

# Causal Mediation in Natural Experiments

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## Abstract

Natural experiments are a cornerstone of applied economics, providing settings for estimating causal effects with a compelling argument for treatment randomisation. Applied researchers often investigate mechanisms with a suggestive investigation of mechanisms, giving shaky evidence for mechanisms behind causal effects. Causal Mediation (CM) provides an alternative, robust framework for identifying and estimating direct and indirect effects (CM effects) for causal mechanisms, but relies on assumptions which are implausible in natural experiment settings. In particular, conventional CM estimates rely on assuming the mediator is as-good-as-randomly assigned. When individuals select into a mediator based on costs and benefits, this assumption fails, undermining causal inference. I develop an alternative approach to credibly estimate CM effects, using control function methods and instrumental variation in take-up costs to avoid unrealistic assumptions. Simulations confirm this approach corrects for bias in conventional CM estimates, providing parametric and semi-parametric methods. This approach gives applied researchers an alternative method to estimate CM effects when an initial treatment is quasi-randomly assigned, but the mediator is not, as is common in natural experiments.

**Keywords:** Direct/indirect effects, quasi-experiment, selection, control function, MTEs.

**JEL Codes:** C21, C31.

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Economists use natural experiments to credibly answer social questions, when an experiment was infeasible. For example, does access to health insurance causally improve health and well-being (Finkelstein, Taubman, Wright, Bernstein, Gruber, Newhouse, Allen, Baicker & Group 2012)? Natural experiments are settings which answer these questions, but give little indication of how these effects came about. Causal Mediation (CM) aims to estimate the mechanisms behind causal effects, by estimating how much of the treatment effect operates through a proposed mediator. For example, do causal gains from access to health insurance come mostly from starting to utilise healthcare more often, or are there other direct effects? This study of mechanisms behind causal effects broadens the economic understanding of social settings studied with natural experiments. This paper shows that the conventional approach to estimating CM effects is inappropriate in a natural experiment setting, provides a theoretical framework for how bias operates, and develops an approach to correctly estimate CM effects under alternative assumptions. These methods contrast the current practice in applied economics of providing suggestive evidence of mechanisms, which does not identify or quantify causal effects.

This paper starts by considering conventional CM methods in a natural experiment setting. Conventional CM methods rely on assuming the initial treatment, and the subsequent mediator, are both ignorable (Imai, Keele & Yamamoto 2010). In a natural experiment setting, however, this may not hold true; winners of the Oregon Health Insurance Experiment wait-list lottery randomly received access to health insurance, but then chose of their own free will how often to use healthcare. Here, conventional CM estimates for the average direct and indirect effects will be contaminated by bias terms, from not accounting for selection-into-healthcare. Indeed, it is unlikely conventional CM estimates in any natural experiment will lead to credible causal effects, unless researchers use another natural experiment to isolate random variation in the mediator (at the same time as using one for the initial treatment).

I consider an alternative approach to estimating CM effects, adjusting for unobserved selection-into-mediator via identifying the marginal treatment effect of the mediator. This solves the identification problem with structural assumptions for selection-into-mediator — mediator monotonicity and selection based on benefits — and requires a valid cost instrument for mediator take-up. While these assumptions are strong, they are plausible in many applied settings. Mediator monotonicity aligns with conventional theories for selection-into-treatment, and is accepted widely in many applications using an instrumental variables research design. Selection based on costs and benefits is central to economic theory, and is the dominant concern for judging empirical designs that identify causal effects. Access to valid instrumental variation is a strong condition, though is important to avoid further modelling assumptions; the most compelling example is using variation in mediator take-up costs as an instrument.

This approach is not perfect in every setting: the structural assumptions are strong, and are tailored to selection-into-mediator concerns pertinent to economic applications. Indeed, this approach provides no safe harbour for estimating CM effects if these structural assumptions do not hold true.

Assuming the mediator is quasi-randomly assigned conveniently ignores selection by assuming either (1) people naïvely made decisions to take or refuse a mediator, or (2) a researcher controlled for everything relevant to this decision. This assumption might be reasonable when studying single-celled organisms in a laboratory — their “decisions” are simple and mechanical. Social scientists, however, study humans who make complex choices based on costs, benefits, and preferences — which are only partially observed by researchers, at best. Assuming a mediator is ignorable in social science contexts is often unrealistic. In practice, the main setting where mediator ignorability becomes credible is when researchers find another natural experiment affecting the mediator — a rare occurrence given how difficult it is to find one source of random variation for a treatment, let alone another independent source for a mediator, at the same time.

The applied economics literature has been hesitant to use explicit CM methods, instead providing suggestive evidence for mechanisms,<sup>1</sup> sometimes accompanied by a practice of controlling for a proposed mediator. Neither of these approaches have a causal interpretation. A new strand of the econometric literature has developed estimators for explicit CM analyses under a variety of strategies. These include overlapping quasi-experimental research designs (Deuchert, Huber & Schelker 2019, Frölich & Huber 2017), functional form restrictions (Heckman & Pinto 2015, Heckman, Pinto & Savelyev 2013), partial identification (Flores & Flores-Lagunes 2009), or a hypothesis test of full mediation through observed channels (Kwon & Roth 2024) — see Huber (2020) for an overview.<sup>2</sup> The new literature has arisen in implicit acknowledgement that suggestive evidence of mechanisms, or a conventional approach to CM, can lead to biased inference and needs alternative methods for credible inference.

I develop a framework showing exactly how selection bias contaminates CM estimates when mediator choices are driven by unobserved gains — settings where none of the natural experiment research designs in the previously cited papers apply (i.e., the mediator is not ignorable). This provides a rigorous warning to applied economists against uncritically applying conventional CM methods to investigate mechanisms in natural experiments — as is common in the applied fields of epidemiology, psychology, and medicine. Selection based

<sup>1</sup>See Blackwell, Ma & Opacic (2024) for an overview of this approach, from the empirical politics literature.

<sup>2</sup>An alternative method to estimate CM effects is ensuring treatment and mediator ignorability holds by a running two randomised controlled trials for both treatment and mediator, at the same time. This set-up has been considered in the literature previously, in theory (Imai, Tingley & Yamamoto 2013) and in practice (Ludwig, Kling & Mullainathan 2011).

on costs and benefits, as in the [Roy \(1951\)](#) model, is at odds with assuming a mediator is randomly assigned in an observational setting, so I import methods grounded in labour economic theory to solve the identification problem.

Identifying CM effects via the marginal treatment effect of the mediator requires mediator take-up respond only positively to the initial treatment (monotonicity), which implies mediator selection follows a selection model. Second, it assumes that mediator take-up is motivated by mediator benefits. Last, it requires a valid instrument for mediator take-up, to avoid relying on parametric assumptions on unobserved selection. This approach to identifying CM effects imports insights from the instrumental variables literature, connecting the influential [Imai et al. \(2010\)](#) approach to CM with the economics literature on selection-into-treatment and marginal treatment effects ([Vytlacil 2002](#), [Heckman & Navarro-Lozano 2004](#), [Heckman & Vytlacil 2005](#), [Florens, Heckman, Meghir & Vytlacil 2008](#), [Kline & Walters 2019](#)). [Frölich & Huber \(2017\)](#) have previously explored identification of CM effects with a control function in the context of two instruments (one each for treatment and mediator) and a continuous mediator. This paper considers the marginal treatment effect of a binary mediator, with a correspondingly different identification analysis and resulting estimation strategies.

This paper proceeds as follows. [Section 1](#) describes the dominant approach in economics for studying mechanisms behind treatment effects, illustrating with data from the Oregon Health Insurance Experiment. [Section 2](#) introduces the formal framework for CM, and develops expressions for bias in CM estimates in natural experiments. [Section 3](#) describes this bias in applied settings with (1) a regression framework, (2) a setting with selection based on costs and benefits. [Section 4](#) purges bias from CM estimates by identifying CM effects via the marginal treatment effect of the mediator, with a control function adjustment. [Section 5](#) demonstrates how to estimate CM effects with this approach, with either parametric or semi-parametric methods, giving supporting simulation evidence. [Section 6](#) returns to the Oregon Health Insurance Experiment, providing credible estimates of access to health insurance effects on self-reported health and well-being mediated through healthcare usage. [Section 7](#) concludes.

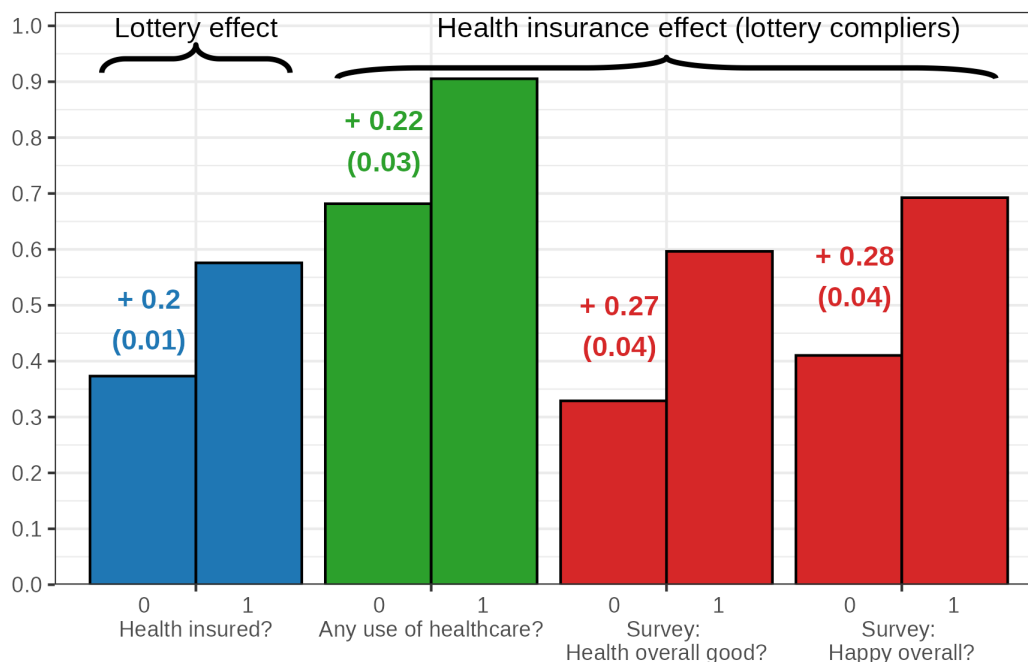
## 1 Mechanisms in the Oregon Health Insurance Experiment

In the United States, healthcare is generally not provided directly by the government. Instead, consumers purchase health insurance to fund healthcare expenses, with the government providing insurance only for elderly individuals (Medicare) and for those with low-incomes

(Medicaid). In 2004, the state of Oregon ceased accepting new applications for Medicaid due to budgetary constraints, and did not reopen applications until 2008. When the state resumed enrolment, approximately 90,000 individuals applied, vastly exceeding the programme's capacity. Oregon therefore allocated the opportunity to apply for Medicaid via a lottery system among those on the wait-list. This wait-list lottery significantly increased healthcare usage, plus self-reported health and happiness, among those who won the wait-list lottery.

**Figure 1:** Effects of Health Insurance in the Oregon Health Insurance Experiment.

Mean Outcome, for each  $z' = 0, 1$ .



**Note:** This figure summarises the relevant results of the Oregon Health Insurance Experiment (Finkelstein et al. 2012). The lottery results show that winning the Medicaid wait-list lottery increased health insurance rate by 19 percentage points. The effect of health insurance shows estimates of the average lottery complier health insurance effect on surveyed outcomes. It uses the wait-list lottery as an instrument for having health insurance, and the Abadie (2003) weighting scheme to estimate the average lottery complier levels,  $\mathbb{E}[Y_i(z', \cdot) | \text{lottery complier}]$  for each  $z' = 0, 1$ , with standard errors calculated by the bootstrap in brackets.

Winning the wait-list lottery increased the average health insurance coverage rate by 19 percentage points (pp). This gave the opportunity to use the lottery as an instrument for having health insurance at all,<sup>3</sup> among the 11,126 people from the wait-list lottery who responded to a survey sent by Finkelstein et al. (2012) one year later.<sup>4</sup> Among the wait-list lottery compliers, having health insurance increased the rate of having visited the doctor

<sup>3</sup>This additionally assumes everyone responded was at least more likely to take health insurance if they won the wait-list lottery, and the lottery did not have other direct effects.

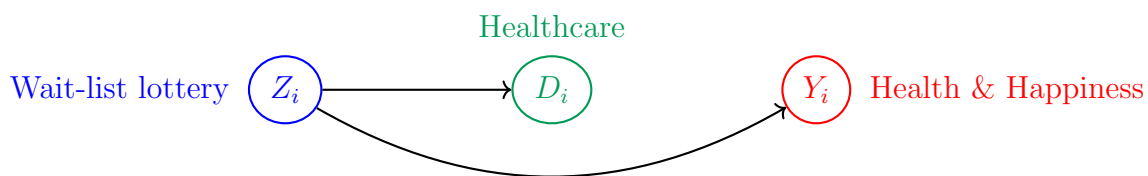
<sup>4</sup>This number restricts to those who gave non-missing answers to all relevant questions.

(either locally or at a hospital) by 25 pp. Furthermore, gaining insurance led to significant improvements in reported well-being: lottery compliers were 25 pp more likely to report being in good overall health, and 31 pp more likely to report being happy overall. These statistics were calculated using the Oregon Health Insurance Experiment replication package (Finkelstein & Baicker 2014), and are summarised in Figure 1.

These results show that health insurance led to large self-reported gains in both health and happiness. The economics, medicine, and health policy literatures have primarily focused on the health benefits — often interpreted as healthcare benefits among lottery compliers, who had previously delayed or avoided care due to cost concerns. However, the original authors also noted other benefits of insurance, including complete elimination of catastrophic out-of-pocket medical debt among the treated. These are plausibly income effects that benefit recipients directly, not only through increased use of healthcare, but also by reducing stress and improving financial security. These plausible direct effects have not been explored in the applied literature.

Accepted practice in applied economics is to investigate mechanisms behind causal effects with suggestive evidence. This involves estimating the average causal effect of health insurance on a proposed mediator (healthcare usage) and separately estimating its effect on the final outcomes (self-reported health and happiness). When both estimates are positive, and the mediator precedes the outcome, it is taken as de facto evidence that the mediator transmits the treatment effect. In the case of the Oregon Health Insurance Experiment, this amounts to concluding that increased healthcare usage mediates the positive effects of health insurance on health and happiness. Figure 2 illustrates this approach, which is also prevalent in other social science fields — see Blackwell et al. (2024), Green, Ha & Bullock (2010).

**Figure 2:** Structural Causal Model for Suggestive Evidence of a Mechanism.



**Note:** This figure shows the structural causal model behind an suggestive analysis for effects of access to health insurance, where arrows represent causal effects — e.g.,  $Z_i \rightarrow D_i$  means  $Z_i$  affects  $D_i$  with no reverse causality. The variables correspond to the causal design in the Oregon Health Insurance Experiment. Finkelstein et al. (2012) use the wait-list lottery as an IV for health insurance (which is not randomly assigned), while this paper focuses on the average effects of winning the wait-list lottery (which is randomly assigned).

There are two main problems with this approach. First, it provides no evidence for the

effect of healthcare on health and well-being, so does not identify the causal mechanism. Claims of identifying the mechanism (even suggestively) would require a hidden assumption that healthcare positively affects health outcomes. While this assumption is not unreasonable, nowhere else in applied economics is a hidden assumption of a positive average causal effect taken at face value — it must be motivated with quantitative evidence. Second, this approach does not quantify the mechanism effects. The proportion of health insurance benefits operating through healthcare could only have a small effect on self-reported outcomes recorded only 12 months later — or possibly a very large effect, it is a priori unclear. In addition, the relevant mediating effect is not the average effect of healthcare usage, but the effect for Oregon residents who were induced to use more healthcare after winning the wait-list lottery. This local effect could differ substantially from a population average, and potentially mislead conclusions about the magnitude or generality of the mechanism. In summary, this approach is not suggestive evidence of mechanisms, it is assumptive evidence of mechanisms; it compels claims about mechanisms behind treatment effects which are not motivated by causal evidence.

Causal Mediation (CM) offers a compelling alternative framework. Unlike suggestive evidence of mechanisms, CM explicitly defines the estimands of interest (the average direct and indirect effects) and lays out clear assumptions under which these quantities are identified. Moreover, it delivers quantitative answers to the key question: how much of a treatment effect operates through a specific mediator mechanism? CM is widely used in fields such as epidemiology, psychology, and medicine where researchers regularly decompose treatment effects into component pathways. However, CM methods have not yet been rigorously examined from an economic perspective to assess their applicability in observational causal research, such as natural experiments.

## 2 Causal Mediation (CM)

CM decomposes causal effects into two channels, through a mediator (indirect effect) and through all other paths (direct effect). To develop notation, write  $Z_i = 0, 1$  for a binary treatment,  $D_i = 0, 1$  a binary mediator, and  $Y_i$  a continuous outcome.<sup>5</sup>  $D_i, Y_i$  are a sum of

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<sup>5</sup>This paper exclusively focuses on the binary case. See [Huber, Hsu, Lee & Lettry \(2020\)](#) or [Frölich & Huber \(2017\)](#) for a discussion of CM with continuous treatment and/or mediator, and the assumptions required.



their potential outcomes,

$$\begin{aligned} D_i &= (1 - Z_i)D_i(0) + Z_iD_i(1), \\ Y_i &= (1 - Z_i)Y_i(0, D_i(0)) + Z_iY_i(1, D_i(1)). \end{aligned}$$

Assume treatment  $Z_i$  is ignorable.<sup>6</sup>

$$Z_i \perp\!\!\!\perp D_i(z'), Y_i(z, d'), \text{ for } z', z, d' = 0, 1$$

There are only two average effects which are identified without additional assumptions.

1. The average first-stage refers to the effect of the treatment on mediator,  $Z_i$  on  $D_i$ :

$$\mathbb{E}[D_i | Z_i = 1] - \mathbb{E}[D_i | Z_i = 0] = \mathbb{E}[D_i(1) - D_i(0)].$$

It is common in the economics literature to assume that  $Z_i$  influences  $D_i$  in at most one direction,  $\Pr(D_i(0) \leq D_i(1)) = 1$  — monotonicity (Imbens & Angrist 1994). I assume mediator monotonicity (and its conditional variant) holds throughout to simplify notation.

2. The Average Treatment Effect (ATE) refers to the effect of the treatment on outcome,  $Z_i$  on  $Y_i$ , and is also known as the average total effect or intent-to-treat effect in social science settings, or reduced-form effect in the instrumental variables literature:

$$\mathbb{E}[Y_i | Z_i = 1] - \mathbb{E}[Y_i | Z_i = 0] = \mathbb{E}[Y_i(1, D_i(1)) - Y_i(0, D_i(0))].$$

$Z_i$  affects outcome  $Y_i$  directly, and indirectly via the  $D_i(Z_i)$  channel, with no reverse causality. Figure 3 visualises the design, where the direction arrows denote the causal direction. CM aims to decompose the ATE of  $Z_i$  on  $Y_i$  into these two separate pathways:

$$\text{Average Direct Effect (ADE): } \mathbb{E}[Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i))],$$

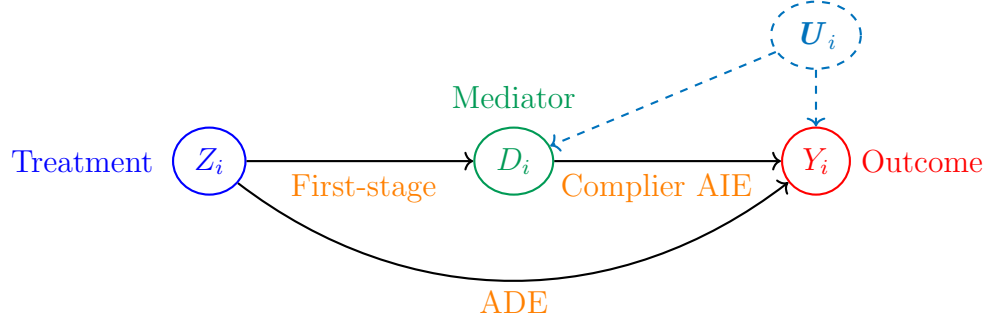
$$\text{Average Indirect Effect (AIE): } \mathbb{E}[Y_i(Z_i, D_i(1)) - Y_i(Z_i, D_i(0))].$$

Estimating the AIE answers the following question: how much of the causal effect  $Z_i$  on  $Y_i$  goes through the  $D_i$  channel? When studying the health gains of health insurance (Finkelstein et al. 2012), the AIE represents how much of the effect comes from using the hospital more often. Estimating the ADE answers the following equation: how much is left over after accounting for the  $D_i$  channel?<sup>7</sup> For the health insurance example, how much of the

<sup>6</sup>This assumption can hold conditional on covariates. To simplify notation in this section, leave the conditional part unsaid, as it changes no part of the identification framework.

<sup>7</sup>In a non-parametric setting it is not necessary that  $\text{ADE} + \text{AIE} = \text{ATE}$ . See Imai et al. (2010) for this point in full.



**Figure 3:** Structural Causal Model for CM.

**Note:** This figure shows the structural causal model behind CM. The Complier AIE refers to the AIE local to  $D_i(Z_i)$  compliers, so that  $AIE = \text{average first-stage} \times \text{Complier AIE}$ .  $U_i$  represents this paper's focus on the case that  $D_i$  is not ignorable by showing an unobserved confounder. [Subsection 3.1](#) defines  $U_i$  in an applied setting.

health insurance effect is a direct effect, other than increased healthcare usage — e.g., income effects of lower medical debt, or less worry over health shocks. The Instrumental Variables (IV) approach assumes this direct effect is zero for everyone (the exclusion restriction). CM is a similar, yet distinct, framework attempting to explicitly model the direct effect, and not assuming it is zero.

The ADE and AIE are not separately identified without further assumptions.

## 2.1 Identification of CM Effects

The conventional approach to estimating direct and indirect effects assumes both  $Z_i$  and  $D_i$  are ignorable, conditional on a vector of control variables  $\mathbf{X}_i$ .

**Definition 1.** *Sequential Ignorability* ([Imai et al. 2010](#))

$$Z_i \perp\!\!\!\perp D_i(z'), Y_i(z, d') \mid \mathbf{X}_i, \quad \text{for } z', z, d' = 0, 1 \quad (1)$$

$$D_i \perp\!\!\!\perp Y_i(z', d') \mid \mathbf{X}_i, Z_i = z', \quad \text{for } z', d' = 0, 1. \quad (2)$$

Sequential ignorability assumes that the initial treatment  $Z_i$  is ignorable conditional on  $\mathbf{X}_i$  (as has already been assumed above). It then also assumes that, after  $Z_i$  is assigned, that  $D_i$  is ignorable conditional on  $\mathbf{X}, Z_i$  (hereafter, mediator ignorability). If [1\(1\)](#) and [1\(2\)](#) hold, then the ADE and AIE are identified by two-stage mean differences conditioning on  $\mathbf{X}_i$ .<sup>8</sup>

<sup>8</sup>In addition, a common support condition for both  $Z_i, D_i$  (across  $\mathbf{X}_i$ ) is necessary. [Imai et al. \(2010\)](#) show a general identification statement; I show identification in terms of two-stage regression, notation for which is more familiar in economics. [Appendix A.1](#) states the [Imai et al. \(2010\)](#) identification result, and then develops the two-stage regression notation which holds as a consequence of sequential ignorability.

$$\begin{aligned}
& \mathbb{E}_{D_i, \mathbf{X}_i} \left[ \underbrace{\mathbb{E}[Y_i | Z_i = 1, D_i, \mathbf{X}_i] - \mathbb{E}[Y_i | Z_i = 0, D_i, \mathbf{X}_i]}_{\text{Second-stage regression, } Y_i \text{ on } Z_i \text{ holding } D_i, \mathbf{X}_i \text{ constant}} \right] = \underbrace{\mathbb{E}[Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i))]}_{\text{Average Direct Effect (ADE)}} \\
& \mathbb{E}_{Z_i, \mathbf{X}_i} \left[ \underbrace{\left( \mathbb{E}[D_i | Z_i = 1, \mathbf{X}_i] - \mathbb{E}[D_i | Z_i = 0, \mathbf{X}_i] \right)}_{\text{First-stage regression, } D_i \text{ on } Z_i} \times \underbrace{\left( \mathbb{E}[Y_i | Z_i, D_i = 1, \mathbf{X}_i] - \mathbb{E}[Y_i | Z_i, D_i = 0, \mathbf{X}_i] \right)}_{\text{Second-stage regression, } Y_i \text{ on } D_i \text{ holding } Z_i, \mathbf{X}_i \text{ constant}} \right] \\
& \quad = \underbrace{\mathbb{E}[Y_i(Z_i, D_i(1)) - Y_i(Z_i, D_i(0))]}_{\text{Average Indirect Effect (AIE)}}
\end{aligned}$$

I refer to the estimands on the left-hand side as CM estimands, which are typically estimated by a composition of two-stage Ordinary Least Squares (OLS) estimates (Imai et al. 2010). While this is the most common approach in the applied literature, I do not assume the linear model for my identification analysis. Linearity assumptions are not necessary for identification, and it suffices to note that heterogeneous treatment effects and non-linear confounding can bias OLS estimates of CM estimands in the same manner that is well documented elsewhere (see e.g., Angrist 1998, Słoczyński 2022). This section focuses on problems that plague CM by selection-on-observables, regardless of estimation method.

## 2.2 Non-identification of CM Effects

Applied research often uses a natural experiment to study settings where treatment  $Z_i$  is ignorable, justifying assumption 1(1). Rarely do they also have access to an additional, overlapping natural experiment to isolate random variation in  $D_i$  — to justify mediator ignorability 1(2). One might consider conventional CM methods in such a setting to learn about the mechanisms behind the causal effect  $Z_i$  on  $Y_i$ , without the problems associated with suggestive evidence of mechanisms. This approach leads to biased estimates, and further contaminates inference regarding direct and indirect effects.

**Theorem 1.** *Absent an identification strategy for the mediator, CM estimates are at risk of selection bias. If 1(1) holds, and 1(2) does not, then CM estimands are contaminated by selection bias and group differences. Proof: see Appendix A.2.*

Below I present the relevant selection bias and group difference terms, omitting the conditional on  $\mathbf{X}_i$  notation for brevity.

For the direct effect: CM estimand = ADE + selection bias + group differences.<sup>9</sup>

$$\begin{aligned} & \mathbb{E}_{D_i} \left[ \mathbb{E} [Y_i | Z_i = 1, D_i] - \mathbb{E} [Y_i | Z_i = 0, D_i] \right] \\ &= \mathbb{E} [Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i))] \\ &+ \mathbb{E}_{D_i=d'} \left[ \mathbb{E} [Y_i(0, D_i(Z_i)) | D_i(1) = d'] - \mathbb{E} [Y_i(0, D_i(Z_i)) | D_i(0) = d'] \right] \\ &+ \mathbb{E}_{D_i=d'} \left[ \left( 1 - \Pr(D_i(1) = d') \right) \left( \mathbb{E} [Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i)) | D_i(1) = 1 - d'] \right) \right. \\ &\quad \left. - \mathbb{E} [Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i)) | D_i(1) = d'] \right] \end{aligned}$$

For the indirect effect: CM estimand = AIE + selection bias + group differences.

$$\begin{aligned} & \mathbb{E}_{Z_i} \left[ \left( \mathbb{E} [D_i | Z_i = 1] - \mathbb{E} [D_i | Z_i = 0] \right) \times \left( \mathbb{E} [Y_i | Z_i, D_i = 1] - \mathbb{E} [Y_i | Z_i, D_i = 0] \right) \right] \\ &= \mathbb{E} [Y_i(Z_i, D_i(1)) - Y_i(Z_i, D_i(0))] \\ &+ \Pr(D_i(1) = 1, D_i(0) = 0) \left( \mathbb{E} [Y_i(Z_i, 0) | D_i = 1] - \mathbb{E} [Y_i(Z_i, 0) | D_i = 0] \right) \\ &+ \Pr(D_i(1) = 1, D_i(0) = 0) \times \\ &\quad \left[ \left( 1 - \Pr(D_i = 1) \right) \left( \mathbb{E} [Y_i(Z_i, 1) - Y_i(Z_i, 0) | D_i = 1] \right. \right. \\ &\quad \left. \left. - \mathbb{E} [Y_i(Z_i, 1) - Y_i(Z_i, 0) | D_i = 0] \right) \right. \\ &\quad \left. - \left( \frac{1 - \Pr(D_i(1) = 1, D_i(0) = 0)}{\Pr(D_i(1) = 1, D_i(0) = 0)} \right) \left( \mathbb{E} [Y_i(Z_i, 1) - Y_i(Z_i, 0) | D_i(1) = 0 \text{ or } D_i(0) = 1] \right) \right. \\ &\quad \left. - \mathbb{E} [Y_i(Z_i, 1) - Y_i(Z_i, 0)] \right] \end{aligned}$$

The selection bias terms come from systematic differences between the groups taking or refusing the mediator ( $D_i = 1$  versus  $D_i = 0$ ), differences not fully unexplained by  $\mathbf{X}_i$ . These selection bias terms would equal zero if the mediator had been ignorable 1(2), but do not necessarily average to zero if not. In the Oregon Health Insurance Experiment, the wait-list gave random variation in the treatment (health insurance) but there was not a similar natural experiment for healthcare usage; correspondingly, the selection-on-observables approach to CM has selection bias.

The group differences represent the fact that a matching approach gives an average effect on the treated group, which is systematically different from the average effect if selection-on-observables does not hold. These terms are a non-parametric framing of the bias from controlling for intermediate outcomes, previously studied only in a linear setting (i.e., bad

<sup>9</sup>The bias terms here mirror those in Heckman, Ichimura, Smith & Todd (1998), Angrist & Pischke (2009) for a single  $D_i$  on  $Y_i$  treatment effect, when  $D_i$  is not ignorable:

$$\mathbb{E} [Y_i | D_i = 1] - \mathbb{E} [Y_i | D_i = 0] = \text{ATE} + \underbrace{\left( \mathbb{E} [Y_i(., 0) | D_i = 1] - \mathbb{E} [Y_i(., 0) | D_i = 0] \right)}_{\text{Selection Bias}} + \underbrace{\Pr(D_i = 0) (\text{ATT} - \text{ATU})}_{\text{Group-differences Bias}}.$$

controls in [Cinelli, Forney & Pearl 2024](#), or M-bias in [Ding & Miratrix 2015](#)).

The AIE group differences term is longer, because the indirect effect is comprised of the effect of  $D_i$  local to  $D_i(Z_i)$  compliers.

$$\begin{aligned} \text{AIE} &= \mathbb{E} [Y_i(Z_i, D_i(1)) - Y_i(Z_i, D_i(0))] \\ &= \mathbb{E} [D_i(1) - D_i(0)] \underbrace{\mathbb{E} [Y_i(Z_i, 1) - Y_i(Z_i, 0) \mid D_i(0) = 0, D_i(1) = 1]}_{\text{Average } D_i \text{ on } Y_i \text{ effect among } D_i(Z_i) \text{ compliers}} \end{aligned}$$

It is important to acknowledge the mediator compliers here, because the AIE is the treatment effect going through the  $D_i(Z_i)$  channel, thus only refers to individuals pushed into mediator  $D_i$  by initial treatment  $Z_i$ . If we had been using a population average effect for  $D_i$  on  $Y_i$ , then this is losing focus on the definition of the AIE; it is not about the causal effect  $D_i$  on  $Y_i$ , it is about the causal effect  $D_i(Z_i)$  on  $Y_i$ .

The group difference bias term arises because the selection-on-observables approach assumes that this complier average effect is equal to the population average effect, which does not hold true if the mediator is not ignorable. This distinction between average effects and complier average effects in the AIE is skipped over by the “controlled indirect effect” definition of [Pearl \(2013\)](#).

### 3 CM in Applied Settings

Unobserved confounding is particularly problematic when studying the mechanisms behind treatment effects. For example, in studying health gains from the Oregon wait-list lottery, we might expect that health gains came about because those who won access to Medicaid started visiting their healthcare provider more often, when in past they avoided it over financial concerns. Applying conventional CM methods to investigate this expectation would be dismissing unobserved confounders for how often individuals visit healthcare providers, leading to biased results.

The wider population does not have one uniform bill of health; many people are born predisposed to ailments, due to genetic variation or other unrelated factors. These conditions can exist for years before being diagnosed. People with severe underlying conditions may visit healthcare providers more often than the rest of the population, to investigate or begin treating the ill-effects. It stands to reason that people with more severe underlying conditions may gain more from more often attending healthcare providers once given health insurance. These underlying causes cannot be controlled for by researchers, as we cannot hope to observe and control for health conditions that are yet to even be diagnosed. This means underlying health conditions are an unobserved confounder, and will bias estimates of the ADE and AIE

in this setting.

In this section, I further develop the issue of selection on unobserved factors in a general CM setting. First, I show the non-parametric bias terms from [Section 2](#) can be written as omitted variables bias in a random coefficients regression framework. Second, I show how selection bias operates in a basic model for selection-into-mediator based on costs and benefits.

### 3.1 Regression Framework

Inference for CM effects can be written in a regression framework with random coefficients, showing how correlation between unobserved error terms and the mediator disrupts identification.

Start by writing potential outcomes  $Y_i(\cdot, \cdot)$  as a sum of observed and unobserved factors, following the notation of [Heckman & Vytlacil \(2005\)](#). For each  $z', d' = 0, 1$ , put  $\mu_{d'}(z'; \mathbf{X}_i) = \mathbb{E}[Y_i(z', d') | \mathbf{X}_i]$  and the corresponding error terms,  $U_{d',i} = Y_i(z', d') - \mu_{d'}(z'; \mathbf{X}_i)$ , so we have the following expressions:

$$Y_i(Z_i, 0) = \mu_0(Z_i; \mathbf{X}_i) + U_{0,i}, \quad Y_i(Z_i, 1) = \mu_1(Z_i; \mathbf{X}_i) + U_{1,i}.$$

With this notation, observed data  $Z_i, D_i, Y_i, \mathbf{X}_i$  have the following random coefficient outcome formulae — which characterise direct effects, indirect effects, and selection bias.

$$D_i = \theta + \bar{\pi}Z_i + \zeta(\mathbf{X}_i) + \eta_i, \tag{3}$$

$$Y_i = \alpha + \beta D_i + \gamma Z_i + \delta Z_i D_i + \varphi(\mathbf{X}_i) + \underbrace{(1 - D_i) U_{0,i} + D_i U_{1,i}}_{\text{Correlated error term.}} \tag{4}$$

This is not consequence of linearity assumptions; the outcome formulae allow for unconstrained heterogeneous treatment effects, because the coefficients are random. If either  $Z_i, D_i$  were continuously distributed, then this representative would not necessarily hold true. First-stage (3) is identified, with  $\theta + \zeta(\mathbf{X}_i)$  the intercept, and  $\bar{\pi}$  the first-stage average compliance rate (conditional on  $\mathbf{X}_i$ ). Second-stage (4) has the following definitions, and is not identified thanks to omitted variables bias. See [Appendix A.3](#) for the derivation.

- (a)  $\alpha = \mathbb{E}[\mu_0(0; \mathbf{X}_i)]$  and  $\varphi(\mathbf{X}_i) = \mu_0(0; \mathbf{X}_i) - \alpha$  are the intercept terms.
- (b)  $\beta = \mu_1(0; \mathbf{X}_i) - \mu_0(0; \mathbf{X}_i)$  is the AIE conditional on  $Z_i = 0, \mathbf{X}_i$ .
- (c)  $\gamma = \mu_0(1; \mathbf{X}_i) - \mu_0(0; \mathbf{X}_i)$  is the ADE conditional on  $D_i = 0, \mathbf{X}_i$ .

- (d)  $\delta = \mu_1(1; \mathbf{X}_i) - \mu_0(1; \mathbf{X}_i) - (\mu_1(0; \mathbf{X}_i) - \mu_0(0; \mathbf{X}_i))$  is the average interaction effect conditional on  $\mathbf{X}_i$ .
- (e)  $(1 - D_i)U_{0,i} + D_iU_{1,i}$  is the disruptive error term.

The ADE and AIE are averages of the random coefficients:

$$\begin{aligned} \text{ADE} &= \mathbb{E}[\gamma + \delta D_i], \\ \text{AIE} &= \mathbb{E}\left[\pi(\beta + \delta Z_i + \tilde{U}_i)\right], \quad \text{with } \tilde{U}_i = \underbrace{\mathbb{E}[U_{1,i} - U_{0,i} \mid \mathbf{X}_i, D_i(0) = 0, D_i(1) = 1]}_{\text{Unobserved complier gains}}. \end{aligned}$$

The ADE is a simple sum of the coefficients, while the AIE includes a group differences term because it only refers to  $D_i(Z_i)$  compliers.

By construction,  $\mathbf{U}_i := (U_{0,i}, U_{1,i})$  is an unobserved confounder. The regression estimates of  $\beta, \gamma, \delta$  in second-stage (4) give unbiased estimates only if  $D_i$  is also conditionally ignorable:  $D_i \perp\!\!\!\perp \mathbf{U}_i$ . If not, then estimates of CM effects suffer from omitted variables bias from failing to adjust for the unobserved confounder,  $\mathbf{U}_i$ .

### 3.2 Selection on Costs and Benefits

CM is at risk of bias because  $D_i \perp\!\!\!\perp \mathbf{U}_i$  is unlikely to hold in applied settings. A separate identification strategy could disrupt the selection-into- $D_i$  based on unobserved factors, and lend credibility to the mediator ignorability assumption. Without it, bias will persist, given how we conventionally think of selection-into-treatment.

Consider a model where individual  $i$  selects into a mediator based on costs and benefits (in terms of outcome  $Y_i$ ), after  $Z_i, \mathbf{X}_i$  have been assigned. In a natural experiment setting, an external factor has disrupted individuals selecting  $Z_i$  by choice (thus  $Z_i$  is ignorable), but it has not disrupted the choice to take mediator (thus  $D_i$  is not ignorable). In the Oregon Health Insurance Experiment, the treatment variation comes from the wait-list lottery, while healthcare usage was not subject to a similar lottery. Write  $C_i$  for individual  $i$ 's costs of taking mediator  $D_i$ , and  $\mathbb{1}\{\cdot\}$  for the indicator function. The Roy model has  $i$  taking the mediator if the benefits exceed the costs,

$$D_i(z') = \mathbb{1}\left\{ \underbrace{C_i}_{\text{Costs}} \leq \underbrace{Y_i(z', 1) - Y_i(z', 0)}_{\text{Benefits}} \right\}, \quad \text{for } z' = 0, 1. \quad (5)$$

The Roy model provides an intuitive framework for analysing selection mechanisms because it captures the fundamental economic principle of decision-making based on costs and benefits in terms of the outcome under study (Roy 1951, Heckman & Honore 1990). In the

Oregon Health Insurance Experiment, this models choice to visit the doctor in terms of health and well-being benefits relative to costs.<sup>10</sup> This makes the Roy model useful as a base case for CM, where selection-into-mediator may be driven by private information (unobserved by the researcher).

By using the Roy model as a benchmark, I explore the practical limits of the mediator ignorability assumption. Decompose the costs into its mean and an error term,  $C_i(Z_i) = \mu_C(Z_i; \mathbf{X}_i) + U_{C,i}$ , to show Roy-selection in terms of unobserved and observed factors,

$$D_i(z') = \mathbb{1} \{U_{C,i} - (U_{1,i} - U_{0,i}) \leq \mu_1(z'; \mathbf{X}_i) - \mu_0(z'; \mathbf{X}_i) - \mu_C(z'; \mathbf{X}_i)\}, \quad \text{for } z' = 0, 1.$$

If selection follows a Roy model, and the mediator is ignorable, then unobserved benefits can play no part in selection. The only driver of selection are individuals' differences in costs (and not benefits). If there are any selection-into- $D_i$  benefits unobserved to the researcher, then mediator ignorability cannot hold.

**Proposition 1.** *Suppose mediator selection follows a Roy model (5), and selection is not fully explained by costs and observed gains. Then mediator ignorability does not hold.*

This is an equivalence statement: selection based on costs and benefits is only consistent with mediator ignorability if the researcher observed every single source of mediator benefits. See [Appendix A.4](#) for the proof. This means that the vector of control variables  $\mathbf{X}_i$  must be incredibly rich. Together,  $\mathbf{X}_i$  and unobserved cost differences  $U_{C,i}$  must explain selection-into- $D_i$  one hundred percent. In the Roy model framework, however, individuals make decisions about mediator take-up based on gains — whether the researcher observes them or not. The unobserved gains are unlikely to be fully captured by an observed control set  $\mathbf{X}_i$ , except in very special cases.

In practice, the only way to believe in the mediator ignorability assumption is to study a setting where the researcher has two causal research designs, one for treatment  $Z_i$  and another for mediator  $D_i$ , at the same time. An unmotivated note saying “we conduct an informal mechanism analysis by controlling for this variable” or “we assume the mediator satisfies selection-on-observables” does not cut it here, and will lead to biased inference in applied settings.

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<sup>10</sup>If the choice is considers over a sum of outcomes, then a simple extension to a utility maximisation model maintains this same framework with expected costs and benefits. See [Heckman & Honore \(1990\)](#), [Eisenhauer, Heckman & Vytlačil \(2015\)](#).



## 4 Solving Identification with a Control Function (CF)

If your goal is to estimate CM effects, and you could control for unobserved selection terms  $U_{0,i}, U_{1,i}$ , then you would. This ideal (but infeasible) scenario would yield unbiased estimates for the ADE and AIE. A Control Function (CF) approach takes this insight seriously, providing conditions to model the implied confounding by  $U_{0,i}, U_{1,i}$ , and then controlling for it.

The main problem is that second-stage regression equation (4) is not identified, because  $U_{0,i}, U_{1,i}$  are unobserved, and lead to omitted variables bias.

$$\begin{aligned} \mathbb{E}[Y_i | Z_i, D_i, \mathbf{X}_i] &= \alpha + \beta D_i + \gamma Z_i + \delta Z_i D_i + \varphi(\mathbf{X}_i) \\ &\quad + \underbrace{(1 - D_i) \mathbb{E}[U_{0,i} | D_i = 0, \mathbf{X}_i] + D_i \mathbb{E}[U_{1,i} | D_i = 1, \mathbf{X}_i]}_{\text{Unobserved confounding.}} \end{aligned} \quad (6)$$

The CF approach models the contaminating terms in (6), avoiding the bias from omitting them in regression estimates. CF methods were first devised to correct for sample selection problems (Heckman 1974), and were extended to a general selection problem of the same form as Equation (6) (Heckman 1979). The approach works in the following manner: (1) assume that the variable of interest follows a selection model, where unexplained first-stage selection informs unobserved second-stage confounding; (2) extract information about unobserved confounding from the first-stage; and (3) incorporate this information as control terms in the second-stage equation to adjust for selection-into-mediator. Identification in CF methods typically relies on an external instrument or distributional assumptions; the identification strategy here focuses exclusively on the case that an instrument is available. By explicitly accounting for the information contained in the first-stage selection model, CF methods enable consistent estimation of causal effects in the second-stage even when selection is driven by unobserved factors (Florens et al. 2008).

In the example of analysing health gains from the Oregon Health Insurance Experiment, a CF approach addresses the unobserved confounding by modelling unobserved effects of underlying health conditions. It does so by assuming that unobserved selection-into-healthcare use is informative for underlying health conditions, assuming people with more severe underlying conditions visit the doctor more often than those without. Then it uses this information in the second-stage estimation of how much the effect goes through increased healthcare usage, estimating the ADE and AIE after controlling for this confounding.

## 4.1 Re-identification of CM Effects

The following assumptions are sufficient to model the correlated error terms, identifying  $\beta, \gamma, \delta$  in the second-stage regression (4), and thus both the ADE and AIE.

**Assumption CF–1.** Mediator monotonicity, conditional on  $\mathbf{X}_i$ .

$$\Pr(D_i(0) \leq D_i(1) \mid \mathbf{X}_i) = 1.$$

Assumption CF–1 is the monotonicity condition first used in an IV context (Imbens & Angrist 1994). Here, it is assuming that people respond to treatment,  $Z_i$ , by consistently taking or refusing the mediator  $D_i$  (always or never-mediators), or taking the mediator  $D_i$  if and only if assigned to the treatment  $Z_i = 1$  (mediator compliers). There are no mediator defiers.

The main implication of Assumption CF–1 is that selection-into-mediator can be written as a selection model with ordered threshold crossing values that describe selection-into- $D_i$  (Vytlacil 2002).

$$D_i(z') = \mathbb{1} \{V_i \leq \psi(z'; \mathbf{X}_i)\}, \quad \text{for } z' = 0, 1$$

where  $V_i$  is a latent variable with continuous distribution and conditional cumulative density function  $F_V(\cdot \mid \mathbf{X}_i)$ , and  $\psi(\cdot; \mathbf{X}_i)$  collects observed sources of mediator selection.  $V_i$  could be assumed to follow a known distribution; the canonical Heckman selection model assumes  $V_i$  is normally distributed (a “Heckit” model). The identification strategy here applies to the general case that the distribution of  $V_i$  is unknown, without parametric restrictions.

I focus on the equivalent transformed model of Heckman & Vytlacil (2005),

$$D_i(z') = \mathbb{1} \{U_i \leq \pi(z'; \mathbf{X}_i)\}, \quad \text{for } z' = 0, 1$$

where  $U_i := F_V(V_i \mid \mathbf{X}_i)$  follows a uniform distribution, and  $\pi(z'; \mathbf{X}_i) = F_V(\psi(z'; \mathbf{X}_i)) = \Pr(D_i = 1 \mid Z_i = z', \mathbf{X}_i)$  is the mediator propensity score.  $U_i$  are the unobserved mediator take-up costs. Note the maintained assumption that treatment  $Z_i$  is ignorable conditional on  $\mathbf{X}_i$  implies  $Z_i \perp\!\!\!\perp U_i$  conditional on  $\mathbf{X}_i$ .

This selection model setup is equivalent to the monotonicity condition, and is importing a well-known equivalence result from the IV literature to the CM setting. The main conceptual difference is not assuming  $Z_i$  is a valid instrument for identifying the  $D_i$  on  $Y_i$  effect among compliers; it is using the selection model representation to correct for selection bias. See Appendix A.5 for a validation of the general Vytlacil (2002) equivalence result in a CM setting, with conditioning covariates  $\mathbf{X}_i$ .

**Assumption CF-2.** Selection on mediator benefits.

$$\text{Cov}(U_i, U_{0,i}), \text{Cov}(U_i, U_{1,i}) \neq 0.$$

Assumption CF-2 is stating that unobserved selection in mediator take-up ( $U_i$ ) informs second-stage confounding, when refusing or taking the mediator ( $U_{0,i}$  and  $U_{1,i}$ ). If there is unobserved confounding in  $Y_i$ , then it can be measured in  $D_i$ .

This is a strong assumption, and will not hold in all examples. If people had been deciding to take  $D_i$  by a Roy model, then this assumption holds because  $V_i = U_{C,i} - (U_{1,i} - U_{0,i})$ . Individuals could be making decisions based on other outcomes, but as long as mediator benefits guide at least part of this decision (i.e., bounded away from zero), then this assumption will hold.

For notation purposes, suppose the vector of control variables  $\mathbf{X}_i$  has at least two entries; denote  $\mathbf{X}_i^{\text{IV}}$  as one entry in the vector, and  $\mathbf{X}_i^-$  as the remaining.

**Assumption CF-3.** Mediator take-up cost instrument.

$$\mathbf{X}_i^{\text{IV}} \text{ satisfies } \frac{\partial}{\partial \mathbf{X}_i^{\text{IV}}} \left\{ \mu_1(z', \mathbf{X}_i) - \mu_0(z', \mathbf{X}_i) \right\} = 0 < \frac{\partial}{\partial \mathbf{X}_i^{\text{IV}}} \left\{ \mathbb{E}[D_i(z') | \mathbf{X}_i] \right\}, \text{ for } z' = 0, 1.$$

Assumption CF-3 is requiring at least one control variable guides selection-into- $D_i$  — an IV. It assumes an instrument exists, which satisfies an exclusion restriction (i.e., not impacting mediator gains  $\mu_1 - \mu_0$ ), and has a non-zero influence on the mediator (i.e., strong IV first-stage). The exclusion restriction is untestable, and must be guided by domain-specific knowledge; IV first-stage strength is testable, and must be justified with data by methods common in the IV literature.

This assumption identifies the mediator propensity score separately from the direct and indirect effects, avoiding indeterminacy in the second-stage outcome equation. While not technically required for identification, it avoids relying entirely on an assumed distribution for unobserved error terms (and bias from inevitably breaking this assumption). The most compelling example of a mediator IV is using data on the cost of mediator take-up as a first-stage IV, if it varies between individuals for unrelated reasons and is strong in explaining mediator take-up.

**Proposition 2.** *If assumptions CF-1, CF-2, CF-3 hold, then second-stage regression equation (4) is identified with a CF adjustment.*

$$\begin{aligned} \mathbb{E}[Y_i | Z_i, D_i, \mathbf{X}_i] &= \alpha + \beta D_i + \gamma Z_i + \delta Z_i D_i + \varphi(\mathbf{X}_i^-) \\ &\quad + \rho_0 (1 - D_i) \lambda_0(\pi(Z_i; \mathbf{X}_i)) + \rho_1 D_i \lambda_1(\pi(Z_i; \mathbf{X}_i)), \end{aligned}$$

where  $\lambda_0, \lambda_1$  are the Control Functions (CFs),  $\rho_0, \rho_1$  are linear parameters, and mediator propensity score  $\pi(z'; \mathbf{X}_i)$  is separately identified in the first-stage (3). Proof: see [Appendix A.6](#).

Again, this set-up required no linearity assumptions, and treatment effects vary, because  $Z_i, D_i$  are categorical and  $\beta, \gamma, \delta, \varphi(\mathbf{X}_i)$  vary with  $\mathbf{X}_i$ . The CFs are functions which measure unobserved mediator gains, for those with unobserved mediator costs above or below a propensity score value. Following the IV notation of [Kline & Walters \(2019\)](#), put  $\mu_V = \mathbb{E} [F_V^{-1}(U_i | \mathbf{X}_i)]$ , to give the following representation for the CFs:

$$\begin{aligned}\lambda_0(p') &= \mathbb{E} [F_V^{-1}(U_i | \mathbf{X}_i) - \mu_V | p' < U_i], \\ \lambda_1(p') &= \mathbb{E} [F_V^{-1}(U_i | \mathbf{X}_i) - \mu_V | U_i \leq p'] = -\lambda_0(p') \left( \frac{1-p'}{p'} \right), \text{ for } p' \in (0, 1).\end{aligned}$$

If we are using the canonical Heckman selection model, we assume the error term follows a normal distribution, so that  $\lambda_0, \lambda_1$  are the inverse Mills ratio. Alternatively,  $\lambda_0, \lambda_1$  could have other definitions following the assumed distribution of the error terms (see e.g, [Wooldridge 2015](#)). If we do not know what distribution class the errors follow, then  $\lambda_0, \lambda_1$  can be estimated separately with semi-parametric methods to avoid relying on parametric assumptions.

**Theorem CF.** If assumptions [CF-1](#), [CF-2](#), [CF-3](#) hold, the ADE and AIE are identified as a function of the parameters in [Proposition 2](#).

$$\begin{aligned}\text{ADE} &= \mathbb{E} [\gamma + \delta D_i], \\ \text{AIE} &= \mathbb{E} \left[ \bar{\pi} \left( \beta + \delta Z_i + \underbrace{(\rho_1 - \rho_0) \Gamma(\pi(0; \mathbf{X}_i), \pi(1; \mathbf{X}_i))}_{\text{Mediator compliers adjustment}} \right) \right]\end{aligned}$$

where  $\Gamma(p, p') = \mathbb{E} [F_V^{-1}(U_i | \mathbf{X}_i) - \mu_V | p < U_i \leq p'] = \frac{p'\lambda_1(p') - p\lambda_1(p)}{p' - p}$  is the average unobserved net gains for those with unobserved costs between  $p < p'$ ,<sup>11</sup> and  $\bar{\pi} = \pi(1; \mathbf{X}_i) - \pi(0; \mathbf{X}_i)$  is the mediator complier score. Proof: see [Appendix A.7](#).

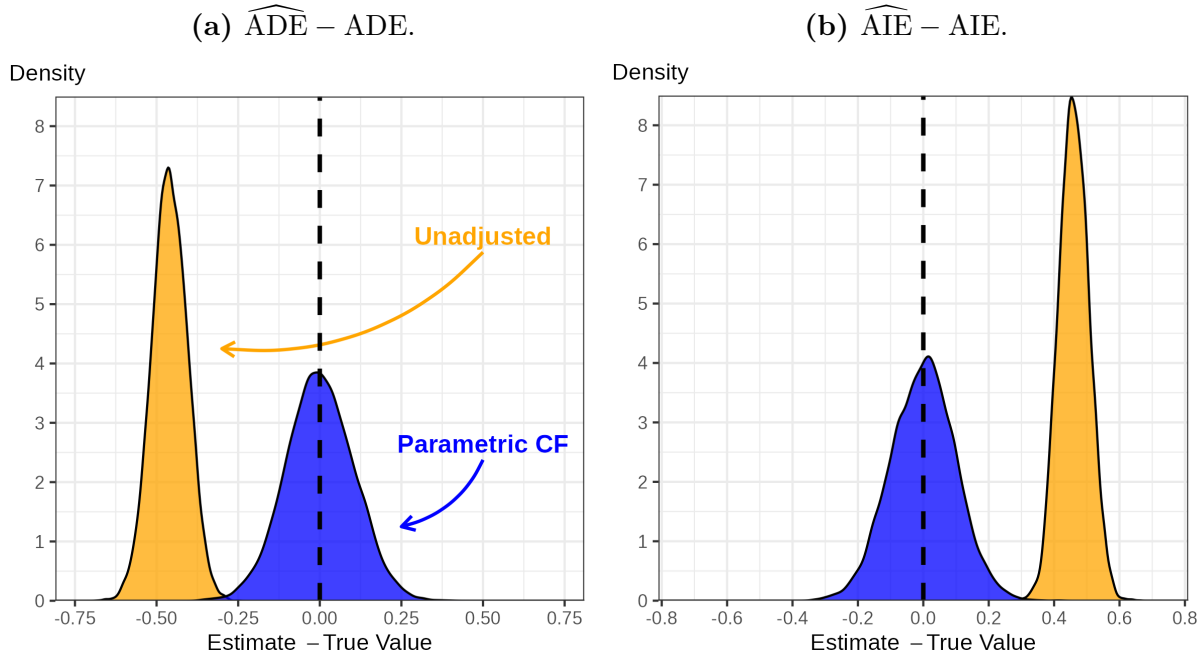
This theorem provides a solution to the identification problem for CM effects when facing selection; rather than assuming away selection problems, it explicitly models them. The ADE is straightforward to calculate as an average of the direct effect parameters, while the AIE also includes an adjustment for unobserved complier gains to the mediator. Again, this is because the AIE only refers to individuals who were induced by treatment  $Z_i$  into taking

<sup>11</sup>The complier adjustment term was first written in this manner by [Kline & Walters \(2019\)](#) for an IV setting.

mediator  $D_i$  (mediator compliers). The CFs allow us to measure both selection bias and complier differences, and thus purge persistent bias in identifying CM effects.

This identification strategy is essentially a Marginal Treatment Effect approach (MTE, Björklund & Moffitt 1987, Heckman & Vytlacil 2005) applied to a CM setting. Just as the local IV approach uses variation in instruments to identify MTEs across the distribution of unobserved treatment take-up costs, this CF approach identifies CM effects across the distribution of unobserved mediator take-up costs. This connection to MTEs provides a conceptual bridge between the literature on structural IV for causal effects and CM.

**Figure 4:** The CF Adjustment Addresses Persistent Bias in Conventional CM Estimates.



**Note:** These figures show the empirical density of point estimates minus the true average effect, for 10,000 different datasets generated from a Roy model with normally distributed error terms (with both correlation and heteroscedasticity, further described in Subsection 5.3). The black dashed line is the true value; orange is the distribution of conventional CM estimates from two-stage OLS (Imai et al. 2010), and blue estimates with a two-stage Heckman selection adjustment.

The ideal instrument  $\mathbf{X}_i^{IV}$  for identification is continuous, and varies  $\pi(z'; \mathbf{X}_i)$  between 0 and 1 for every possible value of  $z'$ ,  $\mathbf{X}_i^-$  (identification at infinity). In practice, it is unlikely to find IV(s) that satisfy this condition. In this case, the Brinch, Mogstad & Wiswall (2017) restricted approach can be used — even with a categorical instrument and no control variables. This amounts to assuming a limited specification for the respective CFs, limiting the number of parameters used to approximate  $\lambda_0, \lambda_1$  to the number of discrete values that  $\pi(z'; \mathbf{X}_i)$  takes minus one. E.g., if there are no control variables and  $\mathbf{X}_i^{IV}$  is binary, then  $\lambda_0, \lambda_1$  can only

be identified up to 3 parameters each.<sup>12</sup> Ultimately, this changes little to the identification strategy, and little to the estimation.

In a simulation with Roy selection-into-mediator based on unobserved error terms, the CF adjustment pushes conventional CM estimates back to the true value. [Figure 4](#) shows how a CF adjustment corrects unadjusted CM effect estimates.

## 5 CF Estimation of CM Effects

A conventional approach to estimating CM effects involves a two-stage approach to estimating the ADE and the AIE: the first-stage ( $Z_i$  on  $D_i$ ), and the second-stage ( $Z_i, D_i$  on  $Y_i$ ). A CF approach is a simple and intuitive addition to this approach: including the CF terms  $\lambda_0, \lambda_1$  in the second-stage regression to address selection-into-mediator.

This section presents two practical estimation strategies. First, I demonstrate how to estimate CM effects with an assumed distribution of error terms, focusing on the Heckman selection model as the leading case. Second, I consider a more flexible semi-parametric approach that avoids distributional assumptions — at the cost of semi-parametrically estimating the corresponding CFs. While both methods effectively address the selection bias issues detailed in previous sections, they differ in their implementation complexity, efficiency, and underlying assumptions.

### 5.1 Parametric CF

A parametric CF solves the identification problem by assuming a distribution for the unobserved error terms in the first-stage selection model, and modelling selection based on this distribution. The Heckman selection model is the most pertinent example, assuming the normal distribution for unobserved errors ([Heckman 1979](#)). A parametric CF using other distributions works in exactly the same manner, replacing the relevant density functions for an alternative distribution as needed. As such, this section focuses exclusively on the Heckman selection model. This estimation approach is the same as in the parametric selection model approach to MTEs, in [Björklund & Moffitt \(1987\)](#).

The Heckman selection model assumes unobserved errors  $V_i$  follow a normal distribution,

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<sup>12</sup>The value of 3 comes from the cases that  $Z_i, \mathbf{X}_i^{\text{IV}}$  each could take 2 values, so  $\pi(z'; \mathbf{X}_i)$  has 4 possible values, giving the semi-parametric identification (and estimation) of each CF only 3 degrees of freedom to work with. If the CFs are instead assumed to have a known distribution (i.e., parametric CF), then those concerns do not matter.

so estimates the first-stage using a probit model.

$$\Pr(D_i = 1 | Z_i, \mathbf{X}_i) = \Phi(\theta + \bar{\pi}Z_i + \boldsymbol{\zeta}'\mathbf{X}_i),$$

where  $\Phi(\cdot)$  is the cumulative density function for the standard normal distribution, and  $\theta, \bar{\pi}, \boldsymbol{\zeta}$  are parameters estimated with maximum likelihood. In the parametric case, an excluded instrument ( $\mathbf{X}_i^{\text{IV}}$ ) is not technically necessary in the first-stage equation — though not including one exposes the method to indeterminacy if the errors are not normally distributed. Thus, it is best practice to use this method with access to an instrument.

From this probit first-stage, construct the inverse Mills ratio terms to serve as the CFs. These terms capture the correlation between unobserved factors influencing both mediator selection and outcomes, when the errors are normally distributed.

$$\lambda_0(p') = \frac{\phi(-\Phi^{-1}(p'))}{\Phi(-\Phi^{-1}(p'))}, \quad \lambda_1(p') = \frac{\phi(\Phi^{-1}(p'))}{\Phi(\Phi^{-1}(p'))}, \quad \text{for } p' \in (0, 1)$$

where  $\phi(\cdot)$  is the probability density function for the standard normal distribution.

Lastly, the second-stage is estimated with OLS, including the CFs with plug in estimates of the mediator propensity score, and  $\boldsymbol{\varphi}'$  a linear approximation of nuisance function  $\varphi(\cdot)$ .

$$\begin{aligned} \mathbb{E}[Y_i | Z_i, D_i, \mathbf{X}_i] &= \alpha + \beta D_i + \gamma Z_i + \delta Z_i D_i + \boldsymbol{\varphi}'\mathbf{X}_i^- \\ &\quad + \rho_0(1 - D_i)\lambda_0(\hat{\pi}(Z_i; \mathbf{X}_i)) + \rho_1 D_i \lambda_1(\hat{\pi}(Z_i; \mathbf{X}_i)) + \varepsilon_i, \end{aligned}$$

where  $\hat{\pi}(z'; \mathbf{X}_i)$  are the predictions from the probit first-stage.

The resulting ADE and AIE estimates are composed from sample estimates of the terms in Theorem [CF](#),

$$\widehat{\text{ADE}} = \hat{\gamma} + \hat{\delta} \bar{D}, \quad \widehat{\text{AIE}} = \hat{\pi} \left( \hat{\beta} + \hat{\delta} \bar{Z} + (\hat{\rho}_1 - \hat{\rho}_0) \frac{1}{N} \sum_{i=1}^N \Gamma(\hat{\pi}(0; \mathbf{X}_i), \hat{\pi}(1; \mathbf{X}_i)) \right)$$

where  $\bar{D} = \frac{1}{N} \sum_{i=1}^N D_i$ ,  $\bar{Z} = \frac{1}{N} \sum_{i=1}^N Z_i$ ,  $\hat{\pi}$  is the estimate of the mean compliance rate, and  $\frac{1}{N} \sum_{i=1}^N \Gamma(\cdot, \cdot)$  is the average of the complier adjustment term as a function of  $\lambda_1$  with  $\hat{\pi}(0; \mathbf{X}_i), \hat{\pi}(1; \mathbf{X}_i)$  values plugged in.

The standard errors for estimates can be computed using the delta method. Specifically, accounting for both the sampling variability in the first-stage estimates of the mediator propensity score as well as the second-stage sampling variability. This approach yields  $\sqrt{N}$ -consistent estimates when the underlying error terms follow a bivariate normal distribution — i.e., when  $\pi(Z_i; \mathbf{X}_i)$  is correctly modelled by the probit first-stage. Errors can also be estimated by the bootstrap, by including estimation of both the first and second-stage within



each bootstrap iteration.

In practice, a parametric CF approach is simple to implement using standard statistical packages. The key advantage is computational simplicity and efficiency, particularly in moderate-sized samples. However, this comes at the cost of strong distributional assumptions. For example, if the error terms deviate substantially from joint normality, the estimates may be biased.<sup>13</sup>

## 5.2 Semi-parametric CF

For settings where researchers are not comfortable specifying a specific distribution for the error terms, a semi-parametric CF will nonetheless consistently estimate CM effects. This method maintains the same identification strategy but avoids assuming a specific error distribution. This estimation approach is used in the modern semi-parametric approach to estimating MTEs, for example in [Brinch et al. \(2017\)](#), [Heckman & Vytlacil \(2007\)](#).

The semi-parametric approach begins with flexible estimation of the first-stage, estimating the mediator propensity score,

$$\Pr(D_i = 1 \mid Z_i, \mathbf{X}_i) = \pi(Z_i; \mathbf{X}_i),$$

where  $\mathbf{X}_i$  must include the instrument(s)  $\mathbf{X}_i^{\text{IV}}$ . This can be estimated using flexible methods, as long as the first-stage is estimated  $\sqrt{N}$ -consistently.<sup>14</sup> An attractive option is the [Klein & Spady \(1993\)](#) semi-parametric binary response model, which avoids relying on an assumed distribution of first-stage errors though requires a linear specification. If it is important to avoid a linear specification, then a probability forest avoids linearity assumptions ([Athey, Tibshirani & Wager 2019](#)) — though is best used for cases with many columns in the  $\mathbf{X}_i$  variables.

The second-stage is estimated with semi-parametric methods. Consider the subsamples of mediator refusers and takers separately,

$$\begin{aligned} \mathbb{E}[Y_i \mid Z_i, D_i = 0, \mathbf{X}_i] &= \alpha + \gamma Z_i + \varphi(\mathbf{X}_i^-) + \rho_0 \lambda_0(\pi(Z_i; \mathbf{X}_i)), \\ \mathbb{E}[Y_i \mid Z_i, D_i = 1, \mathbf{X}_i] &= (\alpha + \beta) + (\gamma + \delta) Z_i + \varphi(\mathbf{X}_i^-) + \rho_1 \lambda_1(\pi(Z_i; \mathbf{X}_i)). \end{aligned}$$

The separated subsamples can be estimated, each individually, with semi-parametric methods. The linear parameters (including a linear approximation  $\varphi'$  of nuisance function  $\varphi(\cdot)$ )<sup>15</sup> can

<sup>13</sup>While this concern is immaterial in an IV setting estimating the LATE ([Kline & Walters 2019](#)), it is pertinent in this setting as the CF extrapolates from IV compliers to mediator compliers.

<sup>14</sup>If an estimate of the first-stage that is not  $\sqrt{N}$ -consistent is used (e.g., a modern machine learning estimator), then the resulting second-stage estimate will not be  $\sqrt{N}$ -consistent.

<sup>15</sup>Appropriate interactions between  $Z_i, D_i$  and  $\mathbf{X}_i$  can also flexibly control for  $\mathbf{X}_i$ , again avoiding linearity

be estimated with OLS, while  $\rho_0\lambda_0$  and  $\rho_1\lambda_1$  take a flexible semi-parametric specification with first-stage estimates  $\hat{\pi}(Z_i; \mathbf{X}_i)$  plugged in. An attractive option is a series estimator, such as a spline specification, as this estimates the function without assuming a functional form but maintains  $\sqrt{n}$ -consistency.

The ADE is estimated by this approach as follows. Take  $\hat{\gamma}$ , the  $D_i = 0$  subsample estimate of  $\gamma$ , and  $(\widehat{\gamma + \delta})$ , the  $D_i = 1$  subsample estimate of  $(\gamma + \delta)$ , to give

$$\widehat{\text{ADE}}^{\text{CF}} = (1 - \overline{D}) \hat{\gamma} + \overline{D} (\widehat{\gamma + \delta}).$$

The AIE is less simple, for two reasons that differ from the parametric CF setting. First, the intercepts for each subsample,  $\alpha$  and  $(\alpha + \beta)$ , are not separately identified from the CFs if the  $\lambda_0, \lambda_1$  functions are flexibly estimated. Second, a semi-parametric specification for the CFs mean  $\rho_0$  and  $\lambda_0$  are no longer separately identified from each other (and same for  $\rho_1, \lambda_1$ ). As such, it is not possible to directly use  $\hat{\lambda}_0, \hat{\lambda}_1$  in estimating the complier adjustment term (as is done in the parametric case).

These problems can be avoided by estimating the AIE using its relation to the ATE. Write  $\widehat{\text{ATE}}$  for the point-estimate of the ATE, and  $\hat{\delta} = (\widehat{\gamma + \delta}) - \hat{\gamma}$  for the point estimate of  $\delta$ , to give the following estimator,

$$\widehat{\text{AIE}}^{\text{CF}} = \widehat{\text{ATE}} - (1 - \overline{Z}) \left( \frac{1}{N} \sum_{i=1}^N \hat{\gamma} + \hat{\delta} \hat{\pi}(1; \mathbf{X}_i) \right) - \overline{Z} \left( \frac{1}{N} \sum_{i=1}^N \hat{\gamma} + \hat{\delta} \hat{\pi}(0; \mathbf{X}_i) \right),$$

where  $\frac{1}{N} \sum_{i=1}^N \hat{\gamma} + \hat{\delta} \hat{\pi}(0; \mathbf{X}_i)$  estimates the ADE conditional on  $Z_i = 0$ ,  $\mathbb{E}[\gamma + \delta D_i(0)]$ , and  $\frac{1}{N} \sum_{i=1}^N \hat{\gamma} + \hat{\delta} \hat{\pi}(1; \mathbf{X}_i)$  estimates the ADE conditional on  $Z_i = 1$ ,  $\mathbb{E}[\gamma + \delta D_i(1)]$ . [Appendix A.8](#) describes the reasoning for this estimator of the AIE, relative to estimates of the ATE and ADE, in further detail.

This semi-parametric approach achieves valid estimation of the CM effects, without specifying the distribution behind unobserved error terms, and achieves desirable properties as long as the first-stage correctly estimates the mediator propensity score, and the structural assumptions hold true. The standard errors for estimates can again be computed using the delta method, or estimated by the bootstrap — again, across both first and second-stages within each bootstrap iteration. Note that relying on propensity score estimation requires assumptions that can be found wanting in real-world settings; a common support condition for the mediator is required, and a semi-/non-parametric first-stage may become cumbersome if there are many control variables or many rows of data.

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assumptions.

### 5.3 Simulation Evidence

The following simulation gives an example to show how these methods work in practice. Suppose data observed to the researcher  $Z_i, D_i, Y_i, \mathbf{X}_i$  are drawn from the following data generating processes, for  $i = 1, \dots, N$ , with  $N = 5,000$  for this simulation.

$$Z_i \sim \text{Binom}(0.5), \quad \mathbf{X}_i^- \sim N(4, 1), \quad \mathbf{X}_i^{\text{IV}} \sim \text{Uniform}(-1, 1), \quad (U_{0,i}, U_{1,i}, U_{C,i}) \sim N(\boldsymbol{\theta}, \boldsymbol{\Sigma})$$

$\boldsymbol{\Sigma}$  is the matrix of parameters which controls the level of confounding from unobserved costs and benefits.<sup>16</sup>

Each  $i$  chooses to take mediator  $D_i$  by a Roy model, with following mean definitions for each  $z', d' = 0, 1$

$$D_i(z') = \mathbb{1}\{C_i \leq Y_i(z', 1) - Y_i(z', 0)\},$$

$$\mu_{d'}(z'; \mathbf{X}_i) = (z' + d' + z'd') + \mathbf{X}_i^-, \quad \mu_C(z'; \mathbf{X}_i) = 3z' + \mathbf{X}_i^- - \mathbf{X}_i^{\text{IV}}.$$

Following [Subsection 3.1](#), these data have the following first and second-stage equations:

$$D_i = \mathbb{1}\{U_{C,i} - (U_{1,i} - U_{0,i}) \leq -3Z_i + \mathbf{X}_i^- - \mathbf{X}_i^{\text{IV}}\},$$

$$Y_i = Z_i + D_i + Z_i D_i + \mathbf{X}_i^- + (1 - D_i)U_{0,i} + D_i U_{1,i}.$$

Treatment  $Z_i$  has a causal effect on outcome  $Y_i$ , and it operates partially through mediator  $D_i$ . Outcome mean  $\mu_{D_i}(Z_i; \mathbf{X}_i)$  contains an interaction term,  $Z_i D_i$ , so while  $Z_i, D_i$  have constant partial effects, the ATE depends on how many  $i$  choose to take the mediator and there is treatment effect heterogeneity.

After  $Z_i$  is assigned,  $i$  chooses to take mediator  $D_i$  by considering the costs and benefits — which vary based on  $Z_i$ , demographic controls  $\mathbf{X}_i$ , and the (non-degenerate) unobserved error terms  $U_{i,0}, U_{1,i}$ . As a result, sequential ignorability does not hold; the mediator is not conditionally ignorable. Thus, a conventional approach to CM does not give an estimate for how much of the ATE goes through mediator  $D$ , but is contaminated by selection bias thanks to the unobserved error terms.

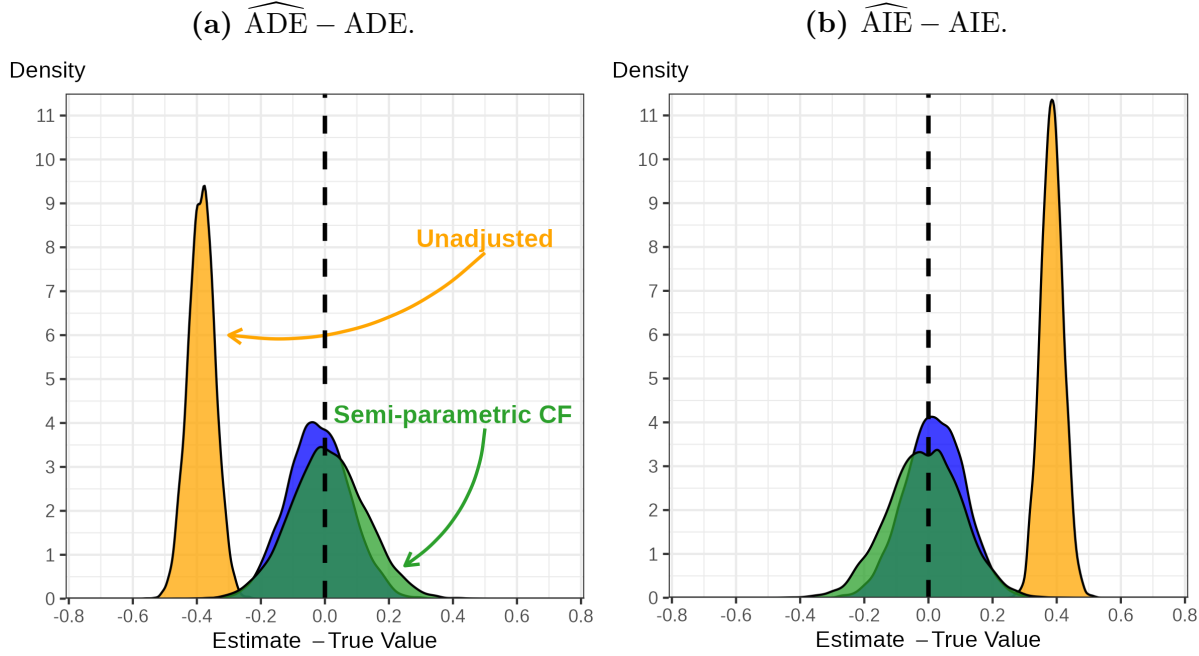
I simulate this data generating process 10,000 times, using  $\boldsymbol{\Sigma} = \begin{pmatrix} 1 & 0.75 & 0 \\ 0.75 & 2.25 & 0 \\ 0 & 0 & 0.25 \end{pmatrix}$ ,<sup>17</sup> and estimate CM effects with conventional CM methods (two-stage OLS) and the introduced CF

<sup>16</sup>The correlation and relative standard deviations for  $U_{0,i}, U_{1,i}$  affect how large selection bias in conventional CM estimates; correlation for these with unobserved costs  $U_{C,i}$  does not particularly matter, though increased variance in unobserved costs makes estimates less precise for both OLS and CF methods.

<sup>17</sup>This choice of parameters has  $\text{Var}(U_{0,i}) = 1$ ,  $\text{Var}(U_{1,i}) = 2.25$ ,  $\text{Corr}(U_{0,i}, U_{1,i}) = 0.5$  so that unobserved errors meaningfully confound conventional CM methods, with notable heteroscedasticity. Unobserved costs are uncorrelated with  $U_{0,i}, U_{1,i}$  (although non-zero correlation would not meaningfully change the results), and  $\text{Var}(U_{C,i}) = 0.25$  maintains uncertainty in unobserved costs.

methods. In this simulation  $\Pr(D_i = 1) = 0.379$ , and 65.77% of the sample are mediator compliers (for whom  $D_i(0) = 0$  and  $D_i(1) = 1$ ). This gives an ATE value of 2.60, ADE 1.38, and AIE 1.22, respectively.<sup>18</sup>

**Figure 5:** Simulated Distribution of CM Effect Estimates, Semi-parametric versus OLS, Relative to True Value.



**Note:** These figures show the empirical density of point estimates minus the true average effect, for 10,000 different datasets generated from a Roy model with correlated uniformly distributed error terms. The black dashed line is the true value; orange is the distribution of conventional CM estimates from two-stage OLS (Imai et al. 2010), and green estimates with a two-stage semi-parametric CF.

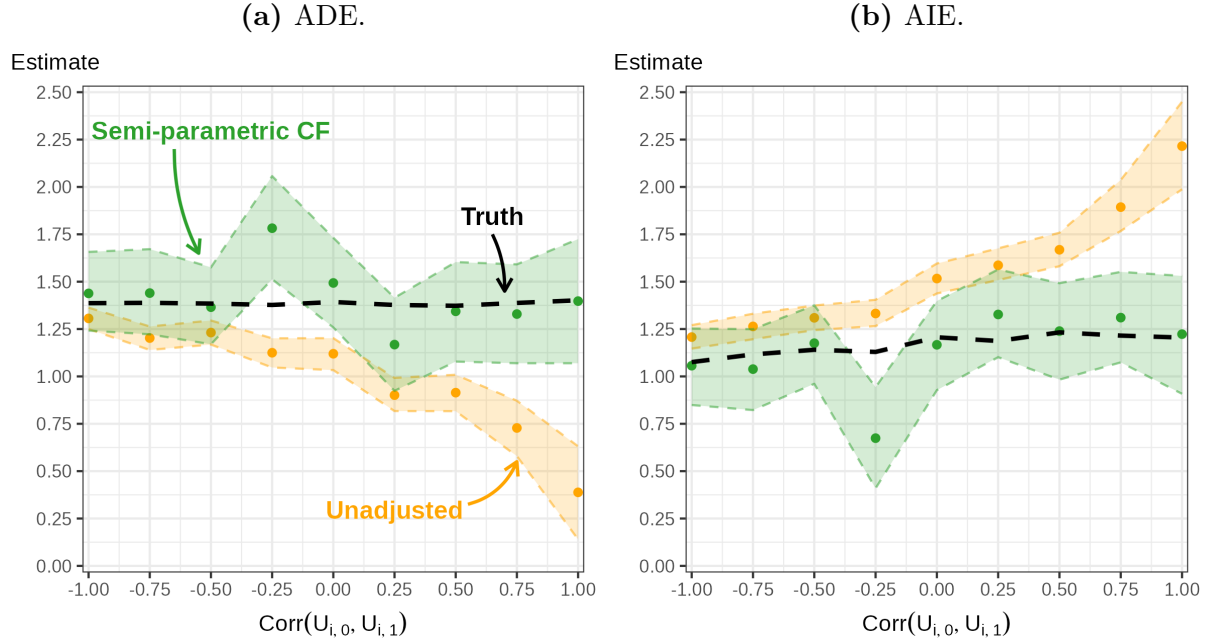
Figure 4 shows how these estimates perform, with a parametric CF approach, relative to the true value. The OLS estimates' distribution do not overlap the true values for any standard level of significance; the distance between the OLS estimates and the true values are the underlying bias terms derived in Theorem 1. The parametric CF approach perfectly reproduces the true values, as the probit first-stage correctly models the normally distributed error terms. The semi-parametric approach (not shown in Figure 4) performs similarly, with a wider distribution; this is to be expected comparing a correctly specified parametric model with a semi-parametric one.

The parametric CF may not be appropriate in setting with non-normal error terms. I simulated the same data again, but transform  $U_{0,i}, U_{1,i}$  to be correlated uniform errors (with the same standard deviations as previously). Figure 5 shows the resulting distribution of

<sup>18</sup>Note that  $\text{ATE} = \text{ADE} + \text{AIE}$  in this setting.  $\Pr(Z_i = 1) = 0.5$  ensures this equality, but it is not guaranteed in general. See Appendix A.8.

point-estimates, relative to the truth, for the parametric and semi-parametric approaches. The parametric CF is slightly off target, showing persistent bias from incorrectly specifying the error term distribution. The semi-parametric approach is centred exactly around the truth, with a slightly higher variance (as is expected).

**Figure 6:** CF Adjusted Estimates Work with Different Error Term Parameters.



**Note:** These figures show the OLS and CF point estimates of the ADE and AIE, for  $N = 5,000$  sample size, varying  $\text{Corr}(U_{0,i}, U_{1,i})$  values with  $\text{Var}(U_{0,i}) = 1, \text{Var}(U_{1,i}) = 1.5$  fixed. The black dashed line is the true value, coloured points are points estimates for the respective data generated, and shaded regions are the 95% confidence intervals from 1,000 bootstraps each. Orange represents OLS estimates, blue the CF approach.

The error terms determine the bias in OLS estimates of the ADE and AIE, so the bias varies for different values of the error-term parameters  $\text{Corr}(U_{0,i}, U_{1,i}) \in [-1, 1]$  and  $\text{Var}(U_{0,i}), \text{Var}(U_{1,i}) \geq 0$ . The true AIE values vary, because  $D_i(Z_i)$  compliers have higher average values of  $U_{1,i} - U_{0,i}$  as  $\text{Corr}(U_{0,i}, U_{1,i})$  increases. Figure 6 shows CF estimates against estimates calculated by standard OLS, showing 95% confidence intervals calculated from 1,000 bootstraps. The point estimates of the CF do not exactly equal the true values, as they are estimates from one simulation (not averages across many generated datasets, as in Figure 5). The CF approach improves on OLS estimates by correcting for bias, with confidence regions overlapping the true values.<sup>19</sup> This correction did not come for free: the standard errors are significantly greater in a CF approach than OLS. In this manner, this simulation shows the pros and cons of using the CF approach to estimating CM effects in

<sup>19</sup>In the appendix, Figure A1 shows the same simulation while varying  $\text{Var}(U_{1,i})$ , with fixed  $\text{Var}(U_{0,i}) = 1, \text{Corr}(U_{0,i}, U_{1,i}) = 0.5$ . The conclusion is the same as for varying the correlation coefficient,  $\rho$ , in Figure 6.

practice.

## 6 CM in the Oregon Health Insurance Experiment

Apply the parametric + semi-parametric CM to the Oregon data.

**Table 1:** CM Effect Estimates, Health Insurance Effects on Health and Happiness.

	First-stage	ATE	ADE	AIE	AIE / ATE
	(1)	(2)	(3)	(4)	(5)
<b>Panel A: Health overall good?</b>					
Unadjusted	4.100 (0.8100)	5.300 (0.8400)	5.400 (0.8500)	-0.080 (0.0460)	-0.015 (0.0095)
Parametric CF	4.20 (0.720)	5.30 (0.880)	3.60 (0.950)	1.40 (0.340)	0.26 (0.085)
Semi-parametric CF	4.20 (0.77)	5.30 (0.90)	1.80 (1.20)	3.50 (0.78)	0.65 (0.20)
<b>Panel B: Happy overall?</b>					
Unadjusted	4.100 (0.810)	5.000 (0.870)	4.800 (0.880)	0.130 (0.050)	0.027 (0.012)
Parametric CF	4.20 (0.72)	5.00 (0.91)	2.20 (1.00)	2.10 (0.44)	0.41 (0.13)
Semi-parametric CF	4.20 (0.77)	5.00 (0.89)	0.86 (1.20)	4.10 (0.85)	0.83 (0.25)

**Note:** This table shows the point estimates (and SEs in brackets) of applying the proposed CM methods to replication data from the Oregon Health Insurance Experiment (Finkelstein & Baicker 2014). These figures were calculated by using the wait-lottery to instrument for having health insurance in 2008, using the Abadie (2003) weighting scheme, as discussed in Appendix A.10. This means the effects identified here are averages among lottery compliers. The First-stage refers to the average effect of health insurance on healthcare usage, ATE average effect of health insurance on surveyed outcomes, ADE and AIE to respective CM effects through and absent healthcare usage. SEs were calculated with 1,000 bootstrap replications.

Note that the wait-list lottery was the random component, and it did not force people to take health insurance, instead randomly assigning access to health insurance. Identifying the direct and indirect effects of Medicaid health insurance explicitly would require further assumptions, and methods to isolation direct and indirect effects among those compliers. See Frölich (2007) for some approaches.

## 7 Summary and Concluding Remarks

This paper has studied a selection-on-observables approach to CM in a natural experiment setting. I have shown the pitfalls of using the most popular methods for estimating direct and indirect effects without a clear case for the mediator being ignorable. Using the Roy model as a benchmark, a mediator is unlikely to be ignorable in natural experiment settings, and the bias terms likely crowd out inference regarding CM effects.

This paper has contributed to the growing CM literature in economics, integrating labour economic theory for selection-into-treatment as a way of judging the credibility of conventional CM analyses. It has drawn on the classic literature, and pointed to already-in-use control function methods as a compelling way of estimating direct and indirect effects in a natural experiment setting. Further research could build on this approach by suggesting efficiency improvements, adjustments for common statistical irregularities (say, cluster dependence), or integrating the control function to the growing double robustness literature on CM (Farbmacher, Huber, Laff ers, Langen & Spindler 2022, Bia, Huber & Laff ers 2024).

This paper does not provide a blanket endorsement for applied researchers to use CM methods. The structural assumptions are strong, and design-based inference requires an instrument for mediator take-up; if the assumptions are broken, then selection-adjusted estimates of CM effects will also be biased, and will not improve on the selection-on-observables approach. And yet, there are likely settings in which the structural assumptions are credible. Mediator monotonicity aligns well with economic theory in many cases, and it is plausible for researchers to study big data settings with external variation in mediator take-up costs. In these cases, this paper opens the door to identifying mechanisms behind treatment effects in natural experiment settings.

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## A Supplementary Appendix

This section is for supplementary information, and validation of presented propositions and theorems. It is not meant for publication.

Any comments or suggestions may be sent to me at [seh325@cornell.edu](mailto:seh325@cornell.edu), or raised as an issue on the Github project, <https://github.com/shoganhennessy/mediation-natural-experiment>.

### A.1 Identification in Causal Mediation

Imai et al. (2010, Theorem 1) states that the ADE and AIE are identified under sequential ignorability, at each level of  $Z_i = 0, 1$ . For  $z' = 0, 1$ :

$$\begin{aligned}\mathbb{E}[Y_i(1, D_i(z')) - Y_i(0, D_i(z'))] &= \int \int \left( \mathbb{E}[Y_i | Z_i = 1, D_i, \mathbf{X}_i] - \mathbb{E}[Y_i | Z_i = 0, D_i, \mathbf{X}_i] \right) dF_{D_i | Z_i=z', \mathbf{X}_i} dF_{\mathbf{X}_i}, \\ \mathbb{E}[Y_i(z', D_i(1)) - Y_i(z', D_i(0))] &= \int \int \mathbb{E}[Y_i | Z_i = z', D_i, \mathbf{X}_i] \left( dF_{D_i | Z_i=1, \mathbf{X}_i} - dF_{D_i | Z_i=0, \mathbf{X}_i} \right) dF_{\mathbf{X}_i}.\end{aligned}$$

I focus on the averages, which are identified by consequence of the above.

$$\begin{aligned}\mathbb{E}[Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i))] &= \mathbb{E}_{Z_i} [\mathbb{E}[Y_i(1, D_i(z')) - Y_i(0, D_i(z')) | Z_i = z']] \\ \mathbb{E}[Y_i(Z_i, D_i(1)) - Y_i(Z_i, D_i(0))] &= \mathbb{E}_{Z_i} [\mathbb{E}[Y_i(z', D_i(1)) - Y_i(z', D_i(0)) | Z_i = z']]\end{aligned}$$

My estimand for the ADE is a simple rearrangement of the above. The estimand for the AIE relies on a different sequence, relying on (1) sequential ignorability, (2) conditional monotonicity. These give (1) identification equivalence of AIE local to compliers conditional on  $\mathbf{X}_i$  and AIE conditional on  $\mathbf{X}_i$ , LAIE = AIE, (2) identification of the complier score.

$$\begin{aligned}\mathbb{E}[Y_i(Z_i, D_i(1)) - Y_i(Z_i, D_i(0)) | \mathbf{X}_i] &= \Pr(D_i(0) = 0, D_i(1) = 1 | \mathbf{X}_i) \mathbb{E}[Y_i(Z_i, 1) - Y_i(Z_i, 0) | D_i(0) = 0, D_i(1) = 1, \mathbf{X}_i] \\ &= \Pr(D_i(0) = 0, D_i(1) = 1 | \mathbf{X}_i) \mathbb{E}[Y_i(Z_i, 1) - Y_i(Z_i, 0) | \mathbf{X}_i] \\ &= \Pr(D_i(0) = 0, D_i(1) = 1 | \mathbf{X}_i) \left( \mathbb{E}[Y_i | Z_i, D_i = 1, \mathbf{X}_i] - \mathbb{E}[Y_i | Z_i, D_i = 0, \mathbf{X}_i] \right) \\ &= \left( \mathbb{E}[D_i | Z_i = 1, \mathbf{X}_i] - \mathbb{E}[D_i | Z_i = 0, \mathbf{X}_i] \right) \left( \mathbb{E}[Y_i | Z_i, D_i = 1, \mathbf{X}_i] - \mathbb{E}[Y_i | Z_i, D_i = 0, \mathbf{X}_i] \right)\end{aligned}$$

Monotonicity is not technically required for the above. Breaking monotonicity would not change the identification in any of the above; it would be the same except replacing the complier score with a complier/defier score,  $\Pr(D_i(0) \neq D_i(1) | \mathbf{X}_i) = \mathbb{E}[D_i | Z_i = 1, \mathbf{X}_i] - \mathbb{E}[D_i | Z_i = 0, \mathbf{X}_i]$ .

## A.2 Bias in Causal Mediation (CM) Estimands

Suppose that  $Z_i$  is ignorable conditional on  $\mathbf{X}_i$ , but  $D_i$  is not.

### A.2.1 Bias in the Average Direct Effect (ADE)

To show that the conventional approach to mediation gives an estimate for the ADE with selection and group difference-bias, start with the components of the conventional estimands. This proof starts with the relevant expectations, conditional on a specific value of  $\mathbf{X}_i$  and  $d' \in \{0, 1\}$ .

$$\begin{aligned}\mathbb{E}[Y_i | Z_i = 1, D_i = d', \mathbf{X}_i] &= \mathbb{E}[Y_i(1, D_i(Z_i)) | D_i(1) = d', \mathbf{X}_i], \\ \mathbb{E}[Y_i | Z_i = 0, D_i = d', \mathbf{X}_i] &= \mathbb{E}[Y_i(0, D_i(Z_i)) | D_i(0) = d', \mathbf{X}_i]\end{aligned}$$

And so,

$$\begin{aligned}& \mathbb{E}[Y_i | Z_i = 1, D_i = d', \mathbf{X}_i] - \mathbb{E}[Y_i | Z_i = 0, D_i = d', \mathbf{X}_i] \\ &= \mathbb{E}[Y_i(1, D_i(Z_i)) | D_i(1) = d', \mathbf{X}_i] - \mathbb{E}[Y_i(0, D_i(Z_i)) | D_i(0) = d', \mathbf{X}_i] \\ &= \mathbb{E}[Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i)) | D_i(1) = d', \mathbf{X}_i] \\ &\quad + \mathbb{E}[Y_i(0, D_i(Z_i)) | D_i(1) = d', \mathbf{X}_i] - \mathbb{E}[Y_i(0, D_i(Z_i)) | D_i(0) = d', \mathbf{X}_i].\end{aligned}$$

The final term is a sum of the ADE, conditional on  $D_i(1) = d'$ , and a selection bias term — difference in baseline outcomes between the (partially overlapping) groups for whom  $D_i(1) = d'$  and  $D_i(0) = d'$ .

To reach the final term, note the following.

$$\begin{aligned}& \mathbb{E}[Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i)) | \mathbf{X}_i] \\ &= \mathbb{E}[Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i)) | D_i(1) = d', \mathbf{X}_i] \\ &\quad + \left(1 - \Pr(D_i(1) = d' | \mathbf{X}_i)\right) \left( \mathbb{E}[Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i)) | D_i(1) = d', \mathbf{X}_i] \right. \\ &\quad \left. - \mathbb{E}[Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i)) | D_i(1) = 1 - d', \mathbf{X}_i] \right)\end{aligned}$$

The second term is the difference between the ADE and LADE local to relevant complier groups.

Collect everything together, as follows.

$$\begin{aligned}
& \mathbb{E}[Y_i | Z_i = 1, D_i = d', \mathbf{X}_i] - \mathbb{E}[Y_i | Z_i = 0, D_i = d', \mathbf{X}_i] \\
&= \underbrace{\mathbb{E}[Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i)) | \mathbf{X}_i]}_{\text{ADE, conditional on } \mathbf{X}_i} \\
&+ \underbrace{\mathbb{E}[Y_i(0, D_i(Z_i)) | D_i(1) = d', \mathbf{X}_i] - \mathbb{E}[Y_i(0, D_i(Z_i)) | D_i(0) = d', \mathbf{X}_i]}_{\text{Selection bias}} \\
&+ \underbrace{\left(1 - \Pr(D_i(1) = d' | \mathbf{X}_i)\right) \left( \mathbb{E}[Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i)) | D_i(1) = 1 - d', \mathbf{X}_i] \right.}_{\text{group difference-bias}} \\
&\quad \left. - \mathbb{E}[Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i)) | D_i(1) = d', \mathbf{X}_i] \right)
\end{aligned}$$

The proof is achieved by applying the expectation across  $D_i = d'$ , and  $\mathbf{X}_i$ .

### A.2.2 Bias in the Average Indirect Effect (AIE)

To show that the conventional approach to mediation gives an estimate for the AIE with selection and group difference-bias, start with the definition of the ADE — the direct effect among compliers times the size of the complier group.

This proof starts with the relevant expectations, conditional on a specific value of  $\mathbf{X}_i$ .

$$\begin{aligned}
& \mathbb{E}[Y_i(Z_i, D_i(1)) - Y_i(Z_i, D_i(0)) | \mathbf{X}_i] \\
&= \Pr(D_i(0) = 0, D_i(1) = 1 | \mathbf{X}_i) \mathbb{E}[Y_i(Z_i, 1) - Y_i(Z_i, 0) | D_i(0) = 0, D_i(1) = 1, \mathbf{X}_i]
\end{aligned}$$

When  $D_i$  is not ignorable, the bias comes from estimating the second term,

$\mathbb{E}[Y_i(Z_i, 1) - Y_i(Z_i, 0) | D_i(0) = 0, D_i(1) = 1, \mathbf{X}_i]$ , the indirect effect among mediator compliers.

Let  $z' \in \{0, 1\}$ . Again, note the mean outcomes in terms of average potential outcomes,

$$\begin{aligned}
\mathbb{E}[Y_i | Z_i = z', D_i = 1, \mathbf{X}_i] &= \mathbb{E}[Y_i(z', 1) | D_i = 1, \mathbf{X}_i], \\
\mathbb{E}[Y_i | Z_i = z', D_i = 0, \mathbf{X}_i] &= \mathbb{E}[Y_i(z', 0) | D_i = 0, \mathbf{X}_i].
\end{aligned}$$

Compose the selection bias term, as follows.

$$\begin{aligned}
& \mathbb{E}[Y_i | Z_i = z', D_i = 1, \mathbf{X}_i] - \mathbb{E}[Y_i | Z_i = z', D_i = 0, \mathbf{X}_i] \\
&= \mathbb{E}[Y_i(z', 1) | D_i = 1, \mathbf{X}_i] - \mathbb{E}[Y_i(z', 0) | D_i = 0, \mathbf{X}_i] \\
&= \mathbb{E}[Y_i(z', 1) - Y_i(z', 0) | D_i = 1, \mathbf{X}_i] + \mathbb{E}[Y_i(z', 0) | D_i = 1, \mathbf{X}_i] - \mathbb{E}[Y_i(z', 0) | D_i = 0, \mathbf{X}_i]
\end{aligned}$$

The final term is a sum of the AIE, among the treated group  $D_i = 1$ , and a selection bias

term — difference in baseline potential outcomes between the groups for whom  $D_i = 1$  and  $D_i = 0$ .

The AIE is the direct effect among compliers times the size of the complier group, so we need to compensate for the difference between the treated group  $D_i = 1$  and complier group  $D_i(0) = 0, D_i(1) = 1$ .

Start with the difference between treated group's average and overall average.

$$\begin{aligned} & \mathbb{E} [Y_i(z', 1) - Y_i(z', 0) \mid D_i = 1, \mathbf{X}_i] \\ &= \mathbb{E} [Y_i(z', 1) - Y_i(z', 0) \mid \mathbf{X}_i] \\ &+ \left(1 - \Pr(D_i = 1 \mid \mathbf{X}_i)\right) \left( \begin{aligned} & \mathbb{E} [Y_i(z', 1) - Y_i(z', 0) \mid D_i = 1, \mathbf{X}_i] \\ & - \mathbb{E} [Y_i(z', 1) - Y_i(z', 0) \mid D_i = 0, \mathbf{X}_i] \end{aligned} \right) \end{aligned}$$

Then the difference between the compliers' average and the overall average.

$$\begin{aligned} & \mathbb{E} [Y_i(z', 1) - Y_i(z', 0) \mid D_i(0) = 0, D_i(1) = 1, \mathbf{X}_i] \\ &= \mathbb{E} [Y_i(z', 1) - Y_i(z', 0) \mid \mathbf{X}_i] \\ &+ \frac{1 - \Pr(D_i(0) = 0, D_i(1) = 1 \mid \mathbf{X}_i)}{\Pr(D_i(0) = 0, D_i(1) = 1 \mid \mathbf{X}_i)} \left( \begin{aligned} & \mathbb{E} [Y_i(z', 1) - Y_i(z', 0) \mid D_i(1) = 0 \text{ or } D_i(0) = 1, \mathbf{X}_i] \\ & - \mathbb{E} [Y_i(z', 1) - Y_i(z', 0) \mid \mathbf{X}_i] \end{aligned} \right) \end{aligned}$$

Collect everything together, as follows.

$$\begin{aligned} & \mathbb{E} [Y_i \mid Z_i = z', D_i = 1, \mathbf{X}_i] - \mathbb{E} [Y_i \mid Z_i = z', D_i = 0, \mathbf{X}_i] \\ &= \underbrace{\mathbb{E} [Y_i(z', 1) - Y_i(z', 0) \mid D_i(1) = 1, D_i(0) = 0, \mathbf{X}_i]}_{\text{AIE among compliers, conditional on } \mathbf{X}_i, Z_i = z'} \\ &+ \underbrace{\mathbb{E} [Y_i(z', 0) \mid D_i = 1, \mathbf{X}_i] - \mathbb{E} [Y_i(z', 0) \mid D_i = 0, \mathbf{X}_i]}_{\text{Selection bias}} \\ &+ \underbrace{\left[ \begin{aligned} & \left(1 - \Pr(D_i = 1 \mid \mathbf{X}_i)\right) \left( \begin{aligned} & \mathbb{E} [Y_i(z', 1) - Y_i(z', 0) \mid D_i = 1, \mathbf{X}_i] \\ & - \mathbb{E} [Y_i(z', 1) - Y_i(z', 0) \mid D_i = 0, \mathbf{X}_i] \end{aligned} \right) \\ & - \frac{1 - \Pr(D_i(0) = 0, D_i(1) = 1 \mid \mathbf{X}_i)}{\Pr(D_i(0) = 0, D_i(1) = 1 \mid \mathbf{X}_i)} \left( \begin{aligned} & \mathbb{E} [Y_i(z', 1) - Y_i(z', 0) \mid D_i(1) = 0 \text{ or } D_i(0) = 1, \mathbf{X}_i] \\ & - \mathbb{E} [Y_i(z', 1) - Y_i(z', 0) \mid \mathbf{X}_i] \end{aligned} \right) \end{aligned} \right]}_{\text{group difference-bias}} \end{aligned}$$

The proof is finally achieved by multiplying by the complier score,  $\Pr(D_i(0) = 0, D_i(1) = 1 \mid \mathbf{X}_i)$   $= \mathbb{E}[D_i \mid Z_i = 1, \mathbf{X}_i] - \mathbb{E}[D_i \mid Z_i = 0, \mathbf{X}_i]$ , then applying the expectation across  $Z_i = z'$ , and  $\mathbf{X}_i$ .



### A.3 A Regression Framework for Direct and Indirect Effects

Put  $\mu_{d'}(z'; \mathbf{X}) = \mathbb{E}[Y_i(z', d') | \mathbf{X}]$  and  $U_{d',i} = Y_i(z', d') - \mu_{d'}(z'; \mathbf{X})$  for each  $z', d' = 0, 1$ , so we have the following expressions:

$$Y_i(Z_i, 0) = \mu_0(Z_i; \mathbf{X}_i) + U_{0,i}, \quad Y_i(Z_i, 1) = \mu_1(Z_i; \mathbf{X}_i) + U_{1,i}.$$

$U_{0,i}, U_{1,i}$  are error terms with unknown distributions, mean independent of  $Z_i, \mathbf{X}_i$  by definition — but possibly correlated with  $D_i$ .  $Z_i$  is conditionally independent of potential outcomes, so that  $U_{0,i}, U_{1,i} \perp\!\!\!\perp Z_i$ .

The first-stage regression of  $Z \rightarrow Y$  has unbiased estimates, since  $Z_i \perp\!\!\!\perp D_i(\cdot) | \mathbf{X}_i$ . Put  $\pi(z'; \mathbf{X}) = \mathbb{E}[D_i(z') | \mathbf{X}]$ , and  $\eta_{z',i} = D_i(z') - \pi(z'; \mathbf{X})$  the first-stage error terms.

$$\begin{aligned} D_i &= Z_i D_i(1) + (1 - Z_i) D_i(0) \\ &= D_i(0) + Z_i [D_i(1) - D_i(0)] \\ &= \underbrace{\pi(0; \mathbf{X}_i)}_{\text{Intercept, } := \theta + \zeta(\mathbf{X}_i)} + \underbrace{Z_i (\pi(1; \mathbf{X}_i) - \pi(0; \mathbf{X}_i))}_{\text{Regressor, } := \bar{\pi} Z_i} + \underbrace{(1 - Z_i) \eta_{0,i} + Z_i \eta_{1,i}}_{\text{Errors, } := \eta_i} \\ \implies \mathbb{E}[D_i | Z_i, \mathbf{X}_i] &= \theta + \bar{\pi} Z_i + \zeta(\mathbf{X}_i). \end{aligned}$$

Since the ignorability assumption gives  $\mathbb{E}[Z_i \eta_{z',i} | \mathbf{X}_i] = \mathbb{E}[Z_i | \mathbf{X}_i] \mathbb{E}[\eta_{z',i} | \mathbf{X}_i] = 0$ , for each  $z' = 0, 1$ . By the same argument  $Z_i$  is also assumed independent of potential outcomes  $Y_i(\cdot, \cdot)$ , so that  $U_{0,i}, U_{1,i} \perp\!\!\!\perp Z_i$ . Thus, the reduced form regression  $Z \rightarrow Y$  also leads to unbiased estimates for the ATE.

The same cannot be said of the regression that estimates direct and indirect effects, without further assumptions.

$$\begin{aligned} Y_i &= Z_i Y_i(1, D_i(1)) + (1 - Z_i) Y_i(0, D_i(0)) \\ &= Z_i D_i Y_i(1, 1) \\ &\quad + (1 - Z_i) D_i Y_i(0, 1) \\ &\quad + Z_i (1 - D_i) Y_i(1, 0) \\ &\quad + (1 - Z_i) (1 - D_i) Y_i(0, 0) \\ &= Y_i(0, 0) \\ &\quad + Z_i [Y_i(1, 0) - Y_i(0, 0)] \\ &\quad + D_i [Y_i(0, 1) - Y_i(0, 0)] \\ &\quad + Z_i D_i [Y_i(1, 1) - Y_i(1, 0) - (Y_i(0, 1) - Y_i(0, 0))] \end{aligned}$$

And so  $Y_i$  can be written as a regression equation in terms of the observed factors and error

terms.

$$\begin{aligned}
Y_i &= \mu_0(0; \mathbf{X}_i) \\
&\quad + D_i [\mu_1(0; \mathbf{X}_i) - \mu_0(0; \mathbf{X}_i)] \\
&\quad + Z_i [\mu_0(1; \mathbf{X}_i) - \mu_0(0; \mathbf{X}_i)] \\
&\quad + Z_i D_i [\mu_1(1; \mathbf{X}_i) - \mu_0(1; \mathbf{X}_i) - (\mu_1(0; \mathbf{X}_i) - \mu_0(0; \mathbf{X}_i))] \\
&\quad + U_{0,i} + D_i (U_{1,i} - U_{0,i}) \\
&= \alpha + \beta D_i + \gamma Z_i + \delta Z_i D_i + \varphi(\mathbf{X}_i) + (1 - D_i) U_{0,i} + D_i U_{1,i}
\end{aligned}$$

With the following definitions:

- (a)  $\alpha = \mathbb{E} [\mu_0(0; \mathbf{X}_i)]$  and  $\varphi(\mathbf{X}_i) = \mu_0(0; \mathbf{X}_i) - \alpha$  are the intercept terms.
- (b)  $\beta = \mu_1(0; \mathbf{X}_i) - \mu_0(0; \mathbf{X}_i)$  is the indirect effect under  $Z_i = 0$
- (c)  $\gamma = \mu_0(1; \mathbf{X}_i) - \mu_0(0; \mathbf{X}_i)$  is the direct effect under  $D_i = 0$ .
- (d)  $\delta = \mu_1(1; \mathbf{X}_i) - \mu_0(1; \mathbf{X}_i) - (\mu_1(0; \mathbf{X}_i) - \mu_0(0; \mathbf{X}_i))$  is the interaction effect.
- (e)  $(1 - D_i) U_{0,i} + D_i U_{1,i}$  is the remaining error term.

This sequence gives us the resulting regression equation:

$$\begin{aligned}
\mathbb{E} [Y_i | Z_i, D_i, \mathbf{X}_i] &= \alpha + \beta D_i + \gamma Z_i + \delta Z_i D_i + \varphi(\mathbf{X}_i) \\
&\quad + (1 - D_i) \mathbb{E} [U_{0,i} | D_i = 0, \mathbf{X}_i] + D_i \mathbb{E} [U_{1,i} | D_i = 1, \mathbf{X}_i]
\end{aligned}$$

Taking the conditional expectation, and collecting for the expressions of the direct and indirect effects:

$$\begin{aligned}
\mathbb{E} [Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i))] &= \mathbb{E} [\gamma + \delta D_i] \\
\mathbb{E} [Y_i(Z_i, D_i(1)) - Y_i(Z_i, D_i(0))] &= \mathbb{E} [\bar{\pi} (\beta + Z_i \delta + \tilde{U}_i)]
\end{aligned}$$

These equations have simpler expressions after assuming constant treatment effects in a linear framework; I have avoided this as having compliers, and controlling for observed factors  $\mathbf{X}_i$  only makes sense in the case of heterogeneous treatment effects.

These terms are conventionally estimated in a simultaneous regression (Imai et al. 2010). If sequential ignorability does not hold, then the regression estimates from estimating the mediation equations (without adjusting for the contaminated bias term) suffer from omitted variables bias.

$$\begin{aligned}
\mathbb{E}_{\mathbf{X}_i} [\mathbb{E} [Y_i | Z_i = D_i = 0, \mathbf{X}_i]] &= \mathbb{E} [\alpha] + \mathbb{E} [U_{0,i} | D_i = 0] \\
\mathbb{E}_{\mathbf{X}_i} [\mathbb{E} [Y_i | Z_i = 0, D_i = 1, \mathbf{X}_i] - \mathbb{E} [Y_i | Z_i = 0, D_i = 0, \mathbf{X}_i]] &= \mathbb{E} [\beta] + (\mathbb{E} [U_{1,i} | D_i = 1] - \mathbb{E} [U_{0,i} | D_i = 0]) \\
\mathbb{E}_{\mathbf{X}_i} [\mathbb{E} [Y_i | Z_i = 1, D_i = 0, \mathbf{X}_i] - \mathbb{E} [Y_i | Z_i = 0, D_i = 0, \mathbf{X}_i]] &= \mathbb{E} [\gamma] + \mathbb{E} [U_{0,i} | D_i = 0] \\
\mathbb{E}_{\mathbf{X}_i} \left[ \mathbb{E} [Y_i | Z_i = 1, D_i = 1, \mathbf{X}_i] - \mathbb{E} [Y_i | Z_i = 1, D_i = 0, \mathbf{X}_i] \right. \\
&\quad \left. - (\mathbb{E} [Y_i | Z_i = 0, D_i = 1, \mathbf{X}_i] - \mathbb{E} [Y_i | Z_i = 0, D_i = 0, \mathbf{X}_i]) \right] = \mathbb{E} [\delta]
\end{aligned}$$

And so the ADE and AIE estimates are contaminated by these bias terms. Additionally, the AIE estimates refers to gains from the mediator among  $D(z)$  compliers (not the entire average), so will be biased when not accounting for  $\tilde{U}_i$ , too.

## A.4 Roy Model and Sequential Ignorability

*Proof of Proposition 1.*

Suppose  $Z_i$  is ignorable, and selection-into- $D_i$  follows a Roy model, with the definitions in Section 3. If selection-into- $D_i$  is degenerate on  $U_{0,i}, U_{1,i}$ :

$$\mathbb{E} [D_i | Z_i, \mathbf{X}_i, U_{1,i} - U_{0,i} = u] = \mathbb{E} [D_i | Z_i, \mathbf{X}_i, U_{1,i} - U_{0,i} = u'], \text{ for all } u, u' \text{ in the range of } U_{1,i} - U_{0,i}.$$

In this case, the control set  $\mathbf{X}_i$  and the costs  $\mu_c, U_{c,i}$  are the only determinants of selection-into- $D_i$  — and,  $U_{0,i}, U_{1,i}$  play no role. This could be achieved by either assuming that unobserved gains are degenerate (the researcher had observed everything in  $\mathbf{X}_i$ ), or selection-into- $D_i$  had been disrupted in some fashion (e.g., by a natural experiment design for  $D_i$ ).

To motivate a contraposition argument, suppose  $D_i$  is ignorable conditional on  $Z_i, \mathbf{X}_i$ . For each  $z', d' = 0, 1$

$$\begin{aligned}
D_i \perp\!\!\!\perp Y_i(z', d') \mid \mathbf{X}_i, Z_i = z' \\
\implies D_i \perp\!\!\!\perp \mu_{d'}(z'; \mathbf{X}_i) + U_{d',i} \mid \mathbf{X}_i, Z_i = z' \\
\implies D_i \perp\!\!\!\perp U_{d',i} \mid \mathbf{X}_i, Z_i = z' \\
\implies D_i \perp\!\!\!\perp U_{1,i} - U_{0,i} \mid \mathbf{X}_i, Z_i = z' \\
\implies \mathbb{E} [D_i | U_{1,i} - U_{0,i} = u', \mathbf{X}_i, Z_i = z'] = \mathbb{E} [D_i | \mathbf{X}_i, Z_i = z'] \\
\text{for all } u' \text{ in the range of } U_{1,i} - U_{0,i}.
\end{aligned}$$

This final implication is that selection-into- $D_i$  is degenerate on  $U_{0,i}, U_{1,i}$ . Thus, a contraposition argument has that if selection-into- $D_i$  is non-degenerate on  $U_{0,i}, U_{1,i}$ , then  $D_i$  is not ignorable.

## A.5 Monotonicity $\implies$ Selection Model, in a CM Setting.

*Proof that (conditional) monotonicity implies a selection model representation in a CM setting. This proof is an applied example of the [Vytlacil \(2002\)](#) equivalence result, now including conditioning covariates  $\mathbf{X}_i$ , and is presented merely as a validation exercise.*

Assume condition monotonicity [CF-1](#) holds, for any treatment values  $z < z'$  and any covariate value  $\mathbf{X}_i = \mathbf{x}$ .

$$\Pr(D_i(z') \geq D_i(z) \mid \mathbf{x}) = 1.$$

For each value of  $\mathbf{X}_i = \mathbf{x}$  and any treatment values  $z < z'$ , we first define:

- $\mathcal{A} = \{i : D_i(z) = D_i(z') = 1\}$ , always-mediators
- $\mathcal{N} = \{i : D_i(z) = D_i(z') = 0\}$ , never-mediators
- $\mathcal{C} = \{i : D_i(z) = 0, D_i(z') = 1\}$ , mediator-compliers.

For any mediator complier  $i \in \mathcal{C}$ , partition the set as follows.

- $\mathcal{Z}_1(i) = \{z' : D_i(z') = 1\}$ , treatment values where  $i$  takes the mediator
- $\mathcal{Z}_0(i) = \{z' : D_i(z') = 0\}$ , treatment values where  $i$  doesn't take the mediator.

Note that having binary  $Z_i = 0, 1$  reduces this to the simple case of  $\mathcal{Z}_0(i) = \{0\}$ , and  $\mathcal{Z}_1(i) = \{1\}$ . The equivalence result holds for continuous values of  $Z_i$ , so continue with the more general  $\mathcal{Z}_0(i), \mathcal{Z}_1(i)$  notation.

By monotonicity, we have

$$\sup_{z' \in \mathcal{Z}_0(i)} \pi(z'; \mathbf{x}) \leq \inf_{z' \in \mathcal{Z}_1(i)} \pi(z'; \mathbf{x}), \quad \text{for any } i \in \mathcal{C}$$

where  $\pi(z'; \mathbf{x}) = \Pr(D_i = 1 \mid Z_i = z', \mathbf{X}_i = \mathbf{x})$  is the mediator propensity score. A simple proof by contradiction verifies this statement ([Vytlacil 2002](#), Lemma 1).

Now we construct  $V_i$  as follows:

$$V_i = \begin{cases} 1, & \text{if } i \in \mathcal{N} \\ 0, & \text{if } i \in \mathcal{A} \\ \inf_{z' \in \mathcal{Z}_1(i)} \pi(z'; \mathbf{x}), & \text{if } i \in \mathcal{C}. \end{cases}$$

Define  $\psi(z'; \mathbf{x}) = \pi(z'; \mathbf{x})$ . Then we can represent  $D_i(z')$  as a selection model,

$$D_i(z') = \mathbb{1} \{ \psi(z'; \mathbf{X}_i) \geq V_i \}, \quad \text{for } z' = 0, 1.$$

We can verify this works:

- For  $i \in \mathcal{A}$ :  $V_i = 0$  and  $\psi(z'; \mathbf{x}) \geq 0$  for all  $z'$ , so  $D_i(z') = 1$
- For  $i \in \mathcal{N}$ :  $V_i = 1$  and  $\psi(z'; \mathbf{x}) \leq 1$  for all  $z'$ , with  $\psi(z'; \mathbf{x}) < 1$  for  $z' \in \mathcal{Z}_0(i)$ , so  $D_i(z') = 0$  for  $z' \in \mathcal{Z}_0(i)$
- For  $i \in \mathcal{C}$ :  $V_i = \inf_{z' \in \mathcal{Z}_1(i)} \pi(z'; \mathbf{x})$ 
  - When  $z' \in \mathcal{Z}_1(i)$ :  $\psi(z'; \mathbf{x}) \geq \inf_{z'' \in \mathcal{Z}_1(i)} \pi(z''; \mathbf{x}) = V_i$ , so  $D_i(z') = 1$
  - When  $z' \in \mathcal{Z}_0(i)$ :  $\psi(z'; \mathbf{x}) < \inf_{z'' \in \mathcal{Z}_1(i)} \pi(z''; \mathbf{x}) = V_i$ , so  $D_i(z') = 0$ .

Therefore, the construction  $D_i(z') = \mathbb{1} \{ \psi(z'; \mathbf{X}_i) \geq V_i \}$  is a valid representation of the selection process under monotonicity.

This selection model can be transformed to one with a uniform distribution, to get the general selection model of Heckman & Vytlacil (2005). Let  $F_V(\cdot | \mathbf{X}_i)$  be the conditional cumulative density function of  $V_i$  given  $\mathbf{X}_i$ . Define

$$U_i = F_V(V_i | \mathbf{X}_i)$$

$$\pi(z'; \mathbf{X}_i) = F_V(\psi(z'; \mathbf{X}_i) | \mathbf{X}_i) = \Pr(D_i = 1 | Z_i = z', \mathbf{X}_i)$$

We can then equivalently represent the mediator choice as the transformed selection model

$$D_i(z') = \mathbb{1} \{ \pi(z'; \mathbf{X}_i) \geq U_i \}, \quad \text{for } z' = 0, 1$$

where  $U_i | \mathbf{X}_i \sim \text{Uniform}(0, 1)$  by the probability integral transformation.

## A.6 Control Function (CF) Identification of the Second-stage

*Proof of Proposition 2.* This proof relies heavily on the notation and reasoning of Kline & Walters (2019) for an IV setting.

By Assumption CF-1 (mediator monotonicity), selection-into-mediator can be represented as a threshold-crossing selection model.

$$D_i(z') = \mathbb{1} \{ \pi(z'; \mathbf{X}_i) \geq U_i \}, \quad \text{for } z' = 0, 1$$

where  $U_i = F_V(V_i | \mathbf{X}_i)$  follows a uniform distribution on  $[0, 1]$ , and  $\pi(z'; \mathbf{X}_i) = \mathbb{E}[D_i | Z_i = z', \mathbf{X}_i]$  is the mediator propensity score.

The threshold crossing selection model represents individuals who refuse the mediator as follows:

$$D_i = 0 \implies \pi(Z_i; \mathbf{X}_i) < U_i$$

Our objective is to determine  $\mathbb{E}[U_{0,i} \mid D_i = 0, Z_i, \mathbf{X}_i]$ , which can then be written as

$$\mathbb{E}[U_{0,i} \mid \pi(Z_i; \mathbf{X}_i) < U_i, Z_i, \mathbf{X}_i].$$

Since  $Z_i$  is ignorable, we have:

$$\mathbb{E}[U_{0,i} \mid \pi(Z_i; \mathbf{X}_i) < U_i, Z_i, \mathbf{X}_i] = \mathbb{E}[U_{0,i} \mid \pi(Z_i; \mathbf{X}_i) < U_i]$$

Assumption [CF-2](#) has  $\text{Cov}(U_i, U_{0,i}) \neq 0$ . This non-zero covariance implies statistical dependence between the selection error and outcome error. This dependence allows us to represent  $U_{0,i}$  using a linear projection. We use  $F_V^{-1}(U_i \mid \mathbf{X}_i)$  rather than  $U_i$  directly in the projection to allow for flexibility in how the selection error affects outcomes. The linear projection can be written as follows

$$U_{0,i} = \rho_0(F_V^{-1}(U_i \mid \mathbf{X}_i) - \mu_V) + \varepsilon_{0,i},$$

where

- $\mu_V = \mathbb{E}[F_V^{-1}(U_i \mid \mathbf{X}_i)]$  is the mean of  $F_V^{-1}(U_i \mid \mathbf{X}_i)$
- $\rho_0 = \frac{\text{Cov}(U_{0,i}, F_V^{-1}(U_i \mid \mathbf{X}_i))}{\text{Var}(F_V^{-1}(U_i \mid \mathbf{X}_i))}$  is the projection coefficient
- $\varepsilon_{0,i}$  is a residual with  $\mathbb{E}[\varepsilon_{0,i} \mid F_V^{-1}(U_i \mid \mathbf{X}_i)] = 0$ .

The coefficient  $\rho_0$  is the slope in the best linear predictor of  $U_{0,i}$  given  $F_V^{-1}(U_i \mid \mathbf{X}_i)$ , and is chosen to ensure that the residual  $\varepsilon_{0,i}$  is uncorrelated with  $F_V^{-1}(U_i \mid \mathbf{X}_i)$ . This property is crucial for the identification strategy, as it isolates the component of  $U_i$  that is related to selection-into- $D_i$ .

The non-zero covariance condition in [CF-2](#) ensures  $\rho_0 \neq 0$ , so is relevant. Since  $U_i$  and  $F_V^{-1}(U_i \mid \mathbf{X}_i)$  are related by a monotonic transformation (the inverse cumulative density function), the covariance  $\text{Cov}(U_i, U_{0,i}) \neq 0$  implies  $\text{Cov}(F_V^{-1}(U_i \mid \mathbf{X}_i), U_{0,i}) \neq 0$ .

Given the linear projection of  $U_{0,i}$  onto  $F_V^{-1}(U_i \mid \mathbf{X}_i)$ , we can compute the conditional expectation:

$$\mathbb{E}[U_{0,i} \mid \pi(Z_i; \mathbf{X}_i) < U_i] = \mathbb{E}[\rho_0(F_V^{-1}(U_i \mid \mathbf{X}_i) - \mu_V) + \varepsilon_{0,i} \mid \pi(Z_i; \mathbf{X}_i) < U_i]$$

Since  $\mathbb{E}[\varepsilon_{0,i} \mid F_V^{-1}(U_i \mid \mathbf{X}_i)] = 0$  by construction, and  $U_i$  is a function of  $F_V^{-1}(U_i \mid \mathbf{X}_i)$ , we have

$$\mathbb{E}[\varepsilon_{0,i} \mid \pi(Z_i; \mathbf{X}_i) < U_i] = 0.$$

Therefore:

$$\mathbb{E}[U_{0,i} | \pi(Z_i; \mathbf{X}_i) < U_i] = \rho_0 \mathbb{E}[F_V^{-1}(U_i | \mathbf{X}_i) - \mu_V | \pi(Z_i; \mathbf{X}_i) < U_i].$$

This gives us the control function representation:

$$\mathbb{E}[U_{0,i} | D_i = 0, Z_i, \mathbf{X}_i] = \rho_0 \lambda_0(\pi(Z_i; \mathbf{X}_i))$$

where  $\lambda_0(p') = \mathbb{E}[F_V^{-1}(U_i | \mathbf{X}_i) - \mu_V | p' < U_i]$ . The control function  $\lambda_0(p')$  captures the expected value of the transformed selection term, conditional on being above the threshold  $p' \in (0, 1)$ .

The same sequence of steps for mediator takers,  $D_i = 1$ , gives the other CF:

$$\mathbb{E}[U_{1,i} | D_i = 1, Z_i, \mathbf{X}_i] = \rho_1 \lambda_1(\pi(Z_i; \mathbf{X}_i)),$$

where  $\lambda_1(p') = \mathbb{E}[F_V^{-1}(U_i | \mathbf{X}_i) - \mu_V | U_i \leq p']$  for  $p' \in (0, 1)$ , and  $\rho_1 = \frac{\text{Cov}(U_{1,i}, F_V^{-1}(U_i | \mathbf{X}_i))}{\text{Var}(F_V^{-1}(U_i | \mathbf{X}_i))}$  is the corresponding projection coefficient.

The relationship between  $\lambda_0(p')$  and  $\lambda_1(p')$  can be derived as:

$$\lambda_1(p') = -\lambda_0(p') \left( \frac{1 - p'}{p'} \right), \text{ for } p' \in (0, 1).$$

This relationship ensures consistency in the CF approach across the  $D_i = 0$  and  $D_i = 1$  groups (Kline & Walters 2019).

Assumption CF-3 (mediator take-up cost instrument  $\mathbf{X}_i^{\text{IV}}$ ) ensures identification of the propensity score function  $\pi(z'; \mathbf{X}_i)$  in the first stage by providing valid instrumental variation. This variation allows us to identify the propensity score, and consequently the control functions  $\lambda_0$  and  $\lambda_1$ .

Combining all elements, the conditional expectation of  $Y_i$  given  $Z_i, D_i, \mathbf{X}_i$  is

$$\begin{aligned} \mathbb{E}[Y_i | Z_i, D_i, \mathbf{X}_i] &= \alpha + \beta D_i + \gamma Z_i + \delta Z_i D_i + \varphi(\mathbf{X}_i) \\ &\quad + (1 - D_i) \mathbb{E}[U_{0,i} | D_i = 0] + D_i \mathbb{E}[U_{1,i} | D_i = 1]. \end{aligned}$$

Substitute the CFs,

$$\begin{aligned} &(1 - D_i) \mathbb{E}[U_{0,i} | Z_i, D_i = 0, \mathbf{X}_i] + D_i \mathbb{E}[U_{1,i} | Z_i, D_i = 1, \mathbf{X}_i] \\ &= (1 - D_i) \rho_0 \lambda_0(\pi(Z_i; \mathbf{X}_i)) + D_i \rho_1 \lambda_1(\pi(Z_i; \mathbf{X}_i)). \end{aligned}$$

This gives the final result,

$$\begin{aligned}\mathbb{E}[Y_i | Z_i, D_i, \mathbf{X}_i] &= \alpha + \beta D_i + \gamma Z_i + \delta Z_i D_i + \varphi(\mathbf{X}_i) \\ &\quad + \rho_0 (1 - D_i) \lambda_0(\pi(Z_i; \mathbf{X}_i)) + \rho_1 D_i \lambda_1(\pi(Z_i; \mathbf{X}_i)).\end{aligned}$$

All parameters —  $\alpha, \beta, \gamma, \delta, \varphi(\cdot), \rho_0, \rho_1$  — are identified once we control for selection bias through the CFs  $\lambda_0, \lambda_1$ , with  $\pi(z'; \mathbf{X}_i)$  identified separately in the first-stage.  $\lambda_0, \lambda_1$  can be assumed to be certain functions (say, the inverse Mills ratio in [Heckman 1979](#)), or treated as non-parametric parameters to be estimated — at cost of the constant and  $\rho_0, \rho_1$  no longer being separately identified from  $\lambda_0, \lambda_1$ , see [Appendix A.8](#).

## A.7 Control Function (CF) Identification of the ADE and AIE

*Proof of Theorem CF.*

Assume [CF-1](#), [CF-2](#), [CF-3](#) hold. Then Proposition 2 has  $\alpha, \beta, \gamma, \delta, \varphi(\cdot), \rho_0, \rho_1$  identified in a regression. The following composes the ADE and AIE from these parameters.

For the ADE,

$$\begin{aligned}\mathbb{E}[\gamma + \delta D_i] &= \mathbb{E}\left[\left(\mu_0(1; \mathbf{X}_i) - \mu_0(0; \mathbf{X}_i)\right) + D_i\left(\mu_1(1; \mathbf{X}_i) - \mu_0(1; \mathbf{X}_i) - (\mu_1(0; \mathbf{X}_i) - \mu_0(0; \mathbf{X}_i))\right)\right] \\ &= \mathbb{E}\left[D_i\left(\mu_1(1; \mathbf{X}_i) - \mu_1(0; \mathbf{X}_i)\right) + (1 - D_i)\left(\mu_0(1; \mathbf{X}_i) - \mu_0(0; \mathbf{X}_i)\right)\right] \\ &= \mathbb{E}\left[D_i\left(Y_i(1, 1) - U_{1,i} - (Y_i(0, 1) - U_{1,i})\right) + (1 - D_i)\left(Y_i(1, 0) - U_{0,i} - (Y_i(0, 0) - U_{0,i})\right)\right] \\ &= \mathbb{E}\left[D_i\left(Y_i(1, 1) - Y_i(0, 1)\right) + (1 - D_i)\left(Y_i(1, 0) - Y_i(0, 0)\right)\right] \\ &= \mathbb{E}[Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i))] \\ &= \text{ADE}.\end{aligned}$$

Identification is similar for the AIE, but also involves the complier adjustment term.

$$\begin{aligned}(\rho_1 - \rho_0) \Gamma(\pi(0; \mathbf{X}_i), \pi(1; \mathbf{X}_i)) &= (\rho_1 - \rho_0) \frac{\pi(1; \mathbf{X}_i) \lambda_1(\pi(1; \mathbf{X}_i)) - \pi(0; \mathbf{X}_i) \lambda_1(\pi(0; \mathbf{X}_i))}{\pi(1; \mathbf{X}_i) - \pi(0; \mathbf{X}_i)} \\ &= (\rho_1 - \rho_0) \mathbb{E}\left[F_V^{-1}(U_i | \mathbf{X}_i) - \mu_V \mid \pi(0; \mathbf{X}_i) < U_i \leq \pi(1; \mathbf{X}_i), \mathbf{X}_i\right] \\ &= (\rho_1 - \rho_0) \mathbb{E}\left[F_V^{-1}(U_i | \mathbf{X}_i) - \mu_V \mid D_i(0) = 0, D_i(1) = 1, \mathbf{X}_i\right] \\ &= \mathbb{E}\left[\rho_1 (F_V^{-1}(U_i | \mathbf{X}_i) - \mu_V) \mid D_i(0) = 0, D_i(1) = 1, \mathbf{X}_i\right] \\ &\quad - \mathbb{E}\left[\rho_0 (F_V^{-1}(U_i | \mathbf{X}_i) - \mu_V) \mid D_i(0) = 0, D_i(1) = 1, \mathbf{X}_i\right] \\ &= \mathbb{E}[U_{1,i} - U_{0,i} \mid D_i(0) = 0, D_i(1) = 1, \mathbf{X}_i].\end{aligned}$$

This complier adjustment was first presented for an IV setting by [Kline & Walters \(2019\)](#).



Collecting for the AIE,

$$\begin{aligned}
& \mathbb{E} \left[ \bar{\pi} \left( \beta + \delta Z_i + (\rho_1 - \rho_0) \Gamma(\pi(0; \mathbf{X}_i), \pi(1; \mathbf{X}_i)) \right) \right] \\
&= \mathbb{E} \left[ \bar{\pi} \left( \left( \mu_1(0; \mathbf{X}_i) - \mu_0(0; \mathbf{X}_i) \right) + Z_i \left( \mu_1(1; \mathbf{X}_i) - \mu_0(1; \mathbf{X}_i) - (\mu_1(0; \mathbf{X}_i) - \mu_0(0; \mathbf{X}_i)) \right) \right) \right] \\
&\quad + \mathbb{E} \left[ \bar{\pi} \mathbb{E} [U_{1,i} - U_{0,i} \mid D_i(0) = 0, D_i(1) = 1, \mathbf{X}_i] \right] \\
&= \mathbb{E} \left[ \bar{\pi} \left( Z_i \left( \mu_1(1; \mathbf{X}_i) - \mu_0(1; \mathbf{X}_i) \right) + (1 - Z_i) \left( \mu_1(0; \mathbf{X}_i) - \mu_0(0; \mathbf{X}_i) \right) \right) \right] \\
&\quad + \mathbb{E} \left[ \bar{\pi} \mathbb{E} [U_{1,i} - U_{0,i} \mid D_i(0) = 0, D_i(1) = 1, \mathbf{X}_i] \right] \\
&= \mathbb{E} \left[ \bar{\pi} \left( \mu_1(Z_i, \mathbf{X}_i) - \mu_0(Z_i, \mathbf{X}_i) + \mathbb{E} [U_{1,i} - U_{0,i} \mid D_i(0) = 0, D_i(1) = 1, \mathbf{X}_i] \right) \right] \\
&= \mathbb{E} [\bar{\pi} \mathbb{E} [\mu_1(Z_i, \mathbf{X}_i) - \mu_0(Z_i, \mathbf{X}_i) + U_{1,i} - U_{0,i} \mid D_i(0) = 0, D_i(1) = 1, \mathbf{X}_i]] \\
&= \mathbb{E} [\mathbb{E} [D_i(1) - D_i(0) \mid \mathbf{X}_i] \mathbb{E} [Y_i(Z_i, 1) - Y_i(Z_i, 0) \mid D_i(0) = 0, D_i(1) = 1, \mathbf{X}_i]] \\
&= \mathbb{E} [\mathbb{E} [Y_i(Z_i, D_i(1)) - Y_i(Z_i, D_i(0)) \mid \mathbf{X}_i]] \\
&= \mathbb{E} [Y_i(Z_i, D_i(1)) - Y_i(Z_i, D_i(0))] \\
&= \text{AIE}.
\end{aligned}$$

## A.8 Semi-parametric Estimation of the AIE

It is difficult to directly use the CFs to compose estimates of the complier adjustment term, because various intercepts lose identification, but also because trusting semi-parametric estimates at individual points across the  $\hat{\lambda}_0(p'), \hat{\lambda}_1(p')$  functions would increase variation more than is necessary.

This can be avoided by noting the relation between the ATE and the conditional ADE and conditional AIE. The following showing how to identify the AIE via relation to the ATE and conditional ADE, and omits the conditional on  $\mathbf{X}_i$  for brevity.

A simple algebraic rearrangement has the following (as first noted in [Imai et al. 2010](#), Section 3.1),

$$\begin{aligned}
\text{ATE} &= \mathbb{E} [Y_i(1, D_i(1)) - Y_i(0, D_i(1))] \\
&= \mathbb{E} [Y_i(1, D_i(1)) - Y_i(0, D_i(1))] + \mathbb{E} [Y_i(0, D_i(1)) - Y_i(0, D_i(0))] \\
&= \underbrace{\mathbb{E} [Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i)) \mid Z_i = 1]}_{\text{ADE conditional on } Z_i=1} + \underbrace{\mathbb{E} [Y_i(Z_i, D_i(1)) - Y_i(Z_i, D_i(0)) \mid Z_i = 0]}_{\text{AIE conditional on } Z_i=0}.
\end{aligned}$$

A similar re-arrangement also has the following,

$$\text{ATE} = \underbrace{\mathbb{E} [Y_i(Z_i, D_i(1)) - Y_i(Z_i, D_i(0)) \mid Z_i = 1]}_{\text{AIE conditional on } Z_i=1} + \underbrace{\mathbb{E} [Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i)) \mid Z_i = 0]}_{\text{ADE conditional on } Z_i=0}.$$

Reverting to the regression notation, to show how the ADE conditional on  $Z_i$  is identified:

$$\begin{aligned} \text{ADE} &= \mathbb{E} [Y_i(1, D_i(Z_i)) - Y_i(0, D_i(Z_i))] \\ &= \mathbb{E} [\gamma + \delta D_i(Z_i)] \\ \implies \text{ADE conditional on } Z_i = 0 &= \mathbb{E} [\gamma + \delta D_i(Z_i) \mid Z_i = 0] \\ &= \mathbb{E} [\gamma + \delta D_i(0)] \\ \text{ADE conditional on } Z_i = 1 &= \mathbb{E} [\gamma + \delta D_i(Z_i) \mid Z_i = 1] \\ &= \mathbb{E} [\gamma + \delta D_i(1)]. \end{aligned}$$

Finally achieve identification of the AIE via the ATE and conditional ADE, as follows,

$$\begin{aligned} \text{AIE} &= \Pr(Z_i = 0) \underbrace{\mathbb{E} [Y_i(Z_i, D_i(1)) - Y_i(Z_i, D_i(0)) \mid Z_i = 0]}_{\text{AIE conditional on } Z_i=0} \\ &\quad + \Pr(Z_i = 1) \underbrace{\mathbb{E} [Y_i(Z_i, D_i(1)) - Y_i(Z_i, D_i(0)) \mid Z_i = 1]}_{\text{AIE conditional on } Z_i=1} \\ &= \Pr(Z_i = 0) [\text{ATE} - (\text{ADE conditional on } Z_i = 1)] \\ &\quad + \Pr(Z_i = 1) [\text{ATE} - (\text{ADE conditional on } Z_i = 0)] \\ &= \text{ATE} - \Pr(Z_i = 0) \mathbb{E} [\gamma + \delta D_i(1)] - \Pr(Z_i = 1) \mathbb{E} [\gamma + \delta D_i(0)]. \end{aligned}$$

The semi-parametric AIE estimate then uses this representation, avoiding directly interacting with the estimated CFs, by plugging in estimates  $\widehat{\Pr}(Z_i = 1) = \bar{Z}$ ,  $\widehat{\text{ATE}}$ , and the estimates from each side of the  $D_i = 0, 1$  separated samples  $\hat{\gamma}, \hat{\delta}$ .

$$\widehat{\text{AIE}}^{\text{CF}} = \widehat{\text{ATE}} - (1 - \bar{Z}) \left( \hat{\gamma} + \frac{1}{N} \sum_{i=1}^N \hat{\delta} \hat{\pi}(1; \mathbf{X}_i) \right) - \bar{Z} \left( \hat{\gamma} + \frac{1}{N} \sum_{i=1}^N \hat{\delta} \hat{\pi}(0; \mathbf{X}_i) \right),$$

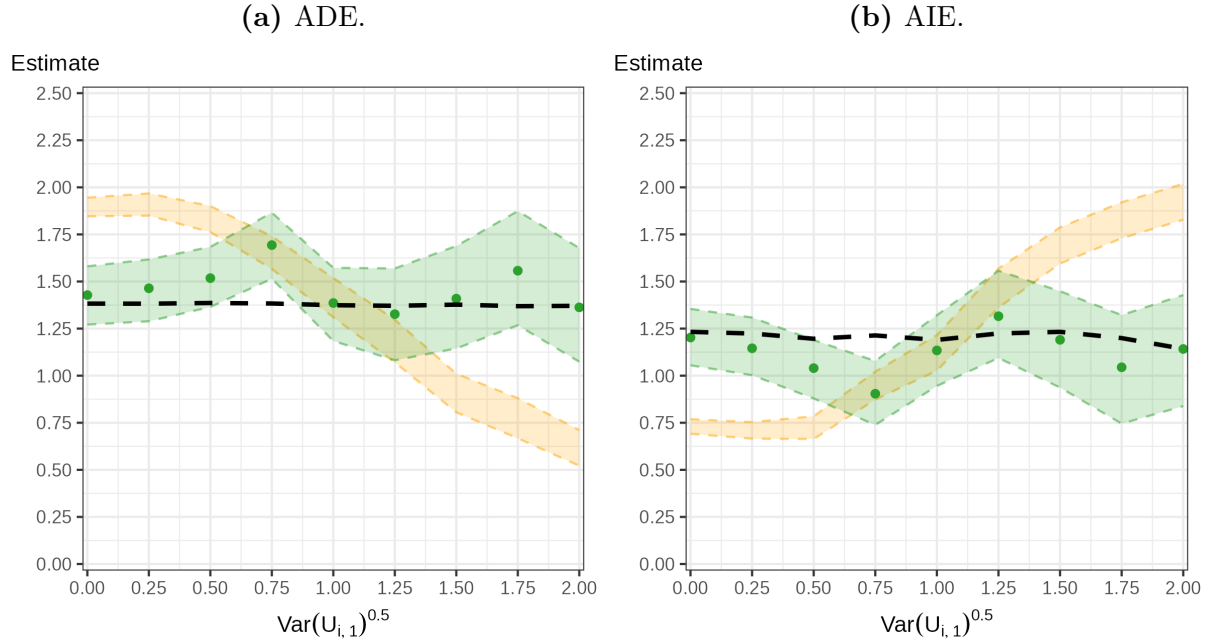
where  $\frac{1}{N} \sum_{i=1}^N \hat{\delta} \hat{\pi}(0; \mathbf{X}_i)$  estimates  $\mathbb{E} [\delta D_i(0)]$ , and  $\frac{1}{N} \sum_{i=1}^N \hat{\delta} \hat{\pi}(1; \mathbf{X}_i)$  estimates  $\mathbb{E} [\delta D_i(1)]$ . Everything involved is a standard point estimate, so their composition will converge to a normal distribution, too. Standard error computation can be achieved by a bootstrap procedure.

## A.9 Implementation and Further Simulation Evidence

A number of statistical packages, for the R language ([R Core Team 2025](#)), made the simulation analysis for this paper possible.

- *Tidyverse* ([Wickham, Averick, Bryan, Chang, McGowan, François, Golemund, Hayes, Henry, Hester, Kuhn, Pedersen, Miller, Bache, Müller, Ooms, Robinson, Seidel, Spinu, Takahashi, Vaughan, Wilke, Woo & Yutani 2019](#)) collected tools for data analysis in the R language.
- *Mgcv* ([Wood, N., Pya & S’afken 2016](#)) allows semi-parametric estimation, using splines, in the R language.
- *Mediate* ([Tingley, Yamamoto, Hirose, Keele & Imai 2014](#)) automates the sequential-ignorability estimates of CM effects ([Imai et al. 2010](#)) in the R language.

**Figure A1:** OLS versus CF Estimates of CM Effects, varying  $\text{Var}(U_{1,i})$  relative to  $\text{Var}(U_{0,i}) = 1$ .



**Note:** These figures show the OLS and control function estimates of the ADE and AIE, for  $N = 1,000$  sample size. The black dashed line is the true value, points are points estimates from data simulated with a given  $\text{Corr}(U_{0,i}, U_{1,i}) = 0.5$ ,  $\text{Var}(U_{0,i}) = 1$ , and  $\text{Var}(U_{1,i})^{\frac{1}{2}}$  varied across  $[0, 2]$ . Shaded regions are the 95% confidence intervals; orange are the OLS estimates, blue the control function approach.

## **A.10 Further Information on the Oregon Health Insurance Experiment**

Brief description on why supplying kappa weights to my methods gets what you want, among lottery compliers. Note, this is the same approach used by [Frölich & Huber \(2017\)](#) to re-weight to treatment compliers in their identification of CM effects with two instruments.