsilon_{i,t}$ are unobservables that also depend on the neighbors' assignments and covariates. As shown in Appendix C.2.3, under the above conditions, such unobservables only depend on γ_N many others, where $\gamma_N^{1/2}$ is the maximum degree of the network (see Assumption 2.1). Also, note that Lemma 2.1 assumes that $\beta_{k,t}$ is exogenous. We guarantee exogeneity with a careful choice of the experimental design discussed in subsequent sections.

Example 2.1. Let $D_{i,t}^{(k)} \sim_{i.i.d.} \text{Bern}(\beta)$ be exogenous, $\mathcal{N}_i = \{j : A_{i,j} = 1\}, A_{i,j} \in \{0, 1\}$, and

$$Y_{i,t} = \alpha_t + D_{i,t}\phi_1 + \frac{\sum_{j\in\mathcal{N}_i} D_{j,t}^{(k)}}{|\mathcal{N}_i|}\phi_2 - \left(\frac{\sum_{j\in\mathcal{N}_i} D_{j,t}}{|\mathcal{N}_i|}\right)^2\phi_3 + \nu_{i,t}, \quad \mathbb{E}[\nu_{i,t}] = 0. \tag{3}$$

Equation (3) states that outcomes depend on the individual treatment, and the percentage of treated neighbors. With some algebra, taking expectations, $Y_{i,t} = \alpha_t + \beta\phi_1 + \beta\phi_2 - \beta\phi_3\kappa - \beta^2\phi_3(1-\kappa) + \varepsilon_{i,t}$ where $\varepsilon_{i,t}$ also depends on neighbors' assignments and $\kappa = \mathbb{E}[1/|\mathcal{N}_i|]$. □

Lemma 2.1 provides the basis for the definition of (utilitarian) welfare.

Definition 2.1 (Welfare)**.** For treatments as assigned in Assumption 2.3 with exogenous β parameter, let welfare be

$$W(\beta) = \int y(x, \beta) dF_X(x). \tag{4}$$

We define welfare as the expected outcome, had treatments been assigned with policy $\pi(\cdot, \beta)$. The expectation is taken over treatment assignments, covariates, and adjacency matrix. We interpret $y(x, \beta)$ the outcome *net of costs*[22], and incorporate the costs in the

[22]This is standard in the literature (Kitagawa and Tetenov, 2018). However, some applications may not have explicit definitions of costs. For these cases, one possible choice of the cost is the opportunity cost, had the treatment being assigned to a population with no externalities. See Section 4.2 for a discussion.

outcome function.[23] We define the welfare-maximizing policy and the marginal effect[24] as

$$\beta^* \in \arg\sup_{\beta \in \mathcal{B}} W(\beta), \quad V(\beta) = \frac{\partial W(\beta)}{\partial \beta}. \tag{5}$$

The marginal effect defines the derivative of the welfare with respect to the vector of parameters β. A useful insight is that, under mild regularity conditions, we can write

$$V(\beta) = \int \Big[\underbrace{\pi(x;\beta)\frac{\partial m(1,x,\beta)}{\partial \beta} + (1-\pi(x;\beta))\frac{\partial m(0,x,\beta)}{\partial \beta}}_{(S)} + \underbrace{\frac{\partial \pi(x;\beta)}{\partial \beta}(m(1,x,\beta) - m(0,x,\beta))}_{(D)} \Big] dF_X(x). \tag{6}$$

The marginal effect depends on the direct effect, weighted by the probability of treatment (D); the marginal spillover effect, i.e., the marginal effect of increasing neighbors' treatment probabilities. Equation (6) follows in the spirit of the total treatment effects decomposition in direct and indirect effects in Hudgens and Halloran (2008).[25]

Remark 2 (Assumption 2.3 and unconstrained optimum)**.** In Theorem 4.5, we compare $W(\beta^*)$ to the unconstrained optimum W_N^*, where treatments can be assigned arbitrary. We show, under additional conditions, that $W(\beta^*) - W_N^* \to 0$, as $N \to \infty$, whenever the treatment costs are the opportunity costs of an intervention with no spillovers. In a cash-transfer program these are the opportunity costs had the cash transfers be given to individuals spread out on a large area instead of individuals in nearby villages. □

2.3 Method's overview and example

We now give an overview and example. Consider a policy-maker who must allocate cash-transfers to *half* of the population. Let $X_i \in \{0,1\}$, equal to one for households farer from the district's center than the median household and zero otherwise, with $P(X_i = 1) = 1/2$. Due to the constraint, the policy-maker assigns treatments $D_{i,t}|X_i = x \sim \text{Bern}(\pi(x,\beta))$, where $\pi(x,\beta) = x\beta + (1-x)(1-\beta)$ is the treatment probability for $x \in \{0,1\}$.[26] Different treatment probabilities for people in remote areas produce different welfare effects, and assigning all treatments to individuals in remote areas is sub-optimal. This is illustrated in Figure 2, where we report $W(\beta)$, calibrated to data from Alatas et al. (2012, 2016).[27]

[23] Namely, we can parametrize $Y_{i,t} = Y_{i,t} - c(X_{i,t};\beta)$, for a cost function $c(\cdot;\beta)$.

[24] Here, we are assuming differentiability. See Assumption 3.1 for explicit conditions.

[25] We also note that in recent work, Hu et al. (2021) have proposed targeting as causal estimand the average indirect effect, which is different from (S) for heterogenous assignments. Also Graham et al. (2010) present peer effects' decompositions in the different context of peer groups' formation.

[26] This follows from the budget constraint where $\beta_1 P(X_i = 1) + \beta_2 P(X_i = 0) = 1/2$, where $\beta_1 = \beta, \beta_2 = 1 - \beta$ are the treatment probabilities for people in remote and closer areas from the center, respectively.

[27] See Appendix D.3 for details.

2.3.1 Single-wave experiment: hypothesis testing

First, we would like to test whether a certain (baseline) intervention β, such as the one currently implemented by the government or NGO, is welfare-optimal. That is, we test

$$H_0 : W(\beta) = W(\beta^*).$$

Its rejection is informative on whether a (small) change to β improves welfare. We use the marginal effect for:

(a) *Hypothesis testing*: assuming that β^* is an interior point, if $\frac{\partial W(\beta)}{\partial \beta} \neq 0$, then H_0 is false;

(b) *Policy update*: estimate the welfare-improving *direction* (increase or decrease β).

Here, (a) tests whether the line's slope in Figure 2 is zero. Similarly, a rejection of one-sided hypotheses $\frac{\partial W(\beta)}{\partial \beta} \leq 0$ suggests to increase β (treat more people in remote areas).

We proceed to construct estimators of the marginal effect. We start from Equation (6). The direct effect (D) can be identified from a single network, taking the difference between treated and untreated outcomes. However, the spillover effect (S) cannot be identified from a single network when unobserved. We instead exploit *two* clusters' variation.

We take two clusters such as two regions. We collect *baseline* ($t = 0$) outcomes and covariates; we then randomize treatments with slightly different probabilities between the regions. In the first region, we treat individuals in remote areas ($X_i = 1$) with probability $\beta + \eta_n$. Here, η_n is a small deterministic number (local perturbation). The remaining individuals are treated with probability $1-\beta-\eta_n$. In the second region, we treat individuals in remote area with probability $\beta - \eta_n$, and the remaining ones with probability $1 - \beta + \eta_n$.

As shown in Figure 2, we can estimate welfare for two different but similar treatment probabilities; the line's slope between the points is approximately equal to the marginal effect. That is, for a suitable choice of η_n (see Theorem 3.1), a consistent marginal effect's estimator is

$$\widehat{V}_{(k,k+1)}(\beta) = \frac{1}{2\eta_n}\left[\bar{Y}_1^{(k)} - \bar{Y}_0^{(k)}\right] - \frac{1}{2\eta_n}\left[\bar{Y}_1^{(k+1)} - \bar{Y}_0^{(k+1)}\right], \tag{7}$$

where $\bar{Y}_t^{(h)}$ is the outcomes' sample average in cluster h at time t, $Y_{i,0}$ is the baseline outcome with no experiment in place yet, and $(k, k+1)$ index the two clusters. The above estimator is a difference-in-differences; we subtract baseline outcomes due to fixed effects.

We can then leverage the marginal effect for hypothesis testing. We take $K = 2G, G > 1$ clusters and 1) we match clusters in G pairs; 2) within each pair g, we obtain an estimate $\widehat{V}_g(\beta)$ for pair g; 3) we construct a (scale-invariant) test statistics using G pairs, which

does not depend on the estimator's variance.[28] As we discuss in Remark 5, pairing clusters guarantees finite clusters asymptotics. Formal details of the algorithm are in Section 3.

Algorithm 1 illustrates the experiment in the cash-transfer example with two clusters. With more than two clusters, we pair clusters and implement Algorithm 1 in each pair.

Algorithm 1 Local perturbation with *two* clusters, β is a scalar

Require: Value β, $K = 2$ with $h \in \{k, k+1\}$, constant $\bar{C}$.

1: $t = 0$ (baseline): either nobody receives treatments or treatments are assigned with $\pi(\cdot;\beta)$ (either case is allowed).

a: Experimenters collect baseline outcomes: for n units in each cluster observe $Y_{i,0}^{(h)}, X_i^{(h)}, h \in \{k, k+1\}$.

2: $t = 1$: experiment starts

a: Based on the target parameter β, assign treatments for $X_i = 1$ as

$$D_{i,1}^{(k)} | \beta, X_i^{(k)} = x \sim \begin{cases} \text{Bern}(\pi(x, \beta + \eta_n)) \text{ if } h = k \\ \text{Bern}(\pi(x, \beta - \eta_n)) \text{ if } h = k+1 \end{cases}, \quad \bar{C} n^{-1/2} < \eta_n < \bar{C} n^{-1/4}.$$

b: For n units in each cluster $h \in \{k, k+1\}$ observe $Y_{i,1}^{(h)}$.

3: Estimate the marginal effect as in Equation (7).

Remark 3 (Inference on treatment effects)**.** Testing H_0 is different from testing hypotheses on treatment effects since these hypotheses do not depend on β^*. However, our design permits us to estimate separately the direct treatment effect and the spillover effect, which may be of independent interest. We estimate the direct effect with a means' difference between treated and controls, *pooled* across clusters. In Theorem 3.3, we show that the estimated direct effect has a negligible bias of order $o(n^{-1/2})$. We estimate the spillover effect by taking the outcomes' difference between clusters (see Section 3.1).[29] □

2.3.2 Multiple-wave experiment: estimation of β^*

We now show how we can estimate β^* using T experimentation periods and $K \geq 2(T+1)$ many clusters. The policy-maker assigns cash transfers and collects outcomes sequentially. Every period t, she assigns treatments in cluster k based on parameters $\beta_{k,t}$, with two goals. First, at time T, she wants to obtain an estimate $\hat{\beta}^*$ which well approximates β^*. Second, she wants to improve the experimental participants' welfare.

[28] See Equation (10). Scale-invariance follows from Ibragimov and Müller (2010).

[29] Also, if the experimenter's goal is precise inference on direct effects only, β can be chosen based on variance considerations, but we can employ our method – which only requires small perturbations to β – to identify spillover effects.

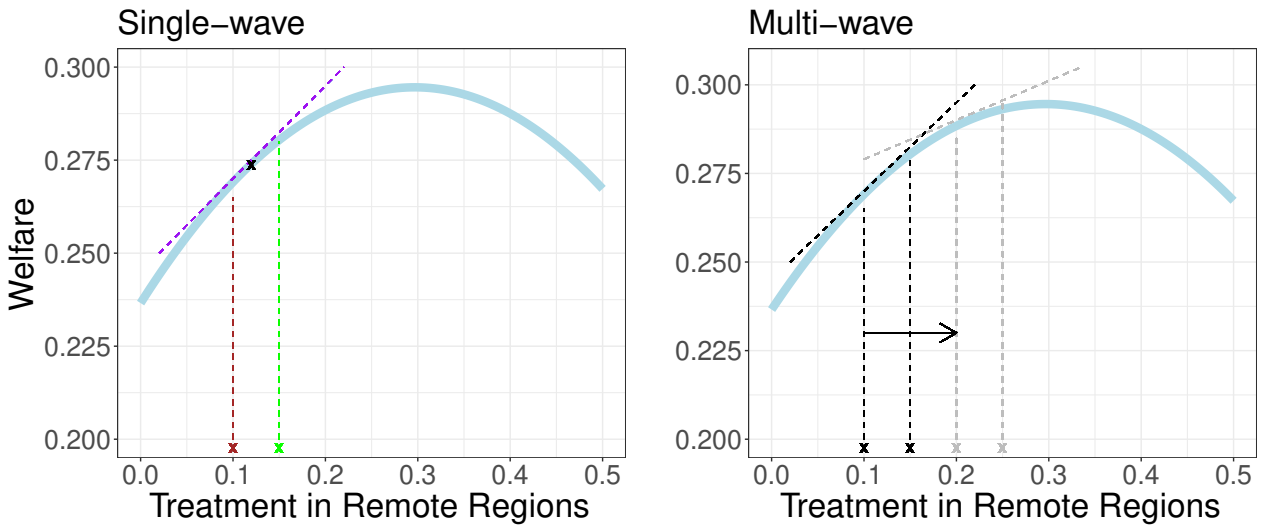


Figure 2: Example of experimental design. The objective function is satisfaction with a cash-transfer program. Half of the population is treated. The left panel is a single-wave experiment with two clusters. In the first cluster, we assign the policy colored in green and the second cluster colored in brown. The right panel is a two-wave experiment. In the first period, we use a pair of clusters to estimate the marginal effect (black color), and we update the policy for a different pair (gray color).

We maximize welfare with the following algorithm: 1) we pair clusters and organize pairs in a circle as in Figure 3; 2) every step t, we estimate the marginal effect within each pair; 3) we then update the policy in a given clusters' pair using the information on the marginal effect from the subsequent pair on the circle. We refer to Step 3 as circular cross-fitting.[30] Step 3 guarantees that the experiment is unconfounded. See Section 4 for details.[31]

The right panel in Figure 2 illustrates the procedure for two waves: we use one pair of clusters to estimate the marginal effect, which we then use for the second pair (and vice-versa, see Algorithm 3). The experiment assumes and leverages the concavity of welfare, generally attained under decreasing marginal effects of neighbors' treatments. Under lack of concavity, the experiment returns a *local* optimum. We relax concavity in Appendix B.2.

In Section 4 we measure the method's performance based on the in-sample and out-of-sample regret, and show that with high-probability

$$W(\beta^*) - W(\hat{\beta}) = \mathcal{O}(1/T), \quad \max_k \frac{1}{T}\sum_{t=1}^{T}\Big[W(\beta^*) - W(\beta_{k,t})\Big] = \mathcal{O}(\log(T)/T).$$

where the first equation indicates the out-of-sample regret, and the second equation indicates

[30] Cross-fitting algorithms were used by Chernozhukov et al. (2018) in the different context of double-machine learning with *i.i.d.* data. Our procedure differs due to the adaptive sampling and the circular structure.

[31] Also, note that our method also extends when treatments can only be assigned once. See Appendix B.5.

the in-sample regret. The regret scales at rate $O(1/K)$, as we choose $2T + 2 = K$. Under additional restrictions, we also derive exponential rates $O(\exp(-c_0 K))$ for the out-of-sample regret, for a constant $c_0 > 0$. The above equations provide strong guarantees on the method's performance as we deploy the estimated policy on a new population (out-of-sample regret) and the experimental participants (in-sample regret).

Remark 4. An alternative approach is to estimate $y(\cdot)$ assigning different policies to clusters and extrapolating the overall effect. We do not consider this alternative for two reasons. For a generic p-dimensional vector β, the out-of-sample regret is either sensitive to the model used for extrapolation or suffers a curse of dimensionality (e.g., when a grid search is employed). Second, this method does not control the *in-sample* regret, i.e., it must incur significant in-sample welfare loss to estimate the response function $y(\cdot)$. Appendix A.4 presents details. □

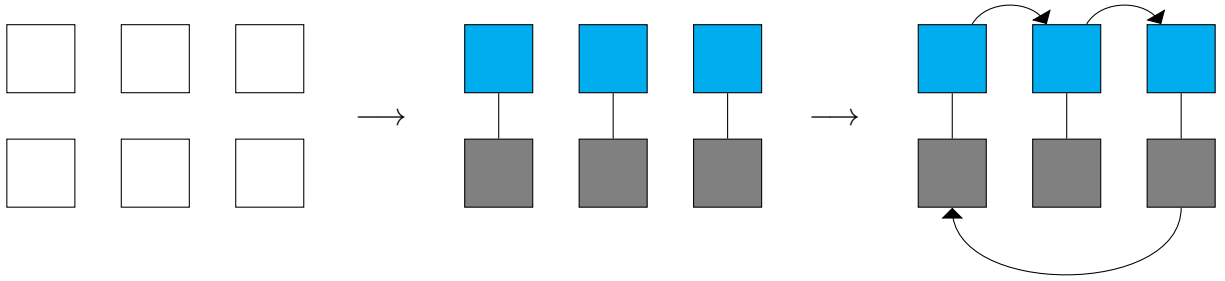


Figure 3: Circular cross-fitting method. Clusters (rectangles) are paired. Within each pair, researchers assign different treatment probabilities to clusters with different colors. Finally, the policy in each pair is updated using information from the consecutive pair.

3 Single-wave experiment

In this section, we turn to the design and analysis of a single-wave experiment.

We consider an experiment to test the following hypothesis.

Definition 3.1 (Testable implication)**.** Let $\beta^* \in \mathcal{B}$ be an interior point. If $W(\beta) = W(\beta^*)$, then

$$H_0 : V^{(j)}(\beta) = 0, \quad \forall j \in \{1, \cdots, l\}, l \leq p. \tag{8}$$

The above implication is at the core of our approach. We can test whether l arbitrary entries of the marginal effect are equal to zero. Rejection implies a lack of global optimality. For expositional convenience, we consider $l = 1$ only (test the first entry being zero). In Appendix B.1 we show how our method generalizes to $l > 1$. We may also test $V^{(j)}(\beta) \leq 0$; for example, for $\pi(x, \beta) = \beta_x$ ($\mathcal{X}$ discrete), one-sided test is informative for whether we

should increase treatment probabilities for individuals with $x = j$ (without assuming that β^* is in the interior). Finally, it is useful to define the vector

$$\underline{e}_j = \left[0, \cdots, 0, 1, 0, \cdots, 0\right], \text{ where } \underline{e}_j \in \{0,1\}^p, \text{ and } \underline{e}_j^{(j)} = 1. \tag{9}$$

Algorithm 2 presents the design. The algorithm pairs clusters. Within each pair, it estimates the first entry of the marginal effect (since here we test $V^{(1)}(\beta) = 0$) using local perturbations – as discussed in Section 2.3. It then constructs a scale-invariant test statistics. Without loss of generality, we index clusters such that each pair contains two consecutive clusters $\{k, k+1\}$ with k being an odd number.

In the following lines, we discuss estimation of marginal and treatment effects and guarantees for inference on H_0.

Algorithm 2 One wave experiment for inference with $l = 1$

Require: Value $\beta \in \mathbb{R}^p$ (exogenous), K clusters, constant $\bar{C}$, size α;

1: Organize clusters into $G = K/2$ pairs with consecutive indexes $\{k, k+1\}$;

2: For each pair $g = \{k, k+1\}$, k is odd, run Algorithm 1, with at $t = 1$,

$$D_{i,1}^{(k)} | \beta, X_i^{(k)} = x \sim \begin{cases} \text{Bern}(\pi(x, \beta + \eta_n \underline{e}_1)) \text{ if } h = k \\ \text{Bern}(\pi(x, \beta - \eta_n \underline{e}_1)) \text{ if } h = k+1 \end{cases}, \quad \bar{C} n^{-1/2} < \eta_n < \bar{C} n^{-1/4},$$

and estimate the marginal effect as in Equation (6).

3: Construct the test statistics

$$\mathcal{T}_n = \frac{\sqrt{G} \bar{V}_n(\beta)}{\sqrt{(G-1)^{-1} \sum_g (\widehat{V}_g(\beta) - \bar{V}_n(\beta))^2}}, \quad \bar{V}_n(\beta) = \frac{1}{G} \sum_g \widehat{V}_g(\beta); \tag{10}$$

here $\widehat{V}_g$ is the marginal effect estimated in pair g.

4: Construct the test as $1\left\{|\mathcal{T}_n| > \text{cv}_{G-1}(\alpha)\right\}$ with size α; $\text{cv}_{G-1}(\alpha)$ is the size α t-test's critical quantile with $G-1$ degrees of freedom.

3.1 Estimation of marginal and treatment effects

The experiment we just described permits us to estimate three quantities of independent interest: the marginal effect, the direct effect, and the spillover effect. These should be reported by researchers once the experiment is concluded. We describe the estimators below.

Equation (6) provides the marginal effect estimator $\widehat{V}_g(\beta)$ for each pair of clusters g. Researchers may report $\bar{V}_n(\beta)$ (Equation 10) in their results – the average across clusters'

pairs. We show below that both $\bar{V}_n(\beta)$ and $\widehat{V}_g(\beta)$ provide a consistent estimate of $V^{(1)}(\beta)$.[32]

The experiment also allows us to estimate the direct effect of the treatment and the (marginal) spillover effect separately, respectively defined as:[33]

$$\Delta(\beta) = \int \Big[m(1,x,\beta) - m(0,x,\beta)\Big] dF_X(x), \quad S_1(d,\beta) = \int \frac{\partial m(d,x,\beta)}{\partial \beta^{(1)}} dF_X(x).$$

The direct effect is the treatment effect, keeping fixed the neighbors' treatment probability. $S_1(\cdot)$, the spillover effect, is the marginal effect of a small change in the first entry of β, keeping fixed individual treatment status.[34] For a given pair of clusters $(k, k+1)$, we estimate

$$\widehat{\Delta}_k(\beta) = \frac{1}{2n} \sum_{h\in\{k,k+1\}} \sum_{i=1}^{n} \Big[\frac{D_{i,1}^{(h)} Y_{i,1}^{(h)}}{\pi(X_i^{(h)}, \beta + \eta_n v_h \underline{e}_1)} - \frac{(1 - D_{i,1}^{(h)}) Y_{i,1}^{(h)}}{1 - \pi(X_i^{(h)}, \beta + \eta_n v_h \underline{e}_1)}\Big], \quad v_h = \begin{cases} 1 \text{ if } h = k \\ -1 \text{ if } h = k+1. \end{cases} \tag{11}$$

The estimator pools observations between the two clusters and takes a difference between treated and control units within each cluster, divided by the probability of treatments. This follows similarly to classical Horvitz-Thompson estimators (Horvitz and Thompson, 1952). Importantly, we divide by the probability of treatments, taking into account the perturbation η_n. We average direct effects across clusters' pairs to obtain a single measure $\bar{\Delta}_n = \frac{1}{G}\sum_g \widehat{\Delta}_g(\beta)$. The indirect effect is estimated as follows:

$$\widehat{S}_{(k,k+1)}(0,\beta) = \frac{1}{2n} \sum_{h\in\{k,k+1\}} \frac{v_h}{\eta_n} \sum_{i=1}^{n} \Big[\frac{Y_{i,1}^{(h)}(1 - D_{i,1}^{(h)})}{1 - \pi(X_i^{(h)}, \beta + v_h \eta_n \underline{e}_1)} - \bar{Y}_0^{(h)}\Big].$$

The estimator takes a difference between the control units between the two clusters, divided by their corresponding treatment probabilities. Researchers may report the between pairs average $\bar{S}_n(0,\beta) = \frac{1}{G}\sum_g \widehat{S}_g(0,\beta)$,[35] which captures spillovers on the control units. Note that in practice, practitioners may also want to report interaction effects between spillovers and individual types, which we discuss in Appendix B.6.

3.2 Theoretical Analysis

Next, we study theoretical guarantees. The following regularity condition is imposed.

[32] Note that our discussion directly extends to estimating each entry of $V(\beta)$ as shown in Appendix B.1.

[33] Here, we are implicitly assuming that $m(\cdot)$ is uniformly bounded to invoke the dominated convergence theorem. See the next subsection for formal assumptions.

[34] Similarly to the marginal effect, our setting also extends to estimating $S_j(\cdot)$ for arbitrary entries of β as in Appendix B.1.

[35] Here, $\hat{S}(1,\beta)$ follows similarly and omitted for brevity.

Assumption 3.1 (Regularity 1)**.** Suppose that for all $x \in \mathcal{X}, d \in \{0,1\}$, $\pi(x,\beta), m(d,x,\beta)$ are uniformly bounded and twice differentiable with bounded derivatives.

Assumption 3.1 imposes smoothness and boundedness restrictions.[36]

Theorem 3.1 (Consistency of the marginal effect)**.** *Suppose that $\varepsilon_{i,t}^{(k)}$ is sub-gaussian. Let Assumptions 2.1, 2.2, 3.1 hold. Let $\text{Var}(\sqrt{n}\hat{V}_{(k,k+1)}(\beta)) = \mathcal{O}(\rho_n)$. Then with probability at least $1-\delta$, for any $\delta \in (0,1)$,*

$$\left|\hat{V}_{(k,k+1)}(\beta) - V^{(1)}(\beta)\right| = \mathcal{O}\Big(\eta_n + \min\Big\{\sqrt{\frac{\gamma_N \log(\gamma_N/\delta)}{n\eta_n^2}}, \sqrt{\frac{\rho_n}{n\eta_n^2\delta}}\Big\}\Big),$$

where $\hat{V}_{(k,k+1)}$ is estimated as in Algorithm 2.

For $\gamma_N \log(\gamma_N)/N^{1/3} = o(1), \eta_n = n^{-1/3}, \hat{V}_{(k,k+1)}(\beta) \to_p V(\beta)$.

The proof is in Appendix C.3.1. Theorem 3.1 shows that we can consistently estimate the marginal effects with two large clusters. Consistency depends on the degree of dependence among unobservables $\varepsilon_{i,t}^{(k)}$ (which also depends on neighbors' assignments). The convergence rate depends on the *minimum* between the maximum degree of the network, which is proportional to $\gamma_N^{1/2}$, and the covariances among unobservables, captured by ρ_n. If either the network has a degree that grows at a slower rate than N (recall that $n/N = O(1)$) or a degree equal to N but vanishing covariances, we can consistently estimate the marginal effects. The theorem also illustrates the trade-off in the choice of the deviation parameter η_n: a larger deviation parameter η_n decreases the variance, but it increases the bias. Appendix B.7 provides an explicit rule of thumb for choosing η_n.

Corollary 1. *Under the conditions in Theorem 3.1, letting $\gamma_N \log(\gamma_N)/n^{1/3} = o(1), \eta_n = n^{-1/3}$, for any K, $\bar{V}_n \to_p V^{(1)}(\beta)$.*

The above corollary illustrates consistency once we pool information from different clusters (with K being finite). Next, we study inference assuming the following condition.

Assumption 3.2 (Regularity 2)**.** Assume that for treatments as assigned in Algorithm 2, for all $k \in \{1, \cdots, K\}$, $\varepsilon_{i,t}^{(k)}$ has bounded fourth moment and for some $\bar{C}_k > 0$, $\rho_n \geq 1$,

$$\text{Var}\Big(\frac{1}{\sqrt{n}}\Big[\bar{Y}_1^{(k)} - \bar{Y}_0^{(k)}\Big]\Big) = \bar{C}_k \rho_n. \tag{12}$$

Assumption 3.2 imposes standard moment bounds and a lower bound on the variance of the estimator, attained under independence and positive dependence.

[36] These restrictions hold for a large set of linear and non-linear functions, assuming that $\mathcal{X}$ is compact. Boundedness is often imposed in the literature, see, e.g., Kitagawa and Tetenov (2018).

Theorem 3.2. *Let Assumptions 2.1, 2.2, 3.1, 3.2 hold. Let* $n^{1/4}\eta_n = o(1), \gamma_N/N^{1/4} = o(1)$, $K < \infty$. *Then for each pair* $(k, k+1)$, *for* $\widehat{V}_{(k,k+1)}$ *estimated as in Algorithm 2,*

$$\mathrm{Var}\Big(\widehat{V}_{(k,k+1)}\Big)^{-1/2}\Big(\widehat{V}_{(k,k+1)} - V^{(1)}(\beta)\Big) \to_d \mathcal{N}(0,1).$$

The proof is in Appendix C.3.2. Theorem 3.2 guarantees asymptotic normality. The theorem assumes that the maximum degree $\gamma_N^{1/2}$ grows at a slower rate than the sample size of order $N^{1/8}$ (and hence $n^{1/8}$ since n is proportional to N). This condition is stronger than what is required for consistency only.[37] Given Theorem 3.2, we conduct inference with scale-invariant test statistics without necessitating estimation of the (unknown) variance.

Corollary 2. *Let the conditions in Theorem 3.2 hold. For* $4 \leq K < \infty$, $\alpha \leq 0.08$,

$$\lim_{n\to\infty} P\Big(|\mathcal{T}_n| \leq \mathrm{cv}_{K/2-1}(\alpha)\Big|H_0\Big) \geq 1-\alpha, \tag{13}$$

where $\mathrm{cv}_{K/2-1}(h)$ *is the size-h critical value of a t-test with* $K/2-1$ *degrees of freedom.*

The proof is in Appendix C.5. The theorem guarantees asymptotically valid inference on H_0 as $n \to \infty$ and K is finite. With $l = 1$ the proof is a direct consequence of Theorem 3.2, combined with properties of pivotal statistics in Ibragimov and Müller (2010).[38] In Appendix B.1 we provide expressions for the test statistics and derivations for $l > 1$.

To our knowledge, this is the first set of results for inference on welfare-maximizing policies with unknown interference.

We conclude this section with a study on the estimated direct and spillover effect.

Theorem 3.3 (Asymptotically neglegible bias of the direct effect)**.** *Let Assumptions 2.1, 2.2, 3.1 hold, and* $\eta_n = o(n^{-1/4})$. *Then for all pairs* $(k, k+1)$, $\mathbb{E}\Big[\widehat{\Delta}_{(k,k+1)}(\beta)\Big] = \Delta(\beta) + o(n^{-1/2})$. *Similarly,* $\mathbb{E}\Big[\bar{\Delta}_n(\beta)\Big] = \Delta(\beta) + o(n^{-1/2})$, *where the second term does not depend on* K.

The proof is in Appendix C.3.3. Theorem 3.3 shows that the bias of the estimated direct effect is asymptotically negligible at a rate faster than the parametric rate $n^{-1/2}$ when *pooling* observations from different clusters. Our insight here is that, with pairing and perturbations of opposite signs, the first-order bias cancels out. This result implies that our experimental design induces a bias that can be ignored for estimation and inference.[39] Given that the bias

[37] We conjecture that weaker restrictions on the degree are possible, as for consistency. We leave their study to future research.

[38] See also Chernozhukov et al. (2018) for a discussion on pivotal inference in the different context of synthetic controls.

[39] Note that $\eta_n = o(n^{-1/4})$ is consistent with requirements in previous theorems.

is asymptotically negligible, we can use existing results for inference on direct effects with a single network (e.g., Sävje et al. 2021). For completeness, we show consistency below.

Corollary 3. *Suppose that $\varepsilon_{i,t}^{(k)}$ is sub-gaussian. Let Assumptions 2.1, 2.2, 3.1 hold, and $\pi(x,\beta) \in (\kappa, 1-\kappa), \kappa \in (0,1)$ for all $x \in \mathcal{X}$. Let $\eta_n = o(n^{-1/4})$. Then with probability at least $1-\delta$, for any $\delta \in (0,1)$, for any K*

$$\left|\bar{\Delta}_n - \Delta(\beta)\right| = \mathcal{O}\Big(\sqrt{\frac{\gamma_N \log(\gamma_N/\delta)}{Kn}}\Big) + o(n^{-1/2}),$$

The proof is in Appendix C.5. The corollary requires strict overlap (standard in the literature on causal inference) and shows that we can attain consistency for $K < \infty$, $n \to \infty$.

The following result is on the bias of the marginal spillover effects estimators.

Theorem 3.4 (Marginal Spillover effects)**.** *Let Assumptions 2.1, 2.2, 3.1 hold. Then for all pairs $(k, k+1)$, $\mathbb{E}\Big[\widehat{S}_{(k,k+1)}(0,\beta)\Big] = S_1(0,\beta) + \mathcal{O}(\eta_n)$.*

The proof is in Appendix C.3.4. Theorem 3.4 shows that the bias converges to zero as $\eta_n \to 0$. The rate is slower than the rate of the direct effect's bias (since $\eta_n > n^{-1/2}$). We obtain a slower rate because the marginal spillover effect depends on *between* clusters variations. Consistency and inference follow verbatim as discussed for the marginal effect.

Remark 5 (Pairing clusters permits finite-clusters asymptotics)**.** Pairing plays a fundamental role in our design with finite K. In the absence of pairing, the bias of the marginal and treatment effects would not converge to zero for finite K. To gain further intuition, consider a uni-dimensional setting ($X = 1$ almost surely), and $\pi(1,\beta) = \beta$. Then

$$\begin{aligned}\mathbb{E}\Big[\widehat{V}_{(k,k+1)}(\beta)\Big] &= \frac{1}{2}\sum_{h\in\{k,k+1\}} \frac{v_h}{\eta_n} y(1;\beta + v_h\eta_n) \\ &= \underbrace{\frac{1}{2}\sum_{h\in\{k,k+1\}} \frac{v_h}{\eta_n} y(1;\beta)}_{(i)} + \underbrace{\frac{\partial y(1;\beta)}{\partial \beta}}_{(ii)} + \mathcal{O}(\eta_n), \quad v_h = \begin{cases} 1 \text{ if } h = k \\ -1 \text{ if } h = k+1. \end{cases}\end{aligned}$$

Component (i) induces a bias, while (ii) is the target estimand. Observe that (i) equals zero because of the paired design: v_h is one for one cluster and minus one for the other cluster. Suppose instead that a paired design was not implemented, and instead we have $v_h \in \{\pm 1\}$ with equal probability.[40] Then (i) would scale to zero at a slow rate $1/\sqrt{K\eta_n^2}$ after averaging across all the clusters, requiring infinite clusters asymptotics. □

[40] Random probabilities assignments are common when estimating treatment effects with saturation design (Baird et al., 2018). Pairing is common in applications in the different context of estimating overall average treatment effects with (different-sized) cluster experiments (Imai et al., 2009).

4 Multi-wave experiment

In this section, we design the adaptive experiment and derive its theoretical properties.

For illustrative purposes, we provide the algorithm for the one-dimensional case $p = 1$, in Algorithm 3, that is when $\beta \in \mathcal{B} = [\mathcal{B}_1, \mathcal{B}_2]$ is a scalar. In Remark 6 and formally in Appendix E we provide the complete algorithm for the p-dimensional case. Theoretical results are for the general p-dimensional case (p is finite). A description is below the algorithmic box.

Algorithm 3 Multiple-wave experiment with β scalar

Require: Starting value β_0, K clusters, $T+1$ periods, constant $\bar{C}$.

1: Create pairs of clusters $\{k, k+1\}, k \in \{1, 3, \cdots, K-1\}$;

2: $t = 0$ (initialization):

a: Assign treatments as $D_{i,0}^{(h)} | X_i^{(h)} = x \sim \text{Bern}(\pi(x, \beta_0))$ for all $h \in \{1, \cdots, K\}$.

b: For n units in each cluster observe $Y_{i,0}^{(h)}, h \in \{1, \cdots, K\}$; initalize $\widehat{V}_{k,t} = 0$, $\check{\beta}_k^0 = \beta_0$.

3: **while** $1 \le t \le T$ **do**

a: Define

$$\check{\beta}_h^t = \begin{cases} P_{\mathcal{B}_1, \mathcal{B}_2 - \eta_n}\left[\check{\beta}_h^{t-1} + \alpha_{h+2,t}\widehat{V}_{h+2,t-1}\right], & h \in \{1, \cdots, K-2\}, \\ P_{\mathcal{B}_1, \mathcal{B}_2 - \eta_n}\left[\check{\beta}_h^{t-1} + \alpha_{1,t}\widehat{V}_{1,t-1}\right], & h \in \{K-1, K\}; \end{cases}$$

where $\alpha_{k,t}$ is the learning rate (see Remark 7), and $P_{\mathcal{B}_1, \mathcal{B}_2 - \eta_n}$ is the projection operator.

b: Assign treatments as (for $\bar{C} n^{-1/2} < \eta_n < \bar{C} n^{-1/4}$)

$$D_{i,t}^{(h)} | X_i^{(h)} = x \sim \text{Bern}(\pi(x, \beta_{h,t})), \quad \beta_{h,t} = \begin{cases} \check{\beta}_h^t + \eta_n \text{ if } h \text{ is odd} \\ \check{\beta}_h^t - \eta_n \text{ if } h \text{ is even} \end{cases} \tag{14}$$

c: For n units in each cluster $h \in \{1, \cdots, K\}$ observe $Y_{i,t}^{(h)}$;

d: For each pair $\{k, k+1\}$, estimate

$$\hat{V}_{k,t} = \hat{V}_{k+1,t} = \frac{1}{2\eta_n}\left[\bar{Y}_t^{(k)} - \bar{Y}_0^{(k)}\right] - \frac{1}{2\eta_n}\left[\bar{Y}_t^{(k+1)} - \bar{Y}_0^{(k+1)}\right].$$

4: **end while**

5: Return $\hat{\beta}^* = \frac{1}{K}\sum_{k=1}^K \check{\beta}_k^T$

The algorithm pairs clusters, and initializes clusters at the same starting value β_0, $\check{\beta}_1^1 = \cdots = \check{\beta}_K^1 = \beta_0$. At $t = 0$, it randomizes treatments independently as

$$t = 0: \quad D_{i,t}^{(k)} | X_i^{(k)} = x \sim \pi(x; \beta_0), \quad \text{for all } (i, k).$$

Here, β_0 is chosen exogenously, e.g., it is the current policy in place. Over each iteration t, we assign treatments based on $\beta_{k,t}$ for cluster k at time t which equals the parameter

$\check{\beta}_k^t$ obtained from a previous iteration plus a positive (negative) perturbation η_n in the first (second) cluster in a pair. The local perturbation follows similarly to what discussed in the previous section; also, by construction, $\check{\beta}_k^t$ is the same for a given pair $(k, k+1)$, where k is odd. We choose $\check{\beta}_k^{t+1}$ via *circular cross-fitting*: we wrap clusters in a circle and update the parameter in a pair of clusters $(k, k+1)$ using information from the subsequent pair (see Figure 3). The algorithm runs over T periods and returns $\hat{\beta}^* = \frac{1}{K}\sum_{k=1}^{K} \check{\beta}_k^{T+1}$.

In our experiment, we update the policy in each clusters' pair with information from a subsequent pair. This approach guarantees that the estimated policy used to randomize treatments in cluster k does not depend on observables and unobservables in that same cluster, assuming that the number of clusters is twice as large as the number of iterations.

Lemma 4.1 (Unconfoundedness)**.** *Let $T/p+1 \leq K/2$. Consider the experimental design in Algorithm E.1 for generic p-dimensions (and Algorithm 3 for $p=1$). Then for any k,*

$$\left(\beta_{k,1}, \cdots, \beta_{k,T}\right) \perp \left\{Y_{i,t}^{(k)}(\mathbf{d}), X_i^{(k)}, \mathbf{d} \in \{0,1\}^N\right\}_{i\in\{1,\cdots,N\}, t\leq T}.$$

The proof is in Appendix C.2.5. Lemma 4.1 shows that the parameters used in the experiment are independent of potential outcomes and covariates in the same cluster. Namely, the circular cross-fitting breaks the dependence due to repeated sampling, which would otherwise confound the experiment.[41] See Remark 8 for a discussion.

Remark 6 (p-dimensional case: Algorithm E.1)**.** The algorithm for the p-dimensional case follows similarly to the uni-dimensional case with a minor change: we consider T/p many *waves*/iterations, each consisting of p periods. Within each wave w, every period, we perturb a single coordinate of $\check{\beta}_k^w$, compute the marginal effect for that coordinate, and repeat over all coordinates $j \in \{1, \cdots p\}$ before making the next policy update to select $\check{\beta}_k^{w+1}$. □

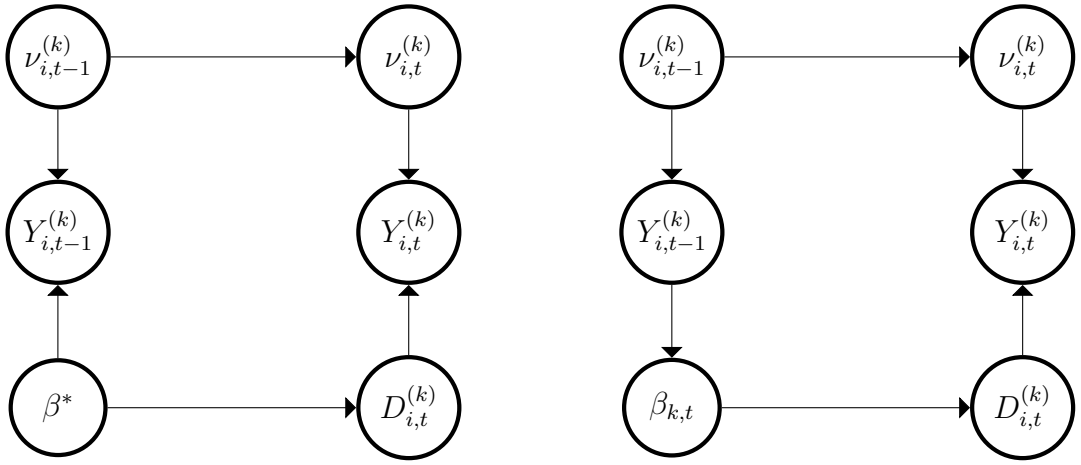
Remark 7 (Learning rate)**.** We are now left to discuss how "large" should be the step size, i.e., if we know that the marginal effect is positive, by how much should we increase the treatment probability? Assuming strong concavity of the objective function, the learning rate $\alpha_{k,t}$ should be of order $1/t$. A more robust choice (see Theorem B.2) is

$$\alpha_{k,t} = \begin{cases} \frac{J}{T^{1/2-v/2}||\hat{V}_{k,t}||} \text{ if } ||\hat{V}_{k,t}||_2^2 > \frac{c}{T^{1-v}} - \epsilon_n, \\ 0 \text{ otherwise} \end{cases}, \tag{15}$$

[41] This setting is different from previous literature on adaptive experimentation, which focuses on settings where repeated sampling does not occur. See, for example, Kasy and Sautmann (2019); Wager and Xu (2021).

for a positive ϵ_n, $\epsilon_n \to 0$, and small constants $1 \geq v, c > 0$.[42] Here, the learning rate divides the estimated marginal effect by its norm (known as gradient norm rescaling, Hazan et al. 2015), and guarantees control of the out-of-sample regret under strict quasi-concavity. This choice is appealing since it guarantees comparable step sizes between different clusters. □

Remark 8 (Why circular cross-fitting? Bias with repeated sampling)**.** Here, we illustrate the source of bias if the circular cross fitting was not employed. Every period, the researcher can only identify the expected outcome of $Y_{i,t}^{(k)}$ conditional on the parameter $\beta_{k,t}$, namely $\widetilde{W}(\beta_{k,t}) = \mathbb{E}_{\beta_{k,t}}[Y_{i,t}^{(k)}|\beta_{k,t}]$. If $\beta_{k,t}$ was chosen exogenously, based on information from a different cluster, then $\mathbb{E}_{\beta_{k,t}}[Y_{i,t}^{(k)}|\beta_{k,t}] = \mathbb{E}_{\beta_{k,t}}[Y_{i,t}^{(k)}] = W(\beta_{k,t})$, where $W(\beta_{k,t})$ defines the expected welfare once we deploy the policy $\beta_{k,t}$ on a new population. However, this is not the case if $\beta_{k,t}$ is estimated using information on $Y_{i,t-1}^{(k)}$. Consider the example where the outcome depends on some auto-correlated unobservables $\nu_{i,t}$ and treatment assignments in Figure 4. The *dependence* structure of Figure 4 implies that: $W(\beta_{k,t}) = \mathbb{E}_{\beta_{k,t}}[Y_{i,t}^{(k)}] \neq \mathbb{E}_{\beta_{k,t}}[Y_{i,t}^{(k)}|\beta_{k,t}] = \widetilde{W}(\beta_{k,t})$, if $\beta_{k,t}$ depends on covariates and unobservables previous outcomes (and so on unobservables $\nu_{i,t}^{(k)}$) in cluster k. Here $W(\beta_{k,t})$ captures the estimand of interest. Instead, $\widetilde{W}(\beta_{k,t})$ denotes what we can identify. Our algorithm breaks such dependence and guarantees unconfounded experimentation as shown in Lemma 4.1. □



Policy on a new population Experiment with repeated sampling

Figure 4: The left-panel shows the dependence structure when a static policy is implemented on a new population (we omit $D_{i,t-1}^{(k)}$ for expositional convenience). The right panel shows the dependence structure of a sequential experiment that uses the same units for policy updates over subsequent periods in the presence of *repeated* sampling.

[42] Formally, we let $\epsilon_n \propto \sqrt{\frac{\gamma_N}{\eta_n^2 n}} + \eta_n$. See Lemma C.12 in the Appendix for further details.

4.1 Theoretical guarantees

Next, we derive theoretical properties. Let $\check{T} = T/p$. We assume the following.

Assumption 4.1. Let : (A) $\varepsilon_{i,t}^{(k)}$ be sub-gaussian; (B) $K \geq 2(T/p+1)$.

Condition (A) states that unobservables have sub-gaussian tails (attained, for example, by bounded random variables); (B) assumes that we have at least twice as many clusters as the number of waves, which guarantees that Lemma 4.1 (unconfoundedness) holds.

In the following results, we impose the following restriction.

Assumption 4.2 (Strong concavity)**.** Assume that $W(\beta)$ is σ-strongly concave over $\mathcal{B}$, for some arbitrary $\sigma > 0$.

Strong concavity is a common feature of objective functions (Bottou et al., 2018). A simple example of strong concavity is Example 2.1, when parameters are calibrated to real-world data as shown in Figure 5.[43] Strong concavity guarantees uniqueness of the optimum. We relax Assumption 4.2 in Appendix B.2.

Theorem 4.2. *Let Assumptions 2.1, 2.2, 3.1, 4.1, 4.2 hold. Take a small* $1/4 > \xi > 0$, $\alpha_{k,w} = J/w$ *for a finite* $J \geq 1/\sigma$. *Let* $n^{1/4-\xi} \geq C\sqrt{p\log(n)\gamma_N T^{B_p}\log(KT)}$, $\eta_n = 1/n^{1/4+\xi}$, *for finite constants* $B_p, C > 0$. *Then with probability at least* $1-1/n$, *for a constant* $\bar{C}' < \infty$, *independent of* (p, n, N, K, T),

$$||\beta^* - \hat{\beta}^*||^2 \leq \frac{p\bar{C}'}{\check{T}}.$$

The proof is in Appendix C.3.5. Theorem 4.2 provides a *small sample* upper bound on the distance between the estimated policy and the optimal one. The bound only depends on T (and not n) since n is assumed to be sufficiently larger than T.

Corollary 4. *Let the conditions in Theorem 4.2 hold. Let* $K = 2(T/p+1)$. *Then with probability at least* $1-1/n$, *for a constant* $C' < \infty$ *independent of* (p, n, N, K, T),

$$W(\beta^*) - W(\hat{\beta}^*) \leq \frac{pC'}{K}.$$

The proof is in Appendix C.5. The above corollary formalizes the out-of-sample regret bound as we choose $K = 2(T/p+1)$. Also, the bound scales polynomially in p, for n sufficiently large.[44]

Researchers may wonder whether the procedure is "harmless" also on the in-sample units. We provide in-sample guarantees in the following theorem.

[43] See Section 5 for details.

[44] This is different from grid-search procedures, where the rate in K would decay (exponentially) in p.

Theorem 4.3 (In-sample regret)**.** *Let the conditions in Theorem 4.2 hold. Then with probability at least* $1-1/n$*, for a constant* $c<\infty$ *independent of* (p,n,N,K,T)*,*

$$\max_{k\in\{1,\cdots,K\}}\frac{1}{\check{T}}\sum_{w=1}^{\check{T}}\left[W(\beta^*)-W(\check{\beta}_k^w)\right]\leq c\frac{p\log(\check{T})}{\check{T}}.$$

The proof is in Appendix C.3.6. Theorem 4.3 guarantees that the cumulative welfare in *each* cluster k, incurred by deploying the current policy $\check{\beta}_k^w$ at wave w (recall that in the general p-dimensional case we have $\check{T}$ many waves), converges to the largest achievable welfare at a rate $\log(T)/T$, also for those units participating in the experiment.[45] This result guarantees that the proposed design is not *harmful* to experimental participants.

We conclude this sub-section deriving a faster (exponential) convergence rate of the out-of-sample regret (but not in-sample regret) with a different choice of the learning rate.

Theorem 4.4 (Out-of-sample regret with larger sample size)**.** *Let Assumptions 2.1, 2.2, 3.1, 4.1, 4.2 hold, with* $W(\beta)$ *being* τ*-smooth, and* $K=2T+2$*. Take a small* $1/4>\xi>0$*,* $\alpha_{k,w}=1/\tau$*. Let* $n^{1/4-\xi}\geq C\sqrt{p\log(n)\gamma_N e^{TB_p}\log(KT)}$*,* $\eta_n=1/n^{1/4+\xi}$*, for finite constants* $B_p,C>0$*. Then with probability at least* $1-1/n$*, for constants* $0<c_0,c_0'<\infty$*, independent of* (n,N,K,T)*,*

$$W(\beta^*)-W(\hat{\beta}^*)\leq c_0\exp(-c_0'K).$$

The proof is in Section C.3.7. The main additional restriction in Theorem 4.4 is that the sample size grows *exponentially* in the number of iterations (instead of polynomially). The theorem leverages properties of the gradient descent under strong concavity and smoothness (Bubeck, 2014), combined with derivations discussed below.

To our knowledge, these are the first regret guarantees under unknown (and partial) interference.

We now contrast with past literature. In the online optimization literature, the rate $1/T$ is common for convex optimization (Bottou et al., 2018), assuming independent units (see Duchi et al., 2018, for out-of-sample regret rates only). Differently, here, because of interference, we leverage between-clusters perturbations. Also, we do not have direct access to the gradient, and related optimization procedures are those of Flaxman et al. (2004); Agarwal et al. (2010), where regret can converge at rate $O(1/T)$ in expectation only, while high-probability bounds are $1/\sqrt{T}$.[46] Here, we exploit within-cluster concentration and between clusters' variation to control large deviations of the estimated gradients and obtain faster

[45] By a first-order Taylor expansion, a corollary is that the bound also holds for $\check{\beta}_k^w\pm\eta_n$ up to an additional factor which scales to zero at rate η_n (and therefore negligible under the conditions imposed on n).

[46] See Theorem 6 in Agarwal et al. (2010) and discussion below.

rates for high-probability bounds. This also allows us to extend out-of-sample guarantees beyond global strong concavity (assumed in the above references) in Appendix B.2. In our derivations, the perturbation parameter depends on the sample size, differently from the references above, and the idea of circular estimation is novel due to repeated sampling. Wager and Xu (2021) derive $1/T$ regret guarantees in the different settings of market pricing, as $n \to \infty$, with independent units and samples each wave. Our results are in a finite sample and do not impose independence or modeling assumptions other than partial interference. Viviano (2019) considers a single study and network, with observed neighbors, instead of a sequential experiment and imposes geometric (VC) restrictions.

These differences require a different set of techniques for derivations. The proof of the theorem (i) uses concentration arguments for locally dependent graphs (Janson, 2004); (ii) it uses the within-cluster and between-clusters variation for consistent estimation of the marginal effect, together with the matching design to guarantee that there is a vanishing bias when estimating marginal spillover effects for $K < \infty$; (iii) it uses a recursive argument to bound the cumulative error obtained through the estimation and circular cross-fitting.

Remark 9 (Relaxing concavity)**.** In Appendix B.2 (Theorem B.2), we derive rates under weaker restrictions than Assumption 4.2 (assuming that the function is strictly quasi-concave and locally strongly concave), using the gradient norm rescaling in Equation (15). □

Remark 10 (Non-adaptive experiment)**.** In Appendix A.4 we introduce an experiment to estimate β^* without adaptive randomization, in the spirit of Section 3. Appendix A.4 shows how we can leverage perturbations and marginal effects as in Algorithm 1 to control the out-of-sample regret without an adaptive experiment, but also formally shows the drawbacks of a non-adaptive experiment, since it does not control the in-sample regret. □

4.2 A comparison with the unconstrained optimum

Next, we study the following question: how does β^* compare to the policy that assigns treatments without restrictions on the policy function? We study this under more restrictive conditions. We omit the super-script k since our argument applies to any cluster. Consider

$$W_N^* - W(\beta^*), \quad W_N^* = \sup_{\mathcal{P}_N(\cdot)\in\mathcal{F}} \frac{1}{N}\sum_{i=1}^{N} \mathbb{E}\Big[\mathbb{E}_{D\sim\mathcal{P}_N(A,X)}[Y_{i,t}|A,X]\Big] \tag{16}$$

where $\mathcal{F}$ denotes the set of conditional distribution of the vector $D \in \{0,1\}^N$, given the network A and the covariates of all observations X. Equation (16) denotes the difference between the expected potential outcomes, evaluated at the global optimum over all possible assignments (conditional on A, X), and the welfare evaluated at the optimal policy β^*.

Assumption 4.3 (Discrete parameter space, assignment and minimum degree)**.** Assume that $X_i \in \mathcal{X}, \mathcal{X} = \{1, \cdots, |\mathcal{X}|\}, |\mathcal{X}| < \infty$. Let $\pi(x, \beta) = \beta_x$, and $\mathcal{B} = [0,1]^{|\mathcal{X}|}$. Assume in addition that $\inf_{x,x',u'} \int l(x, u, x', u') dF_{U|X=x}(u) \geq \kappa$, for some $\kappa \in (0, 1]$.

Assumption 4.3 states that we assign treatments based on finitely many observable types.[47] Each type $x \in \mathcal{X}$ is assigned a different probability β_x, which can take any value between zero and one. Assumption 4.3 also states that conditional on individual's type (X_i, U_i), any other unobserved type U_j can form a connection with individual i with some positive probability, conditional on observables X_j, provided that i and j are connected under the latent space representation (recall Equation 1).[48] The second restriction is on the potential outcomes. Let

$$\begin{aligned} Y_{i,t}(\mathbf{d}_t) &= \Big[\Delta(X_i) - c(X_i)\Big]\mathbf{d}_{i,t} + \mathcal{S}_{i,t}(\mathbf{d}_t) + \nu_{i,t}, \quad \mathbb{E}[\nu_{i,t}|X, A] = 0 \\ \mathcal{S}_{i,t}(\mathbf{d}_t) &= s\Big(\frac{\sum_{j\neq i} A_{i,j}\mathbf{d}_{j,t}1\{X_j = 1\}}{\max\{\sum_{j\neq i} A_{i,j}1\{X_j = 1\}, 1\}}, \cdots, \frac{\sum_{j\neq i} A_{i,j}\mathbf{d}_{j,t}1\{X_j = |\mathcal{X}|\}}{\max\{\sum_{j\neq i} A_{i,j}1\{X_j = |\mathcal{X}|\}, 1\}}\Big). \end{aligned} \tag{17}$$

Here $\Delta(\cdot)$ is the conditional direct treatment effect, and $c(\cdot)$ is the cost of the treatment. The function $s(\cdot)$ captures the spillover effects. The spillover effects depend on the fraction of treated neighbors. Spillovers are heterogeneous in the neighbors' types.

Assumption 4.4 (Costs are opportunity costs of an equal-impact intervention with no spillovers)**.** Assume that $\Delta(x) = c(x)$ for all $x \in \mathcal{X}$.

Assumption 4.4 states that the cost of the treatment is the opportunity cost, had the treatment been assigned to the same individuals which are disconnected. In a cash-transfer program, we may assign treatments to individuals in the same or nearby villages or to individuals spread out on an entire state or continent, without creating spillover effects in the latter case. The cost of the treatment to assign treatments in the same villages is their opportunity cost, i.e., the direct treatment effects, assumed to be the same.[49]

Theorem 4.5. *Let Equation (17) holds, with $s(\cdot)$ twice differentiable with bounded derivatives. Suppose that Assumption 2.1, 2.2, 4.3, 4.4 hold. Then as $N \to \infty, \gamma_N \to \infty$,*

$$\sup_{\mathcal{P}_N(\cdot)\in\mathcal{F}} \frac{1}{N}\sum_{i=1}^{N} \mathbb{E}\Big[\mathbb{E}_{D\sim\mathcal{P}_N(A,X)}[Y_{i,t}|A, X]\Big] - W(\beta^*) \to 0.$$

[47] This is often imposed, see e.g., Manski (2004), Graham et al. (2010).

[48] This condition is consistent with Assumption 2.1, since the assumption states that the expected minimum degree is bounded from below by $\kappa\gamma_N^{1/2}$, which is smaller than the maximum degree $\gamma_N^{1/2}$.

[49] Note that, under such restriction, treating each individual is not optimal only if, for *some* treatment configurations, there are negative spillovers. For instance, giving cash transfers to richer individuals may decrease the average satisfaction with the program.

The proof is in Appendix C.3.8. Theorem 4.5 shows that the assignment mechanism that maximizes welfare, when treatments can be assigned arbitrarily, converges to β^* with individualized treatments. It implies that collecting network information, is, on average not useful to improve welfare in this context. The theorem assumes that the maximum degree converges to infinity, but it may converge at a slower rate than N, consistently with our conditions in previous theorems. This is a novel result in the context of the literature on targeting networked individuals, which depends on the costs' assumption.[50]

Corollary 5. *Let the conditions in Theorem 4.2 and Theorem 4.5 hold. Then for a constant* $C < \infty$ *independent of* (N, n, T, K), $\lim_{N\to\infty,\gamma_N\to\infty} P\Big(W_N^* - W(\hat{\beta}) \leq \frac{C}{T}\Big) = 1$.

5 Calibrated Experiments

In this section, we study the properties of the methodology in numerical studies. We calibrate simulations to data from Cai et al. (2015) and Alatas et al. (2012, 2016), while making simplifying assumptions whenever necessary. In the first calibration, the outcome is insurance adoption, and the treatment is whether an individual received an intensive information session in the experiment. In the second calibration, the treatment is whether a household received a cash transfer, and the outcome is program satisfaction.[51]

Throughout these simulations, we study the problem of choosing a univariate parameter β, which denotes the unconditional treatment probability. In each cluster k, we generate data as

$$Y_{i,t} = \phi_0 + \phi_1 D_{i,t} + \phi_2 S_{i,t} + \phi_3 S_{i,t}^2 - c D_{i,t} + \eta_{i,t}, \quad S_{i,t} = \frac{\sum_{j\neq i} A_{i,j} D_{i,t}}{\max\{1, \sum_{j\neq i} A_{i,j}\}}, \quad \eta_{i,t} \sim_{i.i.d.} \mathcal{N}(0, \sigma^2), \tag{18}$$

where c is the cost of the treatment. We consider two sets of parameters $\Big(\phi_0, \phi_1, \phi_2, \phi_3, \sigma^2\Big)$ calibrated to data from Cai et al. (2015) and Alatas et al. (2012, 2016) respectively. We obtain information on neighbors' treatment directly from data from Cai et al. (2015). For the second application, we merge data from Alatas et al. (2012), and Alatas et al. (2016),

[50] In a different context, Akbarpour et al. (2018) show that for a class of diffusion mechanisms, random seeding is approximately optimal as we choose a few more seeds. Here, we do not study the problem from the perspective of network diffusion but instead focus on an exogenous interference mechanism with heterogeneity, and provide a different set of results. Our result hinges on Assumption 4.4.

[51] The experiment of Cai et al. (2015) contains multiple arms assigned at the household and village level. Here, we only focus on the treatment effects of intensive information sessions, pooling the remaining arms together for simplicity. The experiment of Alatas et al. (2012) contains different arms assigned at the village level, as well as information on cash transfers assigned at the household level. Here, we study the effect of cash transfers only and control for village-level treatments when estimating the parameters of interest.

and use information from approximately one hundred observations whose neighbors' treatments are all observable to estimate the parameters.[52] For either application, we estimate a linear model as in Equation (18) also controlling for additional covariates to guarantee unconfoundedness of the treatment.[53] We consider as cost of treatment $c = \phi_1$, i.e., the opportunity cost of allocating the treatment to a population of disconnected individuals.

We generate clusters with $N = 600$ units, and sample $n \in \{200, 400, 600\}$. We generate a geometric network of the form

$$A_{i,j} = 1\Big\{||U_i - U_j||_1 \leq 2\rho/\sqrt{N}\Big\}, \quad U_i \sim_{i.i.d.} \mathcal{N}(0, I_2),$$

where the parameter ρ governs the density of the network. The geometric formation process and the $1/\sqrt{N}$ follows similarly to simulations in Leung (2020). We consider two networks, a "sparse network" with $\rho = 2$, reported in the main text, and a "dense network", with $\rho = 6$, studied in the Appendix. Throughout our analysis, without loss of generality, we report welfare divided by its maximum $W(\beta^*)$ (i.e., $W(\beta^*) = 1$), and we subtract the intercept ϕ_0 since ϕ_0 does not depend on β.

We conclude with details on estimation. We fix the perturbation parameter $\eta_n = 10\%$;[54] similarly, in the adaptive experiment, we choose the learning rate $10\%/\sqrt{t}$ with gradient norm rescaling as Remark 7. This choice guarantees that for each iteration, we only vary treatment probabilities by at most 10%, and the size of the variation is decreasing over each iteration, in the same spirit of learning rate under strong concavity without norm rescaling.[55] Since the model does not allow for time-varying fixed effects, we estimate marginal effects without baseline outcomes. For the multi-wave experiment, we initialize parameters at a small treatment probability $\beta = 0.2$.[56]

[52] This is different from Figure 2 where we use information from individuals whose 80% or more neighbors are observable. We make this choice in Section 2 to increase precision to estimate heterogeneous effects. This approach introduces a sampling bias in the estimation procedure, which we ignore for simplicity, given that our goal is not the analysis of the original experiment but only calibrating numerical studies.

[53] For Cai et al. (2015) the covariates are gender, age, rice area, literacy level, a coefficient that captures the risk aversion, the baseline disaster probability, education, and a dummy containing information on whether the individual has one to five friends. For Alatas et al. (2012) we control for the education level, village-level treatments, i.e., how individuals have been targeted in a village (i.e., via a proxy variable for income, a community-based method, or a hybrid), the size of the village, the consumption level, the ranking of the individual poverty level, the gender, marital status, household size, the quality of the roof and top (which are indicators of poverty).

[54] Here 10% is consistent with the rule of thumb in Section B.7 which would prescribe values between 7% and 12% as we vary n. In Figure D.7 we report results as we vary η_n.

[55] This choice is preferable to $10\%/\sqrt{T}$ because it allows for larger steps in the initial iterations. A valid alternative choice is also $10\%/t$, corresponding to the one under strong concavity. The latter case has a practical drawback: updates become very small after a few iterations. For a comparison, see Figure D.5.

[56] This choice guarantees that no less than 10% of individuals are treated once we impose the negative perturbation.

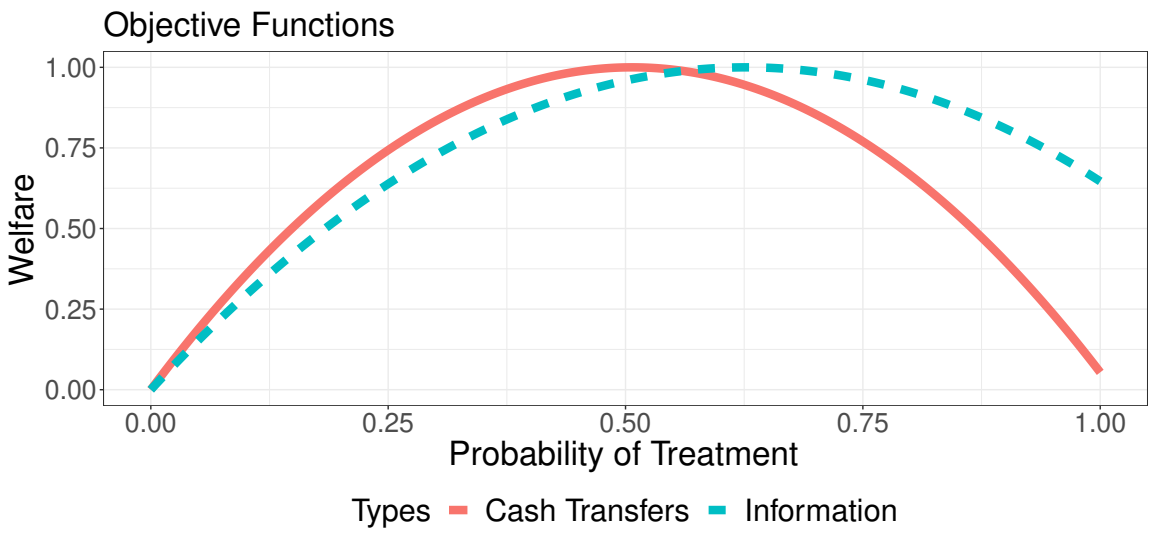


Figure 5: Objective functions (rescaled by $W(\beta^*)$, and minus the intercept ϕ_0) as functions of unconditional treatment probabilities, with cost of treatments $c = \phi_1$.

5.1 One-wave experiment

First, we study the properties of the one wave experiment as we vary the number of clusters K and the sample size from each cluster n. We are interested in testing the one-sided null of whether we should increase the number of treated individuals to increase welfare, i.e.,

$$H_0 : \frac{\partial W(\beta)}{\partial \beta} \leq 0, \quad H_1 = \frac{\partial W(\beta)}{\partial \beta} > 0 \quad \beta \in [0.1, \cdots, \beta^*]. \tag{19}$$

In Figure D.1, in the Appendix, we report the power of the test as a function of the regret for $\rho = 2$. Power is increasing in the regret, the number of clusters, and sample size. However, the marginal improvement in the power from twenty to thirty clusters is small. This result is suggestive of the benefit of the method even with few clusters and a small sample size.

Figure 6 illustrates the benefit of the method as, upon rejection of H_0, we recommend increasing the treatment probability by 5%, as a function of the baseline treatment probability. For example, for $\beta = 0.2$, it indicates the relative welfare increase if we were to increase the treatment probability to 0.25, relative to the status quo where $\beta = 0.2$. The figure shows that the relative improvement can be as large as fifty percentage points compared to the status quo. In addition, while the gain is increasing in the sample size, substantial gains can be obtained even if the sample size from each cluster is as small as $n = 200$. This is particularly relevant for targeting information: in such a case, differences in power across different sample sizes only occur when the regret is small (see Figure D.1), or equivalently, when we are already close to the optimum. As a result, once we take welfare effects into account,

larger sample size may lead to small (negligible) improvements in welfare. As this example suggests, when the goal is welfare maximization, power analysis may be complemented by the welfare analysis discussed here.

Results are robust as we increase the density of the network with $\rho = 6$ (see Appendix D).In Table 1 we report the size of the test, which provides supportive evidence to the theoretical findings.

In Appendix D.4 we discuss an M-Turk experiment which we conducted to diffuse information to increase vaccination. Results are collected in Figure D.9 and show significant welfare improvements.

5.2 Multiple-wave experiment

Finally, we study the performance of the adaptive experiment. We let $T \in \{5, 10, 15, 20\}$. In Table 2 we report the welfare improvement of the proposed method with respect to a grid search method that samples observations from an equally spaced grid between $[0.1, 0.9]$ with a size equal to the number of clusters (i.e., $2T$). We consider the best competitor between the one that maximizes the estimated welfare obtained from a correctly specified quadratic function and the one that chooses the treatment with the largest value within the grid. The panel at the top of Table 2 reports the out-of-sample welfare improvement. The improvement is positive and up to three percentage points for targeting information and up to sixty percentage points for targeting cash transfers. Improvements are generally larger for larger T. In one instance only, for $T = 5$ and a small sample size $n = 200$, we observe a negative effect for targeting information of two percentage points. The panel at the bottom of Table 2 reports positive and large improvements for the in-sample welfare across all the designs, worst-case across clusters. For the worst-case regret, we fix the number of clusters to $K = 40$ for our method and study the properties as a function of the number of iterations.[57] The improvements are twice as large for targeting information and thirty percentage points larger for targeting cash transfers. These are often increasing in T with a few exceptions.[58]

These results illustrate large benefits for *both* estimating policies to be implemented on a new population and to maximize participants' welfare. Figure 7 reports the out-of-sample regret of the method (and Figure D.2 the in-sample regret). The regret converges to zero as we increase the number of iterations. The larger sample size guarantees a smaller regret. In Appendix D.1, D.2, we report results for $\rho = 6$, consistent with findings in the main text. In

[57] Fixing $K = 40$ allows us to change the number of iterations up to $T = 20$ while keeping fixed the number of clusters.

[58] The reason improvements are not always increasing in T is that uniform concentration may deteriorate for large T and small n as we consider the worst-case welfare across clusters.

Appendix D.3 we provide simulations with covariates using data from Alatas et al. (2012).

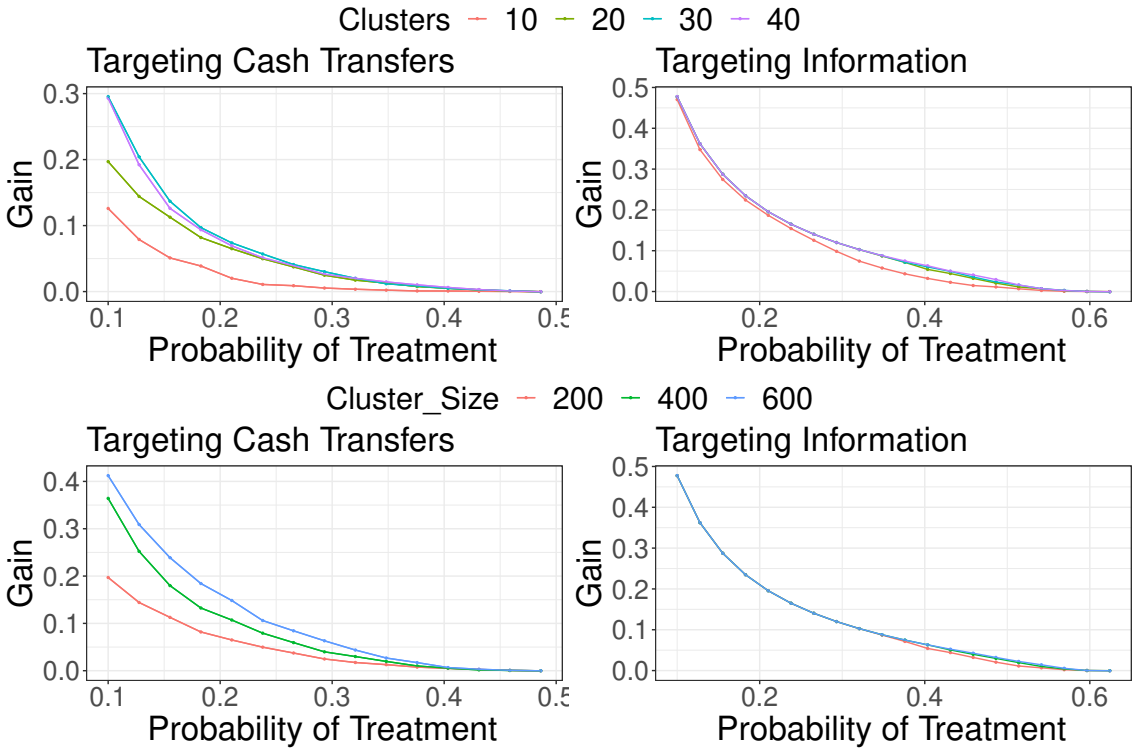


Figure 6: One-wave experiment. $\rho = 2$. Expected percentage increase in welfare from increasing the probability of treatment β by 5% upon rejection of H_0. Here, the x-axis reports $\beta \in [0.1, \cdots, \beta^* - 0.05]$. The panels at the top fix $n = 400$ and vary the number of clusters. The panels at the bottom fix $K = 20$ and vary n.

Table 1: One wave experiment. 200 replications. Coverage for testing H_0 (size is 5%). First panel corresponds to $\rho = 2$, and second panel to $\rho = 6$.

	Information				Cash Transfer			
$K =$	10	20	30	40	10	20	30	40
$n = 200$	0.905	0.950	0.905	0.900	0.920	0.940	0.915	0.895
$n = 400$	0.980	0.960	0.900	0.925	0.980	0.960	0.895	0.930
$n = 600$	0.975	0.970	0.955	0.945	0.970	0.995	0.960	0.935
$n = 200$	0.925	0.880	0.880	0.900	0.925	0.940	0.905	0.905
$n = 400$	0.980	0.940	0.920	0.920	0.980	0.960	0.900	0.925
$n = 600$	0.975	0.890	0.930	0.995	0.975	0.995	0.950	0.915

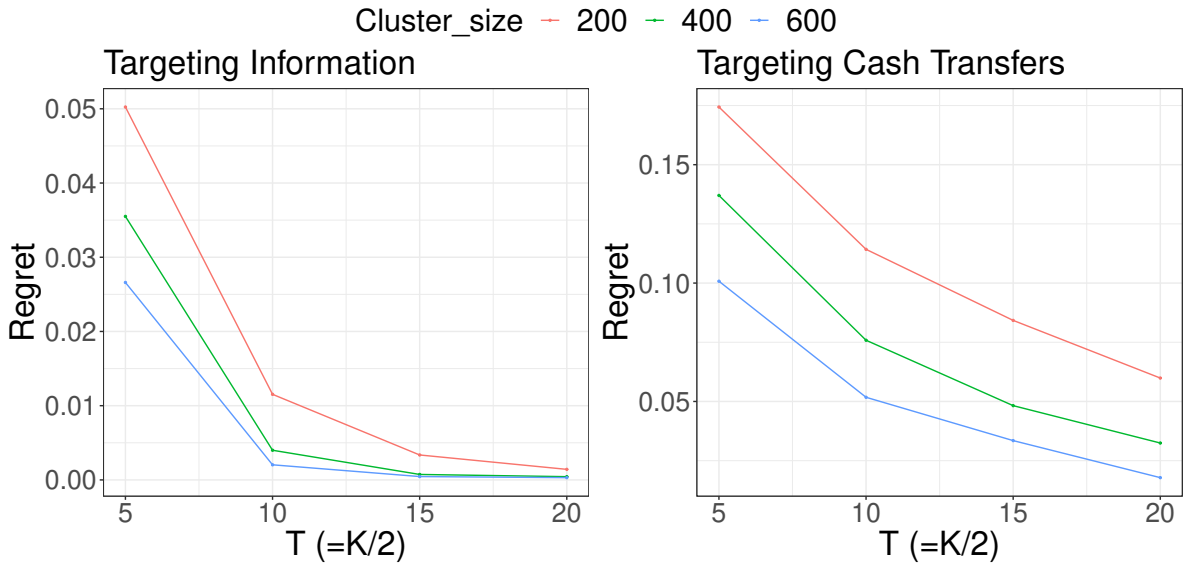


Figure 7: Adaptive experiment $\rho = 2$. 200 replications. The panel reports the out-of-sample regret of the method as a function of the number of iterations.

Table 2: Multiple-wave experiment. 200 replications. Relative improvement in welfare with respect to best competitor for $\rho = 2$. The panel at the top reports the out-of-sample regret and the one at the bottom the worst case in-sample regret across clusters.

	Information				Cash Transfer			
$T =$	5	10	15	20	5	10	15	20
$n = 200$	-0.026	0.014	0.043	0.033	0.295	0.390	0.528	0.322
$n = 400$	0.0003	0.026	0.026	0.035	0.462	0.444	0.589	0.563
$n = 600$	0.002	0.035	0.021	0.022	0.485	0.489	0.622	0.644
$n = 200$	1.103	1.370	1.451	1.447	0.254	0.276	0.305	0.323
$n = 400$	1.400	1.616	1.667	1.626	0.282	0.329	0.367	0.379
$n = 600$	1.546	1.771	1.828	1.751	0.279	0.335	0.364	0.368

6 Dynamic treatment effects: an overview

One assumption that we impose throughout the text is the absence of carry-overs. This section briefly discusses an extension with dynamic treatments. For simplicity, we omit covariates and assume that $X_i = 1$, and defer to Appendix A.3 formal details.

In the presence of carry-over effects, outcomes also depend on past treatments. For

treatments assigned with exogenous parameters $(\beta_{k,1}, \cdots, \beta_{k,t})$ as in Definition 2.3, we let

$$Y_{i,t}^{(k)} = \Gamma(\beta_{k,t}, \beta_{k,t-1}) + \varepsilon_{i,t}^{(k)}, \quad \mathbb{E}_{\beta_{k,1:t}}\left[\varepsilon_{i,t}^{(k)}\right] = 0,$$

for some unknown $\Gamma(\cdot)$, $\varepsilon_{i,t}^{(k)}$. The components $\beta_{k,t}, \beta_{k,t-1}$ capture present and carry-over effects that result from individual and neighbors' treatments in the past two periods. We study the problem of estimating a path of treatment probabilities $(0, \beta_1, \cdots, \beta_T)$ from an experiment, where, in the first period, we assume for simplicity that none of the individuals is treated. We then implement this path on a new population without having access to the outcomes of such a new population. We maximize long-run welfare, defined as follows:

$$\mathcal{W}(\{\beta_s\}_{s=1}^{T^*}) = \sum_{t=1}^{T^*} q^t \Gamma(\beta_t, \beta_{t-1}), \quad \text{for a given horizon } T^*, \text{ and discounting factor } q < 1.$$

The long-run welfare defines the cumulative (discounted) welfare effect obtained from a certain sequence of decisions $(\beta_1, \beta_2, \cdots)$. Our goal is to find the sequence that maximizes the long-run welfare.[59] We start from the following observation: by the first-order conditions

$$\frac{\partial \Gamma(\beta_t, \beta_{t-1})}{\partial \beta_t} + q \frac{\partial \Gamma(\beta_{t+1}, \beta_t)}{\partial \beta_t} = 0, \quad \text{for all } t. \tag{20}$$

Equation (20) shows that the choice of the welfare-maximizing parameter β_{t+1} may depend on the previous two decisions. Using ideas from reinforcement learning (Sutton and Barto, 2018), we parametrize future treatment probabilities based on past treatment probabilities as $\beta_{t+1} = h_\theta(\beta_t, \beta_{t-1}), \theta \in \Theta$, for some given function $h_\theta(\cdot)$, and find the parameter $\theta \in \Theta$, which maximizes welfare. The algorithm estimates the function $\Gamma(\cdot)$ using a single wave experiment and then maximizes

$$\widehat{\theta} \in \arg\max_{\theta \in \Theta} \sum_{t=1}^{T^*} q^t \widehat{\Gamma}(\beta_t, \beta_{t-1}), \quad \beta_t = h_\theta(\beta_{t-1}, \beta_{t-2}) \quad \forall t \geq 1, \quad \beta_0 = \beta_{-1} = 0.$$

Our main insight here is on how to design the experiment to estimate $\widehat{\Gamma}(\cdot)$. If $\widehat{\Gamma}$ is estimated with randomization based on a simple grid-search procedure, the rate of convergence of the regret would be $1/\sqrt{K}$ (see Appendix A.3). However, in Appendix A.3 we show that by using local perturbations – similarly to what was discussed in previous sections – and leveraging information from the estimated gradient, we can achieve a convergence rate of

[59] Alternatively, we may also estimate the same treatment probability each period (see Appendix B.4).

the out-of-sample regret of order $1/K$ (but not faster in n).[60] The single-wave procedure comes at a cost: the rate is specific to the one-dimensional setting and carry-overs over two consecutive periods; in p dimensions, the rate would be of much slower order due to the curse of dimensionality. This is different from the adaptive design discussed in previous sections, where the dimension does not affect the rate in K, and it opens new research questions on the optimal design of experiments in the presence of carry-over effects.

7 Conclusions

This paper makes two main contributions. First, it introduces a single-wave experimental design to estimate the marginal effect of the policy and test for policy optimality. The experiment also allows us to identify and estimate treatment effects, which can be of independent interest. Second, it introduces an adaptive experiment to maximize welfare. We derive asymptotic properties for inference and provide a set of guarantees on the in-sample and out-of-sample regret. To our knowledge, this is the first paper to study inference on marginal effects and adaptive experimentation with unobserved interference.

In a single-wave experiment, we encourage researchers to identify and report estimates of the marginal effects. We show that we can use the information on the marginal effects to conduct hypothesis testing on policy optimality and, ultimately, incorporate uncertainty in decision-making. Future research may explore notions of efficiency in this setting.

Our work opens new questions also from a theoretical perspective. The main assumption is that such clusters are observable before the experiment starts. We leave to future research to study properties when (i) such clusters are not fully disconnected, in the same spirit of Leung (2021); (ii) such clusters need to be estimated, similarly to graph-clustering procedures as in Ugander et al. (2013); (iii) clusters present different distributions, as we discuss in Appendix A.2. Similarly, it may be interesting to study the properties of our method, as the degree of interference is proportional to the sample size. This is theoretically possible, as illustrated in Theorem 3.1, and we leave its comprehensive analysis to future research. An open question is also how to combine policy learning with estimation procedures that impute the network (e.g., Alidaee et al., 2020; Breza et al., 2020; Manresa, 2013).

Finally, an important assumption of this paper and, in general, of the literature on adaptive experimentations is that welfare is a function of observable characteristics. We leave to future research to study whether (and how) we may allow for unobserved utilities.

[60] The idea is as follows: we randomize probabilities $(\beta_1, \beta_2) \in [0,1]^2$ from a coarse grid and use small groups (three) clusters to estimate the partial derivatives at each point. We extrapolate the value of $\Gamma(\cdot)$ throughout the set $[0,1]^2$ with a first-order Taylor approximation around the closest point in the grid, using the information on the estimated marginal effect.

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Appendix to "Policy design in experiments with unknown interference"

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Appendix A Main extensions

This section contains five extensions. Additional extensions are included in Appendix B.

A.1 Estimation with global interference

In this section, we relax the local dependency assumptions. The treatment affects each unit in a cluster k through a global interference mechanism mediated by a random variable $p_t^{(k)}$. For example, we can think of $p_t^{(k)}$ as the average number of treated units in a cluster or to the adjusted price in a particular market or village due to the program. For simplicity, we consider the case where $X_i^{(k)} = 1$, i.e., the policy of interest is a global policy (e.g., the probability of treatment).

We discuss assumptions on the outcome model below.

Assumption A.1 (Global interference)**.** Let treatments be assigned as in Definition 2.3 with exogenous vector of parameters $\beta_{k,1:t}$. Let

$$Y_{i,t}^{(k)} = \alpha_t + \tau_k + g\left(p_t^{(k)}, \beta_{k,t}\right) + \varepsilon_{i,t}^{(k)}, \quad \mathbb{E}_{\beta_{k,1:t}}\left[\varepsilon_{i,t}^{(k)} | p_t^{(k)}\right] = 0,$$

for some function $g(\cdot)$ unknown to the researcher, bounded and twice continuously differentiable with bounded derivatives, and unobservable $p_t^{(k)}$. Assume in addition that $\varepsilon_{i,t}^{(k)} \perp \varepsilon_{j\notin\mathcal{I}_i^{(k)},t}^{(k)} | \beta_{k,1:t}, p_t^{(k)}$ for some set $|\mathcal{I}_i^{(k)}| = \mathcal{O}(\gamma_N)$.

Assumption A.1 states that the outcome within each cluster is a function of a common factor, and treatment assignment rule $\beta_{k,t}$ plus unobservables centered around zero and locally dependent. The factor $p_t^{(k)}$ also depends on the treatment assignments of all individuals. This is formalized below.

Assumption A.2 (Global interference component)**.** Let treatments be assigned as in Definition 2.3. Assume that

$$p_t^{(k)} = q(\beta_{k,t}) + o_p(\eta_n),$$

with $q(\beta)$ being unknown, bounded and twice continuously differentiable in β with uniformly bounded derivatives.

Assumption A.2 states that the factor can be expressed as the sum of two components. The first component $q(\cdot)$ depends on the policy parameter $\beta_{k,t}$ assigned at time t and on the distribution of covariates of all units in a cluster.[61] The second component is a stochastic

[61] Observe that we can equivalently relax Assumption A.2 and assume that $q(\beta, \cdot)$ depends on the empirical distribution of covariates and use basic concentration arguments (Wainwright, 2019) to show that asymptotically the two definitions are equivalent.

component that depends on the realized treatment effects. We illustrate an example below.

Example A.1 (Within cluster average)**.** Suppose that

$$Y_{i,t}^{(k)} = t(\bar{D}_t^{(k)}, \nu_{i,t}), \quad \nu_{i,t}^{(k)} \sim_{i.i.d.} \mathcal{P}_\nu, \quad D_{i,t}^{(k)} \sim_{i.i.d.} \text{Bern}(\beta)$$

where $t(\cdot)$ is some arbitrary (smooth) function. Then $p_t^{(k)} = \bar{D}_t^{(k)}$ i.e., individuals depend on the average exposure in a cluster. We can write

$$Y_{i,t}^{(k)} = t(p_t^{(k)}, \nu_{i,t}^{(k)}) \quad \text{where} \quad p_t^{(k)} = \beta + \underbrace{(\bar{D}_t^{(k)} - \beta)}_{=O_p(n^{-1/2})},$$

which satisfies Assumption A.2 for $\eta_n = n^{-1/3}$ or larger. □

Example A.1 illustrates how we can accommodate global interference mechanism whenever individuals depend on statistics of treatment assignments, such as their average. Example A.1 does not allow for local spillovers between units but only for global interference.[62]

We are interested in the marginal effects defined below.

$$V_g(\beta) = \frac{\partial W_g(\beta)}{\partial \beta}, \quad W_g(\beta) = g(q(\beta), \beta).$$

Estimation of the marginal effect follows similarly to Equation (7). The following theorem guarantees consistency.

Theorem A.1. *Let Assumption A.1, A.2 hold with subgaussian* $\varepsilon_{i,t}^{(k)}$*,* $X = 1$*. Then for* $\hat{V}_{(k,k+1)}$ *estimated as in Algorithm 2, for* k *being odd holds:*

$$\left|\widehat{V}_{(k,k+1)} - V_g(\beta)\right| = \mathcal{O}_p\left(\sqrt{\frac{\gamma_N \log(\gamma_N)}{\eta_n^2 n}} + \eta_n\right) + o_p(1).$$

The proof is in Appendix C.4.1. Theorem A.1 guarantees consistency of the estimated gradient. The experiment can be conducted similarly to what discussed in Section 4 and omitted for the sake of brevity.

[62] Assumption A.2 builds on the model of demand as a function of individual prices in Wager and Xu (2021). The difference is that we do not rely on a specific modeling assumption of market interactions. Instead, we model outcomes as functions of exposures in a certain cluster and exploit the two-clusters variation for consistent estimation, as we discuss below.

A.2 Matching clusters with distributional embeddings in the presence of heterogeneity in the distribution of covariates

In this section, we turn to the problem of matching clusters, allowing for covariates having different distributions in different clusters. In particular, we assume that $X_i^{(k)} \sim_{i.i.d.} F_X^{(k)}$. The main distinction from previous sections is that $F_X^{(k)}$ is cluster-specific. The section works as follows: first, we characterize the bias of the difference in means estimators; second, we propose a matching algorithm that minimizes the worst-case bias.

We start from the simple setting with two clusters k, k' only, and two periods $t \in \{0, 1\}$. Treatments are assigned as follows

$$\begin{aligned} t = 0: \quad & D_{i,0}^{(h)} \sim \pi(X_i^{(h)}; \beta_0), \quad h \in \{k, k'\} \\ t = 1: \quad & D_{i,1}^{(k)} \sim \pi(X_i^{(k)}; \beta), \quad D_{i,1}^{(k')} \sim \pi(X_i^{(k')}; \beta'). \end{aligned} \tag{A.1}$$

Namely, at time $t = 0$, treatments are assigned with a parameter β_0. At time $t = 1$ treatments are assigned with parameter β in cluster k and β' in cluster k'.

The estimand of interest is the difference in the average effects in cluster k, formally

$$\omega_k = \int y(x; \beta) dF_X^{(k)}(x) - \int y(x; \beta') dF_X^{(k)}(x).$$

We study properties of the difference in differences estimator

$$\widehat{\omega}_k(k') = \left[\bar{Y}_1^{(k)} - \bar{Y}_1^{(k')}\right] - \left[\bar{Y}_0^{(k)} - \bar{Y}_0^{(k')}\right],$$

which defines a difference in differences between the two clusters over two consecutive periods.

Our focus is to control the bias of the estimator. This is defined in the following lemma.

Lemma A.2. *Let Assumption 2.1, 2.2, and treatments assigned as in Equation (A.1). Then*

$$\mathbb{E}[\widehat{\omega}_k(k')] - \omega_k = \int \Big(y(x; \beta') - y(x; \beta_0)\Big) d\Big(F_X^{(k)}(x) - F_X^{(k')}(x)\Big).$$

Lemma A.2 shows that the bias depends on the difference between the expectations averaged over two different distributions. Unfortunately, the bias is unknown since it depends on the function $y(\cdot)$, which is not identifiable with finitely many clusters. We therefore bound the worst-case error over a class of functions $x \mapsto [y(x; \beta') - y(x; \beta_0)] \in \mathcal{M}$, with $\mathcal{M}$ defined below. The proof follows directly from Lemma 2.1, and rearrangement.

We start by defining $\mathcal{M}$ be a reproducing kernel Hilbert space (RKHS) equipped with

a norm $||\cdot||_{\mathcal{M}}$.[63] Without loss of generality, we study the worst-case functionals over the unit-ball. Formally, we focus on bounding the worst-case error of the form[64]

$$\sup_{[y(\cdot;\beta')-y(\cdot;\beta_0)]\in\mathcal{M}:||y(\cdot;\beta')-y(\cdot;\beta_0)||_{\mathcal{M}}\leq 1}\left|\omega_k-\mathbb{E}[\widehat{\omega}_k(k')]\right|=\sup_{f\in\mathcal{M}:||f||_{\mathcal{M}}\leq 1}\left\{\int f(x)d(F_X^{(k)}-F_X^{(k')})\right\}. \tag{A.2}$$

The right-hand side is know as the maximum mean discrepancy (MMD), a measure of distances in RKHS (see Muandet et al., 2016, and references therein). It is known that the MMD can be consistently estimated using kernels. In particular, given a particular choice of a kernel $k(\cdot)$, which corresponds to a certain RKHS, we can estimate

$$\begin{aligned}\widehat{\text{MMD}}^2(k,k')&=\frac{1}{n(n-1)}\sum_{i=1}^{n}\sum_{j\neq i}h\Big(X_i^{(k)},X_i^{(k')},X_j^{(k)},X_j^{(k')}\Big),\\ h(x_i,y_i,x_j,y_j)&=k(x_i,x_j)+k(y_i,y_j)-k(x_i,y_j)-k(x_j,y_i).\end{aligned} \tag{A.3}$$

The estimator estimates the squared MMD. It only depends on the kernel function $k(\cdot)$ and hence can be easily constructed in a finite sample without requiring an explicit characterization of RKHS. Here, $\widehat{\text{MMD}}^2(k,k')\to_p\left|\left|\mu_{F_X^{(k)}}-\mu_{F_X^{(k')}}\right|\right|^2_{\mathcal{M}}$ (Sriperumbudur et al., 2012).

We now turn to the problem of matching clusters. We do so using the estimated MMD in Equation (A.3). We first note that the estimator $\widehat{\text{MMD}}^2(k,k')$ only depends on *pre-treatment* variables, and hence can be computed *before* treatments are assigned. As a result, given cluster k, we can match k with the $k'\neq k$ having the smallest estimated MMD. Formally, the following matching algorithm is considered:

- construct
$$k'\in\arg\min_{k\neq k}\widehat{\text{MMD}}^2(k,k'). \tag{A.4}$$
based on the minimum estimated MMD in Equation (A.3).
- Randomize treatments as in Equation (A.1);
- Estimate $\widehat{\omega}_k(k')$.

When instead we want to match without replacement clusters, we can minimize some aggregate measures of error (e.g., the sum of estimated MMD across clusters).

[63] A RKHS is an Hilbert space of functions where all the evaluations functionals are bounded, namely, where for each $f\in\mathcal{M}$, and $x\in\mathcal{X}$, $f(x)\leq C||f||_{\mathcal{M}}$ for a finite constant C. Intuitively, assuming that $[y(\cdot;\beta')-y(\cdot;\beta_0)]\in\mathcal{M}$ imposes smoothness conditions on the average effect as a function of x.

[64] Here Equation (A.2) follows directly from Lemma A.2 and the fact that the integral is a scalar.

A.3 Policy choice with dynamic treatments

This section studies an experimental design when carry-overs occur. For simplicity, we omit covariates and assume that $X_i = 1$.

We start our discussion by introducing the dynamic model. For the sake of brevity, we directly impose a high-level condition on the outcome model.

Assumption A.3 (Dynamic model)**.** For treatments assigned with exogenous parameters $(\beta_{k,1}, \cdots, \beta_{k,t})$ as in Definition 2.3, let the followig hold

$$Y_{i,t}^{(k)} = \Gamma(\beta_t, \beta_{t-1}) + \varepsilon_{i,t}^{(k)}, \quad \mathbb{E}_{\beta_{k,1:t}}\left[\varepsilon_{i,t}^{(k)}\right] = 0,$$

for some unknown $\Gamma(\cdot)$, $\varepsilon_{i,t}^{(k)}$.

The components $\beta_{k,t}, \beta_{k,t-1}$ capture present and carry-over effects that result from individual and neighbors' treatments in the past two periods. Here, we allow for both panels and repeated cross-sections. We study the problem of estimating a path of treatment probabilities $(0, \beta_1, \cdots, \beta_T)$ from an experiment, where, in the first period, we assume for simplicity that none of the individuals is treated. This path is then implemented on a new population without having access to the outcomes of such a new population.

We provide a simple example below.

Example A.2. Suppose that

$$Y_{i,t}^{(k)} = D_{i,t}^{(k)}\phi_1 + \frac{\sum_{j\neq i} A_{i,j}^{(k)} D_{i,t-1}}{\sum_{j\neq i} A_{i,j}^{(k)}}\phi_2 + \nu_{i,t}^{(k)}, \quad D_{i,t}^{(k)} \sim_{i.i.d.} \text{Bern}(\beta_t).$$

That is, individuals depend on their present treatment assignment and on the treatment assignments of the neighbors in the previous period. Let $\nu_{i,t}$ be a zero-mean random variable. The expression simplifies to

$$Y_{i,t}^{(k)} = \beta_t\phi_1 + \beta_{t-1}\phi_2 + \varepsilon_{i,t}^{(k)}$$

where $\varepsilon_{i,t}^{(k)}$ is zero mean, and depends on neighbors' and individual assignments. □

We now define the long-run welfare.

Definition A.1 (Long-run welfare)**.** Given an horizon T^*, define the long-run welfare as follows:

$$\mathcal{W}(\{\beta_s\}_{s=1}^{T^*}) = \sum_{t=1}^{T^*} q^t \Gamma(\beta_t, \beta_{t-1}),$$

for a known discounting factor $q < 1$, where $\beta_0 = 0$.

The long-run welfare deifines the cumulative (discounted) welfare obtained from a certain sequence of decisions $(\beta_1, \beta_2, \cdots)$.

Our goal is to maximize the long-run welfare. Whenever the optimal policy is stationary, i.e., every period we are interested in estimating the same treatment probability, we can apply results to previous sections to this case. This is illustrated in Appendix B.4.

If, instead, we are interested in estimating a *sequence* of decisions, as discussed in the main text, we parametrize future treatment probabilities based on past treatment probabilities as follows

$$\beta_{t+1} = h_\theta(\beta_t, \beta_{t-1}), \quad \theta \in \Theta.$$

The parametrization is imposed for computational convenience. For some arbitrary large T^*, the objective function takes the following form

$$\begin{aligned} \widetilde{W}(\theta) &= \sum_{t=1}^{T^*} q^t \Gamma\Big(\beta_t, \beta_{t-1}\Big), \\ \beta_t &= h_\theta(\beta_{t-1}, \beta_{t-2}) \quad \text{for all } t \geq 1, \quad \beta_0 = \beta_{-1} = 0. \end{aligned} \tag{A.5}$$

Here $\widetilde{W}(\theta)$ denotes the long-run welfare indexed by a given policy's parameter θ. The objective function defines the discounted cumulative welfare induced by the policy h_θ.

Definition A.2 (Non-stationary policy decisions)**.** A non-stationary policy is defined as as a map $h_\theta : \mathcal{B} \times \mathcal{B}, \theta \in \Theta$. Define the non-stationary estimand as follows:

$$h_{\theta^*}(\cdot), \quad \theta^* \in \arg\max_{\theta \in \Theta} \widetilde{W}(\theta).$$

The algorithm estimates the function $\Gamma(\cdot)$ using a single wave experiment, i.e., we use a single period of experimentation. We then use the estimated function $\Gamma(\cdot)$ and its gradient for estimating the optimal policy.

The randomization and estimators are described in Algorithm E.2. We conduct the randomization using two periods of experimentations only. We partition the space $[0, 1]^2$ into a grid $\mathcal{G}$ of equally spaced components (β_1^r, β_2^r) for each triad of clusters r. Within each triad, we induce small deviations to the parameters β. For each triad r, the algorithm returns

$$\widetilde{\Gamma}(\beta_2^r, \beta_1^r), \quad \widehat{g}_1(\beta_2^r, \beta_1^r), \quad \widehat{g}_2(\beta_2^r, \beta_1^r)$$

where the latter two components are the estimated partial derivatives of $\Gamma(\cdot)$, and $\widetilde{\Gamma}(\beta_2^r, \beta_1^r)$ is the within cluster average.

For each pair of parameters (β_2, β_1), we estimate $\widehat{\Gamma}(\beta_2, \beta_1)$ as follows

$$\begin{aligned} &\widehat{\Gamma}(\beta_2, \beta_1) = \widetilde{\Gamma}(\beta_2^r, \beta_1^r) + \widehat{g}_2(\beta_2^r, \beta_1^r)(\beta_2 - \beta_2^r) + \widehat{g}_1(\beta_2^r, \beta_1^g)(\beta_1 - \beta_1^r), \\ &\text{where} \quad (\beta_1^r, \beta_2^r) = \arg \min_{(\tilde{\beta}_1, \tilde{\beta}_2) \in \mathcal{G}} \Big\{ ||\beta_1 - \tilde{\beta}_1||^2 + ||\beta_2 - \tilde{\beta}_2||^2 \Big\}. \end{aligned} \tag{A.6}$$

The idea is as follows: we estimate $\Gamma(\beta_2, \beta_1)$ at (β_2, β_1) using a a first-order Taylor approximation around the closest pairs of parameters in the grid $\mathcal{G}$. Given $\widehat{\Gamma}$, we estimate the welfare-maximizing parameter[65]

$$\widehat{\theta} \in \arg\max_{\theta \in \Theta} \sum_{t=1}^{T^*} q^t \widehat{\Gamma}(\beta_t, \beta_{t-1}), \quad \beta_t = h_\theta(\beta_{t-1}, \beta_{t-2}) \quad \forall t \geq 1, \quad \beta_0 = \beta_{-1} = 0.$$

In the following theorem, we study the behaviour of $\widehat{\theta}$, in terms of out-of-sample regret.

Theorem A.3 (Out-of-sample regret)**.** *Let Assumption A.3 hold. Let $X = 1$, and suppose that $\Gamma(\beta_2, \beta_1)$ is twice differentiable with bounded derivatives. Let treatments be assigned as in Algorithm E.2. Suppose in addition that $\varepsilon_{i,t}^{(k)} \perp \varepsilon_{j \notin \mathcal{I}_i^{(k)}}^{(k)}$ where $|\mathcal{I}_i^{(k)}| \leq \gamma_N$, for some arbitrary γ_N and $\varepsilon_{i,t}^{(k)}$ is sub-gaussian. Let $\gamma_N \log(\gamma_N)/(\eta_n^2 n) = o(1)$. Then*

$$\lim_{n \to \infty} P\Big(\sup_{\theta \in \Theta} \widetilde{W}(\theta) - W(\widehat{\theta}) \leq \frac{\bar{C}}{K} \Big) = 1$$

for a constant $\bar{C}$ independent of K.

The proof is in Appendix C.4.2. To our knowledge, Algorithm E.2 is novel to the literature of experimental design.[66]

Theorem A.3 shows that with probability converging to one as the size of each cluster increases, the regret scales at a rate $1/K$. To gain further intuition on the derivation of the theorem, observe that we can bound

$$\sup_{\theta \in \Theta} \widetilde{W}(\theta) - \widetilde{W}(\widehat{\theta}) \leq 2 \sum_t q^t \times \underbrace{\sup_{(\beta_1, \beta_2) \in [0,1]^2} \Big| \widehat{\Gamma}(\beta_2, \beta_1) - \Gamma(\beta_2, \beta_1) \Big|}_{(A)}.$$

[65] Here, $\widehat{\theta}$ can be obtained using off-the-shelf algorithms. A simple example is running in-parallel multiple gradient descent algorithms initialized over different starting points and choosing the one which leads to the largest objective $\sum_{t=1}^{T^*} q^t \widehat{\Gamma}(\beta_t, \beta_{t-1})$.

[66] We note that optimal dynamic treatments have been studied in the literature on bio-statistics, see, e.g., Laber et al. (2014), while here we consider the different problem of the design of the experiment. Adusumilli et al. (2019) discuss off-line policy estimation in the presence of dynamic budget constraints with *i.i.d.* observations. The authors assume no carry-overs, and do not discuss the problem of experimental design.

To bound (A) observe first that each element in the grid $\mathcal{G}$ has a distance of order $1/\sqrt{K}$ since the grid has two dimensions and $K/3$ components. As a result for any element (β_2, β_1), we can write

$$\Gamma(\beta_2,\beta_1) = \underbrace{\Gamma(\beta_2^r,\beta_1^r)}_{(B)} + \underbrace{\frac{\partial\Gamma(\beta_2^r,\beta_1^r)}{\partial\beta_1^r}(\beta_1-\beta_1^r) + \frac{\partial\Gamma(\beta_2^r,\beta_1^r)}{\partial\beta_2^r}(\beta_2-\beta_2^r)}_{(C)} + \underbrace{\mathcal{O}\Big(||\beta_1-\beta_1^r||^2 + ||\beta_2-\beta_2^r||^2\Big)}_{(D)}$$

where $\beta^r \in \mathcal{G}$ is some value in the grid such that (D) is of order $1/K$. We can then show that $(B)+(C)-\widehat{\Gamma}(\beta_2^r,\beta_1^r)$ converges in probability to zero as n grows which is possible since we use the estimated gradient to construct $\widehat{\Gamma}$. If instead we had not used information on the estimated gradient, the rate would be dominated by (C) which is of order $1/\sqrt{K}$.

However, we note that different from previous sections, the rate $1/K$ is specific to the one-dimensional setting and carry-overs over two consecutive periods. In p dimensions, the rate would be of order $1/K^{2/(p+1)}$ due to the curse of dimensionality.

A.4 Welfare maximization with a non-adaptive experiment using local perturbations

Next, we revisite the non-adaptive experiment in Section 3 and introduce estimators of β^* *without* adaptivity. This sub-section serves two purposes. First, it sheds light on comparisons of the adaptive procedure with grid-search-type methods, showing drawbacks of the grid-search approach in terms of convergence of the regret. Second, it shows how, when an adaptive procedure is not available, we can still use information from the marginal effect estimated as we propose in Algorithm 1, to improve rates of convergence in the number of clusters.

The algorithm that we propose is formally discussed in Algorithm E.4 in Appendix E and works as follows. First, we construct a fine grid $\mathcal{G}$ of the parameter space $\mathcal{B}$ (with p dimensions), with equally spaced parameters under the $l2$-norm. Second, we pair clusters, and we assign a *different* parameter β^k for each pair $(k, k+1)$ from the grid $\mathcal{G}$. Third, in each pair, we estimate the gradient $\hat{V}_{(k,k+1)} \in \mathbb{R}^p$, by perturbing, sequentially for $T = p$ periods, one coordinate at a time of the parameter β^k.[67] We estimate welfare using a first-order Taylor expansion

$$\widehat{W}(\beta) = \bar{W}^{k^*(\beta)} + \hat{V}^{\top}_{(k^*(\beta),k^*(\beta)+1)}(\beta - \beta^{k^*(\beta)}), \quad \hat{\beta}^{ow} = \arg\max_{\beta\in\mathcal{B}} \widehat{W}(\beta), \tag{A.7}$$

[67] Sequentiality here is for notational convenience only, and can be replaced by $T = 1$, but with $2p$ clusters allocated to each coordinate.

where $k^*(\beta) = \arg \min_{k \in \{1,3,\cdots,K-1\}, \beta^k \in \mathcal{G}} ||\beta^k - \beta||^2$, $\bar{W}^k = \frac{1}{2}\Big[\frac{1}{T}\sum_{t=1}^{T} \bar{Y}_t^k - \bar{Y}_0^k + \frac{1}{T}\sum_{t=1}^{T} \bar{Y}_t^{k+1} - \bar{Y}_0^{k+1}\Big]$.

Here, $\bar{Y}_t^k$ is the average outcome in cluster k at time t, and $\hat{V}_{(k^*,k^*+1)}$ is estimated as in Equation (E.3). Also, $\bar{W}^k$ denotes the average outcome, as we *pool* outcomes from two clusters in the same pair $(k, k+1)$. The estimator in Equation (A.7) uses a first-order Taylor expansion around β, using information from the closest element β^k.

We can now characterize guarantees of the estimator as $n \to \infty$, and $K, p < \infty$.

Theorem A.4. *Suppose that $\varepsilon_{i,t}^{(k)}$ is sub-gaussian. Let Assumptions 2.1, 2.2, 3.1 hold. Let $\eta_n = o(n^{-1/4})$. Let $\gamma_N \log(n\gamma_N K)/(\eta_n^2 n) = o(1)$. Consider $\hat{\beta}^{ow}$ as in Algorithm E.4, with $\mathcal{B} \subseteq [0,1]^p$. Then for a constant $\bar{C} < \infty$ independent of (n, T, K),*

$$\lim_{n\to\infty} P\Big(W(\beta^*) - W(\hat{\beta}^{ow}) \leq \frac{\bar{C}}{K^{2/p}}\Big) = 1.$$

The proof is in Appendix C.3.9. Theorem A.4 showcases two properties of the method. First, for $p = 1$, the rate of convergence is of order $1/K^2$, which is possible *because* we also estimate and leverage the gradient $\hat{V}$. Our insight here is to use local perturbations to recover the gradient directly by choosing pairs of points on the grid that are close enough (but not too close, which we control through the perturbation η_n) so that we can recover $\hat{V}$ at a given point consistently as $n \to \infty$, fixing K. We then augment the estimator of the welfare with $\hat{V}$, since, otherwise, the rate would be slower in K.[68] One drawback of a grid search approach is that, as $p > 1$, the method suffers a curse of dimensionality, and the rate in K decreases as p increases. This is different from the adaptive procedure (e.g., Corollary 4), where the rate in K does not depend on p. A second disadvantage of the grid search is that the method does not control the in-sample regret, formalized below.

Proposition A.5 (Non-vanishing in-sample regret)**.** *There exists a strongly concave $W(\cdot)$, such that, for $p = 1$, $W(\beta^*) - \frac{1}{K}\sum_{k=1}^{K} W(\beta^k) \geq c$, for a constant $c > 0$ independent of (n, K, T).*

Proposition A.5 shows that the grid search method performs poorly for the in-sample regret, which is of interest when optimizing participants' welfare (or costs of the experiment), differently from the adaptive procedure in the main text. Similar reasoning can be used for related procedures to the grid search approach.

[68] By a second-order Taylor expansion, using information from the gradient guarantees that $\widehat{W}(\beta)$ converges to $W(\beta)$ up-to a second-order term of order $O(||\beta - \beta^k||^2)$, instead of a first-order term $O(||\beta - \beta^k||)$.

A.5 Inference and estimation with observed cluster heterogeneity

In this subsection, we discuss an extension to allow for cluster heterogeneity. Consider $\theta_k \in \Theta$ to denote the cluster's type for cluster k, where Θ is a finite space (i.e., there are finitely many cluster types). Let θ_k be observable by the researcher and be non-random (i.e., conditions should be interpreted conditional on $\{\theta_k\}_{k=1}^K$). Consider the following assumptions.

Assumption A.4. Let the following holds:

(A) For each cluster k, let Assumption 2.1 holds, with $F_X, F_{U|X}$ replaced by $F_X(\theta_k), F_{U|X}(\theta_k)$ as a function of θ_k;

(B) Assumption 2.2 with $r(\cdot)$ which also depends on θ_k.

Assumption A.4 allows for both the distribution of covariates and unobservables and potential outcomes to also depend on the cluster's type θ_k. We can now state the following lemma.

Lemma A.6. *Under Assumption A.4, under an assignment in Assumption 2.3 with exogenous (i.e., not data-dependent) $\beta_{k,t}$ the following holds:*

$$Y_{i,t}^{(k)} = y\Big(X_i^{(k)}, \beta_{k,t}, \theta_k\Big) + \varepsilon_{i,t}^{(k)} + \alpha_t + \tau_k, \quad \mathbb{E}_{\beta_{k,t}}\Big[\varepsilon_{i,t}^{(k)} | X_i^{(k)}\Big] = 0, \tag{A.8}$$

for some function $y(\cdot)$ unknown to the researcher. In addition, for some unknown $m(\cdot)$, $\mathbb{E}_{\beta_{k,t}}\Big[Y_{i,t}^{(k)} | D_{i,t}^{(k)} = d, X_i^{(k)} = x\Big] = m(d, x, \beta_{k,t}, \theta_k) + \alpha_t + \tau_k$.

Differently from Lemma 2.1, here the the functions also depend on the cluster's type θ_k. The proof of Lemma A.6 follows verbatim from the one of Lemma 2.1, taking here into account also the (deterministic) cluster's type. In the following lines, we discuss the single-wave for inference and multi-wave experiments for estimation under heterogeneity.

Single wave experiment In the context of a single-wave experiment, we are interested in testing the null hypothesis of whether a *class* of decisions $\beta(\theta), \theta \in \Theta$, which depends on the cluster's type, is optimal. Namely, let

$$W\Big(\beta(\theta), \theta\Big) = \int y(x, \beta(\theta), \theta) dF_X(\theta), \quad \beta : \Theta \mapsto \mathcal{B}, \quad \theta \in \Theta$$

be the welfare corresponding to cluster's type θ, for a decision rule $\beta(\theta)$. Also, let

$$V\Big(\beta(\theta), \theta\Big) = \frac{\partial W(b, \theta)}{\partial b}\Big|_{b=\beta(\theta)}$$

be the marginal effect with respect to changing $\beta(\theta)$ (for fixed θ). Our null hypothesis is

$$H_0 : V\Big(\beta(\theta), \theta\Big) = 0, \quad \forall \theta \in \Theta, \tag{A.9}$$

i.e., the (baseline) policy $\beta(\theta)$ is optimal for all clusters under consideration. The algorithm follows similarly to Algorithm 2 with the following modification: instead of matching arbitrary clusters, we construct pairs such that elements in the same pair $(k, k+1)$ are such that $\theta_k = \theta_{k+1}$. We can now state the following corollary.

Corollary 6. *Suppose that for all $x \in \mathcal{X}, d \in \{0,1\}, b \in \mathcal{B}, \theta \in \Theta, \pi(x,b), m(d,x,\beta,\theta)$ are uniformly bounded and twice differentiable with bounded derivatives. Let Assumption 3.2, A.4 hold. Consider Algorithm 2, with parameter $\beta(\theta)$ as a function of the cluster's type and for each pair of clusters $(k, k+1), k$ being odd, being such that $\theta_k = \theta_{k+1}$. Let $\mathcal{T}_n$ as in Algorithm 2. Then for $4 \leq K < \infty$, $\alpha \leq 0.08$,*

$$\lim_{n\to\infty} P\Big(|\mathcal{T}_n| \leq \mathrm{cv}_{K/2-1}(\alpha)\Big| H_0\Big) \geq 1 - \alpha, \tag{A.10}$$

where $\mathrm{cv}_{K/2-1}(h)$ is the size-h critical value of a t-test with $K/2-1$ degrees of freedom, and H_0 is as defined in Equation (A.9).

Corollary 6 states that the proposed algorithm guarantees asymptotic size control also in the presence of cluster heterogeneity, under the null hypothesis that the policy $\beta(\theta)$ as a function of the cluster's type is optimal for all types. Importantly, here we construct a test statistics using information from *all* clusters.

Consistency of the marginal effect follows verbatim as in Theorem 3.1, assuming that clusters of the same type are matched.

Multi-wave experiment In the following lines, we provide details on how our setting extends to the multi-wave experiment. For the multi-wave experiment, our goal is to find $\beta^*(\theta)$ such that

$$\beta^*(\theta) \in \arg\max_{b\in\mathcal{B}} W(b,\theta), \quad \forall \theta \in \Theta.$$

Similarly to the single-wave experiment, clusters (k, k') of the same type $\theta_k = \theta_{k'}$ are first matched together. Algorithm E.3 returns the optimal policy under two alternative conditions and modifications. The first extension consists of grouping clusters of the same type together, estimating separately $\beta^*(\theta)$ for each $\theta \in \Theta$. In this case the regret bound holds up-to a factor of order $\min_t P(\theta = t)$, with $P(\theta = t)$ denoting the (exact) share of clusters of type t. Intuitively, the same analysis carry-over to this case, as we restrict our attention to each subgroup separately. The second approach instead consists of updating the same policy from

a given pair using information from that *same* pair. The validity of the algorithm in this case relies on time independence and fails otherwise.

Appendix B Additional extensions

B.1 Tests with a p-dimensional vector of marginal effects

In the following lines we extend Algorithm 2 to testing the following null

$$H_0 : V^{(j)}(\beta) = 0, \text{ for some } p \geq l \geq 1,$$

where we consider a generic number of dimensions tested l. We introduce the algorithmic procedure in Algorithm B.1.

Algorithm B.1 One wave experiment for inference

Require: Value $\beta \in \mathbb{R}^p$, K clusters, 2 periods of experimentation, number of tests t.

1: Match clusters into pairs $K/2$ pairs with consecutive indexes $\{k, k+1\}$;
2: $t = 0$ (*baseline*):
   a: Treatments are assigned at some baseline β_0 $D_{i,0}^{(h)} \sim \pi(X_i^{(h)}, \beta_0), h \in \{1, \cdots, K\}$ (e.g., none of the individuals is treated).
   b: Collect baseline values: for n units in each cluster observe $Y_{i,0}^{(h)}, h \in \{1, \cdots, K\}$.
3: $t = 1$ (*experimentation-wave*)
4: Assign each pair of clusters $\{k, k+1\}$ to a coordinate $j \in \{1, \cdots, p\}$ (with the same number of pairs to each coordinate)
5: For each pair $\{k, k+1\}$, k is odd, assigned to coordinate j
   a: Randomize

$$D_{i,1}^{(h)} \sim \begin{cases} \pi(X_i^{(h)}, \beta + \eta_n \underline{e}_j) \text{ if } h = k \\ \pi(X_i^{(h)}, \beta - \eta_n \underline{e}_j) \text{ if } h = k+1 \end{cases}, \quad n^{-1/2} < \eta_n \leq n^{-1/4}$$

   b: For n units in each cluster $h \in \{k, k+1\}$ observe $Y_{i,1}^{(h)}$.
   c: Estimate the marginal effect for coordinate j as

$$\widehat{V}_k = \frac{1}{2\eta_n}\left[\bar{Y}_1^{(k)} - \bar{Y}_0^{(k)}\right] - \frac{1}{2\eta_n}\left[\bar{Y}_1^{(k+1)} - \bar{Y}_0^{(k+1)}\right] \tag{B.1}$$

return

$$\widetilde{V}_n = \left[\widehat{V}_1, \widehat{V}_3, \cdots, \widehat{V}_{K-1}\right] \tag{B.2}$$

We define $\mathcal{K}_j$ the set of pairs in Algorithm B.1 used to estimate the j^{th} entry of $V(\beta)$.

Define

$$\bar{V}_n^{(j)} = \frac{2l}{K} \sum_{k \in \mathcal{K}_j} \widehat{V}_k,$$

the average marginal effect for coordinate j estimated from those clusters is used to estimate the effect of the j^{th} coordinate. We construct

$$Q_{j,n} = \frac{\sqrt{K/(2l)}\bar{V}_n^{(j)}}{\sqrt{(K/(2l)-1)^{-1}\sum_{k \in \mathcal{K}_j}(\widehat{V}_k^{(j)} - \bar{V}_n^{(j)})^2}}, \quad \mathcal{T}_n = \max_{j \in \{1,\cdots,l\}} |Q_{j,n}|, \tag{B.3}$$

where $\mathcal{T}_n$ denotes the test statistics. The choice of the l-infinity norm is motivated by its theoretical properties: the statistics $Q_{j,n}$ follows an unknown distribution as a result of possibly heteroskedastic variances of $\widehat{V}_k$ across different clusters. However, the upper-bound on the critical quantiles of the proposed test-statistic for unknown variance attains a simple expression under the proposed test-statistics. From a conceptual stand-point, the proposed test-statistic is particularly suited when a large deviation occurs over one dimension of the vector.[69] We now introduce the following theorem.

Theorem B.1 (Nominal coverage)**.** *Let Assumptions 2.1, 2.2, 3.2 hold. Let* $n^{1/4}\eta_n = o(1), \gamma_N^2/N^{1/4} = o(1)$*,* $K < \infty$*. Let* $K \geq 4l$*,* H_0 *be as defined in Equation (8). For any* $\alpha \leq 0.08$*,*

$$\lim_{n\to\infty} P\Big(\mathcal{T}_n \leq q_\alpha \Big| H_0\Big) \geq 1-\alpha, \text{ where } q_\alpha = \mathrm{cv}_{K/(2l)-1}\Big(1-(1-\alpha)^{1/l}\Big), \tag{B.4}$$

with $\mathrm{cv}_{K/(2l)-1}(h)$ *denotes the critical value of a two-sided t-test with level h with test-statistic having* $K/(2l)-1$ *degrees of freedom.*

The proof is in Appendix C.4.3.

B.2 Out-of-sample regret with strict quasi-concavity

In the following lines, we provide guarantees on the regret bounds for the adaptive algorithm in Section 4 under quasi-concavity. We replace Assumption 4.2 with the following condition.

Assumption B.1 (Local strong concavity and strict quasi-concavity)**.** Assume that the following conditions hold.

(A) For every $\beta, \beta' \in \mathcal{B}$, such that $W(\beta') - W(\beta) \geq 0$, then $V(\beta)^\top(\beta' - \beta) \geq 0$,

[69] Observe that alternatively, we may also consider randomization tests as discussed in Canay et al. (2017). This is omitted for the sake of brevity.

(B) For every $\beta \in \mathcal{B}$, $||V(\beta)||_2 \geq \mu||\beta - \beta^*||_2$, for a positive constant $\mu > 0$;

(C) $\frac{\partial^2 W(\beta)}{\partial \beta^2}\Big|_{\beta=\beta^*}$ has negative eigenvalues bounded away from zero at β^*, with $\beta^* \in \tilde{\mathcal{B}} \subset \mathcal{B}$ being in the interior of $\mathcal{B}$.

Condition (A) imposes a quasi-concavity of the objective function. The condition is *equivalent* to common definitions of quasi concavity (Hazan et al., 2015). Condition (A) holds when increasing the probability of treating more neighbors has decreasing marginal effects (see, e.g., Figure 5). Condition (B) assumes that the marginal effect only vanishes at the optimum, ruling out regions over which marginal effects remain constant at zero. A notion of strict quasi-concavity can be found in Hazan et al. (2015), where the authors assume (A) and that the gradient vanishes at the optimum only. Condition (C) also imposes that the function has negative definite Hessian at β^* only but not necessarily globally. The above restrictions guarantee strong concavity locally at the optimum, but not necessarily globally. We now introduce out-of-sample guarantees in this setting. In such a case, the choice of the learning rate consists of a gradient norm rescaling, as discussed in Remark 7.

Theorem B.2. *Let Assumptions 2.1, 2.2, 4.1, B.1 hold, and choose $\alpha_{k,w}$ as in Equation (15), for arbitrary $v \in (0,1)$, κ as defined in Equation (C.7), and ϵ_n as in Lemma C.12. Take a small $1/4 > \xi > 0$, and let $n^{1/4-\xi} \geq \bar{C}\sqrt{\log(n) p \gamma_N T^2 e^{B_p T} \log(KT)}$, $\eta_n = 1/n^{1/4+\xi}$, for finite constants $\infty > B_p, \bar{C} > 0$. Then for $T \geq \zeta^{1/v}$, for a finite constant $\zeta < \infty$, with probability at least $1 - 1/n$,*

$$W(\beta^*) - W(\hat{\beta}^*) = \mathcal{O}(\check{T}^{-1+v}).$$

The proof is in Appendix C.4.4.

B.3 Non separable fixed effects

In the following lines, we show how we can leverage direct and marginal spillover effects to identify (and then estimate) the marginal effects when fixed effects are non-separable in time and cluster identity.

Theorem B.3 (Marginal effects with non-separable fixed effects)**.** *Let $X = 1$, and suppose that $m(d, 1, \beta)$ is bounded and twice differentiable with bounded derivatives for $d \in \{0, 1\}$. Let Assumptions 2.1 hold. Suppose that fixed-effects are non-separable, with*

$$Y_{i,t}^{(k)} = m(D_{i,t}^{(k)}, 1, \beta) + \alpha_{k,t} + \varepsilon_{i,t}^{(k)}, \quad \mathbb{E}[\varepsilon_{i,t}^{(k)}] = 0, \quad D_{i,t}^{(k)} \sim_{i.i.d.} \text{Bern}(\beta),$$

and $m(1,1,\beta)$ *being a constant function in* β*. Then*

$$\mathbb{E}\Big[\hat{\Delta}_k(\beta) + \hat{S}(0,\beta)(1-\beta) - (1-\beta)\hat{S}(1,\beta)\Big] = V(\beta) + \mathcal{O}(\eta_n).$$

The proof is in Appendix C.4.5. Theorem B.3 shows that we can use the information on the spillover and direct treatment effects to identify the marginal effects in the presence of non-separable time and cluster fixed effects. The theorem leverages the assumption that spillovers only occur on the control individuals but not the treated. We note that in applications where treatment effects have disproportionately large effects on the outcome of an individual, and the treated outcomes do not depend on the neighbors' assignments, there might be some tension with the assumption of fixed effects assumption. Studying economic settings where the assumption does and does not hold is an interesting future direction.

B.4 Experimental design with dynamic treatments and stationary policies

Here, we return to the case of dynamic treatments of Section A.3 and study optimization with stationary decisions.

We note that in some applications, we may be interested in stationary estimands of the following form

$$\beta^{**} \in \arg\sup_{\beta\in\mathcal{B}} \Gamma(\beta,\beta). \tag{B.5}$$

Equation (B.5) defines the vector of parameters that maximizes welfare, *under the constraint* that the decision remains invariant over time. For this case, we propose the following algorithm.

"Patient" gradient descent The algorithm works as follows. We begin our iteration from the starting value β_0, we evaluate $\Gamma(\beta_0,\beta_0)$, and compute its *total* derivative $\nabla(\beta_0)$. We then update the current policy choice in the direction of the total derivative and wait for one more iteration before making the next update. Formally, the first three iteration consists of the following updates:

$$\Gamma(\beta_0,\beta_0) \Rightarrow \Gamma(\beta_0+\nabla(\beta_0),\beta_0) \Rightarrow \Gamma(\beta_0+\nabla(\beta_0),\beta_0+\nabla(\beta_0)).$$

Estimation and updates The estimation procedure follows similarly to Section 4, with a small modifications: for every period t the policy stays enforced for one more period $t+1$, without necessitating that data are collected over the period $t+1$. At the end of all such

iterations, we construct an estimator $\widehat{\beta}_T^{**}$ as in Section 4.

Theorem B.4. *Let Assumption 2.1, 4.1, A.3 hold. Take a small $\xi > 0$. Let $n^{1/4-\xi} \geq \bar{C}\sqrt{\log(n)\gamma_N T^B e^{B\sqrt{pT}}\log(KT)}$, $\eta_n = 1/n^{1/4+\xi}$, for finite constants $\infty > B, \bar{C} > 0$. Let $T \geq \zeta$, for a finite constant $\zeta < \infty$. Let $\beta \mapsto \Gamma(\beta, \beta)$ satisfying strict strong concavity. Then with probability at least $1 - 1/n$, for some $J > 0$, $\alpha_{k,w} = J/w$*

$$||\beta^{**} - \hat{\beta}_T^{**}||^2 = \mathcal{O}(T^{-1}).$$

The proof is in Appendix C.4.6.

B.5 Staggered adoption

In this section, we sketch the experimental design in the presence of staggered adoption, i.e., when treatments are assigned only once to individuals and post-treatment outcomes are collected once. The algorithm works similarly to what was discussed in Section 4 with one small difference: every period, we only collect information from a given clusters' pair and update the policy for the subsequent pair and proceed in an iterative fashion.

The algorithm is included in Algorithm E.3. For simplicity, we only discuss the case where $\beta \in \mathbb{R}$ (single coordinate), while for the case where $\beta \in \mathbb{R}^p$ the algorithm works similarly with the only difference that every perturbation to each coordinate requires a new pair of clusters (with in total $2p$ times the number of iterations many clusters).

Theorem B.5 (In-sample regret)**.** *Let the conditions in Theorem 4.3 hold and let $\beta \in \mathbb{R}$, with $\check{\beta}^t$ estimated as in Algorithm E.3. Then with probability at least $1 - 1/n$,*

$$\frac{1}{T}\sum_{t=1}^{T}\left[W(\beta^*) - W(\check{\beta}^t)\right] \leq \bar{C}\frac{p\log(T)}{T}$$

for a finite constant $\bar{C} < \infty$.

See Appendix C.4.7 for the proof.

B.6 Estimating conditional and interaction effects

In this subsection, we turn to the problem of estimating heterogeneous and interaction effects, which, in practice, researchers may also want to report.[70]

[70] I thank Ariel Pakes for suggesting to also study interaction effects.

Formally, let $X_i^k \in \{0,1\}$, where X_i^k denote the covariate of individual i in cluster k, assumed to be binary for simplicity (while our results directly extend to an arbitrary finite discrete space). We make two additional simplifying assumptions. First, we assume that $P(X_i = x)$ is known to the researcher (if not, it can be replaced by its sample analog). Second, we consider a class of unconditional assignment rules $\pi(x;\beta) = \beta$ for simplicity (recall Definition 2.3). We are interested in computing

$$\frac{\partial m(d,x,\beta)}{\partial \beta}, \quad d,x \in \{0,1\}, \quad \text{and} \quad \frac{\partial m(d,1,\beta)}{\partial \beta} - \frac{\partial m(d,0,\beta)}{\partial \beta}$$

which denote the marginal effect on treated ($d=1$) or untreated ($d=0$), conditional on a given type $x \in \{0,1\}$, and the interaction effect for a given individual treatment, respectively. We provide a simple example below.

Example B.1 (Linear model)**.** To gain further intuition, consider a model where[71]

$$Y_{i,1} = \gamma_0 + \Delta \times D_{i,1} + \gamma_1 \times \frac{1}{|\mathcal{N}_i|}\sum_{j\neq i} A_{i,j}D_{i,1} + \iota \times X_i \frac{1}{|\mathcal{N}_i|}\sum_{j\neq i} A_{i,j}D_{i,1} + \nu_{i,1}, \quad \mathbb{E}[\nu_{i,1}|A,X,D] = 0$$

where $\mathcal{N}_i = \{j : A_{i,j} = 1\}$ denotes the set of friends of individual i, $A_{i,j} \in \{0,1\}$, and ι captures the interaction effect between covariates and spillovers. The equation above describes an outcome model where individual outcomes depend on the share of treated neighbors. In this case, we can write

$$m(d,x,\beta) = \gamma_0 + \Delta d + \gamma_1 \beta + \iota x \beta \Rightarrow \frac{\partial m(d,x,\beta)}{\partial \beta} = \gamma + \iota x.$$

As a result, we have

$$\frac{\partial m(d,1,\beta)}{\partial \beta} - \frac{\partial m(d,0,\beta)}{\partial \beta} = \iota,$$

which denotes the interaction parameter. □

Note that, while under linearity as in Example B.1, $\frac{\partial m(d,1,\beta)}{\partial \beta} - \frac{\partial m(d,0,\beta)}{\partial \beta}$ is constant in the neighbors' treatment probability β, in the absence of linearity the above difference also depends on the parameter β. Here, we recover the conditional marginal effect $\frac{\partial m(d,x,\beta)}{\partial \beta}$, for a given β, without assumptions on the function $m(d,x,\beta)$.

To estimate the derivative $\frac{\partial m(d,x,\beta)}{\partial \beta}$ we follow a similar approach as in Section 3, while, here, we also condition on individual covariates. We construct our estimator taking a

[71] In the equation below, we omit the superscript k here assuming that the model holds in each cluster k.

difference-in-differences between two clusters $(k, k+1)$, appropriately reweighted as follows

$$\widehat{\frac{\partial m(0,x,\beta)}{\partial \beta}} = \frac{1}{2n\eta_n}\sum_{i=1}^{n}\left\{Y_{i,1}^{k} w_i^k(\beta) - Y_{i,1}^{k+1} w_i^{k+1}(\beta)\right\} - \frac{1}{2\eta_n}\left[\bar{Y}_0^k - \bar{Y}_0^{k+1}\right]$$

where $\bar{Y}_0^k$ denotes the average outcome at $t=0$ (baseline) in cluster k, and

$$w_i^h(\beta) = \frac{(1-D_{i,1}^h)}{1-\beta-v_h\eta_n}\frac{1\{X_i^h = x\}}{P(X_i^h = x)}, \quad v_h = \begin{cases} 1 \text{ if } h = k \\ -1 \text{ otherwise.} \end{cases}$$

The estimator is a Horowitz-Thompson type estimator, where we reweight by the treatment probability and by the individual type. We derive consistency below, where we also assume overlap of the propensity score.

Theorem B.6. *Consider treatments assigned as in Algorithm 1 with $\beta, P(X_i^h = x) \in (c, 1-c)$ for some constant $c \in (0,1)$ for all $x \in \{0,1\}$. Suppose that $\varepsilon_{i,t}^{(k)}$ is sub-gaussian. Let Assumptions 2.1, 2.2, 3.1 hold. Then with probability at least $1-\delta$, for any $\delta \in (0,1), x \in \{0,1\}$,*

$$\left|\widehat{\frac{\partial m(0,x,\beta)}{\partial \beta}} - \frac{\partial m(0,x,\beta)}{\partial \beta}\right| = \mathcal{O}\Big(\eta_n + \sqrt{\frac{\gamma_N \log(\gamma_N/\delta)}{n\eta_n^2}}\Big).$$

Corollary 7 (Consistency). *For $\gamma_N \log(\gamma_N)/N^{1/3} = o(1), \eta_n = n^{-1/3}$, $\widehat{\frac{\partial m(0,x,\beta)}{\partial \beta}} \to_p \frac{\partial m(0,x,\beta)}{\partial \beta}$.*

The proof is in Appendix C.4.8.

Theorem B.6 shows that the estimated marginal effect for individuals with type $X = x$ converges to the population counterpart. A similar reasoning also applies as we take a difference $\frac{\partial m(0,x=1,\beta)}{\partial \beta} - \frac{\partial m(0,x=0,\beta)}{\partial \beta}$ to capture an interaction effect as illustrated in Example B.1. Hence, researchers can also report interaction effects of spillovers and different individual types.

We can also conduct inference similarly to what was discussed for the marginal effect in Theorem 3.2. Formally, we can construct pivotal test statistics similar to the one in Equation (10), which, here, depends on the conditional (instead of unconditional) estimates. Guarantees for inference follow similarly to what was discussed for the marginal effect estimator and are omitted for brevity.

B.7 Selection of η_n: rule of thumb

In this subsection we provide a rule of thumb for selecting η_n. Following Theorem 3.1 and following Lemma C.3 which provides exact constants, we can write with probability at least

$1 - 1/n$

$$\left|\hat{V}_{(k,k+1)} - V(\beta)\right| \leq \sqrt{\frac{2\sigma^2\gamma_N \log(2\gamma_N n)}{n\eta_n^2}} + c\eta_n, \quad c = \left|\left|\frac{\partial^2 m(d,x,\beta)}{\partial \beta^2}\right|\right|_\infty,$$

where the l_∞ is taken with respect to each element of the Hessian, x, β.[72] Note that we cannot directly minimize the upper bound since otherwise the bias and variance would converge to zero at the same rate, and we would violate the condition that $\eta_n = o(n^{-1/4})$ used for inference. Instead, we minimize

$$\min_{\eta_n} \sqrt{\frac{2\sigma^2\gamma_N \log(2\gamma_N n)}{n\eta_n^2}} + c\eta_n/s_n^2, \quad s_n = o(1),$$

where s_n penalizes the bias by an $o(1)$ component and chosen below. It follows that

$$\eta_n^2 = \sqrt{\frac{2s_n\sigma^2\gamma_N \log(2\gamma_N n)}{nc}}.$$

Note that under the conditions for inference, we assume that $\gamma_N/n^{1/4} = o(1)$. To also remove the logarithmic terms (which are asymptotically negligible), let $s_n = \gamma_N/(n^{1/4}\log(n\gamma_N))$, assumed to be $o(1)$ for inference. We can then write the solution to the optimization problem as

$$\gamma_N\sqrt{\frac{2\sigma^2}{c}}n^{-5/16} \approx \gamma_N\sqrt{\frac{2\sigma^2}{c}}n^{-1/3}.$$

Here, we can replace σ^2 and c with some out-of-sample estimates of the outcomes' variance and curvature. Since, in practice, researcher may impose small sample upper bound on the bias, our proposal is

$$\eta_n = \begin{cases} \gamma_N\sqrt{\frac{2\sigma^2}{c}}n^{-1/3} & \text{if } \gamma_N\sqrt{\frac{2\sigma^2}{c}}n^{-1/3} \leq B \\ B \text{ otherwise} \end{cases}$$

where Bc denotes an upper bound on the bias of the estimator imposed by the researcher (e.g., $cB = 0.05$). The problem here is that γ_N is unknown. Therefore, whenever the researcher does not have a good guess for γ_N,we recommend choosing $\eta_n = \sqrt{\frac{2\sigma^2}{c}}n^{-1/3}$ (without the term γ_N) which also leads to valid inference, but slightly smaller small-sample bias (and larger small-sample variance) than the optimal choice.

[72]The constants for the upper bound for the variance follow from Lemma C.3, while the component $c\eta_n$ captures the bias obtained from a second-order Taylor expansion to $m(\cdot)$, noting that only one entry of β is perturbed by η_n in opposite directions for two clusters in a pair as discussed in Algorithm 1.

B.8 Extensions to the network formation model

In this subsection we present two simple extensions of the network formation model.

B.8.1 Network which depends also on non-separable shocks

Consider the following equation:

$$(X_i^{(k)}, U_i^{(k)}) \sim_{i.i.d.} F_{U|X}F_X, \quad A_{i,j}^{(k)} = l\Big(X_i^{(k)}, X_j^{(k)}, U_i^{(k)}, U_j^{(k)}, \omega_{i,j}^{(k)}\Big)1\{i_k \leftrightarrow j_k\} \tag{B.6}$$

where

$$\omega_{i,j}^{(k)} \Big| \left\{\omega_{u,v}^{(k)}\right\}_{(u,v)\neq(i,j),(u,v)\neq(j,i)}, X^{(k)}, U^{(k)}, \nu^{(k)} \sim F_\omega. \tag{B.7}$$

Intuitively, Equation (B.6) states that the connections form also based on unobservables $\omega_{i,j}$ which are drawn independently for each pair (i,j) (note that under Equation (B.7) we can have that $\omega_{i,j} = \omega_{j,i}$).

We can now state the following lemma.

Lemma B.7 (Outcomes)**.** *Under Assumption 2.1, with a network formation as in Equation (B.6), Assumption 2.2, under an assignment in Assumption 2.3 with exogenous (i.e., not data-dependent) $\beta_{k,t}$, Lemma C.4 (and so Lemma 2.1) hold.*

The proof is in Appendix C.4.10. Lemma B.7 implies that our results (and derivations) for inference and estimation directly extend with network formation as in Equation (B.6). One exception is the theorem in Section 4.2 which instead holds under some minor modification to the proof, which we omit for brevity.[73]

B.8.2 Local dependence in network formation

In this subsection, we discuss how Lemma 2.1 extends to networks which depend also on others' unobservables.

Consider the following extension:

$$(X_i^{(k)}, U_i^{(k)}) \sim_{i.i.d.} F_{U|X}F_X, \quad A_{i,j}^{(k)} = l\Big(X_i^{(k)}, X_j^{(k)}, U_i^{(k)}, U_j^{(k)}, U_{v:1\{i_k\sim v_k\}=1}, U_{v':1\{j_k\sim v'_k\}=1}\Big)1\{i_k \leftrightarrow j_k\}. \tag{B.8}$$

[73] In particular, the difference is in one step of the proof where we need to show concentration of $\frac{1}{\gamma_N}\sum_{i=1}^N A_{i,j}$ around its expectation, here also taking into account also the component $\omega_{i,j}$. Under the assumption that $\omega_{i,j}$ are independent for all $j \in \{1,\cdots,N\}$ the argument in the proof, i.e., concentration of the edges conditional on (X_i, U_i) directly follows also for this step, since conditional on X_i, U_i, we still obtain independence of $A_{i,j}$, $j \in \{1,\cdots,N\}$, and we can then follow the same argument in the derivation.

Equation (B.8) states that the connection between individual i and j can depend not only on unobservables of i and j, but also unobservables of all the possible connections of both individual i and individual j. The following lemma holds.

Lemma B.8. *Under Assumption 2.1, where the network formation is described in Equation (B.8), Assumption 2.2, under an assignment in Assumption 2.3 with exogenous (i.e., not data-dependent) $\beta_{k,t}$ the following holds:*

$$Y_{i,t}^{(k)} = y\Big(X_i^{(k)}, \beta_{k,t}\Big) + \varepsilon_{i,t}^{(k)} + \alpha_t + \tau_k, \quad \mathbb{E}_{\beta_{k,t}}\Big[\varepsilon_{i,t}^{(k)} | X_i^{(k)}\Big] = 0, \tag{B.9}$$

for some function $y(\cdot)$ unknown to the researcher. In addition, for some unknown $m(\cdot)$, $\mathbb{E}_{\beta_{k,t}}\Big[Y_{i,t}^{(k)} | D_{i,t}^{(k)} = d, X_i^{(k)} = x\Big] = m(d, x, \beta_{k,t}) + \alpha_t + \tau_k$.

Lemma B.8 states that the outcome can be described as the sum of a conditional expectation function and unobservables.[74] All our results then directly follow, where, however, the bound on the maximum degree must hold at a faster rate (i.e., the maximum degree of dependence is now γ_N^2). The proof of Lemma B.8 follows similarly to the one of Lemma C.4.[75] Importantly, here the adjacency matrix can depend on others' unobservables but not on others' *realized* connections.

Appendix C Derivations

C.1 Notation and definitions

First, we introduce conventions and notation. We define $x \lesssim y$ if $x \leq cy$ for a positive constant $c < \infty$. For K many clusters, we say that

$$\lfloor k \rceil = \begin{cases} k \text{ if } k \leq K \\ k - K \text{ otherwise.} \end{cases}$$

We will refer to $\hat{V}_{(k,k+1)}$ as $\hat{V}_k$ for k is odd for short of notation. Also, we define $\check{V}_{k,s} = \hat{V}_{\lfloor k+2 \rceil, s}$.

The following definition introduces the notion of a dependency graph (see also Janson 2004).

[74] Similarly to Lemma 2.1 such unobservables are locally dependent, while here the dependence structure is of order γ_N^2, since individual may depends on unobservables of friends of friends.

[75] Here, the vector of random variables affecting the outcome is also a function of the vector $U_{v':1\{j \sim v'_k\}=1}$ which is identically distributed across individuals j, and the proof follows similarly as the one of Lemma C.4.

Definition C.1 (Dependency graph)**.** For given random variables $R_1, \cdots, R_n, W_n \in \{0,1\}^{n\times n}$ is a non-random matrix defined as dependecy graph of $(R_1, \cdots, R_n)$ if, for any i, $R_i \perp R_{j:W_n^{(i,j)}=0}$. We denote the dependency neighbors $N_i = \{j : W_n^{(i,j)} = 1\}$. □

Intuitively, a dependency graph denotes a deterministic adjacency matrix with entry (i, j) equal to one if (i, j) are statistical dependent.

Definition C.2. (Proper Cover) Given an adjacency matrix A_n, with n rows and columns, a family $\mathcal{C}_n = \{\mathcal{C}_n(j)\}_j$ of disjoint subsets of $\{1, \cdots, n\}$ is a proper cover of A_n if $\cup_j \mathcal{C}_n(j) = \{1, \cdots, n\}$ and $\mathcal{C}_n(j)$ contains units such that for any pair of elements $\{i, k \in \mathcal{C}_n(j)\}$, then $A_n^{(i,k)} = 0$. □

Namely, a proper cover of A_n defines a set of disjoint sets, where each disjoint set contains some indexes of units that are not neighbors in A_n. Note that a proper cover always exists, since, if A_n is fully connected, then the number of disjoint sets is just n, one for each element.

The size of the smallest proper cover is the chromatic number, defined as $\chi(A_n)$.

Definition C.3. (Chromatic Number) The chromatic number $\chi(A_n)$, denotes the size of the smallest proper cover of A_n. □

In the following lines we define the oracle descent procedure absent of sampling error. Let $\beta \in \mathcal{B} = [\mathcal{B}_1, \mathcal{B}_2]^p$, where $\mathcal{B}_1, \mathcal{B}_2$ are finite. Also, let $P_{\mathcal{B}_1,\mathcal{B}_2}$ be the projection operator onto $\mathcal{B}$.

Definition C.4 (Oracle gradient descent under strong concavity)**.** We define

$$\beta_w^{**} = P_{\mathcal{B}_1,\mathcal{B}_2}\left[\beta_{w-1}^{**} + \alpha_{w-1} V(\beta_{w-1}^{**})\right], \quad \beta_1^{**} = \beta_0, \tag{C.1}$$

with $\alpha_w = \frac{\eta}{w+1}$, equal for all clusters. □

Note that in our proofs, we will refer to the general p-dimensional case for the multi-wave experiment, which uses $\check{T} = T/p$ waves. See Algorithm E.1.

C.2 Lemmas

C.2.1 Preliminary lemmas

Lemma C.1. *(Ross et al., 2011) Let $X_1, ..., X_n$ be random variables such that $\mathbb{E}[X_i^4] < \infty$, $\mathbb{E}[X_i] = 0$, $\sigma^2 = \mathrm{Var}(\sum_{i=1}^n X_i)$ and define $W = \sum_{i=1}^n X_i/\sigma$. Let the collection $(X_1, ..., X_n)$*

have dependency neighborhoods N_i, $i = 1, ..., n$ *and also define* $D = \max_{1 \leq i \leq n} |N_i|$. *Then for* Z *a standard normal random variable, we obtain*

$$d_W(W, Z) \leq \frac{D^2}{\sigma^3} \sum_{i=1}^{n} \mathbb{E}|X_i|^3 + \frac{\sqrt{28} D^{3/2}}{\sqrt{\pi}\sigma^2} \sqrt{\sum_{i=1}^{n} \mathbb{E}[X_i^4]}, \tag{C.2}$$

where d_W *denotes the Wasserstein metric.*

Lemma C.2. *(Brook's Theorem,Brooks (1941)) For any connected undirected graph* G *with maximum degree* Δ, *the chromatic number of* G *is at most* Δ *unless* G *is a complete graph or an odd cycle, in which case the chromatic number is* $\Delta + 1$.

C.2.2 Concentration for local dependency graphs

In the following lemma we study concentration of the average of locally dependent random variables (see also Janson 2004 for concentration with local dependency graphs) (we provide exact constants which are useful to derive the rule of thumb for η_n in Appendix B.7).

Lemma C.3 (Concentration for dependency graphs)**.** *Define* $\{R_i\}_{i=1}^n$ *sub-gaussian random variables with parameter* $\sigma^2 < \infty$, *forming a dependency graph with adjacency matrix* A_n *with maximum degree bounded by* γ_N. *Then with probability at least* $1 - \delta$, *for any* $\delta \in (0, 1)$,

$$\left| \frac{1}{n} \sum_{i=1}^{n} (R_i - \mathbb{E}[R_i]) \right| \leq \sqrt{\frac{2\sigma^2 \gamma_N \log(2\gamma_N/\delta)}{n}}.$$

Proof of Lemma C.3. First, we construct a proper cover $\mathcal{C}_n$ as in Definition C.2, with chromatic number $\chi(A_n)$. We can write

$$\left| \frac{1}{n} \sum_{i=1}^{n} (R_i - \mathbb{E}[R_i]) \right| \leq \sum_{\mathcal{C}_n(j) \in \mathcal{C}_n} \underbrace{\left| \frac{1}{n} \sum_{i \in \mathcal{C}_n(j)} (R_i - \mathbb{E}[R_i]) \right|}_{(A)}.$$

Here, we sum over each subset of index $\mathcal{C}_n(j) \in \mathcal{C}_n$ in the proper cover, and then we sum over each element in the subset $\mathcal{C}_n(j)$. Observe now that by definition of the dependency graph, components in (A) are mutually independent. Using the Chernoff's bound (Wainwright, 2019), we have that with probability at least $1 - \delta$, for any $\delta \in (0, 1)$

$$\left| \sum_{i \in \mathcal{C}_n(j)} (R_i - \mathbb{E}[R_i]) \right| \leq \sqrt{2\sigma^2 |\mathcal{C}_n(j)| \log(2/\delta)},$$

where $|\mathcal{C}_n(j)|$ denotes the number of elements in $\mathcal{C}_n(j)$. As a result, using the union bound, we obtain that with probability at least $1-\delta$, for any $\delta \in (0,1)$

$$\Big|\frac{1}{n}\sum_{i=1}^{n}(R_i - \mathbb{E}[R_i])\Big| \leq \underbrace{\frac{1}{n}\sum_{\mathcal{C}_n(j)\in\mathcal{C}_n}\sqrt{2\sigma^2|\mathcal{C}_n(j)|\log(2\chi(A_n)/\delta)}}_{(B)}.$$

Using concavity of the square-root function, after multiplying and dividing (B) by $\chi(A_n)$, we have

$$\begin{aligned}(B) &\leq \frac{1}{n}\chi(A_n)\sqrt{2\sigma^2\frac{1}{\chi(A_n)}\sum_{\mathcal{C}_n(j)\in\mathcal{C}_n}|\mathcal{C}_n(j)|\log(2\chi(A_n)/\delta)} \\ &= \frac{1}{n}\sqrt{2\sigma^2\chi(A_n)n\log(\chi(A_n)/\delta)}.\end{aligned}$$

The last equality follows by the definition of proper cover. The final result follows by Lemma C.2. □

C.2.3 Proof of Lemma 2.1 and local dependence

Lemma 2.1 is stated as a corollary of Lemma C.4.

Lemma C.4. *Let Assumption 2.1, 2.2 hold. For treatment assigned as in Assumption 2.3 with exogenous parameter $\beta_{k,t}$ in cluster k at time t, Lemma 2.1 hold. Also, $\varepsilon_{i,t}^{(k)} \perp \{\varepsilon_{j,t}^{(k)}\}_{j\notin\mathcal{I}_i^{(k)}}|\beta_{k,t}$ for a set $|\mathcal{I}_i^{(k)}| = \mathcal{O}(\gamma_N)$.*

Proof of Lemma C.4. Under Assumption 2.2, we can write for some function g,

$$r\Big(D_{i,t}^{(k)}, D_{j:A_{i,j}^{(k)}>0,t}^{(k)}, X_i^{(k)}, X_{j:A_{i,j}^{(k)}>0}^{(k)}, A_i^{(k)}, U_i^{(k)}, U_{j:A_{i,j}^{(k)}>0}^{(k)}, \nu_{i,t}^{(k)}\Big) = g(Z_{i,t}^{(k)}).$$

Here, $Z_{i,t}^{(k)}$ depends on $A_i^{(k)}$, i.e., the edges of individual i, and on unobservables and observables of all those individuals such that $A_{i,j}^{(k)} > 0$, namely,

$$Z_{i,t}^{(k)} = \Big[D_{i,t}^{(k)}, X_i^{(k)}, \nu_{i,t}^{(k)}, A_i^{(k)} \otimes \Big(X^{(k)}, U^{(k)}, D_t^{(k)}\Big)\Big].$$

Importantly, under Assumption 2.1, $A_i^{(k)}$ is a function of $\Big\{\Big[X_j^{(k)}, U_j^{(k)}\Big], j : 1\{i_k \leftrightarrow j_k\} = 1\Big\}$, only, and each entry depends on (X_j, U_j, X_i, U_i) through the same function f for each individual. What is important, is that $\sum_j 1\{i_k \leftrightarrow j_k\} = \gamma_N^{1/2}$ for each unit i. Therefore, for some function $\tilde{g}$ (which depends on f in Assumption 2.1), we can equivalently write

$$Z_{i,t}^{(k)} = \tilde{g}(D_{i,t}^{(k)}, \nu_{i,t}^{(k)}, X_i^{(k)}, \tilde{Z}_{i,t}^{(k)}), \quad \tilde{Z}_{i,t}^{(k)} = \Big\{\Big[X_j^{(k)}, U_j^{(k)}, D_{j,t}^{(k)}\Big], j : 1\{i_k \leftrightarrow j_k\} = 1\Big\},$$

where $\tilde{Z}_{i,t}^{(k)}$ is the vector of $\left[X_j^{(k)}, U_j^{(k)}, D_{j,t}^{(k)}\right]$ of all individuals j with $1\{i_k \leftrightarrow j_k\} = 1$.

Now, observe that since $(U_i^{(k)}, X_i^{(k)}) \sim_{i.i.d.} F_{X|U}F_U$, and $\{\nu_{i,t}\}$ are $i.i.d.$ conditionally on $U^{(k)}, X^{(k)}$ (Assumption 2.2) and treatments are randomized independently (Assumption 2.3), we have

$$\left[X_j^{(k)}, U_j^{(k)}, \nu_{j,t}^{(k)}, D_{j,t}^{(k)}\right] \Big| \beta_{k,t} \sim_{i.i.d.} \mathcal{D}(\beta_{k,t})$$

is $i.i.d$ with some distribution $\mathcal{D}(\beta_{k,t})$ which only depends on the exogenous coefficient $\beta_{k,t}$ governing the distribution of $D_{i,t}^{(k)}$ under Definition 2.3. As a result for $\beta_{k,t}$ being exogenous, Lemma 2.1 holds since $\sum_j 1\{i_k \leftrightarrow j_k\} = \gamma_N^{1/2}$ for all i, hence $\tilde{Z}_{i,t}$ are identically distributed across units i.

Similarly, also $\varepsilon_{i,t}^{(k)} | \beta_{k,t}$ is a measurable function of a vector $\left[X_j^{(k)}, U_j^{(k)}, \nu_{j,t}^{(k)}, D_{j,t}^{(k)}\right]_{j:1\{i_k \leftrightarrow j_k\}=1}$.[76] As a result since such vectors are independent conditional on $\beta_{k,t}$, $\varepsilon_{i,t}^{(k)}$ is mutally independent with $\varepsilon_{v,t}^{(k)}$ for all v such that they do not share a common element $\left[X_j^{(k)}, U_j^{(k)}, \nu_{j,t}^{(k)}, D_{j,t}^{(k)}\right]$, that is, such that $\max_j 1\{i_k \leftrightarrow j_k\}1\{v_k \leftrightarrow j_k\} = 0$.

There are at most $\gamma_N^{1/2} + \gamma_N$ many of $\varepsilon_{v,t}^{(k)}$ which can share a common neighbor with $\varepsilon_{i,t}^{(k)}$ ($\gamma_N^{1/2}$ many neighhbors and γ_N many neighbors of the neighbors), which concludes the proof. □

C.2.4 Concentration of the average outcomes

In this subsection, we provide three auxiliary lemmas.

Lemma C.5. *Suppose that treatments are assigned as in Assumption 2.3 with*

$$\begin{aligned} D_{i,0}^{(k)} &\sim \pi(X_i^{(k)}, \beta_0), \quad D_{i,0}^{(k+1)} \sim \pi(X_i^{(k+1)}, \beta_0) \\ D_{i,t}^{(k)} &\sim \pi(X_i^{(k)}, \beta), \quad D_{i,t}^{(k+1)} \sim \pi(X_i^{(k+1)}, \beta') \end{aligned}$$

with exogenous parameters β_0, β, β'. Let Assumption 2.1, 2.2 hold. Then with probability at least $1-\delta$, for any $\delta \in (0,1)$

$$\left|\bar{Y}_t^{(k)} - \bar{Y}_t^{(k+1)} - \bar{Y}_0^{(k)} + \bar{Y}_0^{(k+1)} - \int (y(x,\beta) - y(x,\beta'))dF_X(x)\right| = \mathcal{O}(\sqrt{\frac{\gamma_N \log(\gamma_N/\delta)}{n}}).$$

Proof of Lemma C.5. First, note that by Lemma C.4, we can write

$$\begin{aligned} \mathbb{E}\left[\bar{Y}_t^{(k)} - \bar{Y}_t^{(k+1)}\right] &= \int (y(x,\beta) - y(x,\beta'))dF_X(x) + \tau_k - \tau_{k+1} \\ \mathbb{E}\left[\bar{Y}_0^{(k)} - \bar{Y}_0^{(k+1)}\right] &= \tau_k - \tau_{k+1}. \end{aligned} \tag{C.3}$$

[76] Here for notational convenience convenience only, we are letting $1\{i_k \leftrightarrow i_k\} = 1$.

In addition, by Lemma C.4, $\varepsilon_{i,t}^{(k)}$ (and so $Y_{i,t}^{(k)}$) form a dependency graph with maximum degree bounded by γ_N. The proof completes by invoking Lemma C.3. □

Lemma C.6. *Let $y(x,\beta)$ be twice differentiable with uniformly bounded derivatives for all $x \in \mathcal{X}, \beta \in \mathcal{B}$. Then for all $\beta \in \mathcal{B}$, where $\mathcal{B}$ is a compact space*

$$\begin{aligned}\int \Big[y(x,\beta+\eta_n \underline{e}_j) - y(x,\beta-\eta_n \underline{e}_j)\Big] dF_X(x) &= 2\eta_n \int \frac{\partial y(x,\beta)}{\partial \beta^j} dF_X(x) + \mathcal{O}(\eta_n^2) \\ &= 2\eta_n V^{(j)}(\beta) + \mathcal{O}(\eta_n^2).\end{aligned}$$

Proof of Lemma C.6. We can write

$$\begin{aligned} y(x,\beta+\eta_n \underline{e}_j) &= y(x,\beta) + \frac{\partial y(x,\beta)}{\partial \beta^j}\eta_n + \mathcal{O}(\eta_n^2) \\ y(x,\beta-\eta_n \underline{e}_j) &= y(x,\beta) - \frac{\partial y(x,\beta)}{\partial \beta^j}\eta_n + \mathcal{O}(\eta_n^2) \end{aligned}$$

from the mean-value theorem which guarantees that the first equality holds. The second equality holds by the dominated convergence theorem. □

Lemma C.7. *Let the conditions in Lemma C.5 hold. Let $y(x,\beta)$ be twice differentiable in β with uniformly bounded derivatives for all $x \in \mathcal{X}, \beta \in \mathcal{B}$. Suppose that $\beta = \check{\beta} + \eta_n \underline{e}_j$ and $\beta' = \check{\beta} - \eta_n \underline{e}_j$. Then with probability at least $1-\delta$, for any $\delta \in (0,1)$*

$$\left|\frac{\bar{Y}_t^{(k)} - \bar{Y}_t^{(k+1)} - \bar{Y}_0^{(k)} + \bar{Y}_0^{(k+1)}}{2\eta_n} - V^{(j)}(\check{\beta})\right| = \mathcal{O}\Big(\sqrt{\frac{\gamma_N \log(\gamma_N/\delta)}{\eta_n^2 n}} + \eta_n\Big)$$

Proof of Lemma C.7. Using Lemma C.5 and the triangular inequality, with probability at least $1-\delta$,

$$\begin{aligned}&\left|\frac{\bar{Y}_t^{(k)} - \bar{Y}_t^{(k+1)} - \bar{Y}_0^{(k)} + \bar{Y}_0^{(k+1)}}{2\eta_n} - V^{(j)}(\check{\beta})\right| \\ &\leq \left|\frac{1}{2\eta_n}\int (y(x,\check{\beta}+\eta_n \underline{e}_j) - y(x,\check{\beta}-\eta_n \underline{e}_j)) dF_X(x) - V^{(j)}(\check{\beta})\right| + \mathcal{O}\Big(\sqrt{\frac{\gamma_N \log(\gamma_N/\delta)}{\eta_n^2 n}}\Big).\end{aligned}$$

The component

$$\int (y(x,\check{\beta}+\eta_n \underline{e}_j) - y(x,\check{\beta}-\eta_n \underline{e}_j)) dF_X(x) = 2\eta_n V^{(j)}(\check{\beta}) + \mathcal{O}(\eta_n^2),$$

by Lemma C.6. □

C.2.5 Proof of Lemma 4.1

To prove the claim it suffices to show that $\check{\beta}_k^w$ is independent of potential outcomes and covariates in cluster k for all $w \in \{1, \cdots, \check{T}\}$, since $\beta_{k,t}$ is a deterministic function of some coefficient $\check{\beta}_k^w$ (see Algorithm E.1). Take k to be odd. To show that the claim holds it suffices to show that $\check{\beta}_k^w$ is a function of observables and unobservables only of those units in clusters $k' \notin \{k, k+1\}$. The recursive claim that we want to prove is the following: for all w, $\check{\beta}_k^w$ is exogenous with respect to potential outcomes and covariates in clusters with index $\{h > \lfloor k + 2w + 1 \rceil$ or $h \in \{k, k+1\}\}$. Clearly, for $\check{\beta}_k^1$ the lemma holds, since $\check{\beta}_k^1$ depends on the gradient in the pair $\{\lfloor k+2 \rceil, \lfloor k+3 \rceil\}$ only. Suppose that the lemma holds for all $w \leq \check{T} - 1$. Then consider $\check{\beta}_k^{\check{T}}$. Observe that $\check{\beta}_k^{\check{T}}$ is chosen based on the gradient $\hat{V}_{k+2,\check{T}-1}$ estimated in the previous wave in clusters $\{\lfloor k+2 \rceil, \lfloor k+3 \rceil\}$, and $\check{\beta}_k^{\check{T}-1}$. By the recursive algorithm, $\check{\beta}_k^{\check{T}-1}$ is exogenous with respect to covariates and potential outcomes in clusters with index $\{h > \lfloor k + 2\check{T} - 1 \rceil$ or $h \in \{k, k+1\}\}$, which is possible since $K \geq 2\check{T}$, hence $\lfloor k + 2\check{T} - 1 \rceil < k$. We only need to prove exogeneity of $\hat{V}_{k+2,\check{T}-1}$. The gradient estimated $\hat{V}_{k+2,\check{T}-1}$ is a function of the unobservables and observables at any time $t \leq T$ (where $T = \check{T}p$) in clusters $\{\lfloor k+2 \rceil, \lfloor k+3 \rceil\}$ and the policy $\check{\beta}_{k+2}^{\check{T}-1}$. Since $K \geq 2\check{T}$, again by the recursive algorithm $\check{\beta}_{k+2}^{\check{T}-1}$ is exogenous with respect to potential outcomes and covariates in clusters with index $\{h \geq \lfloor k + 2\check{T} \rceil$ or $h \in \{k, k+1\}\}$ which completes the proof.

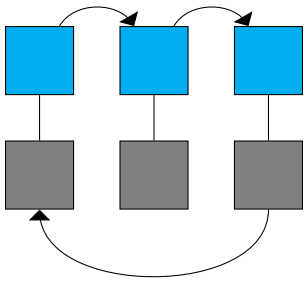


Figure C.1: Idea of the proof. Let $p = 1$. Since we have three clusters pairs (each pair of boxes), by assumption $T = 2$. Then the treatments at $T = 2$ in the first pair are assigned using information from the second pair at $T = 1$. Treatments in the second pair at $T = 1$, depend on information at $T = 0$ in the third pair. Hence, the parameter used at $T = 2$ in the first pair must be independent of covariates and potential outcomes in the first pair of clusters. The same reasoning applies to the other pairs of clusters.

C.2.6 Lemmas for the adaptive experiment with strong concavity

In this section, we discuss theoretical guarantees of the algorithm, assuming the global strong concavity of the objective function $W(\beta)$. The following lemma follows by standard

properties of the gradient descent algorithm (Bottou et al., 2018).

Lemma C.8. *For the learning rate as* $\alpha_w = \eta/(w+1)$*, and* β_w^{**} *as defined in Equation (C.6), under Assumption 3.1, 4.1, 4.2, with* σ*-strong concavity, for* $M \geq \eta \geq 1/\sigma$*, for any* $M \in [1/\sigma, \infty)$*, and let* $L = \max\{2(\mathcal{B}_2 - \mathcal{B}_1)^2, G^2M^2\}$*, with* $G = \sup_\beta ||\frac{\partial W(\beta)}{\partial \beta}||_\infty$*. Then the following holds:*

$$||\beta_w^{**} - \beta^*||^2 \leq \frac{Lp}{w}$$

for a constant $L < \infty$*.*

Proof of Lemma C.8. The proof follows standard arguments of the gradient descent method (Bottou et al., 2018), where, here, we leverage strong concavity and the assumption that the gradient is uniformly bounded. Denote β^* the estimand of interest and recall the definition of β_w^{**} in Equation (C.6). We define ∇_{w-1} the gradient evaluated at β_{w-1}^{**}. From strong concavity, we can write

$$W(\beta^*) - W(\beta_w^{**}) \leq \frac{\partial W(\beta_w^{**})}{\partial \beta}(\beta^* - \beta_w^{**}) - \frac{\sigma}{2}||\beta^* - \beta_w^{**}||_2^2$$
$$W(\beta_w^{**}) - W(\beta^*) \leq \frac{\partial W(\beta^*)}{\partial \beta}(\beta_w^{**} - \beta^*) - \frac{\sigma}{2}||\beta^* - \beta_w^{**}||_2^2.$$

As a result, since $\frac{\partial W(\beta^*)}{\partial \beta} = 0$, we have

$$\Big(\frac{\partial W(\beta^*)}{\partial \beta} - \frac{\partial W(\beta_w^{**})}{\partial \beta}\Big)(\beta^* - \beta_w^{**}) = \frac{\partial W(\beta_w^{**})}{\partial \beta}(\beta^* - \beta_w^{**}) \geq \sigma||\beta_w^{**} - \beta^*||_2^2. \tag{C.4}$$

In addition, we can write:

$$||\beta_w^{**} - \beta^*||_2^2 = ||\beta^* - P_{\mathcal{B}_1,\mathcal{B}_2}(\beta_w^{**} + \alpha_{w-1}\nabla_{w-1})||_2^2 \leq ||\beta^* - \beta_w^{**} - \alpha_{w-1}\nabla_{w-1}||_2^2$$

where the last inequality follows from the fact that $\beta^* \in [\mathcal{B}_1, \mathcal{B}_2]^p$. Observe that we have

$$||\beta^* - \beta_w^{**}||_2^2 \leq ||\beta^* - \beta_{w-1}^{**}||_2^2 - 2\alpha_{w-1}\nabla_{w-1}(\beta^* - \beta_{w-1}^{**}) + \alpha_{w-1}^2||\nabla_{w-1}||_2^2.$$

Using Equation (C.4), we can write

$$||\beta_{w+1}^{**} - \beta^*||_2^2 \leq (1 - 2\sigma\alpha_w)||\beta_w^{**} - \beta^*||_2^2 + \alpha_w^2 G^2 p.$$

We now prove the statement by induction. Clearly, at time $w = 1$, the statement trivially

holds. Consider a general time w. Then using the induction argument, we write

$$||\beta^{**}_{w+1} - \beta^*||_2^2 \leq (1 - 2\frac{1}{w+1})\frac{Lp}{w} + \frac{Lp}{(w+1)^2} \leq (1 - 2\frac{1}{w+1})\frac{Lp}{w} + \frac{Lp}{w(w+1)}$$
$$= (1 - \frac{1}{w+1})\frac{Lp}{w} = \frac{Lp}{w+1}.$$

□

Lemma C.9. *Let Assumption 2.2, 2.1, 4.1 hold. Let α_w be as defined in Lemma C.8. Then with probability at least $1-\delta$, for any $\delta \in (0,1)$, for all $w \geq 1$,*

$$\Big|\Big|P_{\mathcal{B}_1,\mathcal{B}_2-\eta_n}\Big[\sum_{s=1}^{w}\alpha_s\check{V}_{k,s}\Big] - P_{\mathcal{B}_1,\mathcal{B}_2}\Big[\sum_{s=1}^{w}\alpha_s V(\beta^{**}_s)\Big]\Big|\Big|_\infty = \mathcal{O}(P_{\check{T}}(\delta))$$

where $P_1(\delta) = \alpha_1 \times \mathrm{err}(\delta)$ and $P_w(\delta) = B_p\alpha_w P_{w-1}(\delta) + P_{w-1}(\delta) + \alpha_w \mathrm{err}_w(\delta)$, for a finite constant $B_p < \infty$, and $\mathrm{err}_w(\delta) = \mathcal{O}\Big(\sqrt{\gamma_N \frac{\log(p\check{T}K/\delta)}{\eta_n^2 n}} + p\eta_n\Big)$.

Proof of Lemma C.9. Recall that by Lemma C.7 we can write for every k and $w \in \{1, \cdots, \check{T}\}$ (here using the union bound),

$$\check{V}^{(j)}_{k,w} = V^{(j)}(\check{\beta}^w_{k+2}) + \mathcal{O}\Big(\sqrt{\gamma_N \frac{\log(K\check{T}/\delta)}{\eta_n^2 n}} + \eta_n\Big).$$

We now proceed by induction. We first prove the statement, assuming that the constraint is always attained. We then discuss the case of the constraint not being attained. Define (where we suppress the dependence with p for simplicity)

$$B = p\sup_\beta \Big|\Big|\frac{\partial^2 W(\beta)}{\partial \beta^2}\Big|\Big|_\infty.$$

Unconstrained case Consider $w = 1$. Then since we initialize parameters at β_0 (recall that $\beta_0 = \beta^{**}_1$), for all clusters, we can write with probability $1-\delta$, for any $\delta \in (0,1)$,

$$\Big|\Big|\alpha_1\check{V}_{k,1} - \alpha_1 V(\beta_0)\Big|\Big|_\infty = \alpha_1 \mathrm{err}(\delta).$$

Consider $t = 2$, then we obtain for every $j \in \{1, \cdots, p\}$,

$$\alpha_2\check{V}^{(j)}_{k,2} = \alpha_2 V^{(j)}(\check{\beta}^2_{k+2}) + \alpha_2\mathrm{err}(\delta) = \alpha_2 V^{(j)}(\beta^{**}_1 + \alpha_1 V(\beta^{**}_1) + \alpha_1\check{V}_{k,w} - \alpha_1 V(\beta^{**}_1)) + \alpha_2\mathrm{err}(\delta).$$

Using the mean value theorem and Assumption 3.1, we obtain

$$\left|\left|\alpha_2\check{V}_{k,2} - \alpha_2 V(\beta_2^{**})\right|\right|_\infty \leq \alpha_2 \text{err}(\delta) + B\alpha_2\alpha_1\text{err}(\delta)$$
$$\Rightarrow \left|\left|\sum_{w=1}^{2}\alpha_w\check{V}_{k,w} - \sum_{w=1}^{2}\alpha_w V(\beta_w^{**})\right|\right|_\infty \leq \alpha_2\text{err}(\delta) + B\alpha_2\alpha_1\text{err}(\delta) + \alpha_1\text{err}(\delta).$$

Consider now a general w. Then we can write with probability $1-\delta$, for any $\delta \in (0,1)$,

$$\alpha_w\check{V}_{k,w} = \alpha_w V(\check{\beta}_{k+2}^{w-1}) + \alpha_w\text{err}(\delta).$$

Let $\tilde{P}_w^{(j)}(\delta) = \alpha_w\tilde{P}_{w-1}^{(j)}(\delta) + \tilde{P}_{w-1}^{(j)}(\delta) + \alpha_w\text{err}(\delta)$, with $\tilde{P}_1^{(j)}(\delta) = \alpha_1\text{err}(\delta)$, the cumulative error for the jth coordinate. Then, recursively, we have (here, $\tilde{P}_{w-1}(\delta)$ is the vector of cumulative errors)

$$\alpha_w\check{V}_{k,w} = \alpha_w V(\beta_w^{**} + \tilde{P}_{w-1}(\delta)) + \alpha_w\text{err}(\delta).$$

Using the mean value theorem and Assumption 3.1, we obtain

$$\alpha_w\check{V}_{k,w} = \alpha_w V(\beta_w^{**}) + \alpha_w B\max_j \tilde{P}_{w-1}^{(j)}(\delta) + \alpha_w\text{err}(\delta).$$

Therefore, with probability $1 - w\tilde{\delta}$ (using the union bound)

$$\left|\left|\sum_{s=1}^{w}\alpha_s\check{V}_{k,s} - \sum_{s=1}^{w}\alpha_s V(\beta_s^{**})\right|\right|_\infty \leq \left|\left|\alpha_w\check{V}_{k,w} - \alpha_w V(\beta_w^{**})\right|\right|_\infty + \left|\left|\sum_{s=1}^{w-1}\alpha_s\check{V}_{k,s} - \sum_{s=1}^{w-1}\alpha_s V(\beta_s^{**})\right|\right|_\infty$$
$$\leq \alpha_w B P_{w-1}(\tilde{\delta}) + \alpha_w\text{err}(\tilde{\delta}) + P_{w-1}(\tilde{\delta}),$$

where $P_{w-1}(\tilde{\delta})$ captures the largest cumulative error up-to iteration $w-1$ as defined in the statement of the lemma (the log-term as a function of p follows from the union bound). The proof completes once we write $\delta = \tilde{\delta}/w$.

Constrained case Since the statement is true for $w = 1$, we can assume that it is true for all $s \leq w-1$ and prove the statement by induction. Since $\mathcal{B}$ is a compact space, we can

write,

$$\begin{aligned}
&\Big|\Big|P_{\mathcal{B}_1,\mathcal{B}_2-\eta_n}\Big[\sum_{s=1}^{w}\alpha_s\check{V}_{k,s}\Big]-P_{\mathcal{B}_1,\mathcal{B}_2}\Big[\sum_{s=1}^{w}\alpha_s V(\beta_s^{**})\Big]\Big|\Big|_\infty \\
&\leq \Big|\Big|P_{\mathcal{B}_1,\mathcal{B}_2-\eta_n}\Big[\sum_{s=1}^{w}\alpha_s\check{V}_{k,s}\Big]-P_{\mathcal{B}_1,\mathcal{B}_2-\eta_n}\Big[\sum_{s=1}^{w}\alpha_s V(\beta_s^{**})\Big]\Big|\Big|_\infty+\mathcal{O}(p\eta_n) \\
&\leq 2\Big|\Big|\sum_{s=1}^{w}\alpha_s\check{V}_{k,s}-\sum_{s=1}^{w}\alpha_s V(\beta_s^{**})\Big|\Big|_\infty+\mathcal{O}(p\eta_n)
\end{aligned}$$

completing the proof. □

Lemma C.10. *Let the conditions in Lemma C.9 hold. Then with probability at least* $1-\delta$*, for any* $\delta\in(0,1)$*, for all* $w\geq 1, k\in\{1,\cdots,K\}$*,*

$$||\beta^*-\check{\beta}_k^w||_2^2\leq\frac{Lp}{w}+pw^{2B_p}\times\mathcal{O}\Big(\gamma_N\frac{\log(p\check{T}K/\delta)}{\eta_n^2 n}+p^2\eta_n^2\Big),$$

for finite constants $B_p, L<\infty$*.*

Proof. Using the triangular inequality, we can write

$$||\beta^*-\check{\beta}_k^w||_2^2\leq||\beta^*-\beta_w^{**}||_2^2+||\check{\beta}_k^w-\beta_w^{**}||_2^2.$$

The first component on the right-hand side is bounded by Lemma C.8. Using Lemma C.9, we bound the second component with probability at least $1-\delta$, as follows

$$||\check{\beta}_k^w-\beta_w^{**}||_2^2\leq p||\check{\beta}_k^w-\beta_w^{**}||_\infty^2=p\times\mathcal{O}(P_w^2(\delta)).$$

We conclude the proof by explicitely characterizing the rate of $P_w(\delta)$ as defined in Lemma C.9. Following Lemma C.9, we can define recursively $P_w(\delta)$ for any $1\leq w\leq\check{T}$ (recall that $\alpha_w\propto 1/w$) as

$$P_w(\delta)=(1+\frac{B}{w})P_{w-1}(\delta)+\frac{1}{w}\text{err}_n(\delta),\quad P_1(\delta)=\text{err}_n(\delta).$$

where $\text{err}_n=\mathcal{O}(\sqrt{\gamma_N\frac{\log(p\check{T}K/\delta)}{\eta_n^2 n}}+p\eta_n)$. Take, without loss of generality, $B\geq 1$ (if $B<1$), we can find an upper bound with a different $B=1$. Substituting recursively each term, we can write[77]

$$P_w(\delta)\leq\text{err}_n(\delta)\sum_{s=1}^{w}\frac{1}{s}\prod_{j=s}^{w}(\frac{B}{j}+1).$$

[77]The expression is $\leq$ instead of $=$ since the first term in the expression $\text{err}_n(\delta)$ multiplies by $(B/w+1)$ instead of just 1.

We now write

$$\sum_{s=1}^{w} \frac{1}{s} \prod_{j=s}^{w} (\frac{B}{j} + 1) \leq \sum_{s=1}^{w} \frac{1}{s} \exp(\sum_{j=s}^{w} \frac{B}{j}) \leq \sum_{s=1}^{w} \frac{1}{s} \exp\Big(1 + B\log(w) - B\log(s)\Big)$$
$$\lesssim \sum_{s=1}^{w} \frac{1}{s^2} e^{B\log(w)+1} \lesssim w^B,$$

completing the proof.

□

C.3 Proof of the theorems in the main text

C.3.1 Proof of Theorem 3.1

First observe that for any $\delta \in (0,1)$,

$$\mathbb{E}\Big[\hat{V}_k(\beta)\Big] = V^{(1)}(\beta) + \mathcal{O}(\eta_n), \quad P\Big(\Big|\hat{V}_k(\beta) - V^{(1)}(\beta)\Big| > \mathcal{O}\Big(\eta_n + \sqrt{\frac{\gamma_N \log(\gamma_N/\delta)}{n\eta_n^2}}\Big)\Big) \leq \delta,$$

with the proof of the first claim follows similarly as in the proof of Lemma C.6 and the second claim being a direct corollary of Lemma C.7. Finally observe that with probability at least $1-\delta$, for any $\delta \in (0,1)$, we also have

$$\Big|\widehat{V}_k(\beta) - V^{(1)}(\beta)\Big| = \mathcal{O}(\eta_n) + \mathcal{O}\Big(\sqrt{\frac{\rho_n}{\delta n \eta_n^2}}\Big),$$

by Chebishev inequality and the triangular inequality.

C.3.2 Proof of Theorem 3.2

Consider Algorithm 2 for a generic coordinate j. Let β be the target parameter as in Algorithm 2. By Lemma C.6, we have

$$\mathbb{E}[\widehat{V}_k^{(j)}] = V^{(j)}(\beta) + \mathcal{O}(\eta_n).$$

We have

$$\frac{\widehat{V}_k^{(j)} - \mathbb{E}[\widehat{V}_k^{(j)}]}{\sqrt{\mathrm{Var}(\widehat{V}_k^{(j)})}} = \frac{\widehat{V}_k^{(j)} - V^{(j)}(\beta)}{\sqrt{\mathrm{Var}(\widehat{V}_k^{(j)})}} + \mathcal{O}\Big(\frac{\eta_n}{\sqrt{\mathrm{Var}(\widehat{V}_k^{(j)})}}\Big).$$

Observe that under Assumption 3.2,

$$\mathcal{O}\Big(\frac{\eta_n}{\sqrt{\mathrm{Var}(\widehat{V}_k^{(j)})}}\Big) \leq \mathcal{O}(\eta_n^2 \times \sqrt{n}).$$

First, observe that by Lemma C.4, and the fact that covariates are independent, then $Y_{i,t}^{(k)} - Y_{i,0}^{(k)}$ form a locally dependent graph of maximum degree of order $\mathcal{O}(\gamma_N)$. We now invoke Lemma C.1. We can write

$$\begin{aligned}
&d_W\Big(\frac{1}{2\eta_n\sqrt{\mathrm{Var}(\widehat{V}_k^{(j)})}}\Big[\bar{Y}_t^{(k)} - \bar{Y}_0^{(k)}\Big] - \frac{1}{2\eta_n\sqrt{\mathrm{Var}(\widehat{V}_k^{(j)})}}\Big[\bar{Y}_t^{(k+1)} - \bar{Y}_0^{(k+1)}\Big], \mathcal{G}\Big) \\
&\leq \underbrace{\frac{\gamma_N^2}{\sigma^3}\sum_{h\in\{k,k+1\}}\sum_{i=1}^{n}\Big[\mathbb{E}\Big|\frac{Y_{i,t}^{(k)} - Y_{i,0}^{(k)}}{\eta_n n}\Big|^3\Big]}_{(A)} + \underbrace{\frac{\sqrt{28}\gamma_N^{3/2}}{\sqrt{\pi}\sigma^2}\sqrt{\sum_{i=1}^{n}\Big[\mathbb{E}\Big|\frac{Y_{i,t}^{(k)} - Y_{i,0}^{(k)}}{\eta_n n}\Big|^4\Big]}}_{(B)}, \\
&\mathcal{G} \sim \mathcal{N}(0,1), \quad \sigma^2 = \mathrm{Var}\Big(\frac{1}{2\eta_n}\Big[\bar{Y}_t^{(k)} - \bar{Y}_0^{(k)}\Big] - \frac{1}{2\eta_n}\Big[\bar{Y}_t^{(k+1)} - \bar{Y}_0^{(k+1)}\Big]\Big)
\end{aligned}$$

and d_W denotes the Wasserstein metric. We now inspect each argument on the right hand side. Under Assumption 3.2, $\sigma^2 \geq C_k\frac{1}{n\eta_n^2}$ for a constant $C_k > 0$, and the third and fourth moment are bounded. Hence, we have for a constant $C' < \infty$,

$$(A) \leq C'\frac{\gamma_N^2}{n^3\eta_n^3} \times n^{5/2}\eta_n^3 \asymp \frac{\gamma_N^2}{n^{1/2}} \to 0.$$

Similarly, for (B), we have

$$(B) \leq c'\frac{\gamma_N^{3/2} n\eta_n^2}{\eta_n^2 n^{3/2}} \asymp \frac{\gamma_N^{3/2}}{n^{1/2}} \to 0.$$

The proof completes.

C.3.3 Proof of Theorem 3.3

By Lemma 2.1, we can write (we omit the superscript k from $X^{(k)}$ for sake of brevity)

$$\mathbb{E}\Big\{\frac{1}{2n}\sum_{i=1}^{n}\Big[\frac{D_{i,1}^{(k+1)}Y_{i,1}^{(k+1)}}{\pi(X_i,\beta+\eta_n\underline{e}_1)}-\frac{(1-D_{i,1}^{(k+1)})Y_{i,1}^{(k+1)}}{1-\pi(X_i,\beta+\eta_n\underline{e}_1)}\Big]+\frac{1}{2n}\sum_{i=1}^{n}\Big[\frac{D_{i,1}Y_{i,1}^{(k)}}{\pi(X_i,\beta-\eta_n\underline{e}_1)}-\frac{(1-D_{i,1}^{(k)})Y_{i,1}^{(k)}}{1-\pi(X_i,\beta-\eta_n\underline{e}_1)}\Big]\Big\}$$
$$=\frac{1}{2}\mathbb{E}\Big\{\Big[\frac{D_{i,1}^{(k+1)}Y_{i,1}^{(k+1)}}{\pi(X_i,\beta+\eta_n\underline{e}_1)}-\frac{(1-D_{i,1}^{(k+1)})Y_{i,1}^{(k+1)}}{1-\pi(X_i,\beta+\eta_n\underline{e}_1)}\Big]+\frac{1}{2n}\sum_{i=1}^{n}\Big[\frac{D_{i,1}Y_{i,1}^{(k)}}{\pi(X_i,\beta-\eta_n\underline{e}_1)}-\frac{(1-D_{i,1}^{(k)})Y_{i,1}^{(k)}}{1-\pi(X_i,\beta-\eta_n\underline{e}_1)}\Big]\Big\}$$
$$=\underbrace{\frac{1}{2}\int\Big[m(1,x,\beta+\eta_n\underline{e}_1)-m(0,x,\beta+\eta_n\underline{e}_1)+m(1,x,\beta-\eta_n\underline{e}_1)-m(0,x,\beta-\eta_n\underline{e}_1)\Big]dF_X(x)}_{(i)}.$$

The last equality follows from Lemma 2.1 and exogeneity of β. Doing a Taylor expansion to each component around β, we obtain

$$(i)=\int\Big[m(1,x,\beta)-m(0,x,\beta)+\frac{\partial m(1,x,\beta)}{2\partial\beta^1}\eta_n-\frac{\partial m(0,x,\beta)}{2\partial\beta^1}\eta_n-\frac{\partial m(1,x,\beta)}{2\partial\beta^1}\eta_n+\frac{\partial m(0,x,\beta)}{2\partial\beta^1}\eta_n\Big]dF_X(x)$$
$$+O(\eta_n^2)=\int\Big[m(1,x,\beta)-m(0,x,\beta)\Big]dF_X(x)+O(\eta_n^2),$$

which completes the proof, since $\eta_n=o(n^{-1/4})$.

C.3.4 Proof of Theorem 3.4

We are interested in studying

$$\mathbb{E}\Big\{\frac{1}{2n}\sum_{h\in\{k,k+1\}}\frac{v_h}{\eta_n}\sum_{i=1}^{n}\Big[\frac{Y_{i,1}^{(h)}(1-D_{i,1}^{(h)})}{1-\pi(X_i^{(h)},\beta+v_h\eta_n\underline{e}_1)}-\bar{Y}_0^{(h)}\Big]\Big\},\quad v_h=\begin{cases}1 \text{ if } h=k\\ -1 \text{ otherwise.}\end{cases}$$

Using Lemma 2.1, similarly to the derivation of Lemma C.6, we can write the above expression equal to

$$\frac{1}{2\eta_n}\int[m(0,x,\beta+\eta_n\underline{e}_1)-m(0,x,\beta-\eta_n\underline{e}_1)]dF_X(x).$$

Note that from the mean value theorem, and Assumption 3.1

$$m(0,x,\beta+\eta_n\underline{e}_1)-m(0,x,\beta-\eta_n\underline{e}_1)=m(0,x,\beta)-m(0,x,\beta)+2\frac{\partial m(0,x,\beta)}{\partial\beta^1}\eta_n+O(\eta_n^2)$$

which completes the proof.

C.3.5 Proof of Theorem 4.2

Consider Lemma C.10 where we choose $\delta = 1/n$. We can write for each k

$$||\beta^* - \check{\beta}_k^{\check{T}}||_2^2 \leq \frac{pL}{\check{T}} + \mathcal{O}(1/\check{T}),$$

for a finite constant $L < \infty$, since, under the conditions for n stated in the theorem, for finite B, the second component is $\mathcal{O}(1/\check{T})$. Note that

$$||\beta^* - \frac{1}{K}\sum_{k=1}^{K}\check{\beta}_k^{\check{T}}||_2^2 \leq \frac{1}{K}\sum_{k=1}^{K}||\beta^* - \check{\beta}_k^{\check{T}}||_2^2$$

by Jensen's inequality, which completes the proof.

C.3.6 Theorem 4.3

By the mean value theorem and Assumption 3.1, we have

$$\sum_{w=1}^{\check{T}} W(\beta^*) - W(\check{\beta}_k^w) \leq \bar{C}\sum_{w=1}^{\check{T}}||\beta^* - \check{\beta}_k^w||_2^2,$$

for a finite constant $\bar{C} < \infty$, since $\frac{\partial W(\beta^*)}{\partial \beta} = 0$, and the Hessian is uniformly bounded (Assumption 3.1). By Lemma C.10, choosing $\delta = 1/n$, and for n satisfying the conditions in Theorem 4.3, it follows that for all k,

$$\sum_{w=1}^{\check{T}} W(\beta^*) - W(\check{\beta}_k^w) \leq \sum_{w=1}^{\check{T}}\frac{p\kappa'}{w} \lesssim p\log(\check{T})$$

for $\kappa' < \infty$ being a finite constant. The proof completes.

C.3.7 Proof of Theorem 4.4

First, note that for a finite constant c_0, under Assumption 3.1 and Assumption 4.2

$$W(\beta^*) - W(\hat{\beta}^*) \leq c_0||\beta^* - \hat{\beta}||^2 \leq c_0\frac{1}{K}\sum_{k=1}^{K}||\beta^* - \check{\beta}_k^{\check{T}+1}||^2$$

where in the first inequality we used strong concavity (gradient equals zero), and in the second equality we used Jensen's inequality. Define β_w^{**} as in Equation (C.6), where, however, the

learning rate is chosen so that $\alpha_w = 1/\tau$. Using the triangular inequality, we can write

$$||\beta^* - \check{\beta}_k^{\check{T}+1}||_2^2 \leq ||\beta^* - \beta_{\check{T}+1}^{**}||_2^2 + ||\check{\beta}_k^{\check{T}+1} - \beta_{\check{T}+1}^{**}||_2^2.$$

The first component is bounded by Theorem 3.10 in Bubeck (2014) (using the fact that $\mathcal{B}$ is compact) as follows:

$$||\beta^* - \beta_{\check{T}+1}^{**}||_2^2 \leq c_0 \exp(-c_0' 2(\check{T}+1)) = c_0 \exp(-K c_0')$$

for finite constants $0 < c_0, c_0' < \infty$, where we used the fact that $2(\check{T}+1) = K$. Using Lemma C.9, we bound the second component with probability at least $1-\delta$, as follows (for any $w \leq \check{T}+1$)

$$||\check{\beta}_k^w - \beta_w^{**}||_2^2 \leq p||\check{\beta}_k^w - \beta_w^{**}||_\infty^2 = p \times \mathcal{O}(P_w^2(\delta)).$$

We conclude the proof by explicitely characterizing the rate of $P_w(\delta)$ as defined in Lemma C.9. Following Lemma C.9, we can define recursively $P_w(\delta)$ for any $1 \leq w \leq \check{T}$ (recall that $\alpha_w \propto 1/w$) as

$$P_w(\delta) = (1+B)P_{w-1}(\delta) + \text{err}_n(\delta), \quad P_1(\delta) = \text{err}_n(\delta).$$

where $\text{err}_n = \mathcal{O}(\sqrt{\gamma_N \frac{\log(p\check{T}K/\delta)}{\eta_n^2 n}} + p\eta_n)$, and $B > 0$ is a finite constant with depends on p only. Using a recursive argument, we can write

$$P_w(\delta) \leq w(1+B)^w \text{err}_n(\delta).$$

The proof completes as we choose n sufficiently large as stated in the theorem.

C.3.8 Proof of Theorem 4.5

Recall that from Assumption 2.1, the maximum degree is $\gamma_N^{1/2}$. We break the proof into several steps. We will write the model dividing by $1/\sum_{j\neq i} A_{i,j}1\{X_j = x\}$ instead of $1/\max\{\sum_{j\neq i} A_{i,j}1\{X_j = x\}, 1\}$ for notational convenience, but implicitely consider the expression $1/\sum_{j\neq i} A_{i,j}1\{X_j = x\}$ equal to one if $\sum_{j\neq i} A_{i,j}1\{X_j = x\} = 0$.

Upper bound on W_N^* We first provide an upper bound on the largest achievable welfare. Recall that $\Delta = c$. Therefore, we can write

$$W_N^* \leq \frac{1}{N}\sum_{i=1}^N \sup_{\mathcal{P}} \mathbb{E}\Big[\mathbb{E}_{D\sim\mathcal{P}(A,X)}\Big[s\Big(\Big[\frac{\sum_{j\neq i} A_{i,j}D_j 1\{X_j = x\}}{\sum_{j\neq i} A_{i,j}1\{X_j = x\}}\Big]_{x\in\{1,\cdots,|\mathcal{X}|\}}\Big)\Big|A, X\Big]\Big].$$

Let

$$\beta^G = \arg \max_{\beta_1,\cdots,\beta_{|\mathcal{X}|}\in[0,1]^{|\mathcal{X}|}} s\Big(\beta_1,\cdots,\beta_{|\mathcal{X}|}\Big).$$

Note that since $D_j \in \{0,1\}$, we can write

$$\sup_{\mathcal{P}} \mathbb{E}\Big[\mathbb{E}_{D\sim\mathcal{P}(A,X)}\Big[s\Big(\Big[\frac{\sum_{j\neq i} A_{i,j}D_j 1\{X_j = x\}}{\sum_{j\neq i} A_{i,j}1\{X_j = x\}}\Big]_{x\in\{1,\cdots,|\mathcal{X}|\}}\Big)\Big|A,X\Big]\Big] \leq s\Big(\beta_1^G,\cdots,\beta_{|\mathcal{X}|}^G\Big).$$

Lower bound on $W(\beta^*)$ Using the fact that $\mathcal{B} = [0,1]^{|\mathcal{X}|}$, we can write[78]

$$\begin{aligned} W(\beta^*) &= \max_{\beta\in[0,1]^{|\mathcal{X}|}} \mathbb{E}_\beta\Big[s\Big(\Big[\frac{\sum_{j\neq i} A_{i,j}D_j 1\{X_j = x\}}{\sum_{j\neq i} A_{i,j}1\{X_j = x\}}\Big]_{x\in\{1,\cdots,|\mathcal{X}|\}}\Big)\Big] \\ &\geq \mathbb{E}_{\beta^G}\Big[s\Big(\Big[\frac{\sum_{j\neq i} A_{i,j}D_j 1\{X_j = x\}}{\sum_{j\neq i} A_{i,j}1\{X_j = x\}}\Big]_{x\in\{1,\cdots,|\mathcal{X}|\}}\Big)\Big], \end{aligned}$$

where we use the fact that $\beta^G = (\beta_1^G,\cdots,\beta_{|\mathcal{X}|}^G) \in [0,1]^{|\mathcal{X}|}$, and $\Delta(\cdot) = c(\cdot)$. From the mean value theorem

$$\begin{aligned} &\mathbb{E}_{\beta^G}\Big[s\Big(\Big[\frac{\sum_{j\neq i} A_{i,j}D_j 1\{X_j = x\}}{\sum_{j\neq i} A_{i,j}1\{X_j = x\}}\Big]_{x\in\{1,\cdots,|\mathcal{X}|\}}\Big)\Big] \\ &= s(\beta^G) + \mathbb{E}_{\beta^G}\Big\{\frac{\partial s(\beta)}{\partial\beta}\Big|_{\beta\in\Big[\Big[\frac{\sum_{j\neq i} A_{i,j}D_j 1\{X_j=x\}}{\sum_{j\neq i} A_{i,j}1\{X_j=x\}}\Big]_{x\in\{1,\cdots,|\mathcal{X}|\}},\beta^G\Big]} \times \\ &\times \Big(\Big[\frac{\sum_{j\neq i} A_{i,j}D_j 1\{X_j = x\}}{\sum_{j\neq i} A_{i,j}1\{X_j = x\}}\Big]_{x\in\{1,\cdots,|\mathcal{X}|\}} - \beta^G\Big)\Big\}, \end{aligned}$$

where $\frac{\partial s(\cdot)}{\partial\beta}$ is evaluated at a point between the shared of treated neighbors and β^G.

Bound on the difference Combining the two bounds, we can write

$$W_N^* - W(\beta^*) \leq \underbrace{\Big|\mathbb{E}_{\beta^G}\Big\{\frac{\partial s(\beta)}{\partial\beta}\Big|_{\beta\in\Big[\Big[\frac{\sum_{j\neq i} A_{i,j}D_j 1\{X_j=x\}}{\sum_{j\neq i} A_{i,j}1\{X_j=x\}}\Big]_{x\in\{1,\cdots,|\mathcal{X}|\}},\beta^G\Big]} \times \Big(\Big[\frac{\sum_{j\neq i} A_{i,j}D_j 1\{X_j = x\}}{\sum_{j\neq i} A_{i,j}1\{X_j = x\}}\Big]_{x\in\{1,\cdots,|\mathcal{X}|\}} - \beta^G\Big)\Big\}\Big|}_{(I)},$$

where we took the absolute value in the last equation.

[78] $\mathbb{E}_\beta\Big[s\Big(\Big[\frac{\sum_{j\neq i} A_{i,j}D_j 1\{X_j=x\}}{\sum_{j\neq i} A_{i,j}1\{X_j=x\}}\Big]_{x\in\{1,\cdots,|\mathcal{X}|\}}\Big)\Big]$ does not depend on i similarly to Lemma 2.1.

Bound with Cauchy-Schwarz We can now bound (I) as follows.

$$(I) \leq \sup_{\beta} \Big|\Big| \frac{\partial s(\beta)}{\partial \beta} \Big|\Big|_2 \times |\mathcal{X}| \max_{x \in \mathcal{X}} \underbrace{\sqrt{\mathbb{E}_{\beta^G}\Big[\Big(\frac{\sum_{j \neq i} A_{i,j} D_j 1\{X_j = x\}}{\sum_{j \neq i} A_{i,j} 1\{X_j = x\}} - \beta_x^G\Big)^2\Big]}}_{(II)},$$

where we first used Cauchy-Schwarz and then bound the first component by the supremum over β, x and the second component by the largest term over $x \in \mathcal{X}$ times the number of elements $|\mathcal{X}|$.

Bound for (II) we now want to bound (II). We do so fixing $X_j = x$ and show that for all $x \in \mathcal{X}$ (and so also for the maximum) we can obtain a useful bound. Recall that here $\mathbb{E}_{\beta^G}$ indicates that $D_{i,t} | X_i^{(k)} = x \sim_{i.i.d.} \text{Bern}(\beta_x^G)$. Now, note that

$$\mathbb{E}_{\beta^G}\Big[\frac{\sum_{j \neq i} A_{i,j} D_j 1\{X_j = x\}}{\sum_{j \neq i} A_{i,j} 1\{X_j = x\}} \Big| X^{(k)}, A^{(k)}\Big] = \beta_x^G.$$

Also, note that since we take the maximum over $x \in \mathcal{X}$, here β_x is fixed, and therefore, $\text{Var}(\beta_x) = 0$. Therefore, we can write

$$\mathbb{E}_{\beta^G}\Big[\Big(\frac{\sum_{j \neq i} A_{i,j} D_j 1\{X_j = x\}}{\sum_{j \neq i} A_{i,j} 1\{X_j = x\}} - \beta_x^G\Big)^2\Big] = \mathbb{E}_{\beta^G}\Big[\text{Var}\Big(\frac{\sum_{j \neq i} A_{i,j} D_j 1\{X_j = x\}}{\sum_{j \neq i} A_{i,j} 1\{X_j = x\}} \Big| X^{(k)}, A^{(k)}\Big)\Big].$$

In addition,

$$\text{Var}_{\beta^G}\Big(\frac{\sum_{j \neq i} A_{i,j} D_j 1\{X_j = x\}}{\sum_{j \neq i} A_{i,j} 1\{X_j = x\}} \Big| X^{(k)}, A^{(k)}\Big) = \beta_x^G(1 - \beta_x^G) \Big/ \sum_{j \neq i} A_{i,j} 1\{X_j = x\},$$

since treatments are independent conditional on $X^{(k)}$, independent of $A^{(k)}$ conditional on $X^{(k)}$ by construction, and binary. Let $\kappa' = \kappa P(X = x)$, where, without loss of generality, $P(X = x) > 0$ since $|\mathcal{X}| < \infty$, and hence, if $P(X = x) = 0$, we can re-index $s(\cdot)$, for all types except $X = x$ and conduct the same analysis as above, without the case $X = x$. Note

that

$$\mathbb{E}\Big[\beta_x^G(1-\beta_x^G)\Big/\sum_{j\neq i}A_{i,j}1\{X_j=x\}\Big]=\beta_x^G(1-\beta_x^G)\mathbb{E}\Big[1\Big/\sum_{j\neq i}A_{i,j}1\{X_j=x\}\Big]$$
$$\leq \beta_x^G(1-\beta_x^G)P\Big(\sum_{j\neq i}A_{i,j}1\{X_j=x\}<\kappa'\gamma_N^{1/4}\Big)+\beta_x^G(1-\beta_x^G)P\Big(\sum_{j\neq i}A_{i,j}1\{X_j=x\}\geq\kappa'\gamma_N^{1/4}\Big)\frac{1}{\kappa'\gamma_N^{1/4}}$$
$$\leq \underbrace{\beta_x^G(1-\beta_x^G)\,P\Big(\sum_{j\neq i}A_{i,j}1\{X_j=x\}<\kappa'\gamma_N^{1/4}\Big)}_{(III)}+\frac{1}{\kappa'\gamma_N^{1/4}},$$

where in the second inequality we used the fact that we are implicitely using in our notation that $1/\sum_{j\neq i}A_{i,j}1\{X_j=x\}$ equals one if $\sum_{j\neq i}A_{i,j}1\{X_j=x\}=0$ for notational convenience.

Bound for (III) We are left to derive a bound for (III), since the second term converges to zero as $\gamma_N\to\infty$. Define $h_x(X_i,U_i)=P(X=x)\int l(X_i,U_i,x,u)dF_{U|X=x}(u)$. Note that (recall that $1\{i\leftrightarrow j\}$ are fixed)

$$\mathbb{E}[A_{i,j}1\{X_j=x\}|X_i,U_i]=\mathbb{E}[l(X_i,U_i,X_j,U_j)1\{X_j=x\}1\{i\leftrightarrow j\}|X_i,U_i]=h_x(X_i,U_i)1\{i\leftrightarrow j\},$$

since, conditional on (X_i,U_i),the indicator $1\{i\leftrightarrow j\}$ is fixed (exogenous), and $(X_i,U_i)\sim_{i.i.d.}F_XF_{U|X}$. Also, recall that $\sum_j 1\{i\leftrightarrow j\}=\gamma_N^{1/2}$. Hence, only $\gamma_N^{1/2}$ many edges of i can at most be non-zero, while the remaining ones are zero almost surely. Therefore, using Hoeffding's inequality (Wainwright, 2019), and using independence conditional on X_i,U_i,

$$P\Big(\Big|\frac{1}{\gamma_N^{1/2}}\sum_{j\neq i}A_{i,j}1\{X_j=x\}-h_x(X_i,U_i)\Big|\leq\bar{C}\sqrt{\frac{\log(2\gamma_N)}{\gamma_N^{1/2}}}\Big|X_i,U_i\Big)\geq 1-1/\gamma_N,\qquad\text{(C.5)}$$

for a finite constant $\bar{C}<\infty$. Observe that $h_x(X_i,U_i)\geq\kappa'>0,\kappa'=P(X=x)\kappa$ almost surely by assumption. Define the event

$$\mathcal{E}=\Big\{|\sum_{j\neq i}A_{i,j}1\{X_j=x\}-\gamma_N^{1/2}h_x(X_i,U_i)|\leq\bar{C}\sqrt{\log(2\gamma_N)\gamma_N^{1/2}}\Big\}.$$

We can write

$$P\Big(\sum_{j\neq i} A_{i,j}1\{X_j = x\} < \kappa'\gamma_N^{1/4}\Big) = P\Big(\sum_{j\neq i} A_{i,j}1\{X_j = x\} - \gamma_N^{1/2}h_x(X_i,U_i) + \gamma_N^{1/2}h_x(X_i,U_i) < \kappa'\gamma_N^{1/4}\Big)$$
$$\leq P\Big(\gamma_N^{1/2}h_x(X_i,U_i) < \kappa'\gamma_N^{1/4} + |\sum_{j\neq i} A_{i,j}1\{X_j = x\} - \gamma_N^{1/2}h_x(X_i,U_i)|\Big)$$
$$\leq P\Big(\gamma_N^{1/2}h_x(X_i,U_i) < \kappa\gamma_N^{1/4} + |\sum_{j\neq i} A_{i,j}1\{X_j = x\} - \gamma_N^{1/2}h_x(X_i,U_i)|\Big|\mathcal{E}\Big)$$
$$+ P\Big(\gamma_N^{1/2}h_x(X_i,U_i) < \kappa'\gamma_N^{1/4} + |\sum_{j\neq i} A_{i,j}1\{X_j = x\} - \gamma_N^{1/2}h(X_i,U_i)|\Big|\mathcal{E}^c\Big) \times P\Big(\mathcal{E}^c\Big)$$

Note that by Equation (C.5) (which holds conditionally and so also unconditionally)

$$P\Big(\gamma_N^{1/2}h_x(X_i,U_i) < \kappa'\gamma_N^{1/4} + |\sum_{j\neq i} A_{i,j}1\{X_j = x\} - \gamma_N^{1/2}h_x(X_i,U_i)|\Big|\mathcal{E}^c\Big) \times P\Big(\mathcal{E}^c\Big) \leq \frac{1}{\gamma_N} = o(1).$$

Finally, we can write for a finite constant $\bar{C} < \infty$,

$$P\Big(\gamma_N^{1/2}h_x(X_i,U_i) < \kappa'\gamma_N^{1/4} + |\sum_{j\neq i} A_{i,j}1\{X_j = x\} - \gamma_N^{1/2}h_x(X_i,U_i)|\Big|\mathcal{E}\Big)$$
$$\leq P\Big(\gamma_N^{1/2}h_x(X_i,U_i) < \kappa'\gamma_N^{1/4} + \bar{C}\sqrt{\log(2\gamma_N)\gamma_N^{1/2}}\Big|\mathcal{E}\Big)$$
$$\leq P\Big(\inf_{x,x',u'} \gamma_N^{1/2}h_x(x',u') < \kappa'\gamma_N^{1/4} + \bar{C}\sqrt{\log(2\gamma_N)\gamma_N^{1/2}}\Big|\mathcal{E}\Big)$$
$$= 1\Big\{\inf_{x,x',u'} h_x(x',u') < \kappa'\gamma_N^{-1/4} + \bar{C}\sqrt{\log(2\gamma_N)}\gamma_N^{-1/4}\Big\}$$

which equals to zero for N, γ_N large enough, since $\inf_{x,x',u'} h_x(x',u') > 0$.

C.3.9 Proof of Theorem A.4

Recall that $\mathcal{G}$ denotes a finite grid with $K/2$ elements. First, we bound

$$W(\beta^*) - W(\hat{\beta}^{ow}) \leq 2 \sup_{\beta\in[0,1]^p} \Big|W(\beta) - \hat{W}(\beta)\Big|.$$

By the mean value theorem, we can write for any $\beta^k \in \mathcal{G}$

$$W(\beta) = W(\beta^k) + V(\beta^k)^\top(\beta - \beta^k) + O\Big(||\beta^k - \beta||^2\Big),$$

Since we construct $\hat{W}(\beta)$ as in Equation (A.7), we can choose β^k closest to β. In such a case, by construction of the grid, $O\Big(||\beta^k - \beta||^2\Big) = O(1/K^{2/p})$ by construction of the grid.

We can then write (using the fact that $p < \infty$)

$$\sup_{\beta\in[0,1]}\left|W(\beta)-\hat{W}(\beta)\right| \leq \sup_{\beta\in[0,1]^p,k\in\{1,\cdots,K\}}\left|W(\beta^k)+V(\beta^k)^\top(\beta-\beta^k)-\bar{W}^k-\hat{V}_{(k,k+1)}^\top(\beta-\beta^k)\right|+O(1/K^{2/p})$$
$$\leq \sup_{k\in\{1,\cdots,K\}}\left|W(\beta^k)-\bar{W}^k\right|+||V(\beta^k)-\hat{V}_{k,k+1}||_\infty O(1)+O(1/K^{2/p})$$

Observe now that similarly to what discussed in Lemma C.6, where here the first order derivatives cancel out with a second order Taylor expansion, due to the opposite sign of $\pm\eta_n$ in each cluster,

$$2\mathbb{E}\Big[\bar{W}^k\Big] = \int y(x,\beta^k+\eta_n)dF_X(x)+\int y(x,\beta^k-\eta_n)dF_X(x) = 2\int y(x,\beta^k)dF_X(x)+O(\eta_n^2).$$

Using Lemma C.3, we can write for all $k \leq K$, with probability at least $1-\delta$,

$$\bar{W}^k = W(\beta^k)+O\Big(\sqrt{\gamma_N\log(pK\gamma_N/\delta)/n}+\eta_n^2\Big),$$

where we used the union bound over K, p in the expression. Similarly, from Lemma C.7, also using the union bound over K and p, with probability at least $1-\delta$,

$$||\hat{V}_{(k,k+1)}-V(\beta^k)||_\infty = O\Big(\sqrt{\gamma_N\log(Kp\gamma_N/\delta)/(n\eta_n^2)}+\eta_n\Big),$$

which concludes the proof as we choose $\delta = 1/n$, since $\eta_n = o(1)$, and p is finite.

C.4 Proofs for the extensions

C.4.1 Proof of Theorem A.1

Observe that we can write

$$\mathbb{E}\Big[\bar{Y}_t^{(k)}\Big|p_t^{(k)}\Big] = \alpha_t+\tau_k+g\Big(q(\beta+\eta_n)+o_p(\eta_n),\beta+\eta_n\Big),$$
$$\mathbb{E}\Big[\bar{Y}_t^{(k)}\Big|p_t^{(k+1)}\Big] = \alpha_t+\tau_k+g\Big(q(\beta-\eta_n)+o_p(\eta_n),\beta-\eta_n\Big).$$

From a Taylor expansion in its first argument around $q(\beta+\eta_n)$, we obtain

$$g\Big(q(\beta+\eta_n)+o_p(\eta_n),\beta+\eta_n\Big) = g\Big(q(\beta+\eta_n),\beta+\eta_n\Big)+o_p(\eta_n)$$

and similarly once we subctract η_n. Therefore, we obtain

$$\mathbb{E}\Big[\bar{Y}_t^{(k)}\Big|p_t^{(k)}\Big]-\mathbb{E}\Big[\bar{Y}^{(k)}\Big|p_t^{(k+1)}\Big] = \tau_k-\tau_{k+1}+g\Big(q(\beta+\eta_n),\beta+\eta_n\Big)+o_p(\eta_n)-g\Big(q(\beta-\eta_n),\beta-\eta_n\Big).$$

We can now proceed with a Taylor expansion around of the functions $g(\cdot)$ around β to obtain (this follows similarly to Lemma C.6)

$$g\Big(q(\beta+\eta_n),\beta+\eta_n\Big)-g\Big(q(\beta-\eta_n),\beta-\eta_n\Big)=2V_g(\beta)\eta_n+O(\eta_n^2).$$

In addition observe that since at the baseline β_0 is the same for both clusters,

$$\mathbb{E}[Y_0^{(k)}-Y_0^{(k+1)}|p_t^{(k)},p_t^{(k+1)}]=\tau_k-\tau_{k+1}+o_p(\eta_n).$$

The proof concludes from Lemma C.3 and the local dependence assumption in Assumption A.1.

C.4.2 Proof of Theorem A.3

First, we bound

$$\sup_{\theta\in\Theta}\widetilde{W}(\theta)-W(\widehat{\theta})\leq 2\sum_t q^t\times\underbrace{\sup_{(\beta_1,\beta_2)\in[0,1]^2}\Big|\widehat{\Gamma}(\beta_2,\beta_1)-\Gamma(\beta_2,\beta_1)\Big|}_{(A)}\lesssim(A),$$

since $\sum_t q^t<\infty$. To bound (A) observe first that each element in the grid $\mathcal{G}$ has a distance of order $1/\sqrt{K}$, since the grid has two dimensions and $K/3$ components. As a result for any element (β_2,β_1), we can write

$$\Gamma(\beta_2,\beta_1)=\underbrace{\Gamma(\beta_2^r,\beta_1^r)}_{(B)}+\underbrace{\frac{\partial\Gamma(\beta_2^r,\beta_1^r)}{\partial\beta_1^r}(\beta_1-\beta_1^r)+\frac{\partial\Gamma(\beta_2^r,\beta_1^r)}{\partial\beta_2^r}(\beta_2-\beta_2^r)}_{(C)}+\underbrace{\mathcal{O}\Big(||\beta_1-\beta_1^r||^2+||\beta_2-\beta_2^r||^2\Big)}_{(D)}$$

where $\beta^r\in\mathcal{G}$ is some value in the grid such that (B) is of order $1/K$. We can now write

$$\begin{aligned}(A)\leq&\underbrace{\sup_{(\beta_1^r,\beta_2^r)\in\mathcal{G},||\beta_1^r-\beta_1||^2+||\beta_2^r-\beta_2||^2\lesssim 1/K}\Big|\widetilde{\Gamma}(\beta_2^r,\beta_1^r)-\Gamma(\beta_2^r,\beta_1^r)\Big|}_{(i)}\\&+\underbrace{\sup_{(\beta_1^r,\beta_2^r)\in\mathcal{G}}\Big|\widehat{g}_2(\beta_2^r,\beta_1^r)-\frac{\partial\Gamma(\beta_2^r,\beta_1^r)}{\partial\beta_2^r}\Big|\Big(|\beta_2-\beta_2^r|+|\beta_1-\beta_1^r|\Big)}_{(ii)}\\&+\underbrace{\sup_{(\beta_1^r,\beta_2^r)\in\mathcal{G}}\Big|\widehat{g}_1(\beta_2^r,\beta_1^r)-\frac{\partial\Gamma(\beta_2^r,\beta_1^r)}{\partial\beta_1^r}\Big|\Big(|\beta_2-\beta_2^r|+|\beta_1-\beta_1^r|\Big)}_{(iii)}+\mathcal{O}\Big(|\beta_2-\beta_2^r|^2+|\beta_1-\beta_1^r|^2\Big).\end{aligned}$$

We now study each component separately. We start from (i). We observe that under Assumption A.3, by doing a Taylor expansion around (β_1^r, β_2^r), it is easy to observe that we can write

$$\mathbb{E}[\bar{Y}_{t+1}^{(k)}] = \Gamma(\beta_2^r, \beta_1^r) + \mathcal{O}(\eta_n).$$

Therefore by Lemma C.3, and the union bound over K many elements in $\mathcal{G}$ as $\gamma_N \log(\gamma_N K)/n \to 0$, $(i) \to 0$. Consider now (ii). We observe that since $\mathcal{B}$ is compact, we have $\Big(|\beta_2 - \beta_2^r| + |\beta_1 - \beta_1^r|\Big) = \mathcal{O}(1)$. In addition, similarly to what discussed in Lemma C.7, it follows that with probability at least $1-\delta$,

$$\left|\widehat{g}_1(\beta_2^r, \beta_1^r) - \frac{\partial \Gamma(\beta_2^r, \beta_1^r)}{\partial \beta_1^r}\right| = \mathcal{O}\Big(\sqrt{\frac{\gamma_N \log(\gamma_N/\delta)}{\eta_n^2 n}} + \eta_n\Big).$$

Hence, by the union bound as $\frac{\gamma_N \log(\gamma_N K)}{\eta_n^2 n} = o(1)$ $(ii) = o_p(1)$ and similarly (iii). The proof concludes after observing that $|\beta_1^r - \beta_1|^2 + |\beta_2^r - \beta_2|^2 \lesssim 1/K$ by construction of the grid.

C.4.3 Proof of Theorem B.1

Let $\tilde{K} = K/2l$. Take

$$t_z^j = \frac{\frac{1}{\sqrt{z}} \sum_{i=1}^z X_i^j}{\sqrt{(z-1)^{-1} \sum_{i=1}^z (X_i^j - \bar{X}^j)^2}}, \quad X_i^j \sim \mathcal{N}(0, \sigma_i^j).$$

Recall that by Theorem 1 in Ibragimov and Müller (2010), we have that for $\alpha \leq 0.08$

$$\sup_{\sigma_1, \cdots, \sigma_q} P(|t_z| \geq \text{cv}_\alpha) = P(|T_{q-1}| \geq \text{cv}_\alpha),$$

where cv_α is the critical value of a t-test with level α, and T_{z-1} is a t-student random variable with $z-1$ degrees of freedom. The equality is attained under homoskedastic variances (Ibragimov and Müller, 2010). We now write

$$P\Big(\mathcal{T}_n \geq q | H_0\Big) = P\Big(\max_{j \in \{1, \cdots, l\}} |Q_{j,n}| \geq q | H_0\Big) = 1 - P\Big(|Q_{j,n}| \leq q \forall j | H_0\Big) = 1 - \prod_{j=1}^{l} P\Big(|Q_{j,n}| \leq q | H_0\Big),$$

where the last equality follows by between cluster independence. Observe now that by Theorem 3.2 and the fact that the rate of convergence is the same for all clusters (Assumption

3.2) [79], for all j, for some $(\sigma_1, \cdots, \sigma_z)$, $z = \tilde{K}$,

$$\sup_q \left| P\Big(|Q_{j,n}| \leq q | H_0\Big) - P\Big(|t^j_{\tilde{K}}| \leq q\Big) \right| = o(1).$$

As a result, we can write

$$\sup_{\sigma_1,\cdots,\sigma_K} \lim_{n\to\infty} 1 - \prod_{j=1}^{l} P\Big(|Q_{j,n}| \leq q | H_0\Big) = 1 - \prod_{j=1}^{l} \inf_{\sigma^j_1,\cdots,\sigma^j_{\tilde{K}}} P\Big(|t^j_{\tilde{K}}| \leq q\Big),$$

where we used the fact that we use different pairs of (independent) clusters for each entry j. Using the result in Bakirov and Szekely (2006), we have

$$\inf_{\sigma^j_1,\cdots,\sigma^j_{\tilde{K}}} P\Big(|t^j_{\tilde{K}}| \leq q\Big) = P\Big(|T_{\tilde{K}-1}| \leq q | H_0\Big).$$

Therefore,

$$1 - \prod_{j=1}^{l} \inf_{\sigma^j_1,\cdots,\sigma^j_{\tilde{K}}} P\Big(|t^j_{\tilde{K}}| \leq q | H_0\Big) = 1 - P^l\Big(|T_{\tilde{K}-1}| \leq q\Big).$$

Setting the expression equal to α, we obtain

$$1 - P^l\Big(|T_{\tilde{K}-1}| \leq q\Big) = \alpha \Rightarrow P\Big(|T_{\tilde{K}-1}| \geq q\Big) = 1 - (1-\alpha)^{1/l}.$$

The proof completes after solving for q.

C.4.4 Proof of Theorem B.2

In this subsection, we derive the theorem for the gradient descent method under Assumption B.1, for our extension where we relax global strong concavity. The derivation is split into the following lemmas. First define the oracle descent as follows.

Definition C.5 (Oracle gradient descent)**.** We define for positive constants $\infty > \mu, \kappa > 0$, κ as defined in Lemma C.11, arbitrary $v \in (0, 1)$

$$\beta^*_w = \begin{cases} P_{\mathcal{B}_1,\mathcal{B}_2}\Big[\beta^*_{w-1} + \alpha_{w-1} V(\beta^*_{w-1})\Big] \text{ if } ||V(\beta^*_w)||_2 \geq \frac{\kappa}{\mu \tilde{T}^{1/2-v/2}} \\ \beta^*_{w-1} \text{ otherwise} \end{cases} \quad \beta^*_1 = \beta_0, \tag{C.6}$$

for $\alpha_w = \frac{J}{\tilde{T}^{1/2-v/2}||V(\beta^*_{w-1})||}$, $J < 1$. □

[79] Here we use continuity of the Gaussian distribution, and the fact that l is finite.

Lemma C.11 (Adaptive gradient descent for quasi-concave functions and locally strong concave)**.** *Let $\mathcal{B}$ be compact. Define $G = \max\{\sup_{\beta\in\mathcal{B}} 2||\beta||^2, 1\}$. Let Assumption 3.1, 4.1, B.1 hold. Let κ be a positive finite constant, defined as in Equation (C.7). Then for any $v \in (0,1)$, for $\check{T} \geq ((G+1)/J)^{1/v}$, the following holds:*

$$||\beta^*_{\check{T}} - \beta^*||^2 \leq \kappa \check{T}^{-1+v}.$$

Proof of Lemma C.11. To prove the statement, we use properties of gradient descent methods with gradient norm rescaling (Hazan et al., 2015), with modifications to the original arguments to explicitely obtain a rate T^{-1+v} for an arbitrary small v, and account for the formalization of local strong concavity based on the Hessian which we provide in our context.

Preliminaries Clearly, if the algorithm terminates at w, under Assumption B.1 (B), this implies that

$$||\beta^*_w - \beta^*||_2^2 \leq \kappa \check{T}^{-1+v},$$

proving the claim. Therefore, assume that the algorithm did not terminate at time w. This implies that for any $\check{w} \geq 1$, $||\beta^*_{\check{w}} - \beta^*||_2^2 > \kappa \check{T}^{-1+v}$. Define $\epsilon = \check{T}^{-1+v}$ and let ∇_w be the gradient evaluated at β^*_w. For every $\beta \in \mathcal{B}$, define $H(\beta)\Big|_{[\beta^*,\beta]}$ the Hessian evaluated at some point $\tilde{\beta} \in [\beta^*, \beta]$, such that

$$W(\beta) = W(\beta^*) + \frac{1}{2}(\beta - \beta^*)^\top H(\beta)\Big|_{[\beta^*,\beta]}(\beta - \beta^*),$$

which always exists by the mean-value theorem and differentiability of the objective function. Define

$$\frac{1}{2}(\beta - \beta^*)^\top H(\beta)\Big|_{[\beta^*,\beta]}(\beta - \beta^*) = f(\beta) \leq 0,$$

where the inequality follows by definition of β^* (note that $f(\beta)$ also depends on $\tilde{\beta}$, whose dependence we implicitely suppressed).

Claim We claim that

$$-|\lambda_{max}|||\beta - \beta^*||^2 \leq f(\beta) \leq -|\lambda_{\min}|||\beta - \beta^*||^2$$

for constants $\lambda_{\max} > \lambda_{\min} > 0$. The lower bound follows directly by Assumption 3.1, while the upper bound follows directly from Assumption B.1 (C), compactness of $\mathcal{B}$, and continuity of the Hessian. We provide details for the upper bound in the following paragraph.

Proof of the claim on the upper bound We now use a contradiction argument. Suppose that the upper bound does not hold. Then since $\mathcal{B}$ is compact (and hence $||\beta - \beta^*||$ is bounded away from infinity for all $\beta \in \mathcal{B}$), and β^* is unique by (A, B) in Assumption B.1, there must exist a sequence $\beta_s \in \mathcal{B}, \beta_s \to \beta^*$ such that $f(\beta_s) \geq o(||\beta_s - \beta^*||^2)$. Recall that twice continuously differentiability of $W(\beta)$, we have that $H(\beta_s) \to H(\beta^*)$. As a result, we can find, for $s \geq S$, for S large enough, a point in the sequence such that (since p is finite)

$$2f(\beta_s) \leq (\beta_s - \beta^*)^\top H(\beta^*)(\beta_s - \beta^*) + \delta(s)||\beta_s - \beta^*||^2,$$

for $\delta(s) = |\tilde{\lambda}_{\max}(s)|$, where $\tilde{\lambda}_{\max}(s)$ is the maximum eigenvalue of $H(\beta)\Big|_{\beta \in [\beta_s, \beta^*]} - H(\beta^*)$. Note that such decomposition holds by symmetry of $H(\beta)\Big|_{\beta \in [\beta_s, \beta^*]} - H(\beta^*)$. Since $H(\beta^*)$ is negative definite, the above expression is bounded as follows

$$2f(\beta_s) \leq -(|\tilde{\lambda}_{min}| - \delta(s))||\beta_s - \beta^*||^2,$$

where $|\tilde{\lambda}_{min}| > 0$ is the minimum eigenvalue of $H(\beta^*)$ (in absolute value) bounded away from zero by Assumption B.1 (C). By continuity of the Hessian, $\delta(s) \to 0$, and we reach a contradiction. This result implies strong concavity *locally* at the optimum. Here $\lambda_{\max} > \lambda_{\min}$ since the lower bound holds for any finite and large enough $\lambda_{\max}$.

Cases Define

$$\kappa = \frac{|\lambda_{max}|}{|\lambda_{min}|} \geq 1. \tag{C.7}$$

Observe now that if $||\beta_w^* - \beta^*||^2 \leq \epsilon\kappa$, the claim trivially holds. Therefore, consider the case where

$$||\beta_w^* - \beta^*||^2 > \epsilon\kappa.$$

Comparisons within the neighborhood Take $\tilde{\beta} = \beta^* - \sqrt{\epsilon}\frac{\nabla_w}{||\nabla_w||_2}$. Observe that

$$\begin{aligned} W(\tilde{\beta}) - W(\beta_w^*) &= \frac{1}{2}(\tilde{\beta} - \beta^*)^\top H(\tilde{\beta})\Big|_{[\beta^*, \tilde{\beta}]}(\tilde{\beta} - \beta^*) - \frac{1}{2}(\beta_w^* - \beta^*)^\top H(\beta_w^*)\Big|_{[\beta^*, \beta_w^*]}(\beta_w^* - \beta^*) \\ &\geq -|\lambda_{\max}|\epsilon + |\lambda_{\min}|\epsilon\kappa = 0. \end{aligned}$$

As a result, for all $\beta_w^* : ||\beta_w^* - \beta^*||^2 > \epsilon\kappa$, using quasi-concavity

$$\nabla_w^\top(\tilde{\beta} - \beta_w^*) \geq 0 \Rightarrow \nabla_w^\top(\beta^* - \beta_w^*) \geq \sqrt{\epsilon}||\nabla_w||_2 \tag{C.8}$$

Plugging in the above expression in the definition of β_w^* By construction of the algorithm, we write

$$||\beta^* - \beta_{w+1}^*||^2 \leq ||\beta^* - \beta_w^*||^2 - 2\alpha_w J \nabla_w^\top (\beta^* - \beta_w^*) + J^2 \alpha_w^2 ||\nabla_w||^2.$$

By Equation (C.8), we can write

$$||\beta^* - \beta_{w+1}^*||^2 \leq ||\beta^* - \beta_w^*||^2 - 2J\alpha_w \sqrt{\epsilon} ||\nabla_w||_2 + J^2 \alpha_w^2 ||\nabla_w||^2.$$

Plugging in the expression for α_w, and using the fact that $J \leq 1$, we have

$$||\beta^* - \beta_{w+1}^*||^2 \leq ||\beta^* - \beta_w^*||^2 - J\epsilon.$$

Recursive argument Recall that since the algorithm did not terminate, $||\beta^* - \beta_{\check{w}}^*||^2 > \epsilon\kappa$, for all $\check{w} \leq w$. Using this argument recursively, we obtain

$$||\beta^* - \beta_{\check{T}}^*||^2 \leq ||\beta^* - \beta_0||^2 - J \sum_{s=1}^{\check{T}} \epsilon = 2 \max_{\beta \in \mathcal{B}} ||\beta||^2 - J\check{T}^v \leq G + 1 - J\check{T}^v.$$

Whenever $\check{T} > (G/J + 1/J)^{1/v}$, we have a contradiction. The proof completes. □

Lemma C.12. *Let Assumption 2.1, 2.2, 4.1, B.1 hold. Assume that*

$$\epsilon_n \geq \sqrt{p}\Big[\bar{C}\sqrt{\gamma_N \frac{\log(\gamma_N \check{T} K/\delta)}{\eta_n^2 n}} + \eta_n\Big], \quad \frac{1}{4\mu \check{T}^{1/2-v/2}} - \epsilon_n \geq 0$$

for a finite constant $\bar{C} < 0$.

Then with probability at least $1-\delta$, for any $\delta \in (0,1)$ for any $w \leq \check{T}$,

$$\textit{either } (i) \Big|\Big|\check{\beta}_k^w - \beta_w^*\Big|\Big|_\infty = \mathcal{O}(P_w(\delta) + p\eta_n), \textit{ or } (ii) \Big|\Big|\check{\beta}_k^w - \beta^*\Big|\Big|_2^2 \leq \frac{p}{\check{T}^{1-v}}$$

where $P_1(\delta) = \mathrm{err}(\delta)$ and $P_w(\delta) = \frac{2\sqrt{p}}{\nu_n} B_p \frac{1}{\check{T}^{1/2-v/2}} P_{w-1}(\delta) + P_{w-1}(\delta) + \frac{2\sqrt{p}}{\nu_n} \frac{1}{\check{T}^{1/2-v/2}} \mathrm{err}(\delta)$, for a finite constant $B_p < \infty$, and $\mathrm{err}(\delta) = \mathcal{O}\Big(\sqrt{\gamma_N \frac{\log(\gamma_N p \check{T} K/\delta)}{\eta_n^2 n}} + p\eta_n\Big)$, with $\nu_n = \frac{1}{\mu \check{T}^{1/2-v/2}} - 2\epsilon_n$.

Proof of Lemma C.12. First, by Lemma 4.1, the estimated coefficients are exogenous. Hence, by invoking Lemma C.7 and the union bound, we can write for every k and t, $\delta \in (0,1)$,

$$\check{V}_{k,w}^{(j)} = V^{(j)}(\check{\beta}_{k+2}^w) + \mathcal{O}\Big(\sqrt{\gamma_N \frac{\log(\gamma_N K \check{T}/\delta)}{\eta_n^2 n}} + \eta_n\Big).$$

We now proceed by induction. We first prove the statement, assuming that the constraint is always attained. We then discuss the case of the constrained solution. Define

$$B = p \sup_{\beta} \left|\left| \frac{\partial^2 W(\beta)}{\partial \beta^2} \right|\right|_{\infty}.$$

Unconstrained case Consider $w = 1$. Then since all clusters start from the same starting point β_0 recall that ($\beta_1^* = \beta_0$), we can write with probability $1 - \delta$, by the union bound over p (which hence enters in the $\log(p)$ component of err_n) and Lemma C.7

$$\left|\left| \check{V}_{k,1} - V(\beta_1^*) \right|\right|_{\infty} \leq \text{err}(\delta). \tag{C.9}$$

Consider now the case where the algorithm stops. This implies that it must be that $||\check{V}_{k,1}||_2 \leq \frac{1}{\mu \check{T}^{1/2-v/2}} - \epsilon_n$. By Lemma C.7

$$||V(\beta_1^*)||_2 \leq ||\check{V}_{k,1}||_2 + \sqrt{p}\text{err}(\delta) \leq \frac{1}{\mu \check{T}^{1/2-v/2}} - \epsilon_n + \sqrt{p}\text{err}(\delta) \leq \frac{1}{\mu \check{T}^{1/2-v/2}}. \tag{C.10}$$

since $\epsilon_n \geq \sqrt{p}\text{err}(\delta)$. As a result, also the oracle algorithm stops at β_1^* by construction of ϵ_n. Suppose the algorithm does not stop. Then it must be that $||\check{V}_{k,1}|| \geq \frac{1}{\mu \check{T}^{1/2-v/2}} - \epsilon_n$ and

$$||V_1(\beta_1^*)|| \geq \frac{1}{\mu \check{T}^{1/2-v/2}} - \epsilon_n - \sqrt{p}\text{err}_1 \geq \frac{1}{\mu \check{T}^{1/2-v/2}} - 2\epsilon_n := \nu_n > 0.$$

Observe now that

$$\begin{aligned} \left|\left| \frac{\check{V}_{k,1}}{||\check{V}_{k,1}||_2} - \frac{V(\beta_1^*)}{||V(\beta_1^*)||_2} \right|\right|_{\infty} &\leq \left|\left| \frac{\check{V}_{k,1} - V(\beta_1^*)}{||V(\beta_1^*)||_2} \right|\right|_{\infty} + \left|\left| \frac{\check{V}_{k,1}(||\check{V}_{k,1}||_2 - ||V(\beta_1^*)||_2)}{||V(\beta_1^*)||_2 ||\check{V}_{k,1}||_2} \right|\right|_{\infty} \\ &\leq \left|\left| \frac{\check{V}_{k,1} - V(\beta_1^*)}{||V(\beta_1^*)||_2} \right|\right|_{\infty} + \sqrt{p}\left|\left| \frac{\check{V}_{k,1} - V(\beta_1^*)}{||V(\beta_1^*)||_2} \right|\right|_{\infty}. \end{aligned} \tag{C.11}$$

The last inequality follows from the reverse triangular inequalities and standard properties of the norms. Then with probability at least $1 - \delta$, for any $\delta \in (0, 1)$

$$(\text{C.11}) \leq \frac{1}{\nu_n} \times 2\sqrt{p}\text{err}(\delta).$$

completing the claim for $w = 1$. Consider now a general w. Define the error until time $w - 1$ as P_{w-1}.Then for every $j \in \{1, \cdots, p\}$, by Assumption 3.1, we have with probability at least

$1 - w\delta$ (using the union bound),

$$\check{V}_{k,w}^{(j)} = V^{(j)}(\check{\beta}_{k+2}^{w}) + \text{err}(\delta) = V^{(j)}(\beta_w^* + P_w(\delta)) + \text{err}(\delta)$$
$$\Rightarrow \Big|\Big|\check{V}_{k,w} - V(\beta_w^*)\Big|\Big|_\infty \leq BP_w(\delta) + \text{err}(\delta),$$

where the above inequality follows by the mean-value theorem and Assumption 3.1. Suppose now that $||\check{V}_{k,w}||_2 \leq \frac{1}{\mu \check{T}^{1/2-v/2}} - \epsilon_n$. Then for the same argument as in Equation (C.10), we have

$$||V(\check{\beta}_k^w)||_2 \leq \frac{1}{\mu \check{T}^{1/2-v/2}}.$$

Under Assumption B.1 (B) this implies that

$$||\check{\beta}_k^w - \beta^*||_2^2 \leq \frac{1}{\check{T}^{1-v}},$$

which proves the statement. Suppose instead that the algorithm does not stop. Then we can write by the induction argument

$$\Big|\Big|\check{\beta}_k^w + \frac{1}{\check{T}^{1/2-v/2}} \frac{\check{V}_{k,w}}{||\check{V}_{k,w}||_2} - \beta_w^* - \frac{1}{\check{T}^{1/2-v/2}} \frac{V(\beta_w^*)}{||V(\beta_w^*)||_2}\Big|\Big|_\infty \leq P_w(\delta) + \frac{1}{\check{T}^{1/2-v/2}} \underbrace{\Big|\Big|\frac{\check{V}_{k,w}}{||\check{V}_{k,w}||_2} - \frac{V(\beta_w^*)}{||V(\beta_w^*)||_2}\Big|\Big|_\infty}_{(B)}. \tag{C.12}$$

Using the same argument in Equation (C.11), we have with probability at least $1 - \delta$,

$$(B) \leq \frac{2\sqrt{p}}{\nu_n}\Big[\text{err}(\delta) + BP_w(\delta)\Big],$$

which completes the proof for the unconstrained case. The $\check{T}$ component in the error expression follows from the union bound across all $\check{T}$ events.

Constrained case Since the statement is true for $w = 1$, we can assume that it is true for all $s \leq w - 1$ and prove the statement by induction. Since $\mathcal{B}$ is a compact space, we can write

$$\begin{aligned}
&\Big|\Big|P_{\mathcal{B}_1,\mathcal{B}_2-\eta_n}\Big[\sum_{s=1}^{w} \alpha_{k,s}\check{V}_{k,s}\Big] - P_{\mathcal{B}_1,\mathcal{B}_2}\Big[\sum_{s=1}^{w} \alpha_s V(\beta_s^*)\Big]\Big|\Big|_\infty \\
&\leq \Big|\Big|P_{\mathcal{B}_1,\mathcal{B}_2-\eta_n}\Big[\sum_{s=1}^{w} \alpha_{k,s}\check{V}_{k,s}\Big] - P_{\mathcal{B}_1,\mathcal{B}_2-\eta_n}\Big[\sum_{s=1}^{w} \alpha_s V(\beta_s^*)\Big]\Big|\Big|_\infty + p\mathcal{O}(\eta_n) \\
&\leq 2\Big|\Big|\sum_{s=1}^{w} \alpha_{k,s}\check{V}_{k,s} - \sum_{s=1}^{w} \alpha_s V(\beta_s^*)\Big|\Big|_\infty + p\mathcal{O}(\eta_n).
\end{aligned}$$

For the first component in the last inequality, we follow the same argument as above. □

Lemma C.13. *Let the conditions in Lemma C.12 hold. Then with probability at least* $1-\delta$*, for any* $k \in \{1, \cdots, K\}$*, for any* $v \in (0,1), \delta \in (0,1)$*,* $\check{T} \geq \zeta^{1/v}$*, for* $\zeta < \infty$ *being a finite constant*

$$||\beta^* - \check{\beta}_k^{\check{T}}||_2^2 \leq \frac{\kappa}{\check{T}^{1-v}} + \check{T}e^{B_p\sqrt{p}\check{T}} \times \mathcal{O}\Big(\gamma_N \frac{\log(p\gamma_N \check{T}K/\delta)}{\eta_n^2 n} + p^2\eta_n^2\Big),$$

with $\kappa, B_p < \infty$ *being constants independent on* $(n, \check{T})$ *and* ϵ_n *as defined in Lemma C.12.*

Proof. We invoke Lemma C.12. Observe that we only have to check that the result holds for (i) in Lemma C.12, since otherwise the claim trivially holds. Using the triangular inequality, we can write

$$||\beta^* - \check{\beta}_k^{\check{T}}||_2^2 \leq ||\beta^* - \beta_{\check{T}}^*||_2^2 + ||\check{\beta}_k^{\check{T}} - \beta_{\check{T}}^*||_2^2.$$

The first component on the right-hand side is bounded by Lemma C.11, with $\check{T} \geq \zeta^{1/v}, \zeta$ being a constant defined in Lemma C.11.

Using Lemma C.12, we bound with probability at least $1-\delta$, the second component as follows

$$||\check{\beta}_k^{\check{T}} - \beta_{\check{T}}^*||_2^2 \leq p||\check{\beta}_k^{\check{T}} - \beta_{\check{T}}^*||_\infty^2 = p \times \mathcal{O}(P_{\check{T}}^2(\delta)).$$

We conclude the proof by explicitely defining recursively, for all $1 < w \leq \check{T}$,

$$P_w = (1 + \frac{2B_p\sqrt{p}}{\nu_n \check{T}^{1/2-v/2}})P_{w-1} + \frac{1}{\check{T}^{1/2-v/2}}\text{err}_n(\delta), \quad P_1 = \text{err}_n(\delta).$$

where $\text{err}_n(\delta) = \frac{2\sqrt{p}}{\nu_n}\mathcal{O}(\sqrt{\gamma_N \frac{\log(pTK/\delta)}{\eta_n^2 n}} + p\eta_n)$, and $B < \infty$ denotes a finite constant. Using a recursive argument, we obtain

$$P_w = \text{err}_n(\delta) \sum_{s=1}^{w} \alpha_s \prod_{j=s}^{w} (\frac{2B_p\sqrt{p}}{\nu_n \check{T}^{1/2-v/2}} + 1).$$

Recall now that $\nu_n \geq \frac{1}{2\mu\check{T}^{1/2-v/2}}$ as in Lemma C.12. As a result we can bound the above expression as

$$\sum_{s=1}^{w} \alpha_s \prod_{j=s}^{w} (\frac{2B_p\sqrt{p}}{\nu_n \check{T}^{1/2-v/2}} + 1) \leq \sum_{s=1}^{w} \alpha_s \prod_{j=s}^{w} (\frac{8\mu^2\check{T}^{1/2-v/2}B\sqrt{p}}{\check{T}^{1/2-v/2}} + 1) \leq \sum_{s=1}^{w} \alpha_s \exp\Big(\sum_{j=s}^{w} 8\mu^2 B_p\sqrt{p}\Big).$$

Now we have

$$\exp\Big(\sum_{j=s}^{w} 8\mu^2 B_p\sqrt{p}\Big) \leq \exp\Big(8\mu^2(w-s)B_p\sqrt{p}\Big).$$

We now write

$$P_w(\delta) \leq \mathrm{err}_n(\delta) \sum_{s=1}^{w} \alpha_s \exp\left(8\mu^2(w-s)B_p\sqrt{p}\right) \leq \mathrm{err}_n(\delta)\check{T}^{1/2+v} \exp\left(8\mu^2\check{T}B_p\sqrt{p}\right),$$

where we replaced w with $\check{T}$. The proof completes.

□

Corollary 8. *Theorem B.2 holds.*

Proof. Consider Lemma C.12 where we choose $\delta = 1/n$. Observe that we choose $\epsilon_n \leq \frac{1}{4\mu\check{T}^{1/2-v/2}}$, which is attained by the conditions in Lemma C.12 as long as n is small enough such that

$$\sqrt{p}\Big[\bar{C}\sqrt{\log(n)\gamma_N \frac{\log(p\gamma_N\check{T}K)}{\eta_n^2 n}} + \eta_n\Big] \leq \frac{1}{4\mu\check{T}^{1/2-v/2}}$$

attained under the assumptions stated in Lemma C.12. As a result, we have $\nu_n = \frac{1}{4\mu\check{T}^{1/2-v/2}}$. By Lemma C.13 for all k, with probability at least $1 - 1/n$,

$$||\check{\beta}_k^{\check{T}} - \beta^*||^2 \lesssim \frac{p}{\check{T}^{1-v}}.$$

Also, we have

$$||\beta^* - \frac{1}{K}\sum_k \check{\beta}_k^{\check{T}}||_2^2 \leq \frac{1}{K}\sum_k ||\check{\beta}_k^{\check{T}} - \beta^*||^2.$$

The proof concludes by Theorem C.13 and Assumption 3.1, after observing that

$$W(\beta^*) - W(\hat{\beta}^*) \lesssim ||\beta^* - \hat{\beta}^*||_2^2.$$

□

C.4.5 Proof of Theorem B.3

By Lemma 2.1, we can write

$$\mathbb{E}[\hat{\Delta}_k(\beta)] = m(1,1,\beta) - m(0,1,\beta) + O(\eta_n^2).$$

Following the same strategy as in the proof of Theorem 3.4, it is easy to show that

$$\mathbb{E}[\hat{S}(0,\beta)] = \frac{\partial m(0,1,\beta)}{\partial\beta} + \frac{1}{2}\Big[\alpha_{t,k} - \alpha_{t-1,k} - \alpha_{t,k+1} + \alpha_{t-1,k+1}\Big] + \mathcal{O}(\eta_n).$$

Similarly, we can write

$$\mathbb{E}[\hat{S}(1,\beta)] = \frac{\partial m(1,1,\beta)}{\partial \beta} + \frac{1}{2}\Big[\alpha_{t,k} - \alpha_{t-1,k} - \alpha_{t,k+1} + \alpha_{t-1,k+1}\Big] + \mathcal{O}(\eta_n).$$

The proof completes after noticing that $\frac{\partial m(1,1,\beta)}{\partial \beta} = 0$.

C.4.6 Proof of Theorem B.4

The proof follows directly from Lemma C.10, after noticing that every two periods, the function is evaluated at the same vector of parameter $\Gamma(\check{\beta}^w, \check{\beta}^w)$. Therefore, we can apply all our results to the function $\beta \mapsto \Gamma(\beta, \beta)$ which satisfies the same conditions as $W(\beta)$.

C.4.7 Proof of Theorem B.5

The proof mimics the proof of Theorem 4.3.

Consider Lemma C.10 where we choose $\delta = 1/n$. Note that we can directly apply Lemma C.10 also to the gradient estimated with Algorithm E.3, since, by the circular-cross fitting argument, each parameter $\check{\beta}_k^w$ is estimated using sequentially pairs of different clusters as in Algorithm E.3. The rest of the proof follows verbatim from the one of Theorem 4.3 and omitted for brevity.

C.4.8 Proof of Theorem B.6

We first study the expectation of the estimator. We can write

$$\mathbb{E}\Big[\widehat{\frac{\partial m(0,x,\beta)}{\partial \beta}}\Big] = \underbrace{\frac{1}{2n\eta_n}\sum_{i=1}^{n}\mathbb{E}\Big[Y_{i,1}^k w_i^k - Y_{i,1}^{k+1} w_i^{k+1}\Big]}_{(I)} + \underbrace{\frac{1}{2\eta_n}\mathbb{E}\Big[\bar{Y}_0^k - \bar{Y}_0^{k+1}\Big]}_{(II)}.$$

Using Lemma 2.1, and the design in Algorithm 1 (i.e., at $t = 0$, treatments are assigned with the same probability in clusters $(k, k+1)$), we can write

$$(II) = \frac{1}{2\eta_n}(\tau_k - \tau_{k+1}),$$

where τ_k denotes the cluster fixed effect in Lemma 2.1. We can now write, using the law of iterated expectations,

$$(I) = \frac{1}{2n\eta_n}\sum_{i=1}^{n} \underbrace{\mathbb{E}\Big[\Big[Y_{i,1}^k w_i^k \Big| X_i^k, D_{i,1}^k\Big]\Big]}_{(i)} - \underbrace{\mathbb{E}\Big[\Big[Y_{i,1}^{k+1} w_i^{k+1} \Big| X_i^{k+1}, D_{i,1}^{k+1}\Big]\Big]}_{(ii)}.$$

We study (i) and (ii) separately. Using Lemma 2.1, we can write

$$\mathbb{E}\Big[\Big[Y_{i,1}^k w_i^k \Big| X_i^k, D_{i,1}^k\Big]\Big] = \mathbb{E}\Big[\Big[m(D_{i,1}^k, X_i^k, \beta+\eta_n)\frac{(1-D_{i,1}^k)}{1-\beta-v_h\eta_n}\frac{1\{X_i^k = x\}}{P(X_i^k = x)}\Big| X_i^k, D_{i,1}^k\Big]\Big],$$

where we used the assignment mechanism in cluster k at $t = 1$ as in Algorithm 1 which implies that the parameter for assigning treatments in cluster k is $\beta + \eta_n$. We can now write

$$\begin{aligned}
(i) &= \mathbb{E}\Big[\Big[m(0, x, \beta+\eta_n)\frac{(1-D_{i,1}^k)}{1-\beta-v_h\eta_n}\frac{1\{X_i^k = x\}}{P(X_i^k = x)}\Big| X_i^k, D_{i,1}^k\Big]\Big] + \tau_k \\
&= \mathbb{E}\Big[\Big[m(0, x, \beta+\eta_n)\frac{(1-D_{i,1}^k)}{1-\beta-v_h\eta_n}\frac{1\{X_i^k = x\}}{P(X_i^k = x)}\Big| X_i^k\Big]\Big] + \tau_k \\
&= \mathbb{E}\Big[\Big[m(0, x, \beta+\eta_n)\frac{1\{X_i^k = x\}}{P(X_i^k = x)}\Big| X_i^k\Big]\Big] + \tau_k \\
&= \mathbb{E}\Big[m(0, x, \beta+\eta_n)\Big] + \tau_k = m(0, x, \beta+\eta_n) + \tau_k,
\end{aligned}$$

where in the first equality we substituted the input of $m(\cdot)$ with $(0, x)$, since the function multiplies for two indicators $(1 - D_{i,1}^k)1\{X_i^k = x\}$. In the second equality we average over the distribution of $D_{i,1}^k$ and use the assumption that $D_{i,1}^k$ is assigned unconditionally (i.e., does not depend on covariates).[80] In the third equation we average over the distribution of covariates. Using a similar reasoning, we can write

$$(ii) = m(0, x, \beta-\eta_n) + \tau_{k+1}.$$

Using a Taylor expansion around β, and combining (I) and (II), we can write

$$\begin{aligned}
(I) + (II) &= \frac{1}{2\eta_n}\Big[m(0, x, \beta) + \frac{\partial m(0, x, \beta)}{\partial \beta}\eta_n + O(\eta_n^2)\Big] - \frac{1}{2\eta_n}\Big[m(0, x, \beta) + \frac{\partial m(0, x, \beta)}{\partial \beta}(-\eta_n) + O(\eta_n^2)\Big] \\
&= \frac{\partial m(0, x, \beta)}{\partial \beta} + O(\eta_n).
\end{aligned}$$

[80] Note that the reasoning would also hold if $D_{i,1}^k$ was assigned based on covariates by however adjusting by the conditional propensity score.

To obtain the consistency result, we can directly invoke Lemma C.3, where, in the expression of the theorem, we also take into account the rescaling parameter $1/\eta_n$.

C.4.9 Proof of Proposition A.5

By concavity, we can write

$$W(\beta^*) - \frac{1}{K}\sum_{k=1}^{K} W(\beta^k) \geq W(\beta^*) - W(\frac{1}{K}\sum_{k=1}^{K}\beta^k) = W(\beta^*) - W(0.5),$$

which completes the proof, for a suitable choice of $W(\beta^*) - W(0.5)$ (e.g., a quadratic function with $\beta^* = 0.3$).

C.4.10 Proof of Lemma B.7

The proof follows similarly to the proof of Lemma C.4, here taking into account also the component $\omega_{i,j}$. Under Assumption 2.2, we can write for some function g,

$$r\Big(D_{i,t}^{(k)}, D_{j:A_{i,j}^{(k)}>0,t}^{(k)}, X_i^{(k)}, X_{j:A_{i,j}^{(k)}>0}^{(k)}, A_i^{(k)}, U_i^{(k)}, U_{j:A_{i,j}^{(k)}>0}^{(k)}, \nu_{i,t}^{(k)}\Big) = g(Z_{i,t}^{(k)}).$$

Here, $Z_{i,t}^{(k)}$ depends on $A_i^{(k)}$, i.e., the edges of individual i, and on unobservables and observables of all those individuals such that $A_{i,j}^{(k)} > 0$, namely,

$$Z_{i,t}^{(k)} = \Big[D_{i,t}^{(k)}, X_i^{(k)}, \nu_{i,t}^{(k)}, A_i^{(k)} \otimes \Big(X^{(k)}, U^{(k)}, D_t^{(k)}\Big)\Big].$$

Importantly, under Equation (B.6), $A_i^{(k)}$ is a function of $\Big\{\Big[X_j^{(k)}, U_j^{(k)}, \omega_{i,j}^{(k)}\Big], j : 1\{i_k \leftrightarrow j_k\} = 1\Big\}$, only, and each entry depends on $(X_j, U_j, X_i, U_i, \omega_{i,j})$ through the same function l for each individual. What is important, is that $\sum_j 1\{i_k \leftrightarrow j_k\} = \gamma_N^{1/2}$ for each unit i. Therefore, for some function $\tilde{g}$ (which depends on l in Assumption 2.1), we can equivalently write

$$Z_{i,t}^{(k)} = \tilde{g}(D_{i,t}^{(k)}, \nu_{i,t}^{(k)}, X_i^{(k)}, \tilde{Z}_{i,t}^{(k)}), \quad \tilde{Z}_{i,t}^{(k)} = \Big\{\Big[X_j^{(k)}, U_j^{(k)}, D_{j,t}^{(k)}, \omega_{i,j}^{(k)}\Big], j : 1\{i_k \leftrightarrow j_k\} = 1\Big\},$$

where $\tilde{Z}_{i,t}^{(k)}$ is the vector of $\Big[X_j^{(k)}, U_j^{(k)}, D_{j,t}^{(k)}, \omega_{i,j}\Big]$ of all individuals j with $1\{i_k \leftrightarrow j_k\} = 1$.

Now, observe that since $(U_i^{(k)}, X_i^{(k)}) \sim_{i.i.d.} F_{X|U}F_U$, $\omega_{i,j}$ are exogenous and independent as in Equation (B.7), $\{\nu_{i,t}\}$ are *i.i.d.* conditionally on $U^{(k)}, X^{(k)}, \omega^{(k)}$ (Assumption 2.2 and Equation (B.7)) and treatments are randomized as in Assumption 2.3, we have

$$G_{i,j}^{(k)} = \Big[X_j^{(k)}, U_j^{(k)}, \nu_{j,t}^{(k)}, D_{j,t}^{(k)}, \omega_{i,j}^{(k)}\Big]\Big|\beta_{k,t} \sim \mathcal{D}(\beta_{k,t})$$

are distributed with some distribution $\mathcal{D}(\beta_{k,t})$ which only depends on the exogenous coefficient $\beta_{k,t}$ governing the distribution of $D_{i,t}^{(k)}$ under Definition 2.3. Also, $G_{i,j}^{(k)}$ are independent across j (but not i) by the independence assumption of $\omega_{i,j}$. As a result for $\beta_{k,t}$ being exogenous, Lemma 2.1 holds since $\sum_j 1\{i_k \leftrightarrow j_k\} = \gamma_N^{1/2}$ for all i, hence $\tilde{Z}_{i,t}$ are identically (but not independently) distributed across units i, since $\tilde{Z}_{i,t}$ is a vector of γ_N $i.i.d$ random variables, each having the same marginal distribution which does not depend on i (therefore $\tilde{Z}_{i,t}$ has the same joint distribution across i).

To show the local dependence result, note that $G_{i,j}^{(k)}$ is mutually independent of $G_{u,v}^{(k)}$ for all $\{(v,u) : (v,u) \notin \{(i,j),(j,i)\}, u \neq j\}$ given the independent assumption of $\omega_{i,j}$ in Equation (B.7). In addition, $\varepsilon_{i,t}^{(k)}|\beta_{k,t}$ is a measurable function of a vector[81]

$$\left[X_j^{(k)}, U_j^{(k)}, \nu_{j,t}^{(k)}, D_{j,t}^{(k)}, \omega_{i,j}^{(k)}\right]_{j:1\{i_k \leftrightarrow j_k\}=1}.$$

As a result, under mutual independence of $G_{i,j}$ with $G_{u,v}$ for all $\{(v,u) : (v,u) \notin \{(i,j),(j,i)\}, u \neq j\}$, conditional on $\beta_{k,t}$, $\varepsilon_{i,t}^{(k)}$ is mutally independent with $\varepsilon_{v,t}^{(k)}$ for all v such that they are not connected and do not share a common element $\left[X_j^{(k)}, U_j^{(k)}, \nu_{j,t}^{(k)}, D_{j,t}^{(k)}, \omega_{v,j}^{(k)}\right]$, that is, such that $\max_j 1\{i_k \leftrightarrow j_k\}1\{v_k \leftrightarrow j_k\} = 0$ (here $1\{v_k \leftrightarrow v_k\} = 1$ for notational convenience). This holds since the edge between i and v is zero almost surely if $1\{v_k \leftrightarrow i_k\} = 0$.

There are at most $\gamma_N^{1/2} + \gamma_N$ many of $\varepsilon_{v,t}^{(k)}$ which can share a common neighbor with $\varepsilon_{i,t}^{(k)}$ ($\gamma_N^{1/2}$ many neighhbors and γ_N many neighbors of the neighbors), which concludes the proof.

C.5 Proof of the corollaries

In this subsection we provide proofs to the corollary which do not directly follow from the corresponding theorem.

C.5.1 Proof of Corollary 2

The result directly follows from Theorem B.1, here applied to $l = 1$. The reader may refer to the proof of Theorem B.1 for details.

C.5.2 Proof of Corollary 3

The corollary follows from Lemma C.3 and the triangular inequality. Note that the rate is Kn since, after pooling observations from clusters, we can equivalently interpret $\bar{\Delta}_n$ as the estimator from two clusters, each with $K/2 \times n$ observations.

[81] Here for notational convenience convenience only, we are letting $1\{i_k \leftrightarrow i_k\} = 1$.

C.5.3 Proof of Corollary 4

The corollary follows from a second-order Taylor expansion, using the assumption that the Hessian is uniformly bounded.

C.5.4 Proof of Corollary 6

From Theorem 3.2, we know that

$$\frac{(\hat{V}_{k,k+1} - V(\beta(\theta_k), \theta_k))}{\sqrt{\text{Var}(\hat{V}_{k,k+1})}} \to_d \mathcal{N}(0,1),$$

since $\theta_k = \theta_{k+1}$ by construction of the pairs (assumption in the corollary). Therefore, we can directly apply Theorem B.1, here applied to $l = 1$, since, under H_0, each estimated marginal effect (for each type) is centered around zero, and, the theorem does not require that clusters have the same variance. The reader may refer to the proof of Theorem B.1 for details.

Appendix D Numerical studies: additional results

D.1 One-wave experiment

In Figure D.3 we report the power plot for $\rho = 6$. In Figure D.4 we report the welfare gain from increasing β by 5% upon rejection of H_0 for $\rho = 6$. In Figure D.7 we report comparisons for different values of η_n.

D.2 Multiple-wave experiment

In Table D.1 the comparison with competitors for $\rho = 6$. Results are robust as in the main text.

In Figure D.5 we report a comparison among different learning rates, which are the one which rescales by $1/t$, the one that rescales by $1/\sqrt{T}$ and the one that rescales by $1/\sqrt{t}$. The best performing learning rate rescales the step size by a factor of order $1/\sqrt{t}$.

D.3 Calibrated experiment with covariates for cash transfers

In this subsection, we turn to a calibrated experiment where we also control for covariates, as discussed in Section 2.3. We use data from Alatas et al. (2012, 2016). We estimate a function heterogenous in the distance of the household's village from the district's center. We use information from approximately four hundred observations, whose eighty percent or

more neighbors are observed. We let $X_i \in \{0,1\}, X_i = 1$ if the household is far from the district's center than the median household, and estimate

$$Y_i|X_i = x \quad = \quad \phi_0 + \tilde{X}_i\tau + D_i\phi_{1,x} + \frac{\sum_{j\neq i} A_{j,i}D_j}{\max\{\sum_{j\neq i} A_{j,i}, 1\}}\phi_{2,x} + \Big(\frac{\sum_{j\neq i} A_{j,i}D_j}{\max\{\sum_{j\neq i} A_{j,i}, 1\}}\Big)^2\phi_{3,x} + \eta_i, \tag{D.1}$$

where η_i are unobservables centered on zero conditional on $X_i = x$, and $\tilde{X}_i$ denotes controls which also include X_i.[82] Using the estimated parameter, we can then calibrate the simulations as follows.

We let $\eta_{i,t}, \sim \mathcal{N}(0, \sigma^2)$, where σ^2 is the residual variance from the regression. We then generate the network and the covariate as follows:

$$A_{i,j} = 1\Big\{||U_i - U_j||_1 \leq 2\rho/\sqrt{N}\Big\}, \quad U_i \sim_{i.i.d.} \mathcal{N}(0, I_2), \quad X_i = 1\{U_i^{(1)} > 0\}.$$

Here, $U_i^{(1)}$ is continuous and captures a measure of distances. Individuals are more likely to be friends if they have similar distances from the center, and X_i is equal to one if an individual is far from the district's center from the median household. We fix $\rho = 1.5$ to guarantee that the objective's function optimum is approximately equal to the optimum observed from the data (in calibration, the optimum is $\beta \approx 0.26$, while $\beta^* \approx 0.29$ on the data).

We then generate data

$$Y_{i,t}|X_i = x \quad = \quad D_i\hat{\phi}_{1,x} + \frac{\sum_{j\neq i} A_{j,i}D_j}{\max\{\sum_{j\neq i} A_{j,i}, 1\}}\hat{\phi}_{2,x} + \Big(\frac{\sum_{j\neq i} A_{j,i}D_j}{\max\{\sum_{j\neq i} A_{j,i}, 1\}}\Big)^2\hat{\phi}_{3,x} + \eta_{i,t}. \tag{D.2}$$

where we removed covariates that did not interact with the treatment rule (i.e., do not affect welfare computations).

Our policy function is

$$\pi(x;\beta) = x\beta + (1-x)(1-\beta)$$

where β is the probability of treatment for individuals farer from the center. Here, we implicitely imposed a budget constraint $\beta P(X_i = 1) + (1-\beta)P(X_i = 0) = 1/2$, where, by construction $P(X_i = 1) = 1/2$.

We collect results for the one-wave experiment in Figure D.6, D.8 (left-panel), where we report power and the relative improvement from improving by 5% the treatment probability

[82] We also control for the education level, village-level treatments, i.e., how individuals have been targeted in a village (i.e., via a proxy variable for income, a community-based method, or a hybrid), the size of the village, the consumption level, the ranking of the individual poverty level, the gender, marital status, household size, the quality of the roof and top (which are indicators of poverty).

for people in remote areas as discussed in the main text. Welfare improvements (and power) are increasing in the cluster size and the number of clusters. However, such improvements are negligible as we increase clusters from twenty to forty, suggesting that twenty clusters are sufficient to achieve the largest welfare effects.[83] In the right-hand side panel of Figure D.8 we report the out-of-sample regret. The regret is generally decreasing in the number of iterations, especially as the regret is further away from zero. As the regret gets almost zero (0.06%), the regret oscillates around zero as the number of iterations increases due to sampling variation. This behavior is suggestive that for some applications, few iterations (in this case, ten) are sufficient to reach the optimum, up to a small error. In Table D.2, we observe perfect coverage for $n = 600$, and under-coverage by no more than five percentage points in the remaining cases.

D.4 Calibrated simulations to M-Turk experiment for information diffusion to increase vaccination

In this subsection, we study the performance of the methods to increase vaccination against COVID-19 through information diffusion.

Specifically, at the beginning of March 2021 (before the vaccination campaign was extensively implemented), we ran an M-Turk experiment where each individual was assigned either of two arms.[84] A control arm, which consisted of a survey asking basic questions on characteristics of the participants, and a treatment arm. The treatment arm was first assigned simple survey questions. Then, individuals under treatment were assigned three questions about COVID, whose correct answer was rewarded a small economic incentive. The goal of these questions is to increase awareness of the severity of COVID.[85] Each correct answer rewarded a small bonus and was displayed right after the participant submitted her answer to the three questions (before the end of the survey). In addition, at the end of the survey, participants were asked again one of the three questions, whose correct answer rewards a bonus. Participants were made aware of the bonuses and the survey's structure. The scope of the treatment was to increase awareness of the severity of the disease by asking questions and showing the correct answers to facilitate information transmission. At the end

[83] The order of magnitude of the welfare gain is smaller compared to simulations with the unconditional probability since, here, we always treat exactly half of the population. As a result, welfare oscillates between 0.24 and 0.29 only (as opposed to zero to one as in the unconditional case), as shown in Figure 2.

[84] The experiment was certified an IRB exempt by UCSD, Human Research Protections Program.

[85] These were multiple answers questions. The first question asked (i) Which of these events caused more deaths of Covid in the US? (more answers allowed), giving four options (World War I and II, 50 times more than 9/11, US Civil war); What is the percentage of people in the US who had Covid within the last year? (approximately); The number of people infected from Covid in the last year is comparable to ... (giving three options).

of the survey, both controls and treated units were asked when they would have done the vaccine. Our outcome of interest is binary and equals whether individuals would have done the vaccine either as soon as possible or during the spring. We estimate the model with 1035 participants.[86] We estimate treatment effects by running a simple linear regression, where the treatment dummy interacts with the dummy, indicating whether the individual classifies herself as liberal, conservative, or "prefer not to say". We find substantial heterogeneity, with positive effects on early vaccination on liberals only. We consider a model and policy function as in Section D.3, where X, in this case, denotes whether an individual is liberal or conservative (drawn with the same DGP as in Section D.3 for simplicity), and $\rho = 2$ as in the main text. We calibrate $\hat{\phi}_{1,x}$ to the estimated direct treatment effect and fix the percentage of treatment units to fifty percent. A challenge here is that we do not know spillovers ϕ_{2x}, ϕ_{3x}. Therefore, we choose $\phi_{3,x} = r\phi_{2,x}$, where r is estimated from data on information diffusion from Cai et al. (2015) for simplicity.[87] We choose $\phi_{2,x} + r\phi_{2,x} = \max\{\alpha\phi_{1,x}, 0\}$, i.e., total spillovers equal direct effect $\phi_{1,x}$ times a constant $\alpha \in \{0.1, 0.2, 0.3, 0.4\}$ if these are positive, and zero otherwise.

We run our one wave experiment with these calibrations and collect results in Figure D.9. In the figure, we report the relative welfare improvement of increasing treatment probabilities for liberals of ten percent, upon rejection of the null hypothesis H_0 in Equation (19). We note two facts: (i) clusters equal to twenty are sufficient to have sufficient power; (ii) as the size of the spillovers increase, welfare improvements upon rejection increase and similarly power. These results illustrate the benefit of the method, which, here, can lead to up to a twelve percentage points increase in early vaccination. It would be interesting in the future to test these results on real-world experiments run through online platforms.

D.5 Additional figures

[86] We collected information from 2411 participants. We removed 158 observations that had already received the vaccine in March 2020 and 203 observations that took less than thirty seconds and more than five minutes to take the survey. We also removed all those observations which were not living in the US.

[87] Other choices are possible but omitted for brevity.

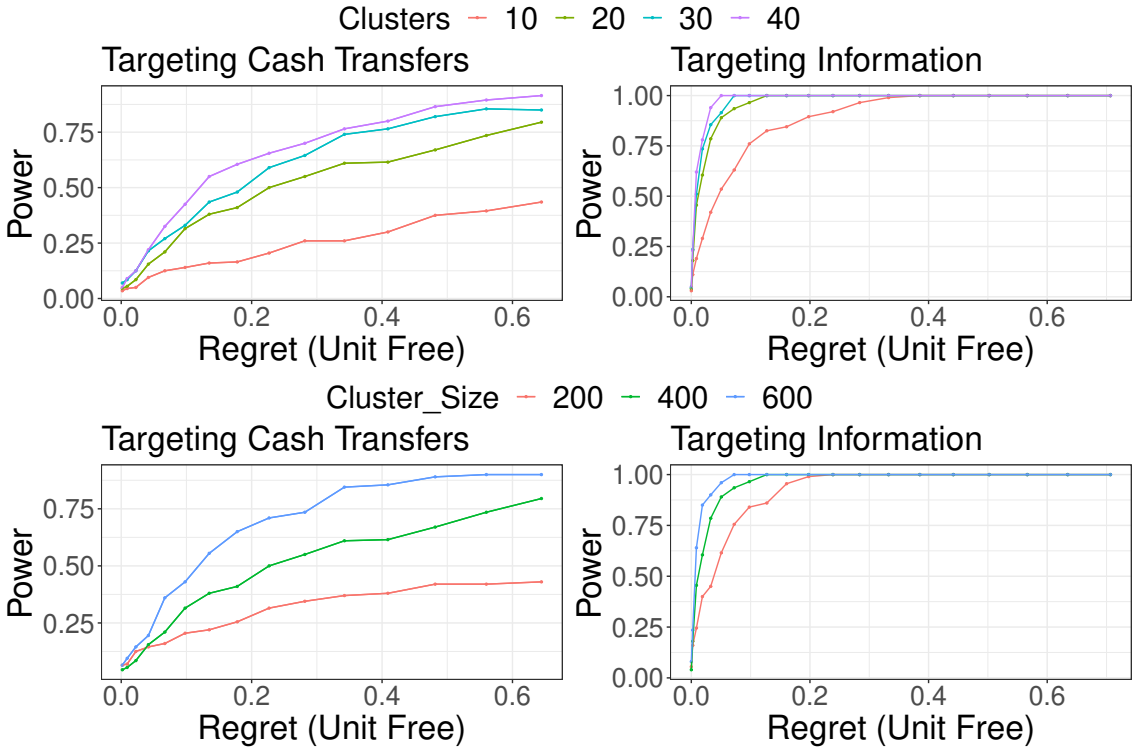


Figure D.1: One-wave experiment in Section 5. 200 replications. Power plot for $\rho = 2$. The panels at the top fix $n = 400$ and varies K. The panels at the bottom fix $K = 20$ and vary n.

Table D.1: Multiple-wave experiment in Section 5. Relative improvement in welfare with respect to best competitor for $\rho = 6$. The panel at the top reports the out-of-sample regret and the one at the bottom the worst case in-sample regret across clusters.

	Information				Cash Transfer			
$T =$	5	10	15	20	5	10	15	20
$n = 200$	-0.023	0.012	0.039	0.029	0.405	0.475	0.542	0.358
$n = 400$	0.001	0.020	0.019	0.029	0.546	0.483	0.602	0.548
$n = 600$	0.0004	0.030	0.022	0.020	0.571	0.481	0.614	0.643
$n = 200$	1.238	1.488	1.546	1.516	0.347	0.415	0.429	0.416
$n = 400$	1.494	1.690	1.704	1.624	0.482	0.576	0.550	0.498
$n = 600$	1.579	1.791	1.809	1.689	0.606	0.689	0.664	0.586

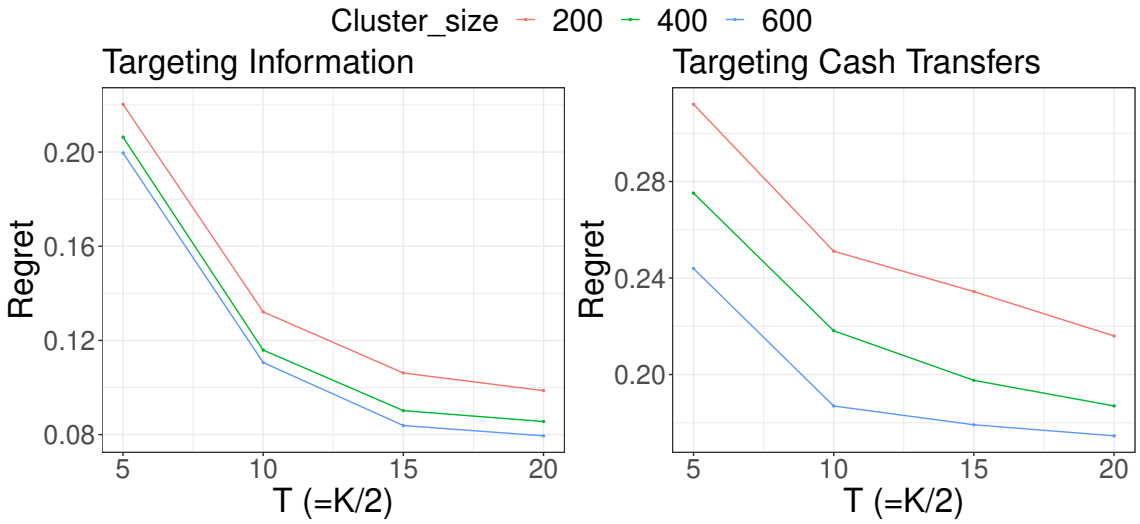


Figure D.2: Multi-wave experiment in Section 5. 200 replications. In-sample regret, average across clusters, for $\rho = 2$.

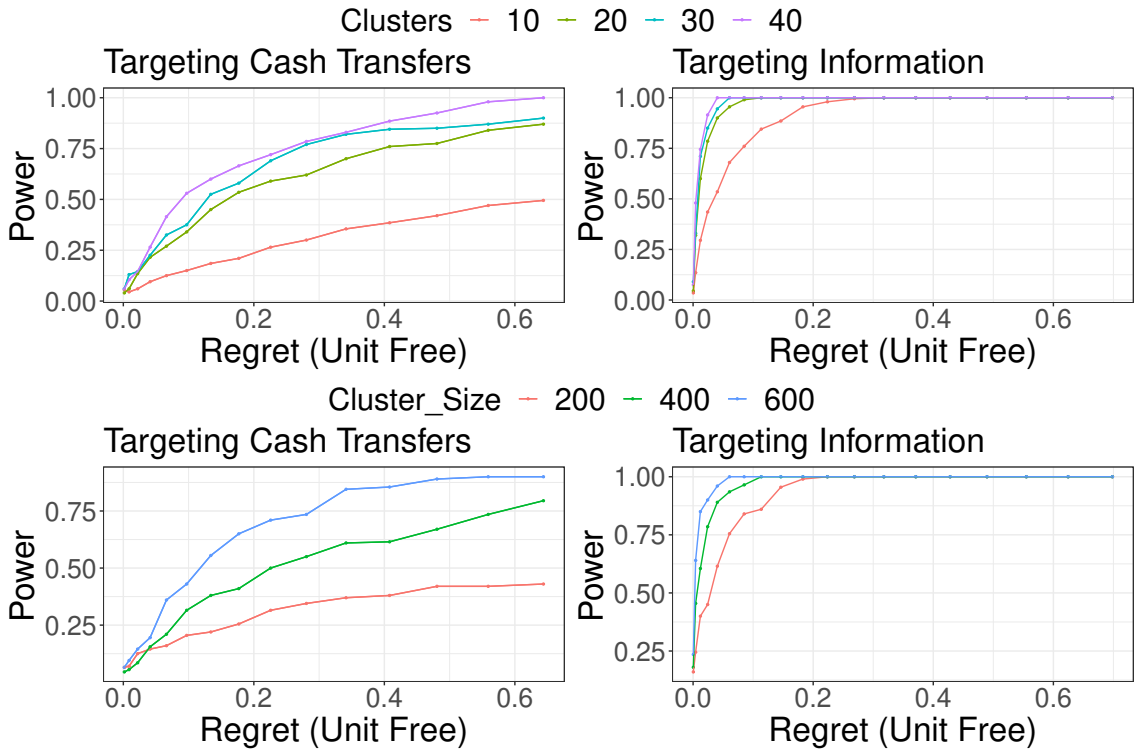


Figure D.3: One-wave experiment in Section 5. Power plot for $\rho = 6$. The panels at the top fix $n = 400$ and varies K. The panels at the bottom fix $K = 20$ and vary n.

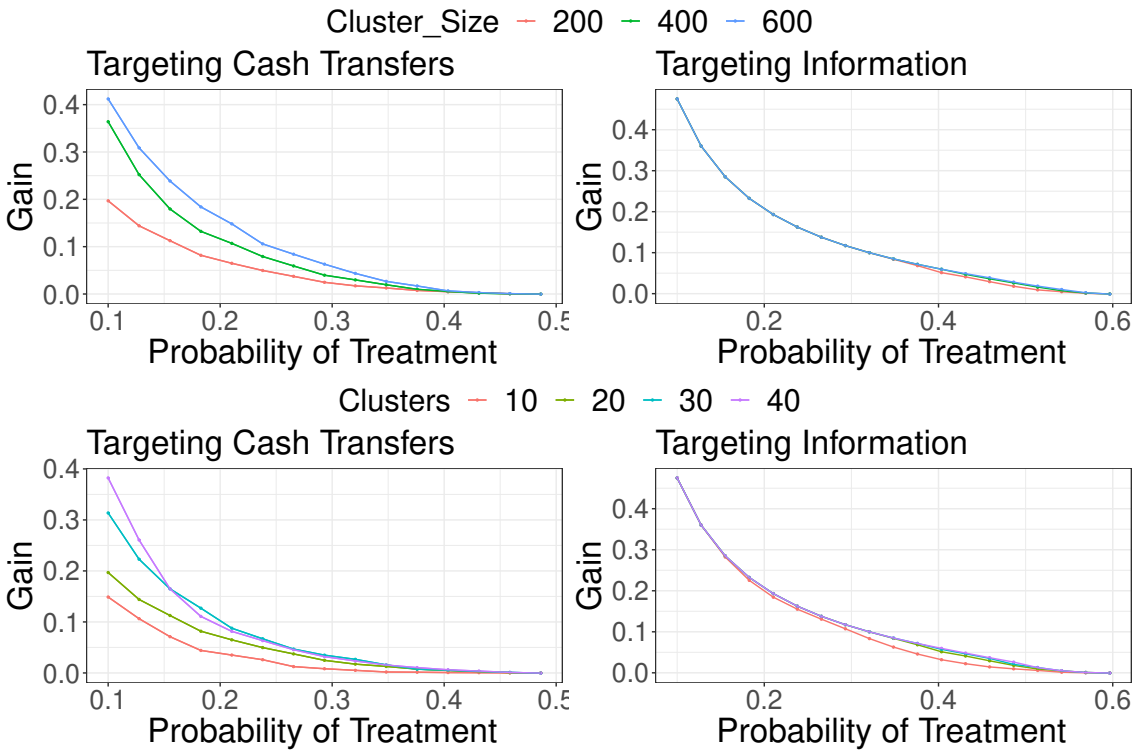


Figure D.4: One-wave experiment in Section 5. $\rho = 6$. Expected percentage increase in welfare from increasing the probability of treatment β by 5% upon rejection of H_0. Here, the x-axis reports $\beta \in [0.1, \cdots, \beta^* - 0.05]$. The panels at the top fix $n = 400$ and varies the number of clusters. The panels at the bottom fix $K = 20$ and vary n.

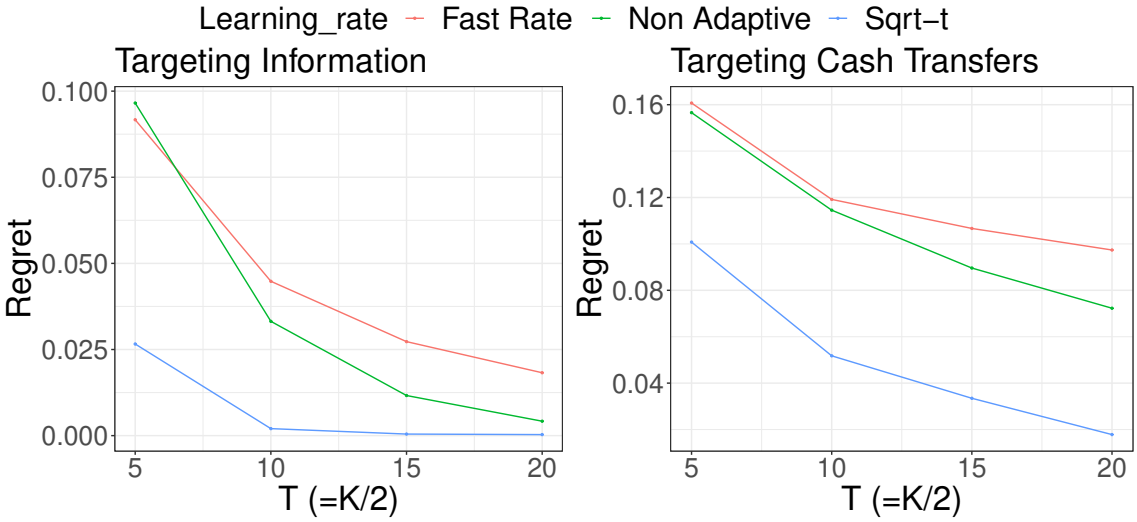


Figure D.5: Comparisons among different learning rates with experiment as in Section 5. 200 replications, $\rho = 2, n = 600, K = 2T$. Fast rate denotes a rescaling of order $1/t$; non-adaptive depends on a rescaling of order $1/\sqrt{T}$; the last one (Sqrt-t) depends on a rescaling of order $1/\sqrt{t}$.

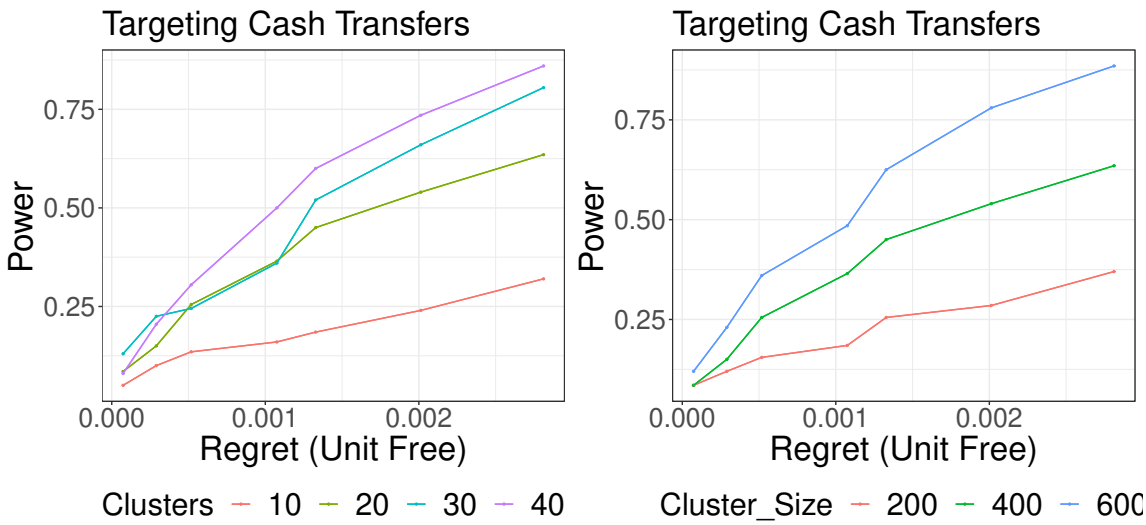


Figure D.6: Single-wave experiment in Section D.3. Power, 200 replications.

Table D.2: Single-wave experiment in Section D.3, 200 replications. Coverage for tests with size 5%.

$K =$	10	20	30	40
$n = 200$	0.940	0.950	0.895	0.900
$n = 400$	0.970	0.940	0.905	0.935
$n = 600$	0.950	0.970	0.950	0.940

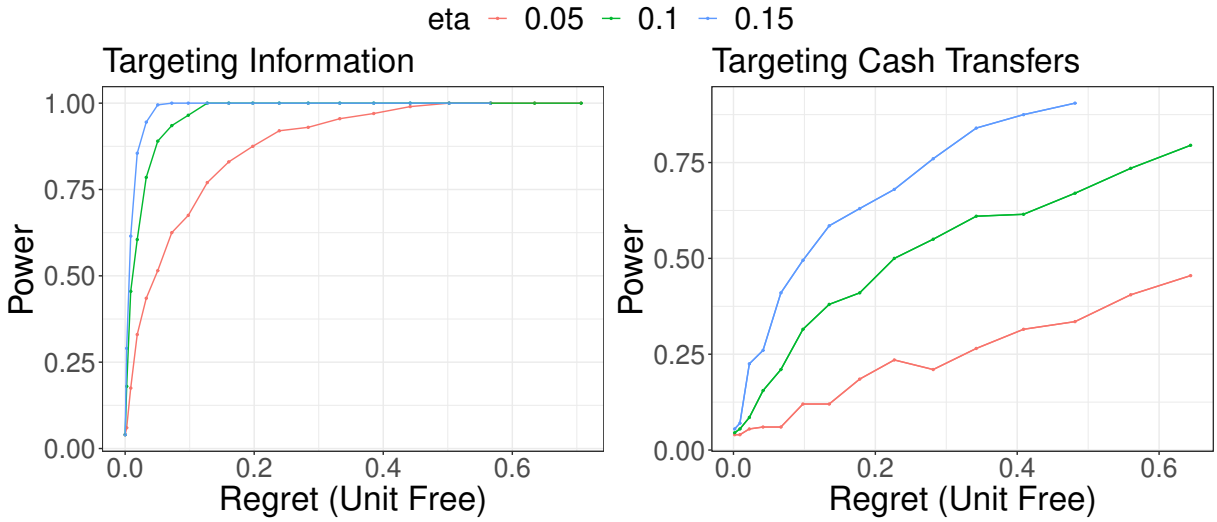


Figure D.7: One wave experiment calibrated to Alatas et al. (2012) and Cai et al. (2015). The plot reports power for different values of η_n varies, with $K = 200, n = 400$, with 200 replications.

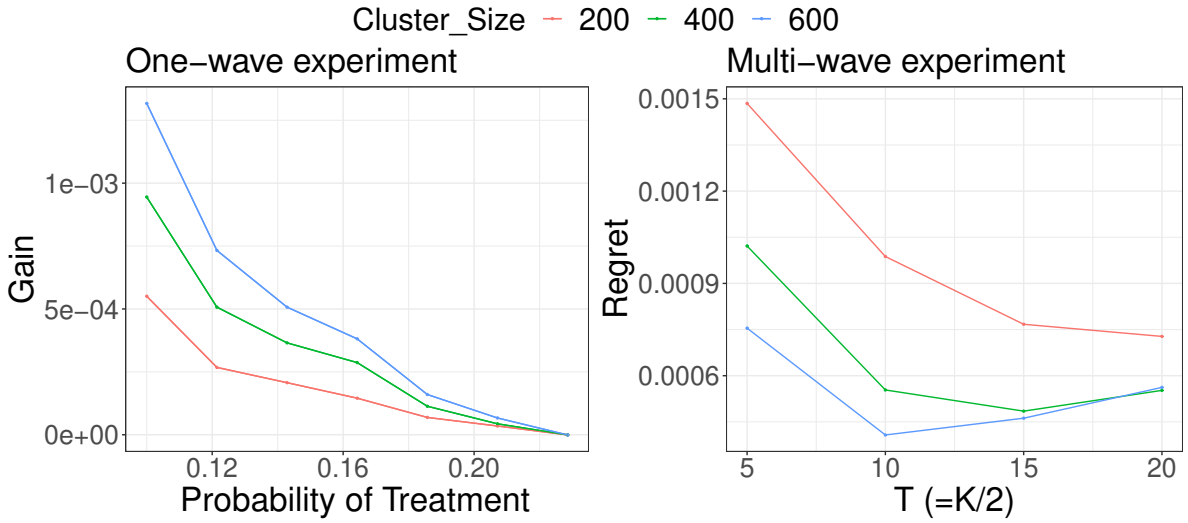


Figure D.8: Experiment in Section D.3. Left-hand side panel reports the expected percentage increase in welfare from increasing the probability of treatment β by 5% to individuals in remote areas upon rejection of H_0. Here, the x-axis reports $\beta \in [0.1, \cdots, \beta^* - 0.05]$. The right-hand side panel reports the in-sample regret. 400 replications.

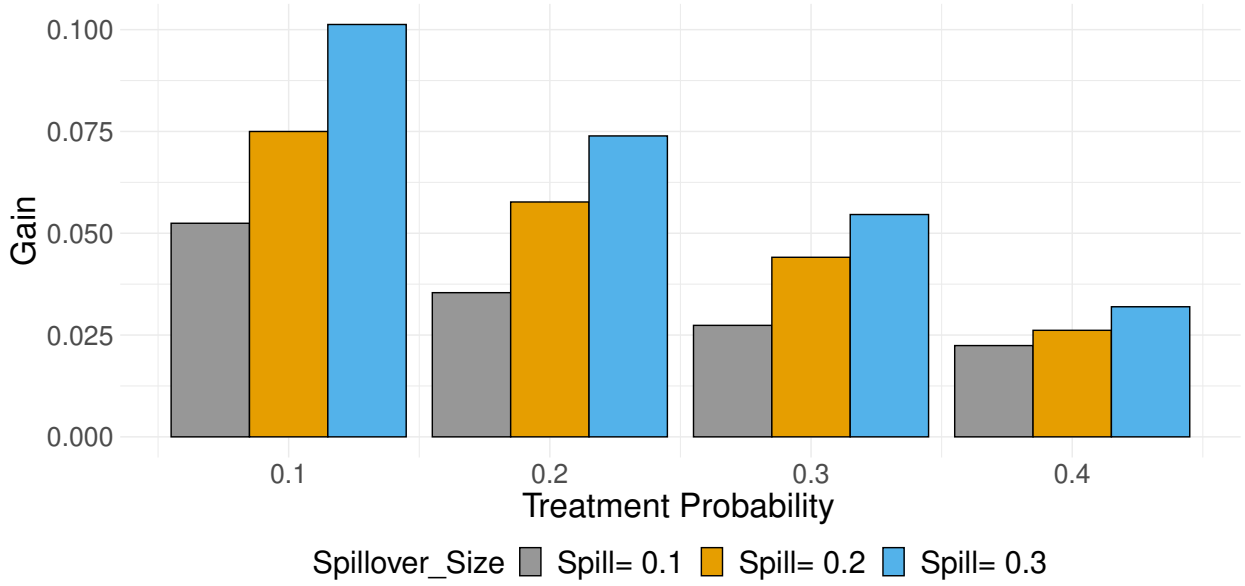


Figure D.9: One wave experiment in Section D.4. Relative welfare improvement for increasing the probability of treatment by ten percent, conditional on rejecting the null hypothesis. 200 replications. Different columns correspond to different levels of spillover effects (captured by α). Here $K = 20$.

Appendix E Additional Algorithms

Algorithm E.1 Adaptive Experiment with Many Coordinates

Require: Starting value $\beta_0 \in \mathbb{R}$, K clusters, $T+1$ periods of experimentation, constant $\bar{C}$.

1: Create pairs of clusters $\{k, k+1\}, k \in \{1, 3, \cdots, K-1\}$;

2: $t = 0 (baseline)$:

a: Assign treatments as $D_{i,0}^{(h)} | X_i^{(h)} = x \sim \pi(x; \beta_0)$ for all $h \in \{1, \cdots, K\}$.

b: For n units in each cluster observe $Y_{i,0}^{(h)}, h \in \{1, \cdots, K\}$.

c: For cluster k initalize a gradient estimate $\widehat{V}_{k,t} = 0$ and initial parameters $\check{\beta}_k^o = \beta_0$.

3: **while** $1 \leq w \leq \check{T} = \frac{T}{p}$ **do**

4: **for each** $j \in \{1, \cdots, p\}$ **do**

a: Define

$$\check{\beta}_h^w = \begin{cases} P_{\mathcal{B}_1, \mathcal{B}_2 - \eta_n}\left[\check{\beta}_h^{w-1} + \alpha_{h+2,w-1} \widehat{V}_{h+2,w-1}\right], & h \in \{1, \cdots, K-2\}, \\ P_{\mathcal{B}_1, \mathcal{B}_2 - \eta_n}\left[\check{\beta}_h^{w-1} + \alpha_{h+2,w-1} \widehat{V}_{1,w-1}\right], & h \in \{K-1, K\}. \end{cases}$$

Here, $P_{\mathcal{B}_1, \mathcal{B}_2 - \eta_n}$ denotes the projection operator onto the set $[\mathcal{B}_1, \mathcal{B}_2 - \eta_n]^p$.

b: Assign treatments as (for a finite constant $\bar{C}$)

$$D_{i,t}^{(h)} | X_{i,t}^{(h)} = x \sim \pi(x, \beta_{h,w}), \quad \beta_{h,w} = \begin{cases} \check{\beta}_h^w + \eta_n \underline{e}_j \text{ if } h \text{ is odd} \\ \check{\beta}_h^t - \eta_n \underline{e}_j \text{ if } h \text{ is even} \end{cases}, \quad \bar{C} n^{-1/2} < \eta_n < \bar{C} n^{-1/4}$$

where $\underline{e}_j$ is the vector of zero, with entry j equal to one (see Equation (9)).

c: For n units in each cluster $h \in \{1, \cdots, K\}$ observe $Y_{i,t}^{(h)}$.

d: For each pair $\{k, k+1\}$, estimate

$$\hat{V}_{k,w}^{(j)} = \hat{V}_{k+1,w}^{(j)} = \frac{1}{2\eta_n}\left[\bar{Y}_t^{(k)} - \bar{Y}_0^{(k)}\right] - \frac{1}{2\eta_n}\left[\bar{Y}_t^{(k+1)} - \bar{Y}_0^{(k+1)}\right].$$

e: $t \leftarrow t + 1$.

5: **end for**

f: $w \leftarrow w + 1$.

6: **end while**

7: Return $\hat{\beta}^* = \frac{1}{K} \sum_{k=1}^{K} \check{\beta}_k^{\check{T}}$

Algorithm E.2 Dynamic Treatment Effects with $\beta \in \mathbb{R}$

Require: Parameter space $\mathcal{B}$, clusters $\{1, \cdots, K\}$, two periods $\{t, t+1\}$, perturbation η_n.

1: Group clusters into groups $r \in \{1, \cdots, K/3\}$ of $\{k, k+1, k+2\}$;

2: Construct a grid of parameters $\mathcal{G} \subset [0,1]^2$ equally spaced on $[0,1]^2$;

3: Assign each parameter $(\beta_1^r, \beta_2^r) \in \mathcal{G}$ to a different triad r.

4: **for** each $r \in \{1, \cdots, K/3\}$ **do**

5: Randomize treatments as follows:

$$\begin{aligned} D_{i,t}^{(k)} | X_i^{(k)}, \beta_1^r, \beta_2^r &\sim \pi(X_i^{(k)}, \beta_2^r) \\ D_{i,t+1}^{(k)} | X_i^{(k)}, \beta_1^r, \beta_2^r &\sim \pi(X_i^{(k)}, \beta_1^r) \\ D_{i,t}^{(k+1)} | X_i^{(k)}, \beta_1^r, \beta_2^r &\sim \pi(X_i^{(k)}, \beta_2^r + \eta_n) \\ D_{i,t+1}^{(k+1)} | X_i^{(k)}, \beta_1^r, \beta_2^r &\sim \pi(X_i^{(k)}, \beta_1^r) \\ D_{i,t}^{(k+2)} | X_i^{(k)}, \beta_1^r, \beta_2^r &\sim \pi(X_i^{(k)}, \beta_2^r) \\ D_{i,t+1}^{(k+2)} | X_i^{(k)}, \beta_1^r, \beta_2^r &\sim \pi(X_i^{(k)}, \beta_1^r + \eta_n) \end{aligned} . \tag{E.1}$$

6: **end for**

7: For each $k \in \{1, 4, \cdots, K-2\}$ estimate

$$\widehat{g_{1,k}} = \frac{\bar{Y}_{t+1}^{(k)} - \bar{Y}_{t+1}^{(k+2)}}{\eta_n}, \quad \widehat{g_{2,k}} = \frac{\bar{Y}_{t+1}^{(k)} - \bar{Y}_{t+1}^{(k+1)}}{\eta_n}, \quad \widetilde{\Gamma}_k = \frac{1}{3} \sum_{h \in \{k, k+1, k+2\}} \bar{Y}_{t+1}^{(h)} \tag{E.2}$$

Algorithm E.3 Adaptive Experiment with staggered adoption

Require: Starting value $\beta \in \mathbb{R}$, K clusters, $T+1$ periods of experimentation.

1: Create pairs of clusters $\{k, k+1\}, k \in \{1, 3, \cdots, K-1\}$;

2: $t = 0$:

a: For n units in each cluster observe the baseline outcome $Y_{i,0}^{(h)}, h \in \{1, \cdots, K\}, \check{\beta}^0 = \beta$.

b: Initalize a gradient estimate $\widehat{V}_t = 0$

3: **while** $1 \leq t \leq T$ **do**

a: Sample without replaceent one pair of clusters $\{k, k+1\}$ not observed in previous iterations;

b: Define

$$\check{\beta}^t = \check{\beta}^{t-1} + \alpha_t \widehat{V}_t;$$

c: Assign treatments as

$$D_{i,t}^{(h)} \sim \pi(1, \beta_t), \quad \beta_t = \begin{cases} \check{\beta}^t + \eta_n \text{ if } h \text{ is even} \\ \check{\beta}^t - \eta_n \text{ if } h \text{ is odd} \end{cases}, \quad n^{-1/2} < \eta_n \leq n^{-1/4}$$

d: For n units in each cluster $h \in \{1, \cdots, K\}$ observe $Y_{i,t}^{(h)}$.

4: **end while**

5: Return $\hat{\beta}^* = \check{\beta}^T$

Algorithm E.4 Welfare maximization with a "non-adaptive" experiment

Require: K clusters, $T = p$ periods of experimentation.

1: Create pairs of clusters $\{k, k+1\}, k \in \{1, 3, \cdots, K-1\}$;

2: $t = 0$:

a: For n units in each cluster observe the baseline outcome $Y_{i,0}^{(h)}, h \in \{1, \cdots, K\}$.

b: Assign each pair $(k, k+1)$ to an element $\beta^k \in \mathcal{G}$, where $\mathcal{G}$ is an equally spaced grid of $\mathcal{B}$

3: **while** $1 \leq t \leq T$ **do**

a: Assign treatments as

$$D_{i,t}^{(h)} \sim \pi(1, \beta^h), \quad \beta^h = \begin{cases} \check{\beta}^h + \eta_n \underline{e}_t \text{ if } h \text{ is even} \\ \check{\beta}^{h-1} - \eta_n \underline{e}_t \text{ if } h \text{ is odd} \end{cases}, \quad n^{-1/2} < \eta_n \leq n^{-1/4}$$

b: For n units in each cluster $h \in \{1, \cdots, K\}$ observe $Y_{i,t}^{(h)}$.

c: Constructs the t entry

$$\hat{V}_{(k,k+1)}^{(t)}(\beta^k) = \frac{1}{2\eta_n}\left[\bar{Y}_t^k - \bar{Y}_0^k\right] - \frac{1}{2\eta_n}\left[\bar{Y}_t^{k+1} - \bar{Y}_0^{k+1}\right] \tag{E.3}$$

for each pair $(k, k+1)$

4: **end while**

5: Return $\hat{\beta}^{ow}$ as in Equation (A.7).