

Supplement to “Tolerating defiance? Local average treatment effects without monotonicity”

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In this supplement, I show that one can estimate quantile treatment effects in the surviving-compliers subpopulation, that one can test the CD condition, and that my results extend to multivariate treatment and instrument, to fuzzy regression discontinuity designs, and to survey nonresponse.

S1. TESTABILITY AND QUANTILE TREATMENT EFFECTS

In this section, I show that random instrument, exclusion restriction, and CD have a testable implication. Then I discuss how the testable implications of random instrument, exclusion restriction, and ND studied in [Huber and Mellace \(2012\)](#), [Kitagawa \(2015\)](#), and [Mourifie and Wan \(forthcoming\)](#) relate to the CD condition.

For any random variable R , let $\mathcal{S}(R)$ denote the support of R . Let U_{dz} be a random variable uniformly distributed on $[0, 1]$ denoting the rank of an observation in the distribution of $Y|D = d, Z = z$. If the distribution of $Y|D = d, Z = z$ is continuous, $U_{dz} = F_{Y|D=d, Z=z}(Y)$. If the distribution of $Y|D = d, Z = z$ is discrete, one can randomly allocate a rank to tied observations by setting

$$U_{dz} = F_{Y|D=d, Z=z}(Y^-) + V(F_{Y|D=d, Z=z}(Y) - F_{Y|D=d, Z=z}(Y^-)),$$

where V is uniformly distributed on $[0, 1]$ and independent of (Y, D, Z) , and $Y^- = \sup\{y \in \mathcal{S}(Y|D = d, Z = z) : y < Y\}$.¹ For every $d \in \{0, 1\}$, let $p_d = \frac{\text{FS}}{P(D=d|Z=d)}$. Notice that both p_0 and p_1 are included between 0 and 1. Finally, let

$$\underline{L} = E(Y|D = 1, Z = 1, U_{11} \leq p_1) - E(Y|D = 0, Z = 0, U_{00} \geq 1 - p_0),$$

$$\overline{L} = E(Y|D = 1, Z = 1, U_{11} \geq 1 - p_1) - E(Y|D = 0, Z = 0, U_{00} \leq p_0).$$

THEOREM S1. *If Assumptions 1, 2, and 5 are satisfied,*

$$\underline{L} \leq W \leq \overline{L}. \tag{29}$$

The intuition of this result is as follows. Under random instrument, exclusion restriction, and CD, $E(Y_1 - Y_0|C_V)$ is point identified: following Theorem 2.1 in the main paper,

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¹The term Y^- should be understood as $-\infty$ if $Y = \inf\{\mathcal{S}(Y|D = d, Z = z)\}$.

it is equal to W . It is also partially identified. The term C_V is included in $\{D = 0, Z = 0\}$, and it accounts for $p_0\%$ of this population. Therefore, $E(Y_0|C_V)$ cannot be larger than the mean of Y_0 of the $p_0\%$ of this population with the largest Y_0 . It also cannot be smaller than the mean of Y_0 of the $p_0\%$ with the lowest Y_0 (see Horowitz and Manski (1995) and Lee (2009)). Combining this with a similar reasoning for $E(Y_1|C_V)$ yields worst-case bounds for $E(Y_1 - Y_0|C_V)$: \underline{L} and \bar{L} . The estimand of $E(Y_1 - Y_0|C_V)$ should lie within its bounds, hence the testable implication.

To implement this test, one can use results from Andrews and Soares (2010). Let θ satisfy the two moment inequality conditions

$$\begin{aligned} 0 &\leq E\left(\frac{YZ}{P(Z=1)} - \frac{Y(1-Z)}{P(Z=0)} - \text{FS}\left(\theta + \frac{YDZ1\{U_{11} \leq p_1\}}{P(D=1, Z=1, U_{11} \leq p_1)}\right.\right. \\ &\quad \left.\left. - \frac{Y(1-D)(1-Z)1\{U_{00} \geq 1-p_0\}}{P(D=0, Z=0, U_{00} \geq 1-p_0)}\right)\right), \\ 0 &\leq E\left(\text{FS}\left(\theta + \frac{YDZ1\{U_{11} \geq 1-p_1\}}{P(D=1, Z=1, U_{11} \geq 1-p_1)} - \frac{Y(1-D)(1-Z)1\{U_{00} \leq p_0\}}{P(D=0, Z=0, U_{00} \leq p_0)}\right)\right. \\ &\quad \left. - \left(\frac{YZ}{P(Z=1)} - \frac{Y(1-Z)}{P(Z=0)}\right)\right). \end{aligned}$$

This defines a moment inequality model with preliminary estimated parameters. This model satisfies all the technical conditions required in Andrews and Soares (2010) if $E(|Y|^{2+\delta}) < +\infty$ for some strictly positive δ , and if $P(Z=0)$, $P(Z=1)$, $P(D=1, Z=1, U_{11} \leq p_1)$, $P(D=1, Z=1, U_{11} \geq 1-p_1)$, $P(D=0, Z=0, U_{00} \leq p_0)$, and $P(D=0, Z=0, U_{00} \geq 1-p_0)$ are all bounded from below by some strictly positive constant ε . Under these two mild restrictions, their Theorem 1 applies, and one can use it to construct a uniformly valid confidence interval for θ . Equation (29) is rejected when 0 does not belong to this confidence interval.

As pointed out in Balke and Pearl (1997) and Heckman and Vytlačil (2005), random instrument, exclusion restriction, and ND also have testable implications. Huber and Mellace (2015), Kitagawa (2015), Machado, Shaikh, and Vytlačil (2013), and Mourifie and Wan (forthcoming) have developed statistical tests of these implications. I now discuss how these tests relate to the CD condition.

The test suggested in Huber and Mellace (2015) is not a test of CD. Huber and Mellace (2015) use the fact that $E(Y_0|NT)$ and $E(Y_1|AT)$ are both point and partially identified under random instrument, exclusion restriction, and ND. For instance, $E(Y|D=0, Z=1)$ is equal to $E(Y_0|NT)$ and $E(Y|D=0, Z=0)$ is equal to $(1-p_0)E(Y_0|NT) + p_0E(Y_0|C)$. This implies

$$\begin{aligned} E(Y|D=0, Z=0, U_{00} \leq 1-p_0) &\leq E(Y|D=0, Z=1) \\ &\leq E(Y|D=0, Z=0, U_{00} \geq p_0). \end{aligned}$$

But under CD, $E(Y|D=0, Z=1)$ is equal to $\frac{P(NT)}{P(NT)+P(F)}E(Y_0|NT) + \frac{P(F)}{P(NT)+P(F)}E(Y_0|F)$, so the previous implication need no longer be true.

The tests suggested in [Kitagawa \(2015\)](#), [Machado, Shaikh, and Vytlačil \(2013\)](#), and [Mourifie and Wan \(forthcoming\)](#) are not tests of CD, but they are tests of the CDM condition introduced hereafter.

ASSUMPTION S1 (Compliers–Defiers for Marginals (CDM)). *There is a subpopulation of C denoted C_F that satisfies*

$$P(C_F) = P(F),$$

$$Y_0|C_F \sim Y_0|F,$$

$$Y_1|C_F \sim Y_1|F.$$

The CDM condition requires that a subgroup of compliers have the same size and the same marginal distributions of Y_0 and Y_1 as defiers. CDM is stronger than CD. On the other hand, it is invariant to the scaling of the outcome, while CD is not. With nonbinary outcomes, working under assumptions invariant to scaling might be desirable so one might then prefer to invoke CDM than CD, despite the fact it is a stronger assumption. Following the logic of Theorem 2.2 in the main paper, a sufficient condition for CDM to hold is that there be more compliers than defiers in each subgroup with the same value of (Y_0, Y_1) :

$$P(F|Y_0, Y_1) \leq P(C|Y_0, Y_1).$$

One can show that under CDM, the marginal distributions of Y_0 and Y_1 are identified for surviving-compliers. The estimands are the same as in [Imbens and Rubin \(1997\)](#):

$$f_{Y_0|C_V}(y_0) = \frac{P(D=0|Z=0)f_{Y|D=0,Z=0}(y_0) - P(D=0|Z=1)f_{Y|D=0,Z=1}(y_0)}{P(D=0|Z=0) - P(D=0|Z=1)},$$

$$f_{Y_1|C_V}(y_1) = \frac{P(D=1|Z=1)f_{Y|D=1,Z=1}(y_1) - P(D=1|Z=0)f_{Y|D=1,Z=0}(y_1)}{P(D=1|Z=1) - P(D=1|Z=0)}.$$

The testing procedures in [Kitagawa \(2015\)](#) and [Mourifie and Wan \(forthcoming\)](#) test whether the right-hand sides of the two previous equations are positive everywhere. Under CDM, these quantities are densities so they must be positive. These procedures are therefore tests of CDM. [Machado, Shaikh, and Vytlačil \(2013\)](#) focus on binary outcomes. Their procedure tests inequalities equivalent to those considered in [Kitagawa \(2015\)](#). Therefore, their procedure is also a test of the CDM condition.

S2. MULTIVARIATE TREATMENT AND INSTRUMENT

Results presented in the main paper extend to applications where treatment is multivariate. Assume that for every $z \in \{0, 1\}$, $D_z \in \{0, 1, 2, \dots, J\}$ for some integer J . Let $(Y_j)_{j \in \{0, 1, 2, \dots, J\}}$ denote the corresponding potential outcomes. Let $C^j = \{D_1 \geq j > D_0\}$ denote the subpopulation of compliers induced to go from a treatment below to above j by the instrument. Let $F^j = \{D_0 \geq j > D_1\}$ denote the subpopulation of defiers induced

to go from above to below j by the instrument. As in Angrist and Imbens (1995), I assume that the c.d.f. of $D|Z = 0$ stochastically dominates that of $D|Z = 1$. This implies that $P(C_j) \geq P(F_j)$. Consider the following CD condition.

ASSUMPTION S2 (Compliers–Defiers for Multivariate Treatment (CDMU)). *For every $j \in \{1, 2, \dots, J\}$, there is a subpopulation of C^j , denoted C_F^j , that satisfies*

$$P(C_F^j) = P(F^j),$$

$$E(Y_j - Y_{j-1}|C_F^j) = E(Y_j - Y_{j-1}|F^j).$$

If CDMU is satisfied, one can show that W is equal to the average causal response,

$$\sum_{j=1}^J w_j E(Y_j - Y_{j-1}|C_V^j),$$

where C_V^j is a subset of C_j of size $P(C_j) - P(F_j)$, and w_j are positive real numbers. This generalizes Theorem 1 in Angrist and Imbens (1995).

Their Theorem 2 considers the case where the instrument is multivariate: $Z \in \{0, 1, 2, \dots, K\}$. Let β denote the coefficient of D in a 2SLS regression of Y on D using the set of dummies $(1\{Z = k\})_{1 \leq k \leq K}$ as instruments. Angrist and Imbens (1995) show that if $D_{z'} \geq D_z$ for every $z' \geq z$, β is equal to a weighted average of average causal responses. Let $C^{jz} = \{D_z \geq j > D_{z-1}\}$ (resp. $F^{jz} = \{D_z \geq j > D_{z-1}\}$) denote the subpopulation of compliers induced to go from a treatment below to above j (resp. above to below j) when the instrument moves from $z - 1$ to z . Consider the following condition:

ASSUMPTION S3 (Compliers–Defiers for Multivariate Treatment and Instrument (CDMU2)). *For every $j \in \{1, 2, \dots, J\}$ and $z \in \{0, 1, 2, \dots, K\}$, there is a subpopulation of C^{jz} , denoted C_F^{jz} , that satisfies*

$$P(C_F^{jz}) = P(F^{jz}),$$

$$E(Y_j - Y_{j-1}|C_F^{jz}) = E(Y_j - Y_{j-1}|F^{jz}).$$

Under CDMU2, one can show that β is equal to a weighted average of average causal responses for surviving-compliers.

S3. FUZZY REGRESSION DISCONTINUITY DESIGNS AND SURVEY NONRESPONSE

My results also extend to many treatment effect models relying on ND type conditions. An important example is fuzzy regression discontinuity (RD) designs studied in Hahn, Todd, and Van der Klaauw (2001). Their Theorem 3 still holds if their Assumption A.3, an “ND at the threshold” assumption, is replaced by a “CD at the threshold” assumption. Let S denote the forcing variable and let s denote the threshold. Let $C^s = \{C, S = s\}$ and $F^s = \{F, S = s\}$, respectively, denote compliers and defiers at the threshold.

ASSUMPTION S4 (Compliers–Defiers at the Threshold (CDT)). *There is a subpopulation of C^s , denoted C_F^s , that satisfies*

$$P(C_F^s|S = s) = P(F^s|S = s),$$

$$E(Y_1 - Y_0|C_F^s, S = s) = E(Y_1 - Y_0|F^s, S = s).$$

Finally, ND type conditions have also been used in the sample selection and survey nonresponse literatures, for instance, in Lee (2009) or Behaghel, Crépon, Gurgand, and Le Barbanchon (2015). My results also extend to these applications: the results in Lee (2009) or Behaghel, Crépon, Gurgand, and Le Barbanchon (2015) remain valid under a weakening of the ND condition similar in spirit to the CD condition.

APPENDIX SA: PROOFS

PROOF OF THEOREM S1. Proving that \underline{L} and \bar{L} are valid bounds for $E(Y_1 - Y_0|C_V)$ relies on an argument similar to the proof of Proposition 1.a in Lee (2009). Due to a concern for brevity, I refer the reader to this paper for this part of the proof. The testable implication directly follows. \square

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