# Difference-in-Differences with Multiple Time Periods and an Application on the Minimum Wage and Employment*

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## Abstract

In this article, we consider identification and estimation of treatment effect parameters using Difference-in-Differences (DID) with (i) multiple time periods, (ii) variation in treatment timing, and (iii) when the "parallel trends assumption" holds potentially only after conditioning on observed covariates. We propose a simple two-step estimation strategy, establish the asymptotic properties of the proposed estimators, and prove the validity of a computationally convenient bootstrap procedure to conduct asymptotically valid simultaneous (instead of pointwise) inference. Our proposed inference procedure naturally adjusts for autocorrelation and other forms of clustering in the data. We also propose a semiparametric data-driven testing procedure to assess the credibility of the DID design in our context. Finally, we illustrate the relevance of our proposed tools by analyzing the effect of the minimum wage on teen employment from 2001-2007. Open-source software is available for implementing the proposed methods.

**JEL:** C14, C21, C23, J23, J38.
**Keywords:** Difference-in-Differences, Multiple Periods, Variation in Treatment Timing, Pre-Testing, Minimum Wage.

# 1    Introduction

Difference-in-Differences (DID) has become one of the most popular designs used to evaluate causal effects of policy interventions. In its canonical format, there are two time periods and two groups: in the first period no one is treated, and in the second period some individuals are treated (the treated group), and some individuals are not (the control group). If, in the absence of treatment, the average outcomes for treated and control groups would have followed parallel paths over time (which is the so-called parallel trends assumption), one can estimate the average treatment effect for the treated subpopulation (ATT) by comparing the average change in outcomes experienced by the treated group to the average change in outcomes experienced by the control group. Most methodological extensions of DID methods focus on this standard two periods, two groups setup, see e.g. Heckman et al. (1997, 1998), Abadie (2005), Athey and Imbens (2006), Qin and Zhang (2008), Bonhomme and Sauder (2011), Botosaru and Gutierrez (2017), de Chaisemartin and D'Haultfœuille (2017), and Callaway et al. (2018).

Many DID empirical applications, however, deviate from the standard DID setup and have more than two time periods and variation in treatment timing. In this article we consider identification and inference procedures for average treatment effects in DID models with ($i$) multiple time periods, ($ii$) variation in treatment timing, and ($iii$) when the parallel trends assumption holds potentially only after conditioning on observed covariates. Importantly, our proposal does not rely on functional form restrictions about the potential outcomes, and allows for treatment effect heterogeneity in terms of observed covariates and time.

We develop our approach in several steps. First, we deviate from the "two-way fixed effects" regression model

$$Y_{it} = \alpha_t + c_i + \beta D_{it} + \theta X_i + \epsilon_{it}, \tag{1.1}$$

where $Y_{it}$ is the outcome of interest, $\alpha_t$ is a time fixed effect, $c_i$ is an "individual/group" fixed effect,[1] $D_{it}$ is a treatment indicator that is equal to one if an individual $i$ is treated at time $t$ and zero otherwise, $X_i$ is a vector of observed characteristics, and $\epsilon_{it}$ is an error term. Therefore,

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[1] Group fixed effects are defined at a different level of aggregation than the covariates $X_i$; otherwise, one cannot separately identity the effect of $c_i$ and $X_i$ on $Y_{it}$.

our proposal differs from Wooldridge (2005), Chernozhukov et al. (2013), Borusyak and Jaravel (2017), Słoczyński (2017), de Chaisemartin and D'Haultfœuille (2018), Goodman-Bacon (2018), Abraham and Sun (2018), and Athey and Imbens (2018). Among other things, these papers consider variations of the regression (1.1) and focus on understanding what causal parameter the $\beta$ coefficient represents under treatment effect heterogeneity. Instead of following this path, we proceed by first defining an easy to interpret parameter of interest, and then provide conditions under which such a parameter is nonparametrically identified. Here, the causal parameters of interest are (functionals of) the average treatment effect for group $g$ at time $t$, where a "group" is defined by when units are first treated. We call these causal parameters *group-time average treatment effects*; in the canonical DID setup, they collapse to the ATT.

The deviation from the two-way fixed effects regression model (1.1) has several advantages. First, our parameter of interest is not determined by the estimation method. On the other hand, the $\beta$ coefficient in (1.1) usually represents a weighted sum of different average treatment effects and these weights are partly determined by the estimation procedure one adopts (usually least squares), see e.g. Goodman-Bacon (2018). Perhaps even more importantly, in settings with treatment effect heterogeneity and variation in treatment timing, these weights can be negative, see e.g. de Chaisemartin and D'Haultfœuille (2018) for a discussion on how the negative weights arise. Hence, inference about the effectiveness of a given policy can be misleading when based on (1.1) unless one further restricts treatment effect heterogeneity. We bypass this difficulty by directly focusing on the causal parameter of interest.

In a second contribution, we note that in some applications there may be "too many" group-time average treatment effects, and researchers may be interested in summarizing them into more aggregated causal parameters. Although there are many ways of combining these treatment effects, we consider aggregation schemes motivated by economic theory that can highlight some particular features of the data. In particular, we propose aggregation procedures depending on whether one is concerned with (*a*) selective treatment timing, i.e., allowing, for example, the possibility that individuals with the largest benefits from participating in a treatment choose to become treated earlier than those with a smaller benefit; (*b*) dynamic treatment effects – where the effect of a

treatment can depend on the length of exposure to the treatment; or $(c)$ calendar time effects – where the effect of treatment may depend on the time period. Overall, like the $\beta$ coefficient in (1.1), our proposed aggregation procedures result in a single estimated "effect" of treatment; however, in contrast with $\beta$, our aggregated parameters have a clear causal interpretation.

A third contribution of this article is to propose estimators and provide asymptotically valid inference procedures for the causal parameters of interest. In the same spirit of Abadie (2005), we consider inverse probability weighted estimators for the treatment effects. We extend Abadie (2005) estimators in two directions. First and most importantly, we account for variation in treatment timing. Second, our proposed estimators are of the Hájek (1971) type, whereas the Abadie (2005) estimator is of the Horvitz and Thompson (1952) type. In other words, our proposed weights are normalized to one, while the weights in Abadie (2005) are not. As discussed in Robins et al. (2007), Hájek-type estimators are sample bounded - i.e., the estimates are enforced to lie within the range of the observed outcomes - whereas Horvitz-Thompson estimators are not. In practice, this modification usually translates to estimators with improved finite sample properties, see e.g. Busso et al. (2014).

In order to conduct asymptotically valid inference, we justify the use of a computationally convenient multiplier-type bootstrap procedure to obtain simultaneous confidence bands for the group-time average treatment effects. Unlike commonly used pointwise confidence bands, our simultaneous confidence bands asymptotically cover the *entire path* of the group-time average treatment effects with probability $1 - \alpha$, and take into account the dependency across different group-time average treatment effects estimators. Thus, our proposed confidence bands are arguably more suitable for visualizing the overall estimation uncertainty than more traditional pointwise confidence intervals. It is also worth mentioning that all our inference procedures naturally account for autocorrelation and can accommodate other types of clustering in a relatively straightforward manner.

Finally, it is worth mentioning that the reliability of the causal interpretation of all aforementioned results relies on the validity of the conditional parallel trends assumption. Unfortunately, such an assumption is fundamentally untestable. On the other hand, we show that provided the

availability of more than two time periods, if one is willing to assume that the conditional parallel trends assumption holds not only in post-treatment but also in pre-treatment periods, the (augmented) conditional parallel trends assumption has testable implications. A fourth contribution of this article is to take advantage of this observation and propose a falsification test based on it.

Our pre-test for the plausibility of the conditional parallel trends assumption is based on the integrated moments approach, completely avoids selecting tuning parameters, and is fully data-driven. To the best of our knowledge, we are the first to note that we can use the conditional moment restrictions to pre-test for the reliability of the conditional parallel trends identification assumption.

We derive the asymptotic null distribution of our falsification test statistic, prove that it is consistent against fixed nonparametric alternatives, and show that critical values can be computed with the assistance of an easy to implement multiplier-type bootstrap. These results build on many papers in the goodness-of-fit literature – see e.g. Bierens (1982), Bierens and Ploberger (1997), Stute (1997), and Escanciano (2006b, 2008); for a recent overview, see González-Manteiga and Crujeiras (2013). However, in contrast with most specification testing proposals, our null hypothesis involves multiple conditional moment restrictions instead of a single conditional moment restriction; see Escanciano (2008) for an exception.

We illustrate the appeal of our method by revisiting findings about the effect of the minimum wage on teen employment. Although classical economic models suggest that a wage floor should result in lower employment, there is a bulk of literature that finds no disemployment effects of the minimum wage – see e.g. Card and Krueger (1994), Dube et al. (2010), among many others. However, another strand of the literature argues that raising the minimum wage leads to lower employment – see e.g. Neumark and Wascher (1992, 2000, 2008), Neumark et al. (2014), and Jardim et al. (2017).

We use data from 2001-2007, where the federal minimum wage was flat at \$5.15 per hour. Using a period where the federal minimum wage is flat allows for a clear source of identification – state level changes in minimum wage policy. However, we also need to confront the issue that states changed their minimum wage policy at different points in time over this period – an issue

not encountered in the case study approach to studying the employment effects of the minimum wage. In addition, for states that changed their minimum wage policy in later periods, we can pre-test the parallel trends assumption which serves as a check of the internal consistency of the models used to identify minimum wage effects.

We consider both an unconditional and conditional DID approach to estimating the effect of increasing the minimum wage on teen employment rates. For the unconditional DID, we find that increases in the minimum wage tend to decrease teen employment, which is in line with most of the work on the minimum wage that uses a similar setup. As Dube et al. (2010) points out, such negative effects may be spurious given potential violations of the common trend assumption. Indeed, when we test for the reliability of the unconditional common trends assumption, we reject it at the usual significance levels. Next, we focus on conditional DID. First, we follow an approach suggested in Dube et al. (2010) and consider a two-way fixed effects regression model with region-year fixed effects. As in Dube et al. (2010), such an estimation strategy finds no adverse effect on teen employment. Nonetheless, one must bear in mind that, as discussed before, such a two-way fixed effects regression may not identify an easy to interpret causal parameter. To circumvent this issue, we use our conditional DID approach and find that increasing the minimum wage does tend to decrease teen employment, though the magnitude of the effects is slightly smaller than in the unconditional case. The contrast of the findings based on two-way fixed effects regression and our proposed method highlights the importance of taking treatment effect heterogeneity into account. On the other hand, when we apply our pre-test for the reliability of the conditional DID setup, we do find evidence against the conditional parallel trends assumption. Thus, one should interpret the findings with care.

The remainder of this article is organized as follows. Section 2 presents our main identification results. We discuss estimation and inference procedures for the treatment effects of interest in Section 3. Section 4 describes our pre-tests for the credibility of the conditional parallel trends assumption. We revisit the effect of minimum wage on employment in Section 5. Section 6 concludes. Proofs as well as additional methodological results are reported in the Supplementary Appendix.

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# 2 Identification

## 2.1 Framework

We first introduce the notation we use throughout the article. We consider the case with $\mathcal{T}$ periods and denote a particular time period by $t$ where $t = 1, \ldots, \mathcal{T}$. In a standard DID setup, $\mathcal{T} = 2$ and no one is treated in period 1. Let $D_t$ be a binary variable equal to one if an individual is treated in period $t$ and equal to zero otherwise. Also, define $G_g$ to be a binary variable that is equal to one if an individual is first treated in period $g$, and define $C$ as a binary variable that is equal to one for individuals in the control group – these are individuals who are never treated so the notation is not indexed by time. For each individual, exactly one of the $G_g$ or $C$ is equal to one. Denote the generalized propensity score as $p_g(X) = P(G_g = 1 | X, G_g + C = 1)$. Note that $p_g(X)$ indicates the probability that an individual is treated conditional on having covariates $X$ and conditional on being a member of group $g$ or the control group. Finally, let $Y_t(1)$ and $Y_t(0)$ be the potential outcome at time $t$ with and without treatment, respectively. The observed outcome in each period can be expressed as $Y_t = D_t Y_t(1) + (1 - D_t) Y_t(0)$.

Given that $Y_t(1)$ and $Y_t(0)$ cannot be observed for the same individual at the same time, researchers often focus on estimating some function of the potential outcomes. For instance, in the standard DID setup, the most popular treatment effect parameter is the average treatment effect on the treated, denoted by[2]

$$ATT = \mathbb{E}[Y_2(1) - Y_2(0) | G_2 = 1].$$

Unlike the two period and two group case, when there are more than two periods and variation in treatment timing, it is not obvious what the main causal parameter of interest should be. We focus on the average treatment effect for individuals first treated in period $g$ at time period $t$, denoted by

$$ATT(g, t) = \mathbb{E}[Y_t(1) - Y_t(0) | G_g = 1].$$

We call this causal parameter the *group-time average treatment effect*. In particular, note that in

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[2]Existence of expectations is assumed throughout.

the classical DID setup, $ATT(2,2)$ collapses to $ATT$.

In this article, we are interested in identifying and making inference about $ATT(g,t)$ and functions of $ATT(g,t)$. Towards this end, we impose the following assumptions.

**Assumption 1** (Sampling). $\{Y_{i1}, Y_{i2}, \ldots Y_{i\mathcal{T}}, X_i, D_{i1}, D_{i2}, \ldots, D_{i\mathcal{T}}\}_{i=1}^n$ *is independent and identically distributed (iid).*

**Assumption 2** (Conditional Parallel Trends). *For all* $t = 2, \ldots, \mathcal{T}$, $g = 2, \ldots, \mathcal{T}$ *such that* $g \leq t$,

$$\mathbb{E}[Y_t(0) - Y_{t-1}(0)|X, G_g = 1] = \mathbb{E}[Y_t(0) - Y_{t-1}(0)|X, C = 1] \ a.s..$$

**Assumption 3** (Irreversibility of Treatment). *For* $t = 2, \ldots, \mathcal{T}$,

$$D_{t-1} = 1 \ implies \ that \ D_t = 1$$

**Assumption 4** (Overlap). *For all* $g = 2, \ldots, \mathcal{T}$, $P(G_g = 1) > 0$ *and for some* $\varepsilon > 0$, $p_g(X) < 1 - \varepsilon$ *a.s..*

Assumption 1 implies that we are considering the case with panel data. The extension to the case with repeated cross sections is developed in Appendix B in the Supplementary Appendix.

Assumption 2, which we refer to as the (conditional) parallel trends assumption throughout the paper, is the crucial identifying restriction for our DID model, and it generalizes the two-period DID assumption to the case where it holds in multiple periods for each group; see e.g. Heckman et al. (1997, 1998), Blundell et al. (2004), and Abadie (2005). It states that, conditional on covariates, the average outcomes for the group first treated in period $g$ and for the control group would have followed parallel paths in the absence of treatment. We require this assumption to hold for all groups $g$ and all time periods $t$ such that $g \leq t$; that is, it holds in all periods after group $g$ is first treated. It is important to emphasize that the parallel trends assumption holds only after conditioning on some covariates $X$, therefore allowing for $X$-specific time trends. All of our analysis continues to go through in the case where an unconditional parallel trends assumption holds by simply setting $X = 1$.

Assumption 3 states that once an individual becomes treated, that individual will also be treated in the next period. With regards to the minimum wage application, Assumption 3 says

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that once a state increases its minimum wage above the federal level, it does not decrease it back to the federal level during the analyzed period. Moreover, this assumption is consistent with most DID setups that exploit the enacting of a policy in some location while the policy is not enacted in another location.

Finally, Assumption 4 states that a positive fraction of the population started to be treated in period $g$, and that, for any possible value of the covariates $X$, there is some positive probability that an individual is not treated.[3] This is a standard covariate overlap condition, see e.g. Heckman et al. (1997, 1998), Blundell et al. (2004), Abadie (2005).

**Remark 1.** *In some applications, eventually all units are treated, implying that $C$ is never equal to one. In such cases one can consider the "not yet treated" $(D_t = 0)$ as a control group instead of the "never treated" $(C = 1)$. We consider this case in Appendix C in the Supplementary Appendix. See also Borusyak and Jaravel (2017) and Abraham and Sun (2018).*

**Remark 2.** *Sometimes, researchers may be interested in different parameters of interests such as the "optimal treatment regime", i.e., the sequence of treatments that maximizes the average potential outcome, see e.g. Han (2018). In order to identify such parameters, one needs to impose alternative identifying assumptions and rely on the availability of instrumental variables. As we show below, by focusing on $ATT(g, t)$, we do not require instruments.*

## 2.2   Group-Time Average Treatment Effects

In this section, we introduce the nonparametric identification strategy for the group-time average treatment effect $ATT(g, t)$. Importantly, we allow for treatment effect heterogeneity, and do not make functional form assumptions about the evolution of potential outcomes.

**Theorem 1.** *Under Assumptions 1 - 4 and for $2 \leq g \leq t \leq \mathcal{T}$, the group-time average treatment*

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[3]In our application on the minimum wage, we must take somewhat more care here as there are some periods where there are no states that increase their minimum wage. In this case, let $\mathcal{G}$ denote the set of first treatment times with $\mathcal{G} \subseteq \{1, \ldots, \mathcal{T}\}$. Then, one can compute $ATT(g, t)$ for groups $g \in \mathcal{G}$ with $g \leq t$. This is a simple complication to deal with in practice, so we consider the notationally more convenient case where there are some individuals treated in all periods (possibly excluding period 1) in the main text of the paper.

*effect for group g in period t is nonparametrically identified, and given by*

$$ATT\left(g,t\right) = \mathbb{E}\left[\left(\frac{G_g}{\mathbb{E}\left[G_g\right]} - \frac{\dfrac{p_g\left(X\right)C}{1 - p_g\left(X\right)}}{\mathbb{E}\left[\dfrac{p_g\left(X\right)C}{1 - p_g\left(X\right)}\right]}\right)\left(Y_t - Y_{g-1}\right)\right]. \tag{2.1}$$

Theorem 1 says that, under Assumptions 1 - 4, a simple weighted average of "long differences" of the outcome variable recovers the group-time average treatment effect. The weights depend on the generalized propensity score $p_g\left(X\right)$, and are normalized to one. The intuition for the weights is simple. One takes observations from the control group and group $g$, omitting other groups and then weights up observations from the control group that have characteristics similar to those frequently found in group $g$ and weights down observations from the control group that have characteristics that are rarely found in group $g$. Such a reweighting procedure guarantees that the covariates of group $g$ and the control group are balanced. Interestingly, in the standard DID setup of two periods only, $\mathbb{E}\left[p_2\left(X\right)C/\left(1 - p_2\left(X\right)\right)\right] = \mathbb{E}\left[G_2\right]$, and the results of Theorem 1 reduce to Lemma 3.1 in Abadie (2005).

To shed light on the role of the "long difference", we give a sketch of how this argument works in the unconditional case, i.e., when $X = 1$. Recall that the key identification challenge is for $\mathbb{E}[Y_t(0)|G_g = 1]$ which is not observed when $g \leq t$. Under the parallel trends assumption,

$$\mathbb{E}[Y_t(0)|G_g = 1] = \mathbb{E}[Y_t(0) - Y_{t-1}(0)|G_g = 1] + \mathbb{E}[Y_{t-1}(0)|G_g = 1]$$

$$= \mathbb{E}[Y_t - Y_{t-1}|C = 1] + \mathbb{E}[Y_{t-1}(0)|G_g = 1]$$

The first term is identified, it is the change in outcomes between $t - 1$ and $t$ experienced by the control group. If $g > t - 1$, then the last term is identified. If not,

$$\mathbb{E}[Y_{t-1}(0)|G_g = 1] = \mathbb{E}[Y_{t-1} - Y_{t-2}|C = 1] + \mathbb{E}[Y_{t-2}(0)|G_g = 1]$$

which holds under the parallel trends assumption. If $g > t - 2$, then every term above is identified. If not, one can proceed recursively in this same fashion until

$$\mathbb{E}[Y_t(0)|G_g = 1] = \mathbb{E}[Y_t - Y_{g-1}|C = 1] + \mathbb{E}[Y_{g-1}|G_g = 1],$$

implying the result for $ATT(g,t)$.

One final thing to consider in this section is the case when the parallel trends assumption holds without needing to condition on covariates. In this case, (2.1) simplifies to

$$ATT(g,t) = \mathbb{E}[Y_t - Y_{g-1}|G_g = 1] - \mathbb{E}[Y_t - Y_{g-1}|C = 1], \tag{2.2}$$

which is simpler than the weighted representation in (2.1) but also implies that all of our results will also cover the unconditional case which is very commonly used in empirical work. We discuss an alternative regression based approach to obtaining $ATT(g,t)$ in Appendix D in the Supplementary Appendix.[4]

## 2.3 Summarizing Group-time Average Treatment Effects

The previous section shows that the group-time average treatment effect $ATT(g,t)$ is identified for $g \leq t$. These are very useful parameters – they allow one to consider how the effect of treatment varies by group and time. However, in some applications there may be many of them, perhaps too many to easily interpret the effect of a given policy intervention. This section considers ways to aggregate group-time average treatment effects into fewer interpretable causal effect parameters. In applications, aggregating the group-time average treatment effects is also likely to increase statistical power, reducing estimation uncertainty.

The two simplest ways of combining $ATT(g,t)$ across $g$ and $t$ are

$$\frac{2}{\mathcal{T}(\mathcal{T}-1)} \sum_{g=2}^{\mathcal{T}} \sum_{t=2}^{\mathcal{T}} \mathbf{1}\{g \leq t\} ATT(g,t) \quad \text{and} \quad \frac{1}{\kappa} \sum_{g=2}^{\mathcal{T}} \sum_{t=2}^{\mathcal{T}} \mathbf{1}\{g \leq t\} ATT(g,t) P(G=g) \tag{2.3}$$

where $\kappa = \sum_{g=2}^{\mathcal{T}} \sum_{t=2}^{\mathcal{T}} \mathbf{1}\{g \leq t\} P(G = g)$ (which ensures that the weights on $ATT(g,t)$ in the second term sum to 1).[5] The first term in (2.3) is just the simple average of $ATT(g,t)$; the second is a weighted average of each $ATT(g,t)$ putting more weight on $ATT(g,t)$ with larger group sizes.

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[4]Unlike the two period, two group case, there does not appear to be any advantage to trying to obtain $ATT(g,t)$ from a regression as it appears to require post-processing the regression output.

[5]Here we use the shorthand notation $P(G = g)$ to denote $P(G_g = 1|G_1 + C = 0)$ . Thus, $P(G = g)$ is the probability that an individual is first treated in period $g$ conditional on not being in the control group or in the group first treated in period 1. Throughout this section, conditional probabilities such as $P(G = g|g \leq t)$ also implicitly condition on not being in the control group or in the group first treated in period 1.

As we argue below, neither of the terms in (2.3) are likely to be "appropriate" summary treatment effect measures, except in the particular case where the effect of treatment is homogeneous across groups and time.

In contrast to our approach in this section, the most common approach to estimating the effect of a binary treatment in a panel data setup is to interpret $\beta$ in the following regression as the average treatment effect

$$Y_{it} = \alpha_t + c_i + \beta D_{it} + \theta X_i + \epsilon_{it},$$

where $\alpha_t$ is a time fixed effect and $c_i$ is an individual/group fixed effect. Interestingly, Wooldridge (2005), Chernozhukov et al. (2013), Borusyak and Jaravel (2017), Goodman-Bacon (2018), Słoczyński (2017), de Chaisemartin and D'Haultfœuille (2018), Abraham and Sun (2018) and Athey and Imbens (2018) have shown that, in general, $\beta$ does not represent an easy to interpret average treatment effect parameter. The results in this section can be used in exactly the same setup to identify a single interpretable average treatment effect parameter and, thus, provide a way to circumvent the issues with the more common approach.

In the following, we consider several common cases that are likely to occur in practice: ($a$) selective treatment timing, ($b$) dynamic treatment effects, and ($c$) calendar time effects. We provide some recommendations on constructing interpretable treatment effect parameters under each of these setups. It is worth mentioning that in each of these cases, $ATT(g, t)$ still provides the average causal effect of the treatment for group $g$ in period $t$; the issue in this section is how to aggregate $ATT(g, t)$ into a smaller number of causal effect parameters.

**Selective Treatment Timing**  In many cases, when to become treated is a choice variable. The parallel trends assumption does place some restrictions on how individuals select when to be treated. In particular, in order for the path of untreated potential outcomes to be the same for a particular group and the control group, the parallel trends assumption does not permit individuals to select into treatment in period $t$ because they anticipate $Y_t(0)$ being small (assuming larger $Y$ is "good"). On the other hand, the parallel trends assumption does not place restrictions on how treated potential outcomes are generated. Thus, our imposed DID assumptions fully allow for

individuals to select into treatment on the basis of expected future values of treated potential outcomes.

While some forms of selective treatment timing are permitted under the parallel trends assumption and do not affect identification of group-time average treatment effects, they do have implications for the "best ways" to combine $ATT(g,t)$ into a single, easy to interpret treatment effect parameter. In particular, when there is selective treatment timing, the period when an individual is first treated may provide information about the size of the treatment effect. In such cases, we propose to summarize the causal effect of a policy by first aggregating $ATT(g,t)$ by group, and then combine group average treatment effects based on the size of each group.

More precisely, we first consider

$$\tilde{\theta}_S(g) = \frac{1}{\mathcal{T} - g + 1} \sum_{t=2}^{\mathcal{T}} \mathbf{1}\{g \leq t\} ATT(g,t).$$

Note that $\tilde{\theta}_S(g)$ is the time-averaged treatment effect for individuals in group $g$, i.e., just a time-average of each available $ATT(g,t)$ for group $g$. Next, in order to further reduce the dimensionality of $\tilde{\theta}_S(g)$, one can average $\tilde{\theta}_S(g)$ across groups to get

$$\theta_S = \sum_{g=2}^{\mathcal{T}} \tilde{\theta}_S(g) P(G = g). \tag{2.4}$$

Note that $\theta_S$ appears to be quite similar to the second term in (2.3). The difference is in the weights. The second term in (2.3) puts more weight on groups that are exposed to treatment longer. The weights in (2.4) only depend on group size, not on the number of post-treatment periods available per group. For example, suppose there is positive selective treatment timing so that individuals who are treated earlier experience larger benefits from being treated than those who are treated later. In the presence of selective treatment timing, the approach in (2.3) would tend to overstate the effect of the treatment due to putting more weight on the groups that are treated the longest, which are precisely the ones that experience the largest benefits of being treated. Thus, we argue that, in the presence of selective treatment timing, $\theta_S$ in (2.4) is a more natural causal parameter than the second term in (2.3).

**Dynamic Treatment Effects**   In other cases, the effect of a policy intervention may depend on the length of exposure to it. To give some examples, Jacobson et al. (1993) argues that workers that are displaced from their jobs tend to have immediate large earnings effects that get smaller over time, and both the immediate effect and the dynamic effect are of interest. In the case of the minimum wage, Meer and West (2016) argue that increasing the minimum wage leads to lower job creation and thus that the effect of the minimum wage on employment is dynamic – one should expect larger effects in subsequent periods than in the initial period.

In the presence of dynamic treatment effects (but not selective treatment timing), we propose to summarize the effects of the policy by first aggregating $ATT(g,t)$ by the length of exposure to treatment (we denote this by $e$), and then (possibly) combining average effects based on length of exposure by averaging over different lengths of exposure. That is, we first consider the parameter

$$\tilde{\theta}_D(e) = \sum_{g=2}^{\mathcal{T}} \sum_{t=2}^{\mathcal{T}} \mathbf{1}\{t - g + 1 = e\} ATT(g,t) P(G = g | t - g + 1 = e),$$

which provides the average effect of treatment for individuals that have been treated for exactly $e$ periods. For example, when $e = 1$, it averages (based on group size) $ATT(g,t)$ for $g = t$ (groups that have been exposed to treatment for exactly one period). Averaging over all possible values of $e$ results in the parameter

$$\theta_D = \frac{1}{\mathcal{T} - 1} \sum_{e=1}^{\mathcal{T}-1} \tilde{\theta}_D(e). \tag{2.5}$$

The primary difference between $\theta_D$, $\theta_S$, and the second term in (2.3) is the weights. Relative to the other parameters, $\theta_D$ puts the most weight on $ATT(g,t)$ when $g$ is much less than $t$, which corresponds to large values of $e$, because there are few groups available for large values of $e$. In the absence of selective treatment timing, these groups are informative about the dynamic effects of treatment for all groups. Hence, we argue that $\theta_D$ is appealing when treatment effects evolve over time.

**Calendar Time Effects**   In other cases, calendar time may matter. For example, graduating during a recession may have a large effect on future earnings, see e.g. Oreopoulos et al. (2012). The case with calendar time effects is similar to the case with dynamic treatment effects. Our

proposed summary treatment effect parameter involves first computing an average treatment effect for all individuals that are treated in period $t$, and then averaging across all periods. Consider the parameter

$$\tilde{\theta}_C(t) = \sum_{g=2}^{\mathcal{T}} \mathbf{1}\{g \le t\} ATT(g,t) P(G = g | g \le t).$$

Here, $\tilde{\theta}_C(t)$ can be interpreted as the average treatment effect in period $t$ for all groups that are treated by period $t$. With $\tilde{\theta}_C(t)$ at hand, one can compute

$$\theta_C = \frac{1}{\mathcal{T} - 1} \sum_{t=2}^{\mathcal{T}} \tilde{\theta}_C(t),$$

which can be interpreted as the average treatment effect when calendar time matters. When calendar time matters, the most weight is put on groups that are treated in the earliest periods. This is because there are fewer groups available to estimate the average treatment effect in period $t$ when $t$ is small relative to the number of groups available to estimate the average treatment effect in period $t$ when $t$ is large.

**Selective Treatment Timing and Dynamic Treatment Effects**  Finally, we consider the case where the timing of treatment is selected and there are dynamic treatment effects. This might very well be the most relevant case in studying the effect of increasing the minimum wage as ($i$) states are not likely to raise their minimum wage during a recession and ($ii$) the effect of the minimum wage takes some time to play out; see e.g. Meer and West (2016).

The fundamental problem with using the dynamic treatment effects approach when there is selective treatment timing is that the composition of the treated group changes when the length of exposure to treatment ($e$) changes. Without selective treatment timing, this does not matter because when an individual first becomes treated does not affect their outcomes. However, with selective treatment timing, changing the composition of the treatment group can have a big effect (See Figure 1 for an example where the dynamic treatment effect is declining with length of exposure to treatment for all groups but ignoring selective treatment timing leads to the opposite (wrong) conclusion – that the effect of treatment is increasing over time.).

To circumvent such an issue, we consider dynamic treatment effects only for $e \le e'$ and for

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Figure 1: Example of Selective Treatment Timing and Dynamic Treatment Effects
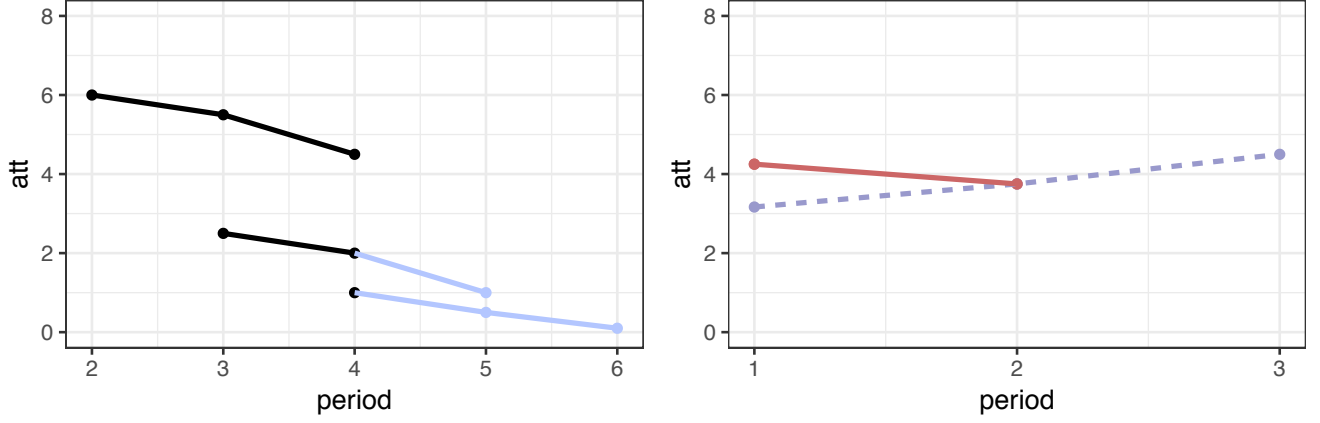
*Notes:* In this example, there are three groups: G2 (first treated in period 2), G3 (first treated in period 3), and G4 (first treated in period 4). Suppose that the last period available in the sample is period 4; thus, the group-time average treatment effect is available in periods 2 through 4 – these are the dark lines in the left panel of the figure. The light lines in the left panel represent group-time average treatment effects that are not observed. Each group experiences a declining dynamic treatment effect, but there is also selective treatment timing. Groups that are treated earlier experience larger effects of the treatment. The right panel (dashed line) plots the dynamic treatment effect ignoring selective treatment timing and allowing the composition of the treated group to change. In particular, this means that group G4 is only included in the average for period 1, and group G3 only is included in the average for periods 1 and 2. In this case, selective treatment timing leads to exactly the wrong interpretation of the dynamic treatment effect – it appears as though the effect of the treatment is increasing. The solid line plots the dynamic treatment effect as suggested in Equation (2.6) that adjusts for selective treatment timing and for $e = 1, 2$ and $e' = 2$.

groups with at least $e'$ periods of post-treatment data available. This setup removes the effect of selective treatment timing by keeping the same set of groups across all values of $e$. For example, one could consider the dynamic effect of treatment over three periods by averaging $ATT(g, t)$ for all the groups that have at least three periods of post-treatment observations while not utilizing $ATT(g, t)$ for groups that have less than three periods of post-treatment observations. Note that there is some trade-off here. Setting $e'$ small results in many groups satisfying the requirement, but in only being able to study the effect of length of exposure to treatment for relatively few periods. Setting $e'$ to be large decreases the number of available groups but allows one to consider the effect of length of exposure to treatment for relatively more periods.

Next, we describe how this proposed summary causal parameter is constructed. Let $\delta_{gt}(e, e') = 1\{t - g + 1 = e\}1\{T - g + 1 \geq e'\}1\{e \leq e'\}$. Here, $\delta_{gt}(e, e')$ is equal to one in the period where

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group $g$ has been treated for exactly $e$ periods, if group $g$ has at least $e'$ post-treatment periods available, and if the length of exposure $e$ is less than the post-treatment periods requirement $e'$.

Then, the average treatment effect for groups that have been treated for $e$ periods and have at least $e'$ post-treatment periods of data available is given by

$$\tilde{\theta}_{SD}(e, e') = \sum_{g=2}^{\mathcal{T}} \sum_{t=2}^{\mathcal{T}} \delta_{gt}(e, e') ATT(g, t) P(G = g | \delta_{gt}(e, e') = 1) \tag{2.6}$$

which is defined for $e \leq e'$. Effectively, we put zero weight on $ATT(g, t)$ for groups that do not meet the minimum required number of periods in order to prevent the composition of groups from changing. Once $\tilde{\theta}_{SD}(e, e')$ is computed, one can further aggregate it to get

$$\theta_{SD}(e') = \frac{1}{\mathcal{T} - e'} \sum_{e=1}^{\mathcal{T} - e'} \tilde{\theta}_{SD}(e, e')$$

which should be interpreted as the average treatment effect for groups with at least $e'$ periods of post-treatment data allowing for dynamic treatment effects and selective treatment timing. Such a causal parameter has the strengths of both $\theta_S$ and $\theta_D$ in (2.4) and (2.5), respectively.

# 3    Estimation and Inference

In this section, we study estimation and inference procedures for estimators corresponding to the estimands introduced in Section 2. Note that the nonparametric identification result in Theorem 1 suggests a simple two-step strategy to estimate $ATT(g, t)$. In the first step, estimate the generalized propensity score $p_g(x) = P(G_g = 1 | X = x, G_g + C = 1)$ for each group $g$, and compute the fitted values for the sample. In the second step, one plugs the fitted values into the sample analogue of $ATT(g, t)$ in (2.1) to obtain estimates of the group-time average treatment effect.

More concisely, we propose to estimate $ATT(g, t)$ by

$$\widehat{ATT}(g, t) = \mathbb{E}_n \left[ \left( \frac{G_g}{\mathbb{E}_n[G_g]} - \frac{\dfrac{\hat{p}_g(X) C}{1 - \hat{p}_g(X)}}{\mathbb{E}_n \left[ \dfrac{\hat{p}_g(X) C}{1 - \hat{p}_g(X)} \right]} \right) (Y_t - Y_{g-1}) \right],$$

where $\hat{p}_g(\cdot)$ is an estimate of $p_g(\cdot)$, and for a generic $Z$, $\mathbb{E}_n[Z] = n^{-1} \sum_{i=1}^{n} Z_i$. As noted in Theorem

[1](#), $ATT(g,t)$ is nonparametrically identified for $2 \leq g \leq t \leq \mathcal{T}$.

With $\widehat{ATT}(g,t)$ in hand, one can use the analogy principle and combine these to estimate the summarized average treatment effect parameters discussed in Section [2.3](#).

In what follows, we consider the case in which one imposes a parametric restriction on $p_g$ and estimates it by maximum likelihood. This is perhaps the most popular approach adopted by practitioners. Nonetheless, under some additional regularity conditions, our results can be extended to allow nonparametric estimators for the $p_g(\cdot)$, see e.g. Abadie (2005), Chen (2007), Chen et al. (2008), Donald and Hsu (2014) and Sant'Anna (2016, 2017). Finally, we note that when propensity score misspecification is a concern, one can use the data-driven specification tests proposed by Sant'Anna and Song (2018).

**Assumption 5.** *For all $g = 2, \ldots, \mathcal{T}$, $(i)$ there exists a known function $\Lambda : \mathbb{R} \to [0,1]$ such that $p_g(X) = P(G_g = 1 | X, G_g + C = 1) = \Lambda(X'\pi_g^0); (ii) \pi_g^0 \in \mathrm{int}(\Pi)$, where $\Pi$ is a compact subset of $\mathbb{R}^k$; $(iii)$ the support of $X$, $\mathcal{X}$, is a subset of a compact set $S$, and $\mathbb{E}[XX'|G_g + C = 1]$ is positive definite; $(iv)$ let $\mathcal{U} = \{x'\pi : x \in \mathcal{X}, \pi \in \Pi\} ; \forall u \in \mathcal{U}, \exists \varepsilon > 0$ such that $\Lambda(u) \in [\varepsilon, 1 - \varepsilon], \Lambda(u)$ is strictly increasing and twice continuously differentiable with first derivatives bounded away from zero and infinity, and bounded second derivative; $(vi) \mathbb{E}[Y_t^2] < \infty$ for all $t = 1, \ldots, \mathcal{T}$.*

Assumption [5](#) is standard in the literature, see e.g. Section 9.2.2 in Amemiya (1985), Example 5.40 in van der Vaart (1998), or Assumption 4.2 in Abadie (2005), and it allows for Logit and Probit models.

Under Assumption [5](#), $\pi_g^0$ can be estimated by maximum likelihood:

$$\hat{\pi}_g = \arg \max_\pi \sum_{i : G_{ig} + C_i = 1} G_{ig} \ln(\Lambda(X_i'\pi)) + (1 - G_{ig}) \ln(1 - \Lambda(X_i'\pi)).$$

Let $\mathcal{W} = (Y_1, \ldots, Y_{\mathcal{T}}, X, G_1, \ldots, G_{\mathcal{T}}, C)'$, $\hat{p}_g(X_i) = \Lambda(X_i'\hat{\pi}_g)$, $\dot{p}_g = \partial p_g(u)/\partial u$, $\dot{p}_g(X) = \dot{p}_g(X'\pi_g^0)$. Under Assumption [5](#), $\hat{\pi}_g$ is asymptotically linear, that is,

$$\sqrt{n}(\hat{\pi}_g - \pi_g^0) = \frac{1}{\sqrt{n}} \sum_{i=1}^n \xi_g^\pi(\mathcal{W}_i) + o_p(1),$$

where

$$\xi_g^\pi (\mathcal{W}) = \mathbb{E} \left[ \frac{(G_g + C) \, \dot{p}_g (X)^2}{p_g (X) (1 - p_g (X))} X X' \right]^{-1} X \frac{(G_g + C) (G_g - p_g (X)) \, \dot{p}_g (X)}{p_g (X) (1 - p_g (X))}, \tag{3.1}$$

see Lemma A.2 in the Appendix.

## 3.1 Asymptotic Theory for Group-Time Average Treatment Effects

Denote the normalized weights by

$$w_g^G = \frac{G_g}{\mathbb{E}[G_g]}, \qquad w_g^C = \frac{p_g (X) C}{1 - p_g (X)} \bigg/ \mathbb{E}\left[ \frac{p_g (X) C}{1 - p_g (X)} \right], \tag{3.2}$$

and define

$$\psi_{gt}(\mathcal{W}_i) = \psi_{gt}^G(\mathcal{W}_i) - \psi_{gt}^C(\mathcal{W}_i), \tag{3.3}$$

where

$$\psi_{gt}^G(\mathcal{W}) = w_g^G \left[ (Y_t - Y_{g-1}) - \mathbb{E}\left[ w_g^G (Y_t - Y_{g-1}) \right] \right],$$

$$\psi_{gt}^C(\mathcal{W}) = w_g^C \left[ (Y_t - Y_{g-1}) - \mathbb{E}\left[ w_g^C (Y_t - Y_{g-1}) \right] \right] + M_{gt}' \, \xi_g^\pi (\mathcal{W}),$$

and

$$M_{gt} = \frac{\mathbb{E}\left[ X \left( \dfrac{C}{1 - p_g (X)} \right)^2 \dot{p}_g (X) \left[ (Y_{it} - Y_{ig-1}) - \mathbb{E}\left[ w_g^C (Y_t - Y_{g-1}) \right] \right] \right]}{\mathbb{E}\left[ \dfrac{p_g (X) C}{1 - p_g (X)} \right]}$$

which is a $k \times 1$ vector, with $k$ the dimension of $X$, and $\xi_g^\pi (\mathcal{W})$ is as defined in (3.1).

Finally, let $ATT_{g \leq t}$ and $\widehat{ATT}_{g \leq t}$ denote the vector of $ATT(g, t)$ and $\widehat{ATT}(g, t)$, respectively, for all $g = 2, \ldots, \mathcal{T}$ and $t = 2, \ldots, \mathcal{T}$ with $g \leq t$. Analogously, let $\Psi_{g \leq t}$ denote the collection of $\psi_{gt}$ across all periods $t$ and groups $g$ such that $g \leq t$.

The next theorem establishes the joint limiting distribution of $\widehat{ATT}_{g \leq t}$.

**Theorem 2.** *Under Assumptions 1-5, for $2 \leq g \leq t \leq \mathcal{T}$,*

$$\sqrt{n}(\widehat{ATT}(g, t) - ATT(g, t)) = \frac{1}{\sqrt{n}} \sum_{i=1}^n \psi_{gt}(\mathcal{W}_i) + o_p(1).$$

*Furthermore,*

$$\sqrt{n}(\widehat{ATT}_{g\leq t} - ATT_{g\leq t}) \xrightarrow{d} N(0, \Sigma)$$

*where* $\Sigma = \mathbb{E}[\Psi_{g\leq t}(\mathcal{W})\Psi_{g\leq t}(\mathcal{W})']$.

Theorem 2 provides the influence function for estimating the vector of group-time average treatment effects $ATT_{g\leq t}$, as well as its limiting distribution. In order to conduct inference, one can show that the sample analogue of $\Sigma$ is a consistent estimator for $\Sigma$, see e.g. Theorem 4.4 in Abadie (2005) which leads directly to standard errors and pointwise confidence intervals.

Instead of following this route, we propose to use a simple multiplier bootstrap procedure to conduct asymptotically valid inference. Our proposed bootstrap leverages the asymptotic linear representations derived in Theorem 2 and inherits important advantages. First, it is easy to implement and very fast to compute. Each bootstrap iteration simply amounts to "perturbing" the influence function by a random weight $V$, and it does not require re-estimating the propensity score in each bootstrap draw. Second, in each bootstrap iteration, there are always observations from each group. This can be a real problem with the traditional empirical bootstrap where there may be no observations from a particular group in some particular bootstrap iteration. Third, computation of simultaneous (in $g$ and $t$) valid confidence bands is relatively straightforward. This is particularly important, since researchers are likely to use confidence bands to visualize estimation uncertainty about $ATT(g, t)$. Unlike pointwise confidence bands, simultaneous confidences bands do not suffer from multiple-testing problems, and are guaranteed to cover all $ATT(g, t)$ with a probability at least $1 - \alpha$. Finally, we note that our proposed bootstrap procedure can be readily modified to account for clustering, see Remark 3 below.

To proceed, let $\widehat{\Psi}_{g\leq t}(\mathcal{W})$ denote the sample-analogue of $\Psi_{g\leq t}(\mathcal{W})$, where population expectations are replaced by their empirical analogue, and the true generalized propensity score, $p_g$, and its derivatives, $\dot{p}_g$, are replaced by their MLE estimates, $\hat{p}_g$ and $\widehat{\dot{p}}_g$, respectively. Let $\{V_i\}_{i=1}^n$ be a sequence of *iid* random variables with zero mean, unit variance and bounded support, independent of the original sample $\{\mathcal{W}_i\}_{i=1}^n$. A popular example involves *iid* Bernoulli variates $\{V_i\}$ with $P(V = 1 - \kappa) = \kappa/\sqrt{5}$ and $P(V = \kappa) = 1 - \kappa/\sqrt{5}$, where $\kappa = (\sqrt{5} + 1)/2$, as suggested by Mammen (1993).

We define $\widehat{ATT}^*_{g\leq t}$, a bootstrap draw of $\widehat{ATT}_{g\leq t}$, via

$$\widehat{ATT}^*_{g\leq t} = \widehat{ATT}_{g\leq t} + \mathbb{E}_n\left[V \cdot \widehat{\Psi}_{g\leq t}(\mathcal{W})\right]. \tag{3.4}$$

The next theorem establishes the asymptotic validity of the multiplier bootstrap procedure proposed above.

**Theorem 3.** *Under Assumptions 1-5,*

$$\sqrt{n}\left(\widehat{ATT}^*_{g\leq t} - \widehat{ATT}_{g\leq t}\right) \xrightarrow[*]{d} N(0, \Sigma),$$

*where $\Sigma = \mathbb{E}[\Psi_{g\leq t}(\mathcal{W})\Psi_{g\leq t}(\mathcal{W})']$ as in Theorem 2, and $\xrightarrow[*]{d}$ denotes weak convergence (convergence in distribution) of the bootstrap law in probability, i.e. conditional on the original sample $\{\mathcal{W}_i\}_{i=1}^n$. Additionally, for any continuous functional $\Gamma(\cdot)$*

$$\Gamma\left(\sqrt{n}\left(\widehat{ATT}^*_{g\leq t} - \widehat{ATT}_{g\leq t}\right)\right) \xrightarrow[*]{d} \Gamma\left(N(0, \Sigma)\right).$$

We now describe a practical bootstrap algorithm to compute studentized confidence bands that cover $ATT(g,t)$ simultaneously over all $g \leq t$ with a prespecified probability $1 - \alpha$ in large samples. This is similar to the bootstrap procedure used in Kline and Santos (2012), Belloni et al. (2017) and Chernozhukov et al. (2017) in different contexts.

**Algorithm 1.** *1) Draw a realization of $\{V_i\}_{i=1}^n$. 2) Compute $\widehat{ATT}^*_{g\leq t}$ as in (3.4), denote its $(g,t)$-element as $\widehat{ATT}^*(g,t)$, and form a bootstrap draw of its limiting distribution as $\hat{R}^*(g,t) = \sqrt{n}\left(\widehat{ATT}^*(g,t) - \widehat{ATT}(g,t)\right)$. 3) Repeat steps 1-2 B times. 4) Compute a bootstrap estimator of the main diagonal of $\Sigma^{1/2}$ such as the bootstrap interquartile range normalized by the interquartile range of the standard normal distribution, $\widehat{\Sigma}^{1/2}(g,t) = (q_{0.75}(g,t) - q_{0.25}(g,t))/(z_{0.75} - z_{0.25})$, where $q_p(g,t)$ is the pth sample quantile of the $\hat{R}^*(g,t)$ in the B draws, and $z_p$ is the pth quantile of the standard normal distribution. 5) For each bootstrap draw, compute $t-test_{g\leq t} = \max_{(g,t)}\left|\hat{R}^*(g,t)\right|\widehat{\Sigma}(g,t)^{-1/2}$. 5) Construct $\hat{c}_{1-\alpha}$ as the empirical $(1-a)$-quantile of the B bootstrap draws of $t-test_{g\leq t}$. 6) Construct the bootstrapped simultaneous confidence band for $ATT(g,t)$, $g \leq t$, as $\widehat{C}(g,t) = [\widehat{ATT}(g,t) \pm \hat{c}_{1-\alpha}\widehat{\Sigma}(g,t)^{-1/2}/\sqrt{n}]$.*

The next corollary to Theorem 3 states that the simultaneous confidence band for $ATT(g,t)$ described in Algorithm 1 has correct asymptotic coverage.

**Corollary 1.** *Under the Assumptions of Theorem 3, for any $0 < \alpha < 1$, as $n \to \infty$,*

$$P\left(ATT(g,t) \in \widehat{C}(g,t) : g \le t\right) \to 1 - \alpha,$$

*where $\widehat{C}(g,t)$ is as defined in Algorithm 1.*

**Remark 3.** Frequently, in DID applications, one wishes to account for clustering, see e.g. Bertrand et al. (2004). This is straightforward to implement with the multiplier bootstrap described above. In the case of county-level minimum wage data, one could allow for clustering at the state level by drawing a scalar $U_s$ $S$ times – where $S$ is the number of states – and setting $V_i = U_s$ for all observations $i$ in state $s$, see e.g. Sherman and Le Cessie (2007), Kline and Santos (2012), Cheng et al. (2013), and MacKinnon and Webb (2016, 2017). Such a cluster-robust bootstrap procedure will lead to reliable inference provided that the number of clusters is "large". Finally, it is important to emphasize that our proposed multiplier-bootstrap procedure automatically accounts for the autocorrelation of the data.

**Remark 4.** In Algorithm 1 we have required an estimator for the main diagonal of $\Sigma$. However, we note that if one takes $\widehat{\Sigma}(g,t) = 1$ for all $(g,t)$, the result in Corollary 1 continues to hold. However, the resulting "constant width" simultaneous confidence band may be of larger length, see e.g. Montiel Olea and Plagborg-Møller (2017) and Freyberger and Rai (2018).

## 3.2   Asymptotic Theory for Summary Parameters

Let $\theta$ generically represent one of the parameters from Section 2.3, including the ones indexed by some variable (for example, $\tilde{\theta}_S(g)$ or $\tilde{\theta}_{SD}(e,e')$). Notice that all of the parameters in Section 2.3 can be expressed as weighted averages of $ATT(g,t)$. Write this generically as

$$\theta = \sum_{g=2}^{\mathcal{T}} \sum_{t=2}^{\mathcal{T}} w_{gt} ATT(g,t)$$

where $w_{gt}$ are some potentially random weights. $\theta$ can be estimated by

$$\hat{\theta} = \sum_{g=2}^{\mathcal{T}} \sum_{t=2}^{\mathcal{T}} \widehat{w}_{gt} \widehat{ATT}(g,t),$$

where $\widehat{w}_{gt}$ are estimators for $w_{gt}$ such that for all $g, t = 2, \ldots, \mathcal{T}$,

$$\sqrt{n} \left( \widehat{w}_{gt} - w_{gt} \right) = \frac{1}{\sqrt{n}} \sum_{i=1}^{n} \xi_{gt}^{w}(\mathcal{W}_i) + o_p\left(1\right),$$

with $\mathbb{E}\left[\xi_{gt}^{w}(\mathcal{W})\right] = 0$ and $\mathbb{E}\left[\xi_{gt}^{w}(\mathcal{W})\xi_{gt}^{w}(\mathcal{W})'\right]$ finite and positive definite. Estimators based on the sample analogue of the weights discussed in Section 2.3 satisfy this condition.

Let

$$l^{w}\left(\mathcal{W}_i\right) = \sum_{g=2}^{\mathcal{T}} \sum_{t=2}^{\mathcal{T}} w_{gt} \cdot \psi_{gt}(\mathcal{W}_i) + \xi_{gt}^{w}(\mathcal{W}_i) \cdot ATT(g,t),$$

where $\psi_{gt}(\mathcal{W})$ are as defined in (3.3).

The following result follows immediately from Theorem 2, and can be used to conduct asymptotically valid inference for the summary causal parameters $\theta$.

**Corollary 2.** *Under Assumptions 1-5,*

$$\sqrt{n}(\hat{\theta} - \theta) = \frac{1}{\sqrt{n}} \sum_{i=1}^{n} l^{w}\left(\mathcal{W}_i\right) + o_p(1)$$

$$\xrightarrow{d} N\left(0, \mathbb{E}\left[l^{w}\left(\mathcal{W}\right)^2\right]\right)$$

Corollary 2 implies that one can construct standard errors and confidence intervals for summary treatment effect parameters based on a consistent estimator of $\mathbb{E}\left[l^{w}\left(\mathcal{W}\right)^2\right]$ or by using a bootstrap procedure like the one in Algorithm 1.

**Remark 5.** As discussed in Remark 3, the validity of the "cluster-robust" multiplier bootstrap procedure relies on the number of clusters being "large". In some applications such a condition may be more plausible when analyzing the aggregated parameter $\theta$ than when analyzing the $ATT(g,t)$ themselves. When there are only a few available clusters, one should look for alternative procedures, see e.g. Conley and Taber (2011) and Ferman and Pinto (2018).

# 4  Pre-testing the Conditional Parallel Trend Assumption

So far, we have discussed how one can nonparametrically identify, and conduct asymptotically valid inference about causal treatment effect parameters using conditional DID models with multiple periods and variation in treatment timing. The credibility of our results crucially relies on the conditional parallel trends assumption stated in Assumption 2. This assumption is fundamentally untestable. However, when one imposes a stronger version of the conditional parallel trends assumption, that is, that Assumption 2 holds *for all* periods $t$, and not only for the periods $g \leq t$, one can assess the reliability of the parallel trends assumption. Relative to Assumption 2, the additional time periods are ones where $g > t$ which are pre-treatment time periods. In this section, we describe how one can construct such a test in our context. Interestingly, our proposed testing procedure exploits more information than simply testing whether $ATT(g, t)$ are equal to zero for all $2 \leq t < g$, and therefore is able to detect a broader set of violations of the stronger conditional parallel trends condition.

Before proceeding, we state the "augmented" conditional parallel trends assumption that allows us to "pre-test" for the conditional parallel trends assumption stated in Assumption 2.

**Assumption 6** (Augmented Conditional Parallel Trends). *For all $t = 2, \ldots, \mathcal{T}$, $g = 2, \ldots, \mathcal{T}$,*

$$\mathbb{E}[Y_t(0) - Y_{t-1}(0)|X, G_g = 1] = \mathbb{E}[Y_t(0) - Y_{t-1}(0)|X, C = 1] \ a.s..$$

In order to understand how such an assumption leads to testable implications, note that, under Assumption 6, for $2 \leq t < g \leq \mathcal{T}$, $\mathbb{E}[Y_t(0)|X, G_g = 1]$ can be expressed as

$$\mathbb{E}[Y_t(0)|X, G_g = 1] = \mathbb{E}[Y_t(0) - Y_{t-1}(0)|X, C = 1] + \mathbb{E}[Y_{t-1}(0)|X, G_g = 1]$$

$$= \mathbb{E}[Y_t - Y_{t-1}|X, C = 1] + \mathbb{E}[Y_{t-1}|X, G_g = 1], \tag{4.1}$$

where the second equality follows since for individuals in group $g$ when $g > t$, $Y_{t-1}(0)$ is observed since treatment did not started yet. Using exactly the same logic, $Y_t(0)$ is also the observed outcome for individuals in group $g$ when $g > t$. Thus, the construction of our test is based on comparing $\mathbb{E}[Y_t(0)|X, G_g = 1]$ in (4.1) to $\mathbb{E}[Y_t|X, G_g = 1]$ for all periods such $2 \leq t < g$: under

Assumption 6 these conditional expectations should be equal.

Formally, the null hypothesis we seek to test is

$$H_0 : \mathbb{E}[Y_t - Y_{t-1}|X, G_g = 1] - \mathbb{E}[Y_t - Y_{t-1}|X, C = 1] = 0 \ a.s. \text{ for all } 2 \leq t < g \leq \mathcal{T}. \qquad (4.2)$$

One option to assess $H_0$ is to nonparametrically estimate each conditional expectation in (4.2), and compare how close their difference is to zero. Such a procedure would involve choosing smoothing parameters such as bandwidths, assuming additional smoothness conditions of these expectations, potentially ruling out discrete covariates $X$, and would also suffer from the "curse of dimensionality" when the dimension of $X$ is moderate.

In order to avoid these potential drawbacks, one can test an *implication* of $H_0$ by using the results of Theorem 1, and compare how close to zero are the estimates of $ATT(g, t)$ for all $2 \leq t < g \leq \mathcal{T}$. Although intuitive, such a procedure does not exploit all the restrictions imposed by $H_0$. For instance, deviations from $H_0$ in opposite directions for different values of $X$ could offset each other, implying that one may fail to reject the plausibility of the conditional parallel trends assumption, even when $H_0$ is violated in some directions. See Remark 7 at the end of this section for more details about this case.

We adopt an alternative approach that avoids all the aforementioned drawbacks: it does not involve choosing bandwidths, does not impose additional smoothness conditions, does not suffer from the "curse of dimensionality," and exploits all the testable restrictions implied by the augmented conditional parallel trends assumption. Our proposal builds on the integrated conditional moments (ICM) approach commonly used in the goodness-of-fit literature; see e.g. Bierens (1982), Bierens and Ploberger (1997), Stute (1997), Stinchcombe and White (1998), and Escanciano (2006a,b, 2008). To the best of our knowledge, we are the first to propose to use ICM to assess the plausibility of the parallel trends assumption, even when there is no treatment timing variation.

Let $w_g^G$ and $w_g^C$ be defined as in (3.2). After some algebra, under Assumptions 1-5, we can rewrite $H_0$ as

$$H_0 : \mathbb{E}\left[ \left( w_g^G - w_g^C \right) (Y_t - Y_{t-1}) |X \right] = 0 \ a.s. \text{ for all } 2 \leq t < g \leq \mathcal{T}, \qquad (4.3)$$

see Lemma A.4 in the Appendix. In fact, by exploiting Lemma 1 in Escanciano (2006$b$), we can further characterize (4.3) as

$$H_0 : \mathbb{E}\left[\left(w_g^G - w_g^C\right)\gamma(X, u)(Y_t - Y_{t-1})\right] = 0 \,\, \forall u \in \Xi \text{ for all } 2 \leq t < g \leq \mathcal{T},  \qquad (4.4)$$

where $\Xi$ is a properly chosen space, and the parametric family $\{\gamma(\cdot, u) : u \in \Xi\}$ is a family of weighting functions such that the equivalence between (4.3) and (4.4) holds. The most popular weighting functions include $\gamma(X, u) = \exp(iX'u)$ as in Bierens (1982) and $\gamma(X, u) = 1\{X \leq u\}$ as in Stute (1997). In the following, to ease the notation, we concentrate our attention on the indicator functions, $\gamma(X, u) = 1\{X \leq u\}$, with $\Xi = \mathcal{X}$, the support of the covariates $X$.

The advantage of the representation in (4.4) is that it resembles the expression for $ATT(g, t)$ in (2.1), and therefore we can use a similar estimation procedure that avoids the use of smoothing parameters. To see this, let

$$J(u, g, t, p_g) = \mathbb{E}\left[\left(w_g^G - w_g^C\right)1(X \leq u)(Y_t - Y_{t-1})\right],$$

and, for each $u$ in the support of $X$, we can estimate $J(u, g, t, p_g)$ by

$$\widehat{J}(u, g, t, \hat{p}_g) = \mathbb{E}_n\left[\left(\frac{G_g}{\mathbb{E}_n[G_g]} - \frac{\dfrac{\hat{p}_g(X)C}{1 - \hat{p}_g(X)}}{\mathbb{E}_n\left[\dfrac{\hat{p}_g(X)C}{1 - \hat{p}_g(X)}\right]}\right)1(X \leq u)(Y_t - Y_{t-1})\right],$$

where $\hat{p}_g$ is a first-step estimator of $p_g$.

With $\widehat{J}(u, g, t, \hat{p}_g)$ in hand, one should reject $H_0$ when it is not "too close" to zero across different values of $u$, $g$, and $t$, $2 \leq t < g \leq \mathcal{T}$. In order to evaluate the distance from $\widehat{J}(u, g, t, \hat{p}_g)$ to zero, we consider the Cramér-von Mises norm,

$$CvM_n = \int_{\mathcal{X}}\left|\sqrt{n}\widehat{J}_{g>t}(u)\right|_M^2 F_{n,X}(du)$$

where $J_{g>t}(u)$ and $\widehat{J}_{g>t}(u)$ denote the vector of $J(u, g, t, p_g)$ and $\hat{J}(u, g, t, \hat{p}_g)$, respectively, for all $g = 2, \ldots, \mathcal{T}$ and $t = 2, \ldots, \mathcal{T}$, such that $2 \leq t < g \leq \mathcal{T}$, $|A|_M$ denotes the weighted seminorm $\sqrt{A'MA}$ for a positive semidefinite matrix $M$ and a real vector $A$, and $F_{n,X}$ is the empirical CDF of $X$. To simplify exposition and leverage intuition, we fix $M$ to be a $(\mathcal{T}-1)^2 \times (\mathcal{T}-1)^2$ diagonal

matrix such that its $(g,t)$-th diagonal element is given by $1\{g > t\}$. As a result, we can write $CvM_n$ as

$$CvM_n = \sum_{g=2}^{\mathcal{T}} \sum_{t=2}^{\mathcal{T}} 1\{g > t\} \int_{\mathcal{X}} \left| \sqrt{n}\widehat{J}(u, g, t, \hat{p}_g) \right|^2 F_{n,X}(du). \tag{4.5}$$

This choice of test statistic is similar to the one used by Escanciano (2008) in a different context. However, one can choose some other $M$ or other norms as well.

The key step to derive the asymptotic properties of $CvM_n$ is to study the process $\sqrt{n}\widehat{J}(u, g, t, \hat{p}_g)$. Here, note that in contrast to $\widehat{ATT}(g,t)$, $\widehat{J}(u, g, t, p_g)$ is infinite dimensional (since it involves a continuum of $u$), and therefore we need to use uniform (instead of pointwise) arguments. Furthermore, we must account for the uncertainty inherited by using the estimated generalized propensity scores $\hat{p}_g$ instead of the unknown true $p_g$. To accomplish this, we build on the existing literature on empirical processes with a first step estimation of the propensity score; see e.g. Donald and Hsu (2014) and Sant'Anna (2017) for applications in the causal inference context. As before, we focus on the case where the $p_g$ is estimated parametrically.

Define

$$\psi_{ugt}^{test}(\mathcal{W}_i) = \psi_{ugt}^{G,test}(\mathcal{W}_i) - \psi_{ugt}^{C,test}(\mathcal{W}_i), \tag{4.6}$$

where

$$\psi_{ugt}^{G,test}(\mathcal{W}) = w_g^G \left[ (Y_t - Y_{t-1})\, 1(X \le u) - \mathbb{E}\left[ w_g^G 1(X \le u)(Y_t - Y_{t-1}) \right] \right],$$

$$\psi_{ugt}^{C,test}(\mathcal{W}) = w_g^C \left[ (Y_t - Y_{t-1})\, 1(X \le u) - \mathbb{E}\left[ w_g^C 1(X \le u)(Y_t - Y_{t-1}) \right] \right] + M_{ugt}^{test\,\prime}\, \xi_g^{\pi}(\mathcal{W}),$$

with $\xi_g^{\pi}(\mathcal{W})$ as defined in (3.1), and

$$M_{ugt}^{test} = \frac{\mathbb{E}\left[ X \left( \dfrac{C}{1 - p_g(X)} \right)^2 \dot{p}_g(X) \left[ 1(X \le u)(Y_t - Y_{t-1}) - \mathbb{E}\left[ w_g^C 1(X \le u)(Y_t - Y_{t-1}) \right] \right] \right]}{\mathbb{E}\left[ \dfrac{p_g(X)\, C}{1 - p_g(X)} \right]}.$$

Let $\Psi_{g>t}^{test}(\mathcal{W}_i; u)$ denote the vector of $\psi_{ugt}^{test}(\mathcal{W}_i)$ across all periods $t$ and groups $g$ such that $2 \le t < g \le \mathcal{T}$.

The next theorem establishes the weak convergence of the process $\sqrt{n}\widehat{J}_{g>t}(u)$ under $H_0$, characterizes the limiting null distribution of $CvM_n$, and shows that our proposed test is consistent.

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From these results, we can conclude that our proposed test controls size, and if Assumption 6 does not hold, our test procedure rejects $H_0$ with probability approaching one as $n$ goes to infinity. Hence, our tests can indeed be used to assess the reliability of our main identification assumption.

**Theorem 4.** *Suppose Assumptions 1-5 hold. Then,*

*1. If Assumption 6 holds, i.e., under the null hypothesis (4.4), as $n \to \infty$,*

$$\sqrt{n}\widehat{J}_{g>t}(u) \Rightarrow \mathbb{G}(u) \ in \ l^\infty(\mathcal{X}),$$

*where $\Rightarrow$ denote weak convergence in the sense of J. Hoffmann-Jørgensen (see e.g. Definition 1.3.3 in van der Vaart and Wellner (1996)), $\mathcal{X}$ is the support of $X$, and $\mathbb{G}$ is a zero-mean Gaussian process with covariance function*

$$V(u_1, u_2) = \mathbb{E}[\Psi_{g>t}^{test}(\mathcal{W}; u_1)\,\Psi_{g>t}^{test}(\mathcal{W}; u_2)'].$$

*In particular, as $n \to \infty$.*

$$CvM_n \xrightarrow{d} \int_{\mathcal{X}} |\mathbb{G}(u)|_M^2\, F_X(du)$$

*2. If Assumption 6 does not hold, i.e., under the negation of the null hypothesis (4.4)*

$$\lim_{n\to\infty} P\left(CvM_n > c_\alpha^{CvM}\right) = 1,$$

*where $c_\alpha^{CvM} = \inf\{c \in [0, \infty) : \lim_{n\to\infty} P(CvM_n > c) = \alpha\}$.*

From Theorem 4, we see that the asymptotic distribution of $CvM_n$ depends on the underlying data generating process (DGP) and standardization is complicated. To overcome this problem, we propose to compute critical values with the assistance of the multiplier bootstrap akin to the one discussed in Theorem 3.

To proceed, let $\widehat{\Psi}_{g>t}^{test}(\cdot; u)$ denote the sample-analogue of $\Psi_{g>t}^{test}(\cdot; u)$, where population expectations are replaced by their empirical analogues, and the true generalized propensity score, $p_g$, and its derivatives, $\dot{p}_g$, are replaced by their MLE estimates, $\hat{p}_g$ and $\widehat{\dot{p}}_g$, respectively. Let

$$\widehat{J}_{g>t}^*(u) = \mathbb{E}_n\left[V \cdot \widehat{\Psi}_{g>t}^{test}(\mathcal{W}; u)\right], \tag{4.7}$$

where $\{V_i\}_{i=1}^n$ is defined as in Section 3. The next algorithm provides a step-by-step procedure to approximate $c_\alpha$, the critical value of our test $CvM_n$.

**Algorithm 2.** 1) *Draw a realization of* $\{V_i\}_{i=1}^n$. 2) *For each* $u \in \mathcal{X}$, *compute* $\widehat{J}_{g>t}^*(u)$ *as in (4.7).* 3) *Compute* $CvM_n^* = \int_{\mathcal{X}} \left| \sqrt{n}\widehat{J}_{g>t}^*(u) \right|_M^2 F_{n,X}(du)$. 4) *Repeat steps 1-3 B times.* 5) *Construct* $\widehat{c}_{1-\alpha}^{CvM}$ *as the empirical* $(1-a)$-*quantile of the B bootstrap draws of* $CvM_n^*$.

The next Theorem establishes the asymptotic validity of the multiplier bootstrap described in Algorithm 2.

**Theorem 5.** *Suppose Assumptions 1-5 hold. Then, under the null hypothesis (4.4) or under fixed alternatives (i.e., the negation of (4.4)),*

$$\sqrt{n}\widehat{J}_{g>t}^*(u) \underset{*}{\Rightarrow} \mathbb{G}(u) \ in \ l^\infty(\mathcal{X}),$$

*where* $\mathbb{G}(u)$ *in* $l^\infty(\mathcal{X})$ *is the same Gaussian process of Theorem 4 and* $\underset{*}{\Rightarrow}$ *indicates weak convergence in probability under the bootstrap law, see* Giné and Zinn (1990). *In particular,*

$$CvM_n^* \underset{*}{\overset{d}{\to}} \int_{\mathcal{X}} |\mathbb{G}(u)|_M^2 F_X(du).$$

**Remark 6.** As discussed in Remark 3, it is straightforward to account for clustering with the multiplier bootstrap described above.

**Remark 7.** As described above, our proposed test $CvM_n$ fully exploits the null hypothesis (4.4), and can detect a broad set of violations against the conditional parallel trends assumption. However, sometimes researchers are also interested in visualizing deviations from the conditional parallel trends assumption, but our proposed Cramér-von Mises test does not directly provide that. In such cases, we note that one can test an implication of the augmented conditional parallel trends assumption, at the cost of losing power against some directions. Namely, under the augmented conditional parallel trends assumptions, $ATT(g,t)$ should be equal to 0 in periods before individuals become treated, that is, when $g > t$. This test is simple to implement in practice though it is distinct from the tests commonly employed in DID with multiple periods and multiple groups (see

e.g. Autor et al. (2007) and Angrist and Pischke (2008)) which we briefly discuss in Appendix D in the Supplementary Appendix. In fact, as formally shown by Abraham and Sun (2018), traditional regression-based tests for (unconditional) pre-trends may be unreliable in settings with treatment effect heterogeneity. Our proposal does not suffer from this drawback.

Let $ATT_{g>t}$ denote the "ATT" in periods before an individual in group $g$ is treated (and also satisfying $2 \leq g$). Using exactly the same arguments as in Section 3, one can establish the limiting distribution of an estimator of $ATT_{g>t}$ (we omit the details for brevity). And one can implement a test of the augmented parallel trends assumption using a Wald-type test. We also found it helpful in the application to obtain the joint limiting distribution of estimators of $ATT_{g \leq t}$ and $ATT_{g>t}$ (once again using the same arguments as in Section 3) and then reporting uniform confidence bands that cover both pre-tests and estimates of $ATT(g, t)$ across all $g = 2, \ldots, \mathcal{T}$ and $t = 2, \ldots, \mathcal{T}$. From these uniform confidence bands, one can immediately infer whether or not the implication of the augmented parallel trends assumption is violated.

# 5 The Effect of Minimum Wage Policy on Teen Employment

In this section, we illustrate the empirical relevance of our proposed methods by studying the effect of the minimum wage on teen employment.

From 1999-2007, the federal minimum wage was flat at $5.15 per hour. In July 2007, the federal minimum wage was raised from $5.15 to $5.85. We focus on county level teen employment in states whose minimum wage was equal to the federal minimum wage at the beginning of the period. Some of these states increased their minimum wage over this period – these become treated groups. Others did not – these are the untreated group. This setup allows us to have more data than local case study approaches. On the other hand, it also allows us to have cleaner identification (state-level minimum wage policy changes) than in studies with more periods; the latter setup is more complicated than ours particularly because of the variation in the federal minimum wage over time. It also allows us to check for internal consistency of identifying assumptions – namely

whether or not the identifying assumptions hold in periods before particular states raised their minimum wages.

We use county-level data on teen employment and other county characteristics. County level teen employment as well as minimum wage levels by state comes from the Quarterly Workforce Indicators (QWI), as in Dube et al. (2016); see Dube et al. (2016) for a detailed discussion of this dataset. Other county characteristics come from the 2000 County Data Book. These include whether or not a county is located in an MSA, county population in 2000, the fraction of population that are white, educational characteristics from 1990, median income in 1997, and the fraction of population below the poverty level in 1997.

For forty-one states, the federal minimum wage was binding in quarter 2 of 1999. We omit two states that raised their minimum wage between then and the first quarter of 2004. We drop several other states for lack of data. We also drop states in the Northern census region because all but two of them had minimum wages higher than the federal minimum wage at the beginning of the period and census region is an important control in the minimum wage literature. We use quarterly employment in the first quarter of each year from 2001 to 2007 for employment among teenagers. Alternatively, we could use more periods of data, but this would come at the cost of losing several states due to lack of data. Also, we choose first quarter employment because it is further away from the federal minimum wage increase in Q3 of 2007. Our final sample includes county level teen employment for 29 states matched with county characteristics.

Our strategy is to divide the observations based on the timing of when a state increased its minimum wage above the federal minimum wage. States that did not raise their minimum wage during this period form the untreated group. We also have groups of states that increased their minimum wage during 2004, 2006, and 2007.[6] Before 2004, Illinois did not have a state minimum wage. In Q1 of 2004, Illinois set a state minimum wage of $5.50 which was 35 cents higher than the federal minimum wage. In Q1 of 2005, Illinois increased its minimum wage to $6.50 where it stayed for the remainder of the period that we consider. No other states changed their minimum wage policy by the first quarter of 2005. In the second quarter of 2005, Florida and Wisconsin

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[6]To be precise, we use only employment data from the first quarter of each year. A state is considered to raise its minimum wage in year $y$ if it raised its minimum wage in Q2, Q3, or Q4 of year $y-1$ or in Q1 of year $y$.

set a state minimum wage above the federal minimum wage. In Q3 of 2005, Minnesota also set a state minimum wage. Florida and Wisconsin each gradually increased their minimum wages over time, while Minnesota's was flat over the rest of the period. These three states constitute the treated group for 2006. West Virginia increased its minimum wage in Q3 of 2006; Michigan and Nevada increased their minimum wages in Q4 of 2006; Colorado, Maryland, Missouri, Montana, North Carolina, and Ohio increased their state minimum wages in Q1 of 2007. These states form the 2007 treated group. Among these there is some heterogeneity in the size of the minimum wage increase. For example, North Carolina only increased its minimum wage to $6.15 though each state increased its minimum wage to strictly more than the new federal minimum wage of $5.85 per hour in Q3 of 2007. At the other extreme, Michigan increased its minimum wage to $6.95 and then to $7.15 by Q2 of 2007.

Figure 2 contains the spatial distribution of state-level minimum wage policy changes in our sample. Dube et al. (2010) argue that differential trends in employment rates across regions bias estimates of the effect of changes in state-level minimum wages. Indeed, Figure 2 shows that states in the Southeast are less likely to increase their minimum wage between 2001 and 2007 than states in the Northeast or Midwest. Table 1 contains the complete details of the exact date when a state changed its minimum wage as well as which states are used in our analysis.

Summary statistics for county characteristics are provided in Table 2. As discussed above, treated counties are much less likely to be in the South. They also have much lower population (on average 53,000 compared to 94,000 for treated counties). The proportion of black residents is much higher in treated counties (on average, 10% compared to 6% for untreated counties). There are smaller differences in the fraction with high school degrees and the poverty rate though the differences are both statistically significant. Treated counties have a somewhat smaller fraction of high school graduates and a somewhat higher poverty rate.

In the following we discuss different sets of results using different identification strategies. In particular, we consider the cases in which one would assume that the parallel trends assumption would hold unconditionally, and when it holds only after controlling for observed characteristics $X$.

Figure 2: The Spatial Distribution of States by Minimum Wage Policy
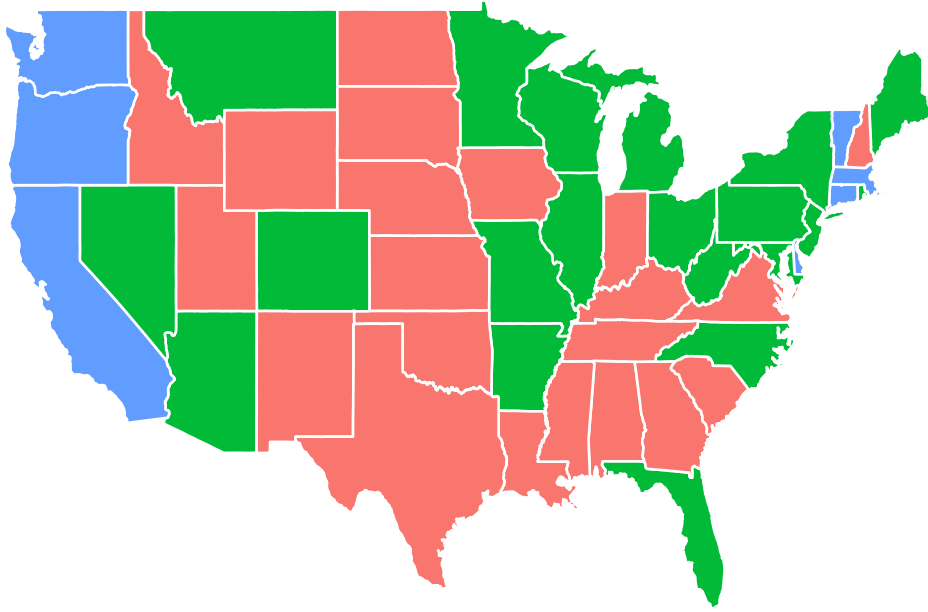


*Notes:* Blue states had minimum wages higher than the federal minimum wage in Q1 of 2000. Green states increased their state minimum wage between Q2 of 2000 and Q1 of 2007. Some of these states are omitted from the main dataset either due to missing data or being located in the Northern census region where there are no states that did not raise their minimum wage between 2000 and 2007 with available data. Otherwise, the green states constitute the treated group. See Table 1 for exact timing of each state's change in the minimum wage. Red states did not increase their minimum wage over the period from 2000 to 2007.

Table 1: Timing of States Raising Minimum Wage

| State | Year-Quarter Raised MW | State | Year-Quarter Raised MW |
|---|---|---|---|
| Alabama | Never Increased | Montana* | 2007-1 |
| Alaska | Always Above | Nebraska* | Never Increased |
| Arizona | 2007-1 | Nevada* | 2006-4 |
| Arkansas | 2006-4 | New Hampshire | Never Increased |
| California | Always Above | New Jersey | 2005-4 |
| Colorado* | 2007-1 | New Mexico* | Never Increased |
| Connecticut | Always Above | New York | 2005-1 |
| Delaware | 1999-2 | North Carolina* | 2007-1 |
| Florida* | 2005-2 | North Dakota* | Never Increased |
| Georgia* | Never Increased | Ohio* | 2007-1 |
| Hawaii | Always Above | Oklahoma* | Never Increased |
| Idaho* | Never Increased | Oregon | Always Above |
| Illinois* | 2004-1 | Pennsylvania | 2007-1 |
| Indiana* | Never Increased | Rhode Island | 1999-3 |
| Iowa* | 2007-2 | South Carolina* | Never Increased |
| Kansas* | Never Increased | South Dakota* | Never Increased |
| Kentucky | Never Increased | Tennessee* | Never Increased |
| Louisiana* | Never Increased | Texas* | Never Increased |
| Maine | 2002-1 | Utah* | Never Increased |
| Maryland* | 2007-1 | Vermont | Always Above |
| Massachusetts | Always Above | Virginia* | Never Increased |
| Michigan* | 2006-4 | Washington | 1999-1 |
| Minnesota* | 2005-3 | West Virginia* | 2006-3 |
| Mississippi | Never Increased | Wisconsin* | 2005-2 |
| Missouri* | 2007-1 | Wyoming | Never Increased |

*Notes*: The timing of states increasing their minimum wage above the federal minimum wage of $5.15 per hour which was set in Q4 of 1997 and did not change again until it increased in Q3 of 2007. States that are ultimately included in the main sample are denoted with a *. States that had minimum wages higher than the federal minimum wage at the beginning of the period are excluded. We also exclude some states who raised their minimum wage very soon after the federal minimum wage increase, some others due to lack of data availability, and those in the Northern Census region. There are 29 states ultimately included in the sample.

Table 2: Summary Statistics for Main Dataset

|  | Treated States | Untreated States | Diff | P-val on Difference |
|---|---|---|---|---|
| Midwest | 0.59 | 0.34 | 0.259 | 0.00 |
| South | 0.27 | 0.59 | -0.326 | 0.00 |
| West | 0.14 | 0.07 | 0.067 | 0.00 |
| Black | 0.06 | 0.10 | -0.042 | 0.00 |
| HS Graduates | 0.59 | 0.55 | 0.327 | 0.00 |
| Population (1000s) | 94.32 | 53.43 | 40.896 | 0.00 |
| Poverty Rate | 0.13 | 0.16 | -0.259 | 0.00 |

*Notes*: Summary statistics for counties located in states that raised their minimum wage between Q2 of 2003 and Q1 of 2007 (treated) and states whose minimum wage was effectively set at the federal minimum wage for the entire period (untreated). The sample consists of 2284 counties. *Sources*: Quarterly Workforce Indicators and 2000 County Data Book

The first set of results comes from using the unconditional parallel trends assumption to estimate the effect of raising the minimum wage on teen employment. The results for group-time average treatment effects are reported in Figure 3 along with a uniform 95% confidence band. All inference procedures use clustered bootstrapped standard errors at the county level, and account for the autocorrelation of the data. The plot contains pre-treatment estimates that can be used to test the parallel trends assumption as well as treatment effect estimates in post-treatment periods.

The group-time average treatment effect estimates provide support for the view that increases on the minimum wage lead to a reduction in teen employment. For 4 out of 7 group-time average treatment effects, there is a clear statistically significant negative effect on employment. The other three are marginally insignificant (and negative). The group-time average treatment effects range from 2.3% lower teen employment to 13.6% lower teen employment. The simple average (weighted only by group size) is 5.2% lower teen employment (see Table 3). A two-way fixed effects model with a post treatment dummy variable also provides similar results, indicating 3.7% lower teen employment due to increasing the minimum wage. In light of the literature on the minimum wage these results are not surprising as they correspond to the types of regressions that tend to find that increasing the minimum wage decreases employment; see the discussion in Dube et al. (2010).

As in Meer and West (2016), there also appears to be a dynamic effect of increasing the minimum wage. For Illinois (the only state in the group that first raised its minimum wage in

Figure 3: Minimum Wage Results under Unconditional DID
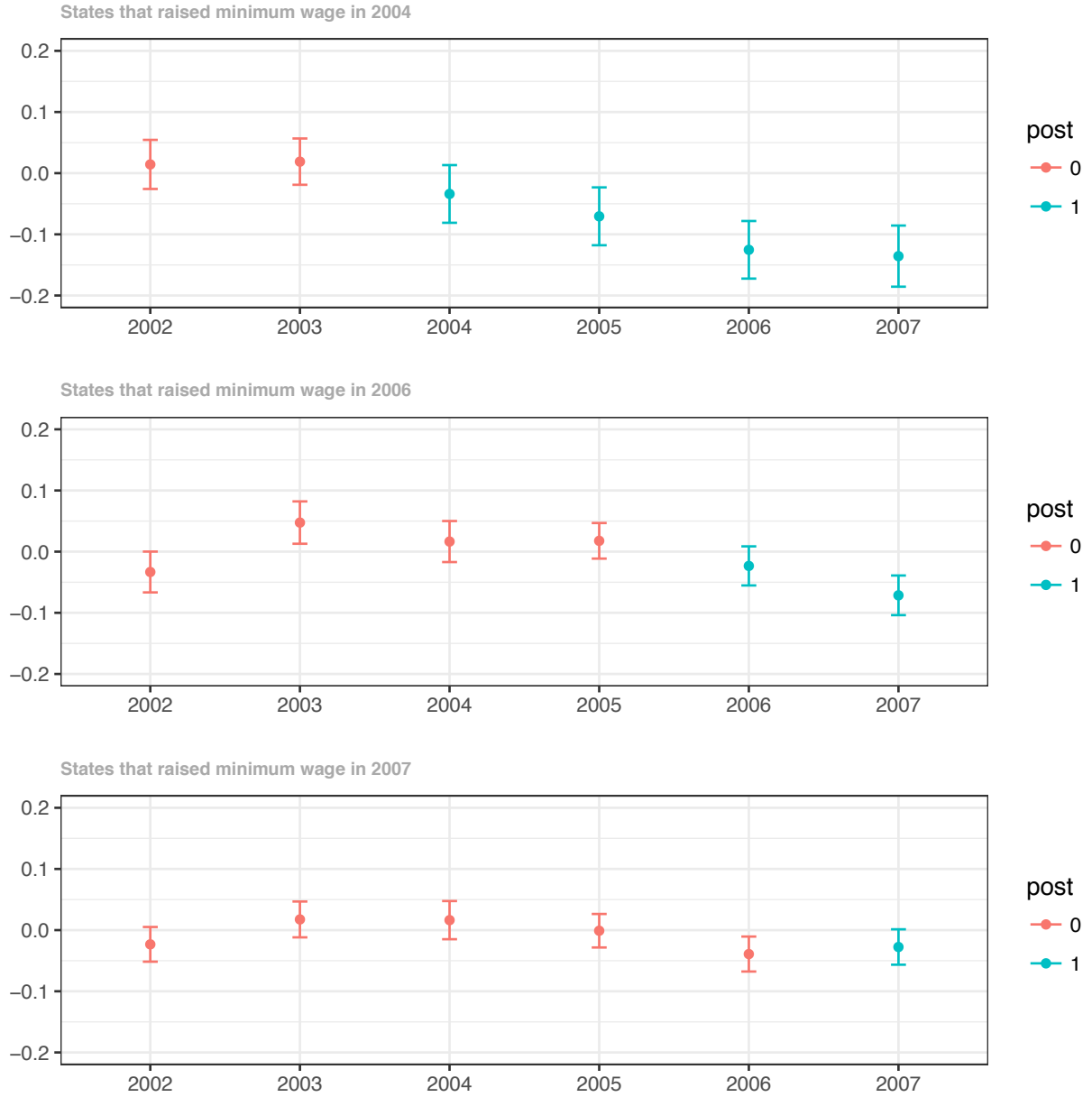
*Notes:* The effect of the minimum wage on teen employment estimated under the Unconditional DID Assumption. Red lines give point estimates and uniform 95% confidence bands for pre-treatment periods allowing for clustering at the county level. Under the null hypothesis of the Unconditional DID Assumption holding in all periods, these should be equal to 0. Blue lines provide point estimates and uniform 95% confidence bands for the treatment effect of increasing the minimum wage allowing for clustering at the county level. The top panel includes states that increased their minimum wage in 2004, the middle panel includes states that increased their minimum wage in 2006, and the bottom panel includes states that increased their minimum wage in 2007. No states raised their minimum wages in other years prior to 2007.

2004), teen employment is 3.4% lower on average in 2004 than it would have been if the minimum wage had not been increased. In 2005, teen employment is estimated to 7.1% lower; in 2006, 12.5% lower; and in 2007, 13.6% lower. For states first treated in 2006, there is a small effect in 2006 – 2.3% lower teen employment; however, it is larger in 2007 – 7.1% lower teen employment.

Table 3 reports aggregated treatment effect measures. Allowing for dynamic treatment effects is perhaps the most useful for our study. These parameters paint largely the same picture as the group-time average treatment effects. The effect of increasing the minimum wage on teen employment appears to be negative and getting stronger the longer states are exposed to the higher minimum wage. In particular, in the first year that a state increases its minimum wage, teen employment is estimated to decrease by 2.7%, in the second year it is estimated to decrease by 7.1%, in the third year by 12.5%, and in the fourth year by 13.6%. Notice that the last two dynamic treatment effect estimates are exactly the same as the estimates coming from Illinois alone because Illinois is the only state that is treated for more than two years. These results are robust to keeping the treated group constant to make sure that selective treatment timing does not bias the results (see the row in Table 3 labeled 'Selectivity and Dynamics'). When we restrict the sample to only include groups with at least two years of exposure to treatment (and only considering the first two periods of exposure which keeps the groups constant across length of exposure), we estimate that the effect of minimum wage increases in the first period of exposure is 2.7% lower teen employment and 7.1% lower teen employment in the second period.[7]

Allowing for calendar time effects or selective treatment timing also is consistent with the idea that states that increased their minimum wage experienced negative effects on teen employment relative to what they would have experienced if they had not increased their minimum wage.

We consider testing the unconditional parallel trends assumption. First, since the confidence bands in Figure 3 are uniform, one can immediately infer that the unconditional parallel trends assumption should be rejected based on the implication of the unconditional parallel trends as-

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[7]Notice that these estimates are exactly the same as in the first two periods for the dynamic treatment effect estimates that do not condition on the group remaining constant. The reason that they are the same for the first period is coincidental; the estimated effect of the minimum wage in 2007 for the group of states first treated in 2007 is 2.76% lower teen employment which just happens to correspond to the estimated effect in the latter case. For the second period, they correspond by construction.

Table 3: Aggregate Treatment Effect Parameters under Unconditional Parallel Trends

| | **Partially Aggregated** | | | | **Single Parameters** |
|---|---|---|---|---|---|
| Standard DID | | | | | -0.037 |
| | | | | | (0.006) |
| Simple Weighted Average | | | | | -0.052 |
| | | | | | (0.006) |
| Selective Treatment Timing | g=2004 | g=2006 | g=2007 | | |
| | -0.091 | -0.047 | -0.028 | | -0.039 |
| | (0.019) | (0.008) | (0.007) | | (0.007) |
| Dynamic Treatment Effects | e=1 | e=2 | e=3 | e=4 | |
| | -0.027 | -0.071 | -0.125 | -0.136 | -0.090 |
| | (0.006) | (0.009) | (0.021) | (0.023) | (0.013) |
| Calendar Time Effects | t=2004 | t=2005 | t=2006 | t=2007 | |
| | -0.034 | -0.071 | -0.055 | -0.050 | -0.052 |
| | (0.019) | (0.02) | (0.009) | (0.006) | (0.013) |
| Selectivity and Dynamics | e=1 | e=2 | | | |
| | -0.027 | -0.071 | | | -0.049 |
| | (0.009) | (0.009) | | | (0.008) |

*Notes:* The table reports aggregated treatment effect parameters under the Unconditional DID Assumption and with clustering at the county level. The row 'Standard DID' reports the coefficient on a post-treatment dummy variable from a two-way fixed effects regression. The row 'Single Weighted Average' reports the weighted average (by group size) of all available group-time average treatment effects as in Equation (2.3). The row 'Selective Treatment Timing' allows for period that a county is first treated to affect its group-time average treatment effect; here, $g$ indexes the year that a county is first treated. The row 'Dynamic Treatment Effects' allows for the effect of the minimum wage to depend on length of exposure; here, $e$ indexes the length of exposure to the treatment. The row 'Calendar Time Effects' allows the effect of the minimum wage to change across years; here, $t$ indexes the year. The row 'Selectivity and Dynamics' allows for the effect of the minimum wage to depend on length of exposure while making sure that the composition of the treatment group does not change with $e$; here, $e$ indexes the length of exposure and the sample consists of counties that have at least two years of exposure to minimum wage increases. The column 'Single Parameters' represents a further aggregation of each type of parameter, as discussed in the text.

sumption that the "ATT" in periods before treatment should be equal to 0. Likewise, our proposed test also rejects the unconditional parallel trends assumption (p-value: 0.000). The estimated uniform confidence bands in Figure 3 also provide some insight into how to think about our pre-tests. For the group first treated in 2004, the parallel trends assumption is not rejected in any period. For the group first treated in 2006, it is rejected in 2003; for the group first treated in 2007, it is rejected in 2006. Interestingly, with the exception of 2006 for the group first treated in 2007, in each of the cases where it is rejected, the placebo estimates are positive.

The second set of results comes from using the conditional parallel trends assumption; that is,

we assume only that counties with *the same characteristics* would follow the same trend in teen employment in the absence of treatment. The county characteristics that we use are region of the country, county population, county median income, the fraction of the population that is white, the fraction of the population with a high school education, and the county's poverty rate. Estimation requires a first step estimation of the generalized propensity score. For each generalized propensity score, we estimate a logit model that includes each county characteristic along with quadratic terms for population and median income.[8] In particular, the conditional results allow for differential trends in teen employment across different regions as well as in the other county characteristics mentioned above. In what follows, all inference procedures use clustered bootstrapped standard errors at the county level.

For comparison's sake, we first estimate the coefficient on a post-treatment dummy variable in a model with individual fixed effects and region-year fixed effects. This is very similar to one of the sorts of models that Dube et al. (2010) finds to eliminate the correlation between the minimum wage and employment. Like Dube et al. (2010), using this specification, we find that the estimated coefficient is small and not statistically different from 0. However, one must have in mind that the approach we proposed in this article is different from the two-way fixed effects regression. In particular, we explicitly identify group-time average treatment effects for different groups and different times, allowing for arbitrary treatment effect heterogeneity as long as the conditional parallel trends assumption is satisfied. Thus, our causal parameters have a clear interpretation. As pointed out by Wooldridge (2005), Chernozhukov et al. (2013), de Chaisemartin and D'Haultfœuille (2018), Borusyak and Jaravel (2017), Goodman-Bacon (2018) and Słoczyński (2017), the same may not be true for two-way fixed effect regressions in the presence of treatment effect heterogeneity.[9]

The results using our approach are available in Figure 4 and Table 4. Interestingly, we find

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[8]Using the propensity score specification tests proposed by Sant'Anna and Song (2018), we fail to reject the null hypothesis that these models are correctly specified at the usual significance levels.

[9]Our approach is also different from that of Dube et al. (2010) in several other ways that are worth mentioning. We focus on teen employment; Dube et al. (2010) considers employment in the restaurant industry. Their most similar specification to the one mentioned above includes census division-time fixed effects rather than region-time fixed effects though the results are similar. Finally, our period of analysis is different from theirs; in particular, there are no federal minimum wage changes over the periods we analyze.

Figure 4: Minimum Wage Results under Conditional DID
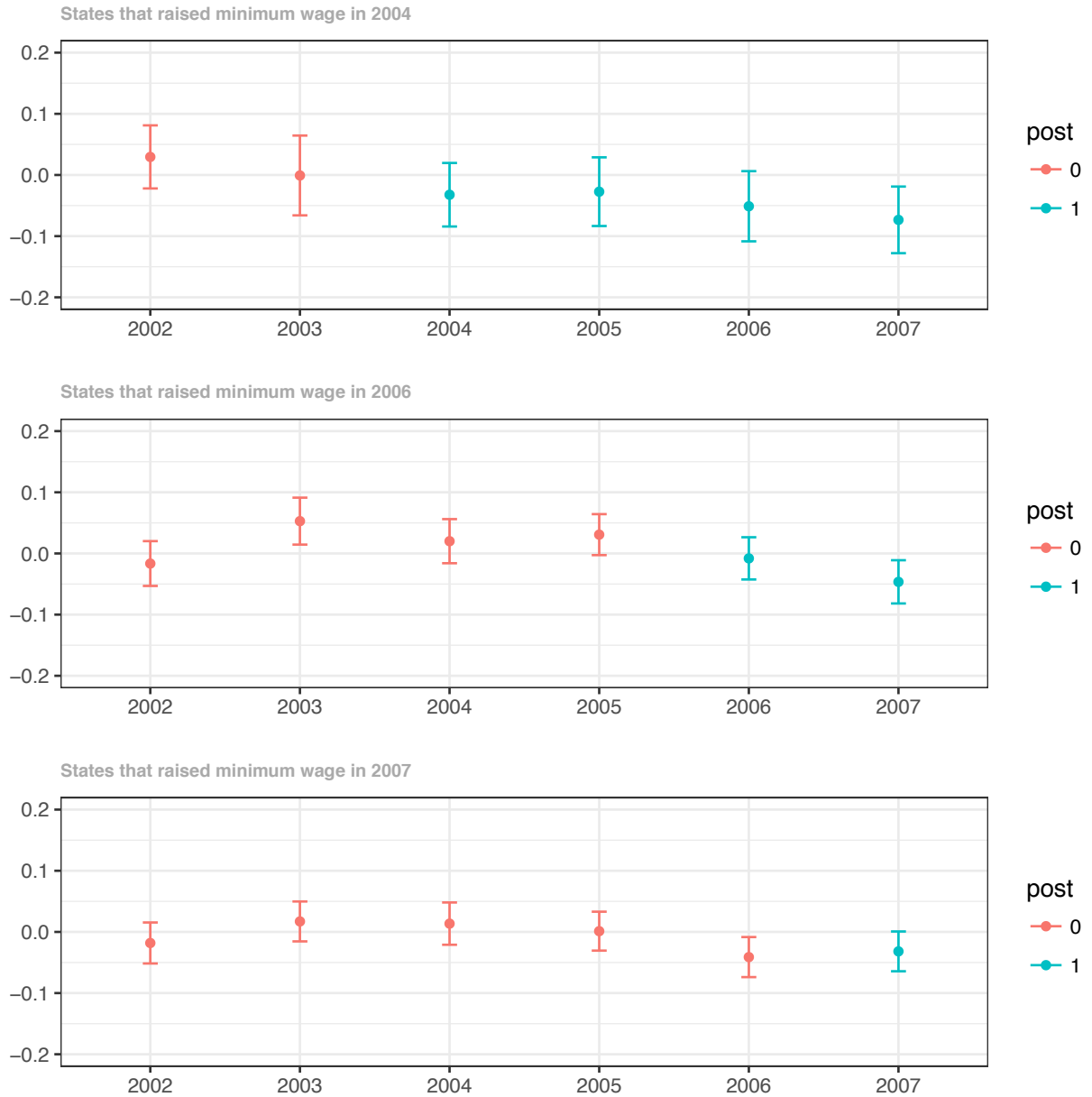
*Notes*: The effect of the minimum wage on teen employment estimated under the Conditional DID Assumption. Red lines give point estimates and uniform 95% confidence bands for pre-treatment periods allowing for clustering at the county level. Under the null hypothesis of the Conditional DID Assumption holding in all periods, these should be equal to 0. Blue lines provide point estimates and uniform 95% confidence bands for the treatment effect of increasing the minimum wage allowing for clustering at the county level. The top panel includes states that increased their minimum wage in 2004, the middle panel includes states that increased their minimum wage in 2006, and the bottom panel includes states that increased their minimum wage in 2007. No states raised their minimum wages in other years prior to 2007.

quite different results using our approach than are suggested by the two-way fixed effect regression approach. In particular, we continue to find evidence that increasing the minimum wage tended to reduce teen employment. The estimated group-time average treatment effects range from 0.8% lower teen employment (not statistically different from 0) in 2006 for the group of states first treated in 2006 to 7.3% lower teen employment in 2007 for states first treated in 2004. Now only 2 of 7 group-time average treatment effects are statistically significant. The pattern of dynamic treatment effects where the effect of minimum wage increases tends to increase with length of exposure is the same as in the unconditional case. Similarly, using our aggregated treatment effect parameters, allowing for dynamic treatment effects, we estimate that increasing the minimum wage led on average to 4.8% lower teen employment. Allowing for dynamic treatment effects and selective treatment timing, we estimate that increasing the minimum wage lowers teen employment by 2.8%.

The evidence of the negative effect of minimum wage increases is somewhat mitigated by the fact that we reject the conditional parallel trends assumption in pre-treatment periods. This is immediately evident from Figure 4 because we can reject that the "ATT" is equal to zero in 2 out of 11 pre-treatment periods. Using the consistent Cramér-von Mises tests discussed in Section 4, we also reject the conditional parallel trends assumption (p-value: 0.000). In addition, we conducted our test of the augmented conditional parallel trends assumption separately for states first treated in 2004 because the pre-treatment "ATT" is not statistically significant in any period for this group. Here, we reject the augmented conditional parallel trends assumption. This is an interesting result because the "visual" test often conducted in empirical work would incorrectly lead the researcher to believe that the conditional parallel trends assumption is valid for states first treated in 2004.

Overall, our results suggests that the minimum wage decreased teen employment in states that increased their minimum wage relative to what it would have been had those states not increased their minimum wage. Nonetheless, our proposed tests indicate that the parallel trends assumption should be rejected in pre-treatment periods, implying that the DID research design may lead to non-reliable conclusions. Perhaps not surprisingly, given the amount of disagreement

Table 4: Aggregate Treatment Effect Parameters under Conditional Parallel Trends

| | Partially Aggregated | | | | Single Parameters |
|---|---|---|---|---|---|
| **Standard DID** | | | | | |
| | | | | | -0.008 |
| | | | | | (0.006) |
| **Simple Weighted Average** | | | | | |
| | | | | | -0.034 |
| | | | | | (0.008) |
| **Selective Treatment Timing** | g=2004 | g=2006 | g=2007 | | |
| | -0.046 | -0.027 | -0.032 | | -0.032 |
| | (0.020) | (0.008) | (0.008) | | (0.007) |
| **Dynamic Treatment Effects** | e=1 | e=2 | e=3 | e=4 | |
| | -0.026 | -0.041 | -0.051 | -0.073 | -0.048 |
| | (0.006) | (0.010) | (0.025) | (0.024) | (0.014) |
| **Calendar Time Effects** | t=2004 | t=2005 | t=2006 | t=2007 | |
| | -0.032 | -0.027 | -0.021 | -0.040 | -0.030 |
| | (0.019) | (0.024) | (0.011) | (0.007) | (0.013) |
| **Selectivity and Dynamics** | e=1 | e=2 | | | |
| | -0.016 | -0.041 | | | -0.028 |
| | (0.009) | (0.010) | | | (0.008) |

*Notes:* The table reports aggregated treatment effect parameters under the Unconditional DID Assumption and with clustering at the county level. The row 'Standard DID' reports the coefficient on a post-treatment dummy variable from a fixed effects regression with individual fixed effects and region-year fixed effects. The row 'Single Weighted Average' reports the weighted average (by group size) of all available group-time average treatment effects as in Equation (2.3). The row 'Selective Treatment Timing' allows for period that a county is first treated to affect its group-time average treatment effect; here, $g$ indexes the year that a county is first treated. The row 'Dynamic Treatment Effects' allows for the effect of the minimum wage to depend on length of exposure; here, $e$ indexes the length of exposure to the treatment. The row 'Calendar Time Effects' allows the effect of the minimum wage to change across years; here, $t$ indexes the year. The row 'Selectivity and Dynamics' allows for the effect of the minimum wage to depend on length of exposure while making sure that the composition of the treatment group does not change with $e$; here, $e$ indexes the length of exposure and the sample consists of counties that have at least two years of exposure to minimum wage increases. The column 'Single Parameters' represents a further aggregation of each type of parameter, as discussed in the text.

in the minimum wage literature, our results should be interpreted with care and are ultimately inconclusive.

# 6  Conclusion

This paper has considered Difference-in-Differences methods in the case where there are more than two periods and individuals can become treated at different points in time – a commonly encountered setup in empirical work in economics. In this setup, we have suggested computing

group-time average treatment effects, $ATT(g,t)$, that are the average treatment effect in period $t$ for the group of individuals first treated in period $g$. Unlike the more common approach of running a regression with a post-treatment dummy variable, $ATT(g,t)$ corresponds to a well defined treatment effect parameter. And once $ATT(g,t)$ has been obtained for different values of $g$ and $t$, they can be aggregated into a single parameter, though the exact implementation depends on the particular case. We view such a flexibility as a plus of our proposed methodology.

Given that our nonparametric identification results are constructive, we proposed to estimate $ATT(g,t)$ using its sample analogue. We established consistency and asymptotic normality of the proposed estimators, and proved the validity of a powerful, but easy to implement, multiplier bootstrap procedure to construct simultaneous confidence bands for $ATT(g,t)$. Importantly, we have also proposed a new pre-test for the reliability of the conditional parallel trends assumption.

We applied our approach to study the effect of minimum wage increases on teen employment. We found some evidence that increasing the minimum wage led to reductions in teen employment and found strikingly different results from the more common approach of interpreting the coefficient on a post-treatment dummy variable as the effect of the minimum wage on employment. However, using the pre-tests developed in the current paper, we found evidence against both the unconditional and conditional parallel trends assumption.

Our results can be extended to other situations of practical interest. For instance, one can combine our proposal with Callaway and Li (2017) in order to consider group-time quantile treatment effects. On a different direction, it is also interesting to extend our inference procedures to incorporate not only the sampling uncertainty as we do in this article, but also the design uncertainty as in Abadie et al. (2017) and Athey and Imbens (2018). In light of our empirical application, we note that it is worth considering DID procedures that relax the conditional parallel trends assumption. A possibility in this direction is to use conditional moment inequalities. More precisely, one could assume that, for all $t = 2, \ldots, \mathcal{T}$, $g = 2, \ldots, \mathcal{T}$, such that $g \leq t$,

$$\mathbb{E}\left[Y_t\left(0\right) - Y_{t-1}\left(0\right) | X, G_g = 1\right] \geq \mathbb{E}\left[Y_t\left(0\right) - Y_{t-1}\left(0\right) | X, C = 1\right] \ a.s.. \tag{6.1}$$

Note that (6.1) implies that

$$ATT(g,t) \leq \mathbb{E}\left[Y_t(1) - Y_{t-1}(0) | X, G_g = 1\right] - \mathbb{E}\left[Y_t(0) - Y_{t-1}(0) | X, C = 1\right] \ a.s..$$

Thus, this one-sided relaxation of the conditional parallel trends assumption suggests that, under (6.1), $\widehat{ATT}(g,t)$ would be an estimator for the upper bound of the $ATT(g,t)$. By combining our pre-test procedure with Andrews and Shi (2013), one would then be able to assess the reliability of (6.1). These extensions are beyond the scope of this article and are left for future research.

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# Supplementary Appendix: Difference-in-Differences with Multiple Time Periods and an Application on the Minimum Wage and Employment

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This supplementary appendix contains $(a)$ the proofs of the results stated in the main text; $(b)$ results for the case where a researcher has access to repeated cross sections data rather than panel data; $(c)$ extensions of our main results when using "not yet treated" observations as a control group; and $(d)$ additional details on group-time average treatment effects under an unconditional parallel trends assumption, paying particular attention to the possibilities of using regressions to estimate group-time average treatment effects.

## Appendix A: Proofs of Main Results

We provide the proofs of our results in this appendix. Before proceeding, we first state and prove several auxiliary lemmas that help us proving our main theorems.

Let

$$ATT_X(g,t) = \mathbb{E}[Y_t(1) - Y_t(0)|X, G_g = 1].$$

**Lemma A.1.** *Under Assumptions 1-4, and for $2 \leq g \leq t \leq \mathcal{T}$,*

$$ATT_X(g,t) = \mathbb{E}[Y_t - Y_{g-1}|X, G_g = 1] - \mathbb{E}[Y_t - Y_{g-1}|X, C = 1] \; a.s..$$

**Proof of Lemma A.1:** In what follows, take all equalities to hold almost surely (a.s.). Notice that for identifying $ATT_X(g,t)$, the key term is $E[Y_t(0)|X, G_g = 1]$. And notice that for $h > s$, $E[Y_s(0)|X, G_h = 1] = E[Y_s|X, G_h = 1]$, which holds because in time periods before an individual

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is first treated, their untreated potential outcomes are observed outcomes. Also, note that, for $2 \leq g \leq t \leq \mathcal{T}$,

$$\mathbb{E}[Y_t(0)|X, G_g = 1] = \mathbb{E}[\Delta Y_t(0)|X, G_g = 1] + \mathbb{E}[Y_{t-1}(0)|X, G_g = 1] \qquad (A.1)$$
$$= \mathbb{E}[\Delta Y_t|X, C = 1] + \mathbb{E}[Y_{t-1}(0)|X, G_g = 1],$$

where the first equality holds by adding and subtracting $E[Y_{t-1}(0)|X, G_g = 1]$ and the second equality holds by Assumption 2. If $g = t - 1$, then the last term in the final equation is identified; otherwise, one can continue recursively in similar way to (A.1) but starting with $\mathbb{E}[Y_{t-1}(0)|X, G_g = 1]$. As a result,

$$\mathbb{E}[Y_t(0)|X, G_g = 1] = \sum_{j=0}^{t-g} \mathbb{E}[\Delta Y_{t-j}|X, C = 1] + \mathbb{E}[Y_{g-1}|X, G_g = 1]$$
$$= \mathbb{E}[Y_t - Y_{g-1}|X, C = 1] + \mathbb{E}[Y_{g-1}|X, G_g = 1]. \qquad (A.2)$$

Combining (A.2) with the fact that, for all $g \leq t$, $\mathbb{E}[Y_t(1)|X, G_g = 1] = \mathbb{E}[Y_t|X, G_g = 1]$ (which holds because observed outcomes for group $g$ in period $t$ with $g \leq t$ are treated potential outcomes), implies the result. $\square$

Next, recall that

$$\hat{\pi}_g = \arg\max_{\pi} \sum_{i:G_{ig}+C_i=1} G_{ig} \ln\left(p_g\left(X_i'\pi\right)\right) + (1 - G_{ig}) \ln\left(1 - p_g\left(X_i'\pi\right)\right),$$

$\dot{p}_g = \partial p_g(u)/\partial u$, $\dot{p}_g(X) = \dot{p}_g\left(X'\pi_g^0\right)$, and $\pi_g^0$ is the true, unknown vector of parameter indexed the generalized propensity score $p_g(X) = \mathbb{E}[G_g = 1|X, G_g + C = 1]$.

**Lemma A.2.** *Under Assumption 5,*

$$\sqrt{n}\left(\hat{\pi}_g - \pi_g^0\right) = \frac{1}{\sqrt{n}} \sum_{i=1}^{n} \xi_g^{\pi}\left(\mathcal{W}_i\right) + o_p(1),$$

*where*
$$\xi_g^{\pi}(\mathcal{W}) = \mathbb{E}\left[\frac{(G_g + C)\dot{p}_g(X)^2}{p_g(X)(1 - p_g(X))}XX'\right]^{-1} X \frac{(G_g + C)(G_g - p_g(X))\dot{p}_g(X)}{p_g(X)(1 - p_g(X))}.$$

**Proof of Lemma A.2:** Let $n_{gc} = \sum_{i=1}^{n}(C_i + G_{ig})$. Under Assumption 5, from Theorem 5.39 and Example 5.40 in van der Vaart (1998), we have

$$\sqrt{n_{gc}}\left(\hat{\pi}_g - \pi_g^0\right)$$

$$= \frac{1}{\sqrt{n_{gc}}} \sum_{i:G_{ig}+C_i=1} \left( \mathbb{E}\left[ \frac{\dot{p}_g(X)^2}{p_g(X)(1-p_g(X))} XX' \middle| G_g + C = 1 \right]^{-1} X_i \frac{(G_{ig}-p_g(X_i))\dot{p}_g(X_i)}{p_g(X_i)(1-p_g(X_i))} \right) + o_p(1)$$

$$= \frac{\mathbb{E}[G_g+C]}{\sqrt{n_{gc}}} \sum_{i=1}^{n} \left( \mathbb{E}\left[ \frac{(G_g+C)\dot{p}_g(X)^2}{p_g(X)(1-p_g(X))} XX' \right]^{-1} X_i \frac{(G_{ig}+C_i)(G_{ig}-p_g(X_i))\dot{p}_g(X_i)}{p_g(X_i)(1-p_g(X_i))} \right) + o_p(1)$$

$$= \frac{\mathbb{E}_n[G_g+C]}{\sqrt{n_{gc}}} \sum_{i=1}^{n} \xi_g^\pi(\mathcal{W}_i) + o_p(1)$$

$$= \frac{\sqrt{n_{gc}}}{n} \sum_{i=1}^{n} \xi_g^\pi(\mathcal{W}_i) + o_p(1).$$

Thus,

$$\sqrt{n}\left(\hat{\pi}_g - \pi_g^0\right) = \frac{\sqrt{n}}{\sqrt{n_{gc}}} \sqrt{n_{gc}}\left(\hat{\pi}_g - \pi_g^0\right)$$

$$= \frac{1}{\sqrt{n}} \sum_{i=1}^{n} \xi_g^\pi(\mathcal{W}_i) + o_p(1),$$

and the proof is complete. □

For an arbitrary $\pi$, let $p_g(x;\pi) = p_g(x'\pi)$, $\dot{p}_g(x;\pi) = \dot{p}_g(x'\pi)$, for all $g = 2,\ldots,\mathcal{T}$. Define the classes of functions,

$$\mathcal{H}_{1,g} = \left\{ (x,c) \mapsto c\frac{p_g(x;\pi)}{1-p_g(x;\pi)} : \pi \in \Pi_g \right\},$$

$$\mathcal{H}_{2,g} = \left\{ (x,c,y_t,y_{g-1}) \mapsto c\frac{p_g(x;\pi)(y_t-y_{g-1})}{1-p_g(x;\pi)} : \pi \in \Pi_g \right\}$$

$$\mathcal{H}_{3,g} = \left\{ (x,c,y_t,y_{g-1}) \mapsto x\frac{c\,\dot{p}_g(x;\pi)(y_t-y_{g-1})}{(1-p_g(x;\pi))^2} : \pi \in \Pi_g \right\},$$

$$\mathcal{H}_{4,g} = \left\{ (x,c) \mapsto x\frac{c\,\dot{p}_g(x;\pi)}{(1-p_g(x;\pi))^2} : \pi \in \Pi_g \right\},$$

$$\mathcal{H}_{5,g} = \left\{ (x,c,g_g) \mapsto x\frac{(g_g+c)(g_g-p_g(x;\pi))\dot{p}_g(x;\pi)}{p_g(x;\pi)(1-p_g(x;\pi))} : \pi \in \Pi_g \right\}.$$

**Lemma A.3.** *Under Assumptions 1 and 5, for all $g = 2,\ldots,\mathcal{T}$, $t = 2,\ldots,\mathcal{T}$, the classes of functions $\mathcal{H}_{j,g}$, $j = \{1,2\ldots,5\}$, are Donsker.*

**Proof of Lemma A.3:** This follows from Example 19.7 in van der Vaart (1998).

**Lemma A.4.** *Under Assumptions 1 and 5, the null hypothesis*

$$H_0 : \mathbb{E}[Y_t - Y_{t-1}|X, G_g = 1] - \mathbb{E}[Y_t - Y_{t-1}|X, C = 1] = 0 \; a.s. \textit{for all } 2 \leq t < g \leq \mathcal{T},$$

*can be equivalently written as*

$$H_0 : \mathbb{E}\left[\left(\frac{G_g}{\mathbb{E}\left[G_g\right]} - \frac{\dfrac{p_g\left(X\right)C}{1 - p_g\left(X\right)}}{\mathbb{E}\left[\dfrac{p_g\left(X\right)C}{1 - p_g\left(X\right)}\right]}\right)\left(Y_t - Y_{t-1}\right)\middle|X\right] = 0 \ a.s. \text{for all } 2 \le t < g \le \mathcal{T}.$$

**Proof of Lemma A.4:** First note that

$$\mathbb{E}[Y_t - Y_{t-1}|X, G_g = 1] = \mathbb{E}[G_g\left(Y_t - Y_{t-1}\right)|X, G_g = 1]$$
$$= \mathbb{E}\left[\frac{G_g}{\mathbb{E}\left[G_g|X\right]}\left(Y_t - Y_{t-1}\right)\middle|X\right].$$

Analogously,

$$\mathbb{E}[Y_t - Y_{t-1}|X, C = 1] = \mathbb{E}\left[\frac{C}{\mathbb{E}\left[C|X\right]}\left(Y_t - Y_{t-1}\right)\middle|X\right],$$

implying that

$$\mathbb{E}[Y_t - Y_{t-1}|X, G_g = 1] - \mathbb{E}[Y_t - Y_{t-1}|X, C = 1] = 0 \ \ a.s. \text{for all } 2 \le t < g \le \mathcal{T}.$$
$$\Longleftrightarrow$$
$$\mathbb{E}\left[\left(\frac{G_g}{\mathbb{E}\left[G_g|X\right]} - \frac{C}{\mathbb{E}\left[C|X\right]}\right)\left(Y_t - Y_{t-1}\right)\middle|X\right] = 0 \ a.s. \text{for all } 2 \le t < g \le \mathcal{T}.$$

Given that under Assumptions 4 and 5, $\mathbb{E}\left[G_g + C|X\right] > 0 \ a.s.$, we have that

$$\mathbb{E}\left[\left(\frac{G_g}{\mathbb{E}\left[G_g|X\right]} - \frac{C}{\mathbb{E}\left[C|X\right]}\right)\left(Y_t - Y_{t-1}\right)\middle|X\right] = 0 \ a.s. \text{for all } 2 \le t < g \le \mathcal{T}$$

if and only if

$$\mathbb{E}\left[\mathbb{E}\left[G_g + C|X\right]\left(\frac{G_g}{\mathbb{E}\left[G_g|X\right]} - \frac{C}{\mathbb{E}\left[C|X\right]}\right)\left(Y_t - Y_{t-1}\right)\middle|X\right] = 0 \ a.s. \text{for all } 2 \le t < g \le \mathcal{T}. \quad (\text{A.3})$$

By noticing that

$$p_g\left(X\right) = \frac{\mathbb{E}\left[G_g|X\right]}{\mathbb{E}\left[G_g + C|X\right]}, \ 1 - p_g\left(X\right) = \frac{\mathbb{E}\left[C|X\right]}{\mathbb{E}\left[G_g + C|X\right]},$$

and that both of these are bounded away from zero under Assumption 5, we can rewrite (A.3) as

$$\mathbb{E}\left[\left(\frac{G_g}{\mathbb{E}\left[G_g\right]} - \frac{\dfrac{p_g\left(X\right)C}{1 - p_g\left(X\right)}}{\mathbb{E}\left[\dfrac{p_g\left(X\right)C}{1 - p_g\left(X\right)}\right]}\right)\left(Y_t - Y_{t-1}\right)\middle|X\right] = 0 \ a.s. \text{for all } 2 \le t < g \le \mathcal{T},$$

since

$$\mathbb{E}\left[\frac{p_g\left(X\right)C}{\left(1-p_g(X)\right)}\right] = \mathbb{E}\left[\frac{\mathbb{E}\left[G_g|X,C+G_g=1\right]C}{\mathbb{E}\left[C|X,C+G_g=1\right]}\right]$$

$$= \mathbb{E}\left[\frac{\mathbb{E}\left[G_g|X\right]C}{\mathbb{E}\left[C|X\right]}\right]$$

$$= \mathbb{E}\left[\frac{\mathbb{E}\left[G_g|X\right]\mathbb{E}\left[C|X\right]}{\mathbb{E}\left[C|X\right]}\right]$$

$$= \mathbb{E}\left[\mathbb{E}\left[G_g|X\right]\right]$$

$$= \mathbb{E}\left[G_g\right]. \tag{A.4}$$

This completes the proof.□

Now, we are ready to proceed with the proofs of our main theorems.

**Proof of Theorem 1:** Given the result in Lemma A.1,

$$ATT(g,t) = \mathbb{E}[ATT_X(g,t)|G_g=1]$$

$$= \mathbb{E}\left[\mathbb{E}[Y_t - Y_{g-1}|X,G_g=1] - \mathbb{E}[Y_t - Y_{g-1}|X,C=1]\Big|G_g=1\right]$$

$$:= \mathbb{E}[A_X|G_g=1] - \mathbb{E}[B_X|G_g=1],$$

and we consider each term separately. For the first term

$$\mathbb{E}[A_X|G_g=1] = \mathbb{E}[Y_t - Y_{g-1}|G_g=1]$$

$$= \mathbb{E}\left[\frac{G_g}{\mathbb{E}\left[G_g\right]}(Y_t - Y_{g-1})\right]. \tag{A.5}$$

For the second term, by repetition of the law of iterated expectations, we have

$$\mathbb{E}[B_X|G_g=1] = \mathbb{E}\left[\mathbb{E}[Y_t - Y_{g-1}|X,C=1]\Big|G_g=1\right]$$

$$= \mathbb{E}\left[G_g\mathbb{E}[C(Y_t - Y_{g-1})|X,C=1]\Big|G_g=1\right]$$

$$= \mathbb{E}\left[G_g\mathbb{E}\left[\frac{C}{(1-p_g(X))}(Y_t - Y_{g-1})\Big|X,G_g+C=1\right]\Big|G_g=1\right]$$

$$= \frac{\mathbb{E}\left[G_g\mathbb{E}\left[\frac{C}{(1-p_g(X))}(Y_t - Y_{g-1})\Big|X,G_g+C=1\right]\Big|G_g+C=1\right]}{\mathbb{E}\left[G_g|G_g+C=1\right]}$$

$$= \frac{\mathbb{E}\left[\mathbb{E}\left[\frac{p_g\left(X\right)C}{(1-p_g(X))}(Y_t - Y_{g-1})\Big|X,G_g+C=1\right]\Big|G_g+C=1\right]}{\mathbb{E}\left[G_g|G_g+C=1\right]}$$

$$= \mathbb{E}\left[G_g\right]^{-1}\mathbb{E}\left[\mathbb{E}\left[G_g+C|X\right]\mathbb{E}\left[\frac{p_g\left(X\right)C}{(1-p_g(X))}(Y_t - Y_{g-1})\Big|X,G_g+C=1\right]\right]$$

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$$= \mathbb{E}\left[G_g\right]^{-1} \mathbb{E}\left[\mathbb{E}\left[\frac{p_g\left(X\right)C}{\left(1 - p_g(X)\right)}(Y_t - Y_{g-1})\middle| X\right]\right]$$

$$= \mathbb{E}\left[G_g\right]^{-1} \mathbb{E}\left[\frac{p_g\left(X\right)C}{\left(1 - p_g(X)\right)}(Y_t - Y_{g-1})\right]$$

$$= \frac{\mathbb{E}\left[\dfrac{p_g\left(X\right)C}{\left(1 - p_g(X)\right)}(Y_t - Y_{g-1})\right]}{\mathbb{E}\left[\dfrac{p_g\left(X\right)C}{\left(1 - p_g(X)\right)}\right]}, \tag{A.6}$$

where (A.6) follows from (A.4). The proof is completed by combining (A.5) and (A.6). $\square$

**Proof of Theorem 2:** Remember that

$$\widehat{ATT}(g,t) = \mathbb{E}_n\left[\frac{G_g}{\mathbb{E}_n\left[G_g\right]}(Y_t - Y_{g-1})\right] - \mathbb{E}_n\left[\frac{\dfrac{\hat{p}_g\left(X\right)C}{1 - \hat{p}_g\left(X\right)}}{\mathbb{E}_n\left[\dfrac{\hat{p}_g\left(X\right)C}{1 - \hat{p}_g\left(X\right)}\right]}(Y_t - Y_{g-1})\right],$$

$$:= \widehat{ATT}_g(g,t) - \widehat{ATT}_C(g,t),$$

and

$$ATT(g,t) = \mathbb{E}\left[\frac{G_g}{\mathbb{E}\left[G_g\right]}(Y_t - Y_{g-1})\right] - \mathbb{E}\left[\frac{\dfrac{p_g\left(X\right)C}{1 - p_g\left(X\right)}}{\mathbb{E}\left[\dfrac{p_g\left(X\right)C}{1 - p_g\left(X\right)}\right]}(Y_t - Y_{g-1})\right]$$

$$:= ATT_g(g,t) - ATT_C(g,t).$$

In what follows we will separately show that, for $2 \leq g \leq t \leq \mathcal{T}$,

$$\sqrt{n}\left(\widehat{ATT}_g(g,t) - ATT_g(g,t)\right) = \frac{1}{\sqrt{n}}\sum_{i=1}^{n}\psi_{gt}^G(\mathcal{W}_i) + o_p\left(1\right), \tag{A.7}$$

and

$$\sqrt{n}\left(\widehat{ATT}_C(g,t) - ATT_C(g,t)\right) = \frac{1}{\sqrt{n}}\sum_{i=1}^{n}\psi_{gt}^C(\mathcal{W}_i) + o_p\left(1\right). \tag{A.8}$$

Then,

$$\sqrt{n}(\widehat{ATT}(g,t) - ATT(g,t)) = \frac{1}{\sqrt{n}}\sum_{i=1}^{n}\psi_{gt}(\mathcal{W}_i) + o_p(1)$$

hold from (A.7) and (A.8), and the asymptotic normality result follows from the application of the multivariate central limit theorem.

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Let $\beta_g = \mathbb{E}[G_g]$ and $\widehat{\beta}_g = \mathbb{E}_n[G_g]$, and note that

$$\sqrt{n}\left(\widehat{\beta}_g - \beta_g\right) = \frac{1}{\sqrt{n}}\sum_{i=1}^{n}(G_{ig} - \mathbb{E}[G_g]).$$

Then, for all $2 \le g \le t \le \mathcal{T}$, by the continuous mapping theorem,

$$
\begin{aligned}
\sqrt{n}\left(\widehat{ATT}_g(g,t) - ATT_g(g,t)\right) &= \frac{1}{\widehat{\beta}_g}\sqrt{n}\left(\mathbb{E}_n\left[G_g\left(Y_t - Y_{g-1}\right)\right] - \mathbb{E}\left[G_g\left(Y_t - Y_{g-1}\right)\right]\right) \\
&\quad - \mathbb{E}\left[G_g\left(Y_t - Y_{g-1}\right)\right]\sqrt{n}\left(\frac{1}{\beta_g} - \frac{1}{\widehat{\beta}_g}\right) \\
&= \frac{1}{\beta_g}\frac{1}{\sqrt{n}}\sum_{i=1}^{n}\left(G_{ig}\left(Y_{it} - Y_{ig-1}\right) - \mathbb{E}\left[G_g\left(Y_t - Y_{g-1}\right)\right]\right) \\
&\quad - \frac{\mathbb{E}\left[G_g\left(Y_t - Y_{g-1}\right)\right]}{\beta_g^2}\sqrt{n}\left(\widehat{\beta}_g - \beta_g\right) + o_p(1) \\
&= \frac{1}{\sqrt{n}}\sum_{i=1}^{n}\left(\frac{G_{ig}\left(Y_{it} - Y_{ig-1}\right)}{\beta_g} - \frac{G_{ig}\mathbb{E}\left[G_g\left(Y_t - Y_{g-1}\right)\right]}{\beta_g^2}\right) + o_p(1) \\
&= \frac{1}{\sqrt{n}}\sum_{i=1}^{n}\frac{G_{ig}\left(\left(Y_{it} - Y_{ig-1}\right) - ATT_g(g,t)\right)}{\beta_g} + o_p(1) \\
&:= \frac{1}{\sqrt{n}}\sum_{i=1}^{n}\psi_{gt}^{G}(\mathcal{W}_i) + o_p(1),
\end{aligned}
$$

concluding the proof of (A.7).

Next we focus on (A.8). For an arbitrary function $g$, let

$$w(g) = \frac{g(X)\,C}{1 - g(X)},$$

and note that

$$
\begin{aligned}
\sqrt{n}\left(\widehat{ATT}_C(g,t) - ATT_C(g,t)\right) &= \frac{1}{\mathbb{E}_n\left[w\left(\hat{p}_g\right)\right]}\sqrt{n}\left(\mathbb{E}_n\left[w\left(\hat{p}_g\right)\left(Y_t - Y_{g-1}\right)\right] - \mathbb{E}\left[w\left(p_g\right)\left(Y_t - Y_{g-1}\right)\right]\right) \\
&\quad - \frac{\mathbb{E}\left[w\left(p_g\right)\left(Y_t - Y_{g-1}\right)\right]}{\mathbb{E}_n\left[w\left(\hat{p}_g\right)\right]\mathbb{E}\left[w\left(p_g\right)\right]}\sqrt{n}\left(\mathbb{E}_n\left[w\left(\hat{p}_g\right)\right] - \mathbb{E}\left[w\left(p_g\right)\right]\right) \\
&:= \frac{1}{\mathbb{E}_n\left[w\left(\hat{p}_g\right)\right]}\cdot\sqrt{n}A_n\left(\hat{p}_g\right) - \frac{ATT_C(g,t)}{\mathbb{E}_n\left[w\left(\hat{p}_g\right)\right]}\cdot\sqrt{n}B_n\left(\hat{p}_g\right).
\end{aligned}
$$

From Assumption 5, Lemmas A.2 and A.3, and the continuous mapping theorem,

$$
\begin{aligned}
\frac{1}{\mathbb{E}_n\left[w\left(\hat{p}_g\right)\right]} &= \frac{1}{\mathbb{E}\left[w\left(p_g\right)\right]} + o_p(1), \\
\frac{ATT_C(g,t)}{\mathbb{E}_n\left[w\left(\hat{p}_g\right)\right]} &= \frac{ATT_C(g,t)}{\mathbb{E}\left[w\left(p_g\right)\right]} + o_p(1).
\end{aligned}
$$

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Thus,

$$\sqrt{n}\left(\widehat{ATT}_C(g,t) - ATT_C(g,t)\right) = \frac{1}{\mathbb{E}\left[w\left(p_g\right)\right]} \cdot \sqrt{n}A_n\left(\hat{p}_g\right)$$
$$- \frac{ATT_C(g,t)}{\mathbb{E}\left[w\left(p_g\right)\right]} \cdot \sqrt{n}B_n\left(\hat{p}_g\right) + o_p\left(1\right) \quad \text{(A.9)}$$

Applying a classical mean value theorem argument,

$$A_n\left(\hat{p}_g\right) = \mathbb{E}_n\left[w\left(p_g\right)\left(Y_t - Y_{g-1}\right)\right] - \mathbb{E}\left[w\left(p_g\right)\left(Y_t - Y_{g-1}\right)\right]$$
$$+ \mathbb{E}_n\left[X\left(\frac{C}{1 - p_g\left(X;\bar{\pi}_g\right)}\right)^2 \dot{p}_g\left(X;\bar{\pi}_g\right)\left(Y_{it} - Y_{ig-1}\right)\right]'\left(\hat{\pi}_g - \pi_g^0\right),$$

where $\bar{\pi}$ is an intermediate point that satisfies $\left|\bar{\pi}_g - \pi_g^0\right| \leq \left|\hat{\pi}_g - \pi_g^0\right|$ a.s. Thus, by Assumption 5, Lemmas A.2 and A.3, and the Classical Glivenko-Cantelli's theorem,

$$A_n\left(\hat{p}_g\right) = \mathbb{E}_n\left[w\left(p_g\right)\left(Y_t - Y_{g-1}\right) - \mathbb{E}\left[w\left(p_g\right)\left(Y_t - Y_{g-1}\right)\right]\right] \quad \text{(A.10)}$$
$$+ \mathbb{E}\left[X\left(\frac{C}{1 - p_g\left(X\right)}\right)^2 \dot{p}_g\left(X\right)\left(Y_{it} - Y_{ig-1}\right)\right]'\left(\hat{\pi}_g - \pi_g^0\right) + o_p\left(n^{-1/2}\right). \quad \text{(A.11)}$$

Analogously,

$$B_n\left(\hat{p}_g\right) = \mathbb{E}_n\left[w\left(p_g\right) - \mathbb{E}\left[w\left(p_g\right)\right]\right]$$
$$+ \mathbb{E}\left[X\left(\frac{C}{1 - p_g\left(X\right)}\right)^2 \dot{p}_g\left(X\right)\right]'\left(\hat{\pi}_g - \pi_g^0\right) + o_p\left(n^{-1/2}\right). \quad \text{(A.12)}$$

Then, (A.9), (A.10), (A.12) and Lemma A.2 yield (A.8), concluding the proof.$\square$

**Proof of Theorem 3:** Note that, by the conditional multiplier central limit theorem, see Lemma 2.9.5 in van der Vaart and Wellner (1996), as $n \to \infty$,

$$\frac{1}{\sqrt{n}}\sum_{i=1}^n V_i \cdot \Psi_{g \leq t}(\mathcal{W}_i) \overset{d}{\to} N(0, \Sigma), \quad \text{(A.13)}$$

where $\Sigma = \mathbb{E}[\Psi_{g \leq t}(\mathcal{W})\Psi_{g \leq t}(\mathcal{W})']$. Thus, to conclude the proof that

$$\sqrt{n}\left(\widehat{ATT}^*_{g \leq t} - \widehat{ATT}_{g \leq t}\right) \overset{d}{\underset{*}{\to}} N(0, \Sigma),$$

it suffices to show that, for all $2 \leq g \leq t \leq \mathcal{T}$,

$$\frac{1}{\sqrt{n}}\sum_{i=1}^n V_i \cdot \left[\widehat{\psi}_{gt}(\mathcal{W}_i) - \psi_{gt}(\mathcal{W}_i)\right] = o_{p^*}\left(1\right).$$

Towards this, note that

$$\frac{1}{\sqrt{n}} \sum_{i=1}^{n} V_i \cdot \left[\widehat{\psi}_{gt}(\mathcal{W}_i) - \psi_{gt}(\mathcal{W}_i)\right] = \frac{1}{\sqrt{n}} \sum_{i=1}^{n} V_i \cdot \left[\widehat{\psi}_{gt}^{G}(\mathcal{W}_i) - \psi_{gt}^{G}(\mathcal{W}_i)\right] \tag{A.14}$$
$$- \frac{1}{\sqrt{n}} \sum_{i=1}^{n} V_i \cdot \left[\widehat{\psi}_{gt}^{C}(\mathcal{W}_i) - \psi_{gt}^{C}(\mathcal{W}_i)\right],$$

where

$$\widehat{\psi}_{gt}^{G}(\mathcal{W}) = \frac{G_g}{\mathbb{E}_n[G_g]}\left[(Y_t - Y_{g-1}) - \widehat{ATT}_g(g,t)\right],$$

and

$$\widehat{\psi}_{gt}^{C}(\mathcal{W}) = \frac{w(\hat{p}_g)}{\mathbb{E}_n[w(\hat{p}_g)]}\left[(Y_{it} - Y_{ig-1}) - \widehat{ATT}_C(g,t)\right] + \widehat{M}_{gt}'\,\widehat{\xi}_g^{\pi}(\mathcal{W}),$$

with

$$w(\hat{p}_g) = \frac{\hat{p}_g(X)\,C}{1 - \hat{p}_g(X)},$$

$$\widehat{M}_{gt} = \frac{\mathbb{E}_n\left[X\left(\frac{C}{1-\hat{p}_g(X)}\right)^2 \widehat{p}_g(X)\left[(Y_{it} - Y_{ig-1}) - \widehat{ATT}_g(g,t)\right]\right]}{\mathbb{E}_n[w(\hat{p}_g)]},$$

$$\widehat{\xi}_g^{\pi}(\mathcal{W}) = \mathbb{E}_n\left[\frac{(G_g + C)\,\hat{p}_g(X)^2}{\hat{p}_g(X)(1 - \hat{p}_g(X))}XX'\right]^{-1} X\frac{(G_g + C)(G_g - \hat{p}_g(X))\,\widehat{p}_g(X)}{\hat{p}_g(X)(1 - \hat{p}_g(X))}.$$

We will show that each term in (A.14) is $o_{p^*}(1)$. For the first term in (A.14), we have

$$\frac{1}{\sqrt{n}} \sum_{i=1}^{n} V_i \cdot \left[\widehat{\psi}_{gt}^{G}(\mathcal{W}_i) - \psi_{gt}^{G}(\mathcal{W}_i)\right]$$
$$= \left[\frac{1}{\mathbb{E}_n[G_g]} - \frac{1}{\mathbb{E}[G_g]}\right]\frac{1}{\sqrt{n}}\sum_{i=1}^{n} V_i \cdot G_{ig}\left(Y_{it} - Y_{ig-1}\right)$$
$$- \left[\widehat{ATT}_g(g,t) - ATT_g(g,t)\right]\frac{1}{\sqrt{n}}\sum_{i=1}^{n} V_i \cdot G_{ig},$$
$$= o_{p^*}(1), \tag{A.15}$$

where the last equality follows from the results in Theorem 1, together with the law of large numbers, continuous mapping theorem, and Lemma 2.9.5 in van der Vaart and Wellner (1996).

For the second term in (A.14), we have

$$\frac{1}{\sqrt{n}} \sum_{i=1}^{n} V_i \cdot \left[\widehat{\psi}_{gt}^{C}(\mathcal{W}_i) - \psi_{gt}^{C}(\mathcal{W}_i)\right]$$
$$= \frac{1}{\mathbb{E}_n[w(\hat{p}_g)]}\frac{1}{\sqrt{n}}\sum_{i=1}^{n} V_i \cdot (w_i(\hat{p}_g) - w_i(p_g))(Y_{it} - Y_{ig-1})$$

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$$+ \left( \frac{1}{\mathbb{E}_n \left[ w \left( \hat{p}_g \right) \right]} - \frac{1}{\mathbb{E} \left[ w \left( p_g \right) \right]} \right) \frac{1}{\sqrt{n}} \sum_{i=1}^{n} V_i \cdot w_i \left( p_g \right) \left( Y_{it} - Y_{ig-1} \right)$$

$$+ \left( \widehat{M}_{gt} - M_{gt} \right) \frac{1}{\sqrt{n}} \sum_{i=1}^{n} V_i \cdot \xi_g^{\pi} \left( \mathcal{W}_i \right).$$

$$+ \widehat{M}_{gt} \frac{1}{\sqrt{n}} \sum_{i=1}^{n} V_i \cdot \left( \widehat{\xi}_g^{\pi} \left( \mathcal{W}_i \right) - \xi_g^{\pi} \left( \mathcal{W}_i \right) \right)$$

$$:= A_{1n} + A_{2n} + A_{3n} + A_{4n}.$$

From Lemma A.3, we have that $\mathcal{H}_{1,g}$, $\mathcal{H}_{2,g}$, $\mathcal{H}_{3,g}$ and $\mathcal{H}_{5,g}$ are Donsker, and by Assumption 5, $\mathbb{E} \left[ w \left( p_g \right) \right]$ it is bounded away from zero. Thus, by a stochastic equicontinuity argument, Glivenko-Cantelli's theorem, continuous mapping theorem, and Theorem 2.9.6 in van der Vaart and Wellner (1996),

$$A_{1n} = o_{p^*} \left( 1 \right), \ \ A_{2n} = o_{p^*} \left( 1 \right), \ \ A_{3n} = o_{p^*} \left( 1 \right), \ \text{and} \ A_{4n} = o_{p^*} \left( 1 \right),$$

implying that

$$\frac{1}{\sqrt{n}} \sum_{i=1}^{n} V_i \cdot \left[ \widehat{\psi}_{gt}^C (\mathcal{W}_i) - \psi_{gt}^C (\mathcal{W}_i) \right] = o_{p^*} \left( 1 \right). \tag{A.16}$$

From (A.13)-(A.16), it follows that

$$\sqrt{n} \left( \widehat{ATT}_{g \leq t}^* - \widehat{ATT}_{g \leq t} \right) \overset{d}{\underset{*}{\to}} N(0, \Sigma).$$

Finally, by the continuous mapping theorem, see e.g. Theorem 10.8 in Kosorok (2008), for any continuous functional $\Gamma(\cdot)$

$$\Gamma \left( \sqrt{n} \left( \widehat{ATT}_{g \leq t}^* - \widehat{ATT}_{g \leq t} \right) \right) \overset{d}{\underset{*}{\to}} \Gamma \left( N(0, V) \right),$$

concluding our proof. $\square$

**Proof of Theorem 4:** In order to prove the first part of Theorem 4, we first show that, under $H_0$, for all $2 \leq t < g \leq \mathcal{T}$,

$$\widehat{J}(u, g, t, \hat{p}_g) = \mathbb{E}_n \left[ \psi_{ugt}^{test} (\mathcal{W}_i) \right] + o_p \left( n^{-1/2} \right),$$

Towards this end, we write

$$\widehat{J}(u, g, t, \hat{p}_g) = \mathbb{E}_n \left[ \frac{G_g}{\mathbb{E}_n \left[ G_g \right]} 1(X \leq u) \left( Y_t - Y_{t-1} \right) \right]$$

$$- \mathbb{E}_n \left[ \frac{\dfrac{\hat{p}_g \left( X \right) C}{1 - \hat{p}_g \left( X \right)}}{\mathbb{E}_n \left[ \dfrac{\hat{p}_g \left( X \right) C}{1 - \hat{p}_g \left( X \right)} \right]} 1(X \leq u) \left( Y_t - Y_{t-1} \right) \right]$$

$$:= \widehat{J}_G(u, g, t, \hat{p}_g) - \widehat{J}_C(u, g, t, \hat{p}_g),$$

and analyze each term separately.

As in the proof of Theorem 1, let $\beta_g = \mathbb{E}[G_g]$ and $\widehat{\beta}_g = \mathbb{E}_n[G_g]$. Applying a classical mean value theorem argument, uniformly in $u \in \mathcal{X}$,

$$\widehat{J}_G(u, g, t, \hat{p}_g) = \mathbb{E}_n \left[ \frac{G_g}{\beta_g} 1(X \le u) (Y_t - Y_{t-1}) \right]$$
$$- \frac{\mathbb{E}_n [G_g 1(X \le u) (Y_t - Y_{t-1})]}{\bar{\beta}_g^2} \cdot \mathbb{E}_n [G_g - \mathbb{E}[G_g]].$$

where $\bar{\beta}_g$ is an intermediate point that satisfies $\left| \bar{\beta}_g - \beta_g \right| \le \left| \widehat{\beta}_g - \beta_g \right|$ a.s.. Define the class of functions

$$\mathcal{H}_{6,g} = \left\{ (x, g_g, y_t, y_{t-1}) \mapsto g_g (y_t - y_{t-1}) 1 \{x \le u\} : u \in \mathcal{X} \right\}.$$

By Example 19.11 in van der Vaart (1998), $\mathcal{H}_{6,g}$ is Donsker under Assumption 5. Furthermore,

$$\mathbb{E}_n [G_g - \mathbb{E}[G_g]] = O_p \left( n^{-1/2} \right).$$

Thus, by the Glivenko-Cantelli's theorem and the continuous mapping theorem, uniformly in $u \in \mathcal{X}$,

$$\widehat{J}_G(u, g, t, \hat{p}_g) = \mathbb{E}_n \left[ \frac{G_g}{\mathbb{E}[G_g]} 1(X \le u) (Y_t - Y_{t-1}) \right]$$
$$- \frac{J_G(u, g, t, p_g)}{\mathbb{E}[G_g]} \cdot \mathbb{E}_n [G_g - \mathbb{E}[G_g]] + o_p \left( n^{-1/2} \right)$$
$$= \mathbb{E}_n \left[ w_g^G \left( (Y_t - Y_{t-1}) 1(X \le u) - \mathbb{E} \left[ w_g^G 1(X \le u) (Y_t - Y_{t-1}) \right] \right) \right] + J_G(u, g, t, p_g)$$

(A.17)

$$+ o_p \left( n^{-1/2} \right),$$

where

$$J_G(u, g, t, p_g) = \mathbb{E} \left[ \frac{G_g}{\mathbb{E}[G_g]} 1(X \le u) (Y_t - Y_{t-1}) \right].$$

We analyze $\widehat{J}_C(u, g, t, \hat{p}_g)$ next. Applying a classical mean value theorem argument, uniformly in $u \in \mathcal{X}$,

$$\widehat{J}_C(u, g, t, \hat{p}_g) = \widehat{J}_C(u, g, t, p_g)$$
$$+ \frac{\mathbb{E}_n \left[ X \dfrac{C \, \dot{p}_g (X; \bar{\pi}_g)}{(1 - p_g (X; \bar{\pi}_g))^2} 1(X \le u) (Y_t - Y_{t-1}) \right]'}{\mathbb{E}_n \left[ \dfrac{p_g (X; \bar{\pi}_g) \, C}{1 - p_g (X; \bar{\pi}_g)} \right]} \left( \hat{\pi}_g - \pi_g^0 \right)$$

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$$-\frac{\mathbb{E}_n\left[X\dfrac{C\,\dot{p}_g\left(X;\bar{\pi}_g\right)}{\left(1-p_g\left(X;\bar{\pi}_g\right)\right)^2}\right]'}{\mathbb{E}_n\left[\dfrac{p_g\left(X;\bar{\pi}_g\right)C}{1-p_g\left(X;\bar{\pi}_g\right)}\right]}\frac{\mathbb{E}_n\left[\dfrac{p_g\left(X;\bar{\pi}_g\right)C}{1-p_g\left(X;\bar{\pi}_g\right)}1(X\le u)\left(Y_t-Y_{t-1}\right)\right]}{\mathbb{E}_n\left[\dfrac{p_g\left(X;\bar{\pi}_g\right)C}{1-p_g\left(X;\bar{\pi}_g\right)}\right]}\left(\hat{\pi}_g-\pi_g^0\right)$$

where $\bar{\pi}$ is an intermediate point that satisfies $\left|\bar{\pi}_g-\pi_g^0\right|\le\left|\hat{\pi}_g-\pi_g^0\right|$ *a.s*, and

$$\widehat{J}_C(u,g,t,p_g)=\frac{\mathbb{E}_n\left[\dfrac{p_g\left(X\right)C}{1-p_g\left(X\right)}1(X\le u)\left(Y_t-Y_{t-1}\right)\right]}{\mathbb{E}_n\left[\dfrac{p_g\left(X\right)C}{1-p_g\left(X\right)}\right]}.$$

Define the classes of functions

$$\mathcal{H}_{7,g}=\left\{(x,c,y_t,y_{t-1})\mapsto\frac{p_g\left(x;\pi\right)}{1-p_g\left(x;\pi\right)}c\left(y_t-y_{t-1}\right)1\left\{x\le u\right\}:\pi\in\Pi_g,u\in\mathcal{X}\right\},$$

$$\mathcal{H}_{8,g}=\left\{(x,c,y_t,y_{t-1})\mapsto x\frac{\dot{p}_g\left(x;\pi\right)c\left(y_t-y_{t-1}\right)1\left\{x\le u\right\}}{\left(1-p_g\left(x;\pi\right)\right)^2}:\pi\in\Pi_g,u\in\mathcal{X}\right\},$$

$$\mathcal{H}_{9,g}=\left\{(x,c)\mapsto\frac{cp_g\left(x;\pi\right)}{1-p_g\left(x;\pi\right)}:\pi\in\Pi_g,\right\},$$

$$\mathcal{H}_{10,g}=\left\{(x,c)\mapsto x\frac{\dot{p}_g\left(x;\pi\right)c}{\left(1-p_g\left(x;\pi\right)\right)^2}:\pi\in\Pi_g\right\}.$$

By Examples 19.7, 19.11, and 19.20 in van der Vaart (1998), all these classes of functions are Donsker under Assumption 5. Thus, by the Glivenko-Cantelli's theorem, continuous mapping theorem, and Lemma A.2, uniformly in $u\in\mathcal{X}$,

$$\widehat{J}_C(u,g,t,\hat{p}_g)=\widehat{J}_C(u,g,t,p_g)+M_{ugt}^{test}{}'\left(\hat{\pi}_g-\pi_g^0\right)+o_p\left(n^{-1/2}\right),\tag{A.18}$$

for every $g,t$.

Denote

$$\hat{\beta}_g^C=\mathbb{E}_n\left[\frac{p_g\left(X\right)C}{1-p_g\left(X\right)}\right],\ \ \beta_g^C=\mathbb{E}\left[\frac{p_g\left(X\right)C}{1-p_g\left(X\right)}\right].$$

Applying a classical mean value theorem argument, we have

$$\widehat{J}_C(u,g,t,p_g)=\frac{\mathbb{E}_n\left[\dfrac{p_g\left(X\right)C}{1-p_g\left(X\right)}1(X\le u)\left(Y_t-Y_{t-1}\right)\right]}{\mathbb{E}\left[\dfrac{p_g\left(X\right)C}{1-p_g\left(X\right)}\right]}$$

$$-\frac{\mathbb{E}_n\left[\dfrac{p_g\left(X\right)C}{1-p_g\left(X\right)}1(X\le u)\left(Y_t-Y_{t-1}\right)\right]}{\left(\bar{\beta}_g^C\right)^2}\cdot\mathbb{E}_n\left[\frac{p_g\left(X\right)C}{1-p_g\left(X\right)}-\mathbb{E}\left[\frac{p_g\left(X\right)C}{1-p_g\left(X\right)}\right]\right]$$

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where $\bar{\beta}_g^C$ is an intermediate point that satisfies $\left|\bar{\beta}_g^C - \beta_g^C\right| \leq \left|\widehat{\beta}_g^C - \beta_g^C\right|$ $a.s.$. Since $\mathcal{H}_{7,g}$ is a Donsker Class of functions and

$$\mathbb{E}_n\left[\frac{p_g(X)\,C}{1 - p_g(X)} - \mathbb{E}\left[\frac{p_g(X)\,C}{1 - p_g(X)}\right]\right] = O_p\left(n^{-1/2}\right),$$

we have that, by the Glivenko-Cantelli's theorem and the continuous mapping theorem, uniformly in $u \in \mathcal{X}$,

$$
\begin{aligned}
\widehat{J}_C(u, g, t, p_g) &= \mathbb{E}_n\left[w_g^C\,(Y_t - Y_{t-1})\,\mathbf{1}(X \leq u)\right] \\
&\quad - \frac{\mathbb{E}\left[w_g^C\,(Y_t - Y_{t-1})\,\mathbf{1}(X \leq u)\right]}{\mathbb{E}\left[\dfrac{p_g(X)\,C}{1 - p_g(X)}\right]} \cdot \mathbb{E}_n\left[\frac{p_g(X)\,C}{1 - p_g(X)} - \mathbb{E}\left[\frac{p_g(X)\,C}{1 - p_g(X)}\right]\right] + o_p\left(n^{-1/2}\right) \\
&= \mathbb{E}_n\left[w_g^C\,\left((Y_t - Y_{t-1})\,\mathbf{1}(X \leq u) - \mathbb{E}\left[w_g^C\mathbf{1}(X \leq u)\,(Y_t - Y_{t-1})\right]\right)\right] + J_C(u, g, t, p_g) \\
&\quad (A.19)
\end{aligned}
$$

$$+ o_p\left(n^{-1/2}\right).$$

Hence, from (A.17), (A.18), (A.19), and the asymptotic linear representation of $\left(\hat{\pi}_g - \pi_g^0\right)$ in Lemma A.2, for every $g, t$,

$$\widehat{J}(u, g, t, \hat{p}_g) = \mathbb{E}_n\left[\psi_{ugt}^{test}(\mathcal{W})\right] + \left(J_G(u, g, t, p_g) - J_C(u, g, t, p_g)\right) + o_p\left(n^{-1/2}\right) \quad (A.20)$$

By noticing that under $H_0$, $J_G(u, g, t, p_g) = J_C(u, g, t, p_g)$ for all $u \in \mathcal{X}$, $(g, t)$ such that $2 \leq t < g \leq \mathcal{T}$, we have that, under $H_0$, uniformly in $u \in \mathcal{X}$, for all $2 \leq t < g \leq \mathcal{T}$

$$\widehat{J}(u, g, t, \hat{p}_g) = \mathbb{E}_n\left[\psi_{ugt}^{test}(\mathcal{W}_i)\right] + o_p\left(n^{-1/2}\right).$$

In order to show that $\sqrt{n}\widehat{J}_{g>t}(u) \Rightarrow \mathbb{G}(u)$ $in$ $l^\infty(\mathcal{X})$, it suffices to show that the class of functions

$$\mathcal{H}_{10} = \left\{(x, g_g, c, y_t, y_{t-1}) \mapsto \psi_{ugt}^{test} : u \in \mathcal{X},\ 2 \leq t < g \leq \mathcal{T}\right\}$$

is Donsker. This follows straightforwardly from the previously discussed Donsker results and Example 19.20 in van der Vaart (1998). Finally,

$$CvM_n \xrightarrow{d} \int_{\mathcal{X}} |\mathbb{G}(u)|_M^2\, F_X(du)$$

follows from the continuous mapping theorem,

$$\sup_{u \in \mathcal{X}} |F_{n,X}(u) - F_X(u)| = o_{a.s.}(1),$$

and the Helly-Bray Theorem.

Next, we study the behavior of $CvM_n$ under $H_1$. First, note that under $H_1$, for some $u \in \mathcal{X}$,

and some $(g, t)$, $2 \leq t < g \leq \mathcal{T}$,

$$J(u, g, t, p_g) \neq 0 \,.$$

Thus, from (A.20), under $H_1$, uniformly in $u \in \mathcal{X}$,

$$\sqrt{n} \widehat{J}_{g>t}(u) = O_p\left(n^{1/2}\right),$$

implying that $CvM_n$ diverges to infinity under $H_1$. Because $c_\alpha^{CvM} = O(1)$ a.s., as $n \to \infty$,

$$P\left(CvM_n > c_\alpha^{CvM}\right) \to 1,$$

concluding the proof of Theorem 4. $\square$

**Proof of Theorem 5:** In the proof of Theorem 4, we have shown that

$$\mathcal{H}_{11} = \left\{(x, g_g, c, y_t, y_{t-1}) \mapsto \psi_{ugt}^{test} : u \in \mathcal{X}, \ 2 \leq t < g \leq \mathcal{T}\right\}$$

is a Donsker class of functions. Then, by the conditional multiplier functional central limit theorem, see Theorem 2.9.6, in van der Vaart and Wellner (1996), as $n \to \infty$,

$$\frac{1}{\sqrt{n}} \sum_{i=1}^n V_i \cdot \Psi_{g>t}^{test}(\mathcal{W}_i) \underset{*}{\Rightarrow} \mathbb{G}(u) \ in \ l^\infty(\mathcal{X}),$$

where $\mathbb{G}(u)$ *in* $l^\infty(\mathcal{X})$ is the same Gaussian process of Theorem 4 and $\underset{*}{\Rightarrow}$ indicates weak convergence in probability under the bootstrap law. Thus, to conclude the proof it suffices to show that, for all $2 \leq t < g \leq \mathcal{T}$, uniformly in $u \in \mathcal{X}$,

$$\frac{1}{\sqrt{n}} \sum_{i=1}^n V_i \cdot \left[\widehat{\psi}_{ugt}^{test}(\mathcal{W}_i) - \psi_{ugt}^{test}(\mathcal{W}_i)\right] = o_{p^*}(1). \tag{A.21}$$

The proof of (A.21) follows exactly the same steps as the proof of Theorem (3), and is therefore omitted.$\square$

# Appendix B: Additional Results for Repeated Cross Sections

In this section we extend our results to the case with repeated cross sections data instead of panel data. Here we assume that for each individual in the pooled sample, we observe $(Y, G_1, \ldots, G_\mathcal{T}, C, T, X)$ where $T \in \{1, \ldots, \mathcal{T}\}$ denotes the time period when that individual is observed. Let $T_t = 1$ if an observation is observed at time $t$, and zero otherwise.

We assume that random samples are available for each time period.

**Assumption B.1.** *Conditional of $T = t$, the data are independent and identically distributed from the distribution of $(Y_t, G_1, \ldots, G_\mathcal{T}, C, X)$, for all $t = 1, \ldots, \mathcal{T}$.*

Assumption B.1 implies that our sample consists of random draws from the mixture distribution

$$F_M(y, g_1, \ldots, g_\mathcal{T}, c, t, x) = \sum_{t=1}^{\mathcal{T}} \lambda_t \cdot F_{Y, G_1, \ldots, G_\mathcal{T}, C, X|T}(y, g_1, \ldots, g_\mathcal{T}, c, x|t),$$

where $\lambda_t = P(T_t = 1)$. Notice that, once one conditions on the time period, then expectations under the mixture distribution correspond to population expectations. Also, because $X$, $G_g$, and $C$ are observed for all individuals, one can use draws from the mixture distribution to estimate the generalized propensity score. With some abuse of notation, we then use $p_g(X)$ as a short notation for $\mathbb{E}_M[G_g|X, G_g + C = 1]$, where $\mathbb{E}_M[\cdot]$ denotes expectations with respect to $F_M(\cdot)$.

Define the stabilized weights

$$w_{treat}(a, b) = T_b \cdot G_a / \mathbb{E}_M[T_b \cdot G_a],$$
$$w_{cont}(a, b) = \frac{T_b \cdot p_a(X) C}{1 - p_a(X)} \bigg/ \mathbb{E}_M\left[\frac{T_b \cdot p_a(X) C}{1 - p_a(X)}\right],$$

where $a, b = 1, 2, \ldots \mathcal{T}$.

**Theorem B.1.** *Under Assumption B.1 and Assumptions 2-4 in the main text, for $2 \leq g \leq t \leq \mathcal{T}$, the group-time average treatment effect for group $g$ in period $t$ is nonparametrically identified, and given by*

$$ATT(g, t) = \mathbb{E}_M\left[(w_{treat}(g, t) - w_{treat}(g, g - 1)) \cdot Y\right] -$$
$$\mathbb{E}_M\left[(w_{cont}(g, t) - w_{cont}(g, g - 1)) \cdot Y\right].$$

**Proof of Theorem B.1**: By the law of iterated expectations, Assumption B.1 and Assumption 3 in the main text, for all $2 \leq g \leq t \leq \mathcal{T}$,

$$\mathbb{E}_M[w_{treat}(g, t) \cdot Y] = \frac{\mathbb{E}_M[T_t G_g \cdot Y]}{\mathbb{E}_M[T_t G_g]}$$
$$= \frac{\mathbb{E}[G_g \cdot Y|T_t = 1]}{\mathbb{E}[G_g|T_t = 1]}$$
$$= \mathbb{E}[Y|T_t = 1, G_g = 1]$$
$$= \mathbb{E}[Y_t(1)|G_g = 1].$$

To complete the proof of Theorem B.1, we must show that

$$\mathbb{E}_M\left[(w_{treat}(g, g - 1,) + w_{cont}(g, t) - w_{cont}(g, g - 1)) \cdot Y\right] = \mathbb{E}[Y_t(0)|G_g = 1]. \tag{B.1}$$

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Towards this, from Assumption B.1 and proceeding as in Lemma A.1, we get

$$
\begin{aligned}
\mathbb{E}[Y_t(0)|X, G_g = 1] ={}& \mathbb{E}[Y(0)|X, G_g = 1, T_t = 1] \\
={}& \mathbb{E}[Y|X, G_g = 1, T_{g-1} = 1] \\
&+ \mathbb{E}[Y|X, C = 1, T_t = 1] - \mathbb{E}[Y|X, C = 1, T_{g-1} = 1].
\end{aligned} \tag{B.2}
$$

From the above result, it follows that

$$
\begin{aligned}
\mathbb{E}[Y_t(0)|X, G_g = 1] ={}& \mathbb{E}\left[\mathbb{E}[Y|X, G_g = 1, T_{g-1} = 1]\,\middle|\, G_g = 1, T_{g-1} = 1\right] \\
&+ \mathbb{E}\left[\mathbb{E}[Y|X, C = 1, T_t = 1]\,\middle|\, G_g = 1, T_t = 1\right] \\
&- \mathbb{E}\left[\mathbb{E}[Y|X, C = 1, T_{g-1} = 1]\,\middle|\, G_g = 1, T_{g-1} = 1\right].
\end{aligned} \tag{B.3}
$$

We consider each term separately. For the first term of (B.3),

$$
\begin{aligned}
\mathbb{E}\left[\mathbb{E}[Y|X, G_g = 1, T_{g-1} = 1]\,\middle|\, G_g = 1, T_{g-1} = 1\right] &= \mathbb{E}[Y|G_g = 1, T_{g-1} = 1] \\
&= \mathbb{E}_M\left[w_{treat}\left(t, g\right) \cdot Y\right].
\end{aligned} \tag{B.4}
$$

Let $\mathbb{E}[Y|X, C = 1, T_t = 1] = A_{C=1, T_t=1}(X)$, and note that, by repeated application of the law of iterated expectations as in the proof of Theorem 1, we have that for the second term of (B.3),

$$
\begin{aligned}
\mathbb{E}\left[A_{C=1, T_t=1}(X)\,\middle|\, G_g = 1, T_t = 1\right] &= \mathbb{E}\left[G_g|T_t = 1\right]^{-1} \mathbb{E}\left[\frac{p_g(X)C}{(1 - p_g(X))}Y\,\middle|\, T_t = 1\right] \\
&= \mathbb{E}_M\left[G_g \cdot T_t\right]^{-1} \mathbb{E}_M\left[\frac{T_t \cdot p_g(X)C}{(1 - p_g(X))}Y\right] \\
&= \mathbb{E}_M\left[w_{cont}\left(g, t\right) \cdot Y\right],
\end{aligned} \tag{B.5}
$$

where the last equality follows from $p_g(X) := \mathbb{E}_M\left[G_g|X, G_g + C = 1\right]$, and

$$
\begin{aligned}
\mathbb{E}_M\left[\frac{T_t \cdot p_g(X)C}{(1 - p_g(X))}\right] &= \mathbb{E}_M\left[T_t \cdot \frac{\mathbb{E}_M\left[G_g|X, C + G_g = 1\right]C}{\mathbb{E}_M\left[C|X, C + G_g = 1\right]}\right] \\
&= \mathbb{E}_M\left[T_t \cdot \frac{\mathbb{E}_M\left[G_g|X\right]C}{\mathbb{E}_M\left[C|X\right]}\right] \\
&= \mathbb{E}_M\left[\frac{\mathbb{E}_M\left[G_g|X\right]\mathbb{E}_M\left[C|X\right]}{\mathbb{E}_M\left[C|X\right]}\right] \\
&= \mathbb{E}_M\left[T_t \cdot \mathbb{E}\left[G_g|X\right]\right] \\
&= \mathbb{E}_M\left[T_t \cdot G_g\right].
\end{aligned}
$$

Following analogous steps, we get that, for the third term of (B.3),

$$
\mathbb{E}\left[A_{C=1, T_{g-1}=1}(X)\,\middle|\, G_g = 1, T_{g-1} = 1\right] = \mathbb{E}_M\left[w_{cont}\left(g, g-1\right) \cdot Y\right]. \tag{B.6}
$$

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Then, (B.1) follows by combining (B.4), (B.5) and (B.6). The proof of Theorem B.1 is therefore completed. $\square$

The identification results in Theorem B.1 suggest a simple two-step estimation procedure for the $ATT(g,t)$ with repeated cross-section data. Similar to the panel data case discussed in the main text, we propose to estimate $ATT(g,t)$ by

$$\widehat{ATT}(g,t) = \mathbb{E}_n\left[(\widehat{w}_{treat}(g,t) - \widehat{w}_{treat}(g,g-1)) \cdot Y\right] -$$
$$\mathbb{E}_n\left[(\widehat{w}_{cont}(g,t;\hat{p}) - \widehat{w}_{cont}(g,g-1;\hat{p})) \cdot Y\right].$$

where $\hat{p}_g(\cdot)$ is an estimate of $p_g(\cdot)$, and for $a,b = 1,2,\ldots\mathcal{T}$,

$$\widehat{w}_{treat}(a,b) = T_b \cdot G_a / \mathbb{E}_n[T_b \cdot G_a],$$
$$\widehat{w}_{cont}(a,b;\hat{p}) = \frac{T_b \cdot \hat{p}_a(X) C}{1 - \hat{p}_a(X)} \bigg/ \mathbb{E}_n\left[\frac{T_b \cdot \hat{p}_a(X) C}{1 - \hat{p}_a(X)}\right].$$

Next, we show that $\widehat{ATT}(g,t)$ is $\sqrt{n}$-consistent, admits an asymptotically linear representation, and is asymptotically normal. These results are analogous to Theorem 2 in the main text. Let $ATT_{g\le t}$ and $\widehat{ATT}_{g\le t}$ denote the vector of $ATT(g,t)$ and $\widehat{ATT}(g,t)$, respectively, for all $g = 2,\ldots,\mathcal{T}$ and $t = 2,\ldots,\mathcal{T}$ with $g \le t$. Define

$$\psi_{g,t}^{rc}(\mathcal{W}_i) = \left(\psi_{g,t}^{rc,G}(\mathcal{W}_i) - \psi_{g,g-1}^{rc,G}(\mathcal{W}_i)\right) - \left(\psi_{g,t}^{rc,C}(\mathcal{W}_i) - \psi_{g,g-1}^{rc,C}(\mathcal{W}_i)\right),$$

where, for $g,t = 1,2,\ldots\mathcal{T}$,

$$\psi_{g,t}^{rc,G}(\mathcal{W}) = w_{treat}(g,t)\ \left[Y - \mathbb{E}_M[w_{treat}(g,t) \cdot Y]\right],$$
$$\psi_{g,t}^{rc,C}(\mathcal{W}) = w_{cont}(g,t)\ \left[Y - \mathbb{E}_M[w_{cont}(g,t) \cdot Y]\right] + M_{g,t}^{rc}{}'\ \xi_g^{\pi}(\mathcal{W}),$$

and

$$M_{g,t}^{rc} = \frac{\mathbb{E}_M\left[X\left(\frac{T_t \cdot C}{1 - p_g(X)}\right)^2 \dot{p}_g(X) \cdot [Y - \mathbb{E}[w_{cont}(g,t) \cdot Y]]\right]}{\mathbb{E}_M\left[\frac{T_t \cdot p_g(X) C}{1 - p_g(X)}\right]},$$

which is a $k \times 1$ vector, with $k$ the dimension of $X$, and $\xi_g^{\pi}(\mathcal{W})$ is as defined in (3.1) in the main text. Finally, let $\Psi_{g\le t}^{rc}$ denote the collection of $\psi_{g,t}^{rc}$ across all periods $t$ and groups $g$ such that $g \le t$.

**Theorem B.2.** *Under Assumption B.1 and Assumptions 2-5 in the main text, for $2 \le g \le t \le \mathcal{T}$,*

$$\sqrt{n}(\widehat{ATT}(g,t) - ATT(g,t)) = \frac{1}{\sqrt{n}}\sum_{i=1}^{n}\psi_{g,t}^{rc}(\mathcal{W}_i) + o_p(1).$$

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*Furthermore,*

$$\sqrt{n}(\widehat{ATT}_{g\leq t} - ATT_{g\leq t}) \xrightarrow{d} N(0, \Sigma^{rc})$$

*where* $\Sigma^{rc} = \mathbb{E}_M[\Psi^{rc}_{g\leq t}(\mathcal{W})\Psi^{rc}_{g\leq t}(\mathcal{W})']$.

**Proof of Theorem B.2:** The proof of Theorem B.2 follows the same steps as the Proof of Theorem 2. From Theorem B.1, for each $2 \leq g \leq t \leq \mathcal{T}$ we can write

$$
\begin{aligned}
\sqrt{n}(\widehat{ATT}\,(g,t) &- ATT\,(g,t)) \\
&= \sqrt{n}\,(\mathbb{E}_n\,[\widehat{w}_{treat}\,(g,t)\cdot Y] - \mathbb{E}_M\,[w_{treat}\,(g,t)\cdot Y]) \\
&\quad - \sqrt{n}\,(\mathbb{E}_n\,[\widehat{w}_{treat}\,(g,g-1)\cdot Y] - \mathbb{E}_M\,[w_{treat}\,(g,g-1)\cdot Y]) \\
&\quad - \sqrt{n}\,(\mathbb{E}_n\,[\widehat{w}_{cont}\,(g,t;\hat{p})\cdot Y] - \mathbb{E}_M\,[w_{cont}\,(g,t)\cdot Y]) \\
&\quad + \sqrt{n}\,(\mathbb{E}_n\,[\widehat{w}_{cont}\,(g,g-1;\hat{p})\cdot Y] - \mathbb{E}_M\,[w_{cont}\,(g,g-1)\cdot Y])\,.
\end{aligned} \tag{B.7}
$$

We analyze each term separately. First, note that, for each $2 \leq g \leq t \leq \mathcal{T}$,

$$\sqrt{n}\,(\mathbb{E}_n\,[T_t \cdot G_g] - \mathbb{E}_M\,[T_t \cdot G_g]) = \frac{1}{\sqrt{n}}\sum_{i=1}^{n}(T_{it} \cdot G_{ig} - \mathbb{E}\,[T_t \cdot G_g])\,.$$

Then, by the continuous mapping theorem,

$$\sqrt{n}\,(\mathbb{E}_n\,[\widehat{w}_{treat}\,(g,t)\cdot Y] - \mathbb{E}_M\,[w_{treat}\,(g,t)\cdot Y]) = \frac{1}{\sqrt{n}}\sum_{i=1}^{n}\psi^{rc,G}_{g,t}(\mathcal{W}_i) + o_p\,(1)\,. \tag{B.8}$$

Analogously,

$$\sqrt{n}\,(\mathbb{E}_n\,[\widehat{w}_{treat}\,(g,g-1)\cdot Y] - \mathbb{E}_M\,[w_{treat}\,(g,g-1)\cdot Y]) = \frac{1}{\sqrt{n}}\sum_{i=1}^{n}\psi^{rc,G}_{g,g-1}(\mathcal{W}_i) + o_p\,(1)\,. \tag{B.9}$$

Next we focus on $\sqrt{n}\,(\mathbb{E}_n\,[\widehat{w}_{cont}\,(g,t;\hat{p})\cdot Y] - \mathbb{E}_M\,[w_{cont}\,(g,t)\cdot Y])$. To simplify notation, write

$$w_{a,b}\,(p) = \frac{T_b \cdot p_a\,(X)\,C}{1 - p_a\,(X)},$$

and note that $\widehat{w}_{cont}\,(g,t;\hat{p}) = w_{g,t}\,(\hat{p})\,/\mathbb{E}_n\,[w_{g,t}\,(\hat{p})]$ and $w_{cont}\,(g,t;p) = w_{g,t}\,(p)\,/\mathbb{E}_M\,[w_{g,t}\,(p)]$. Then,

$$
\begin{aligned}
\sqrt{n}\,(\mathbb{E}_n\,[\widehat{w}_{cont}\,&(g,t;\hat{p})\cdot Y] - \mathbb{E}_M\,[w_{cont}\,(g,t)\cdot Y]) \\
&= \frac{1}{\mathbb{E}_n\,[w_{g,t}\,(\hat{p})]}\sqrt{n}\,(\mathbb{E}_n\,[w_{g,t}\,(\hat{p})\cdot Y] - \mathbb{E}_M\,[w_{g,t}\,(p)\cdot Y]) \\
&\quad - \frac{\mathbb{E}_M\,[w_{g,t}\,(p)\cdot Y]}{\mathbb{E}_n\,[w_{g,t}\,(\hat{p})]\,\mathbb{E}_M\,[w_{g,t}\,(p)]}\sqrt{n}\,(\mathbb{E}_n\,[w_{g,t}\,(\hat{p})] - \mathbb{E}_M\,[w_{g,t}\,(p)])
\end{aligned}
$$

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$$:= \frac{1}{\mathbb{E}_n \left[ w_{g,t} \left( \hat{p} \right) \right]} \cdot \sqrt{n} A^{rc}_{n,\ g,t} \left( \hat{p}_g \right) - \frac{\mathbb{E}_M \left[ w_{cont} \left( g, t \right) \cdot Y \right]}{\mathbb{E}_n \left[ w_{g,t} \left( \hat{p} \right) \right]} \cdot \sqrt{n} B^{rc}_{n,\ g,t} \left( \hat{p}_g \right).$$

From Assumption 5, Lemmas A.2 and A.3, and the continuous mapping theorem,

$$\frac{1}{\mathbb{E}_n \left[ w_{g,t} \left( \hat{p} \right) \right]} = \frac{1}{\mathbb{E}_M \left[ w_{g,t} \left( p \right) \right]} + o_p \left( 1 \right),$$

$$\frac{\mathbb{E}_M \left[ w_{cont} \left( g, t \right) \cdot Y \right]}{\mathbb{E}_n \left[ w_{g,t} \left( \hat{p} \right) \right]} = \frac{\mathbb{E}_M \left[ w_{cont} \left( g, t \right) \cdot Y \right]}{\mathbb{E}_M \left[ w_{g,t} \left( p \right) \right]} + o_p \left( 1 \right).$$

Thus,

$$\sqrt{n} \left( \mathbb{E}_n \left[ \widehat{w}_{cont} \left( g, t; \hat{p} \right) \cdot Y \right] - \mathbb{E}_M \left[ w_{cont} \left( g, t \right) \cdot Y \right] \right)$$

$$= \frac{1}{\mathbb{E}_M \left[ w_{g,t} \left( p \right) \right]} \cdot \sqrt{n} A^{rc}_{n,\ g,t} \left( \hat{p}_g \right)$$

$$- \frac{\mathbb{E}_M \left[ w_{cont} \left( g, t \right) \cdot Y \right]}{\mathbb{E}_M \left[ w_{g,t} \left( p \right) \right]} \cdot \sqrt{n} B^{rc}_{n,\ g,t} \left( \hat{p}_g \right) + o_p \left( 1 \right) \quad \text{(B.10)}$$

Applying a classical mean value theorem argument,

$$A^{rc}_{n,\ g,t} \left( \hat{p}_g \right) = \mathbb{E}_n \left[ w_{g,t} \left( p \right) \cdot Y \right] - \mathbb{E}_M \left[ w_{g,t} \left( p \right) \cdot Y \right]$$

$$+ \mathbb{E}_n \left[ X \left( \frac{T_t \cdot C}{1 - p_g \left( X; \bar{\pi}_g \right)} \right)^2 \dot{p}_g \left( X; \bar{\pi}_g \right) \cdot Y \right]' \left( \hat{\pi}_g - \pi^0_g \right),$$

where $\bar{\pi}$ is an intermediate point that satisfies $\left| \bar{\pi}_g - \pi^0_g \right| \leq \left| \hat{\pi}_g - \pi^0_g \right|$ *a.s.* Thus, by Assumption 5, Lemmas A.2 and A.3, and the Glivenko-Cantelli's theorem,

$$A^{rc}_{n,\ g,t} \left( \hat{p}_g \right) = \mathbb{E}_n \left[ w_{g,t} \left( p \right) Y - \mathbb{E}_M \left[ w_{g,t} \left( p \right) Y \right] \right]$$

$$+ \mathbb{E}_M \left[ X \left( \frac{T_t \cdot C}{1 - p_g \left( X \right)} \right)^2 \dot{p}_g \left( X \right) \cdot Y \right]' \left( \hat{\pi}_g - \pi^0_g \right) + o_p \left( n^{-1/2} \right). \quad \text{(B.11)}$$

Analogously,

$$B_n \left( \hat{p}_g \right) = \mathbb{E}_n \left[ w_{g,t} \left( p \right) - \mathbb{E}_M \left[ w_{g,t} \left( p \right) \right] \right]$$

$$+ \mathbb{E}_M \left[ X \left( \frac{T_t \cdot C}{1 - p_g \left( X \right)} \right)^2 \dot{p}_g \left( X \right) \right]' \left( \hat{\pi}_g - \pi^0_g \right) + o_p \left( n^{-1/2} \right). \quad \text{(B.12)}$$

Combining (B.10), (B.11), (B.12) with Lemma A.2 yield

$$\sqrt{n} \left( \mathbb{E}_n \left[ \widehat{w}_{cont} \left( g, t; \hat{p} \right) \cdot Y \right] - \mathbb{E}_M \left[ w_{cont} \left( g, t \right) \cdot Y \right] \right) = \frac{1}{\sqrt{n}} \sum_{i=1}^{n} \psi^{rc,C}_{g,t} \left( \mathcal{W}_i \right) + o_p \left( 1 \right). \quad \text{(B.13)}$$

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Using the same arguments, we conclude that

$$\sqrt{n}\left(\mathbb{E}_n\left[\widehat{w}_{cont}\left(g,g-1;\hat{p}\right)\cdot Y\right] - \mathbb{E}_M\left[w_{cont}\left(g,g-1\right)\cdot Y\right]\right) = \frac{1}{\sqrt{n}}\sum_{i=1}^{n}\psi_{g,g-1}^{rc,C}(\mathcal{W}_i) + o_p\left(1\right). \quad (B.14)$$

Hence, from (B.7), (B.8), (B.9), (B.13) and (B.14), we conclude that, for each $2 \le g \le t \le \mathcal{T}$,

$$\sqrt{n}(\widehat{ATT}\left(g,t\right) - ATT\left(g,t\right)) = \frac{1}{\sqrt{n}}\sum_{i=1}^{n}\psi_{g,t}^{rc}(\mathcal{W}_i) + o_p(1).$$

The proof is then completed by applying the multivariate central limit theorem.□

Based on the above results, one can conclude that estimation and inference procedures for $ATT\left(g,t\right)$ in the case of repeated cross sections is similar to what we did in the case with panel data. In fact, one simply needs to adjust the weights slightly. In order to conduct asymptotically valid simultaneous inference, one can leverage on the asymptotic linear representation in Theorem B.2, and use a multiplier bootstrap procedure analogous to the one in Theorem 3. The proof of the bootstrap validity in the repeated cross section case follows exactly the same steps as in Theorem 3 and is therefore omitted.

# Appendix C: Analysis with "Not yet Treated" as a Control Group

In this appendix, we discuss the case where one considers the "not yet treated" instead of the "never treated" as a control group. This case is particularly relevant in applications when eventually (almost) all units are treated, though the timing of the treatment differs across groups. To carry this analysis, we make the following assumptions.

**Assumption C.1.** $\{Y_{i1}, Y_{i2}, \ldots Y_{i\mathcal{T}}, X_i, D_{i1}, D_{i2}, \ldots, D_{i\mathcal{T}}\}_{i=1}^{n}$ *is independent and identically distributed (iid).*

**Assumption C.2.** *For all $t = 2, \ldots, \mathcal{T}$, $g = 2, \ldots, \mathcal{T}$ such that $g \le t$,*

$$\mathbb{E}[Y_t(0) - Y_{t-1}(0)|X, G_g = 1] = \mathbb{E}[Y_t(0) - Y_{t-1}(0)|X, D_t = 0] \ a.s..$$

**Assumption C.3.** *For $t = 2, \ldots, \mathcal{T}$,*

$$D_{t-1} = 1 \ implies \ that \ D_t = 1$$

**Assumption C.4.** *For all $t = 2, \ldots, \mathcal{T}$, $g = 2, \ldots, \mathcal{T}$, $P\left(G_g = 1\right) > 0$ and $P\left(D_t = 1|X\right) < 1$ a.s..*

Assumptions C.1 and C.3 are the same as Assumptions 1 and 3 in the main text. Assumptions C.2 and C.4 are the analogue of Assumptions 2 and 4, but using those "not yet treated" ($D_t = 0$) as a control group instead of the "never treated" ($C = 0$ or $D_{\mathcal{T}} = 0$). Note that Assumption C.4 rules out the case in which eventually everyone is treated; in these time periods, there is no "control group" available, and therefore the data itself is not informative about the average treatment effect when $D_t = 1$ $a.s.$. In these cases, one should concentrate their attention only to the time periods such that $P(D_t = 1|X) < 1$ $a.s.$.

Remember that

$$ATT_X(g, t) = \mathbb{E}[Y_t(1) - Y_t(0)|X, G_g = 1].$$

Next lemma states that, under Assumptions C.1-C.4, we can identify $ATT_X(g,t)$ for $2 \leq g \leq t \leq \mathcal{T}$. This is the analogue of Lemma A.1.

**Lemma C.1.** *Under Assumptions C.1-C.4, and for $2 \leq g \leq t \leq \mathcal{T}$,*

$$ATT_X(g, t) = \mathbb{E}[Y_t - Y_{g-1}|X, G_g = 1] - \mathbb{E}[Y_t - Y_{g-1}|X, D_t = 0] \ a.s..$$

**Proof of Lemma C.1:** In what follows, take all equalities to hold almost surely (a.s.). Notice that for identifying $ATT_X(g,t)$, the key term is $E[Y_t(0)|X, G_g = 1]$. And notice that for $h > s$, $E[Y_s(0)|X, G_s = 1] = E[Y_s|X, G_h = 1]$, which holds because in time periods before an individual is first treated, their untreated potential outcomes are observed outcomes. Also, note that, for $2 \leq g \leq t \leq \mathcal{T}$,

$$\mathbb{E}[Y_t(0)|X, G_g = 1] = \mathbb{E}[\Delta Y_t(0)|X, G_g = 1] + \mathbb{E}[Y_{t-1}(0)|X, G_g = 1] \quad \text{(C.1)}$$
$$= \mathbb{E}[\Delta Y_t|X, D_t = 0] + \mathbb{E}[Y_{t-1}(0)|X, G_g = 1],$$

where the first equality holds by adding and subtracting $E[Y_{t-1}(0)|X, G_g = 1]$ and the second equality holds by Assumption C.2. If $g = t-1$, then the last term in the final equation is identified; otherwise, one can continue recursively in similar way to (C.1) but starting with $\mathbb{E}[Y_{t-1}(0)|X, G_g = 1]$. As a result,

$$\mathbb{E}[Y_t(0)|X, G_g = 1] = \sum_{j=0}^{t-g} \mathbb{E}[\Delta Y_{t-j}|X, D_t = 0] + \mathbb{E}[Y_{g-1}|X, G_g = 1]$$
$$= \mathbb{E}[Y_t - Y_{g-1}|X, D_t = 0] + \mathbb{E}[Y_{g-1}|X, G_g = 1]. \quad \text{(C.2)}$$

Combining (C.2) with the fact that, for all $g \leq t$, $\mathbb{E}[Y_t(1)|X, G_g = 1] = \mathbb{E}[Y_t|X, G_g = 1]$ (which holds because observed outcomes for group $g$ in period $t$ with $g \leq t$ are treated potential outcomes), implies the result. $\square$

With the result of Lemma C.1 at hands, we proceed to show that the $ATT(g,t)$ is nonparametrically identified under Assumptions C.1 - C.4 and for $2 \leq g \leq t \leq \mathcal{T}$. The following Theorem

is the analogue Theorem 1 .

**Theorem C.1.** *Under Assumptions C.1 - C.4 and for $2 \leq g \leq t \leq \mathcal{T}$, the group-time average treatment effect for group g in period t is nonparametrically identified, and given by*

$$ATT\left(g,t\right) = \mathbb{E}\left[\left(\frac{G_g}{\mathbb{E}\left[G_g\right]} - \frac{\dfrac{P\left(G_g = 1|X\right)\left(1 - D_t\right)}{1 - P\left(D_t = 1|X\right)}}{\mathbb{E}\left[\dfrac{P\left(G_g = 1|X\right)\left(1 - D_t\right)}{1 - P\left(D_t = 1|X\right)}\right]}\right)\left(Y_t - Y_{g-1}\right)\right]. \quad (\text{C.3})$$

**Proof of Theorem C.1**: Given the result in Lemma C.1,

$$
\begin{aligned}
ATT(g,t) &= \mathbb{E}[ATT_X(g,t)|G_g = 1] \\
&= \mathbb{E}\left[\mathbb{E}[Y_t - Y_{g-1}|X, G_g = 1] - \mathbb{E}[Y_t - Y_{g-1}|X, D_t = 0]\Big|G_g = 1\right] \\
&:= \mathbb{E}[A_X|G_g = 1] - \mathbb{E}[B_X^{n.yet}|G_g = 1],
\end{aligned}
$$

and we consider each term separately. For the first term

$$
\begin{aligned}
\mathbb{E}[A_X|G_g = 1] &= \mathbb{E}[Y_t - Y_{g-1}|G_g = 1] \\
&= \mathbb{E}\left[\frac{G_g}{\mathbb{E}\left[G_g\right]}(Y_t - Y_{g-1})\right]. \quad (\text{C.4})
\end{aligned}
$$

For the second term, by repetition of the law of iterated expectations, we have

$$
\begin{aligned}
\mathbb{E}[B_X^{n.yet}|G_g = 1] &= \mathbb{E}\left[\mathbb{E}[Y_t - Y_{g-1}|X, D_t = 0]\Big|G_g = 1\right] \\
&= \mathbb{E}\left[G_g \mathbb{E}[(1 - D_t)(Y_t - Y_{g-1})|X, D_t = 0]\Big|G_g = 1\right] \\
&= \mathbb{E}\left[G_g \mathbb{E}\left[\frac{(1 - D_t)}{1 - P\left(D_t = 1|X\right)}(Y_t - Y_{g-1})\Big|X\right]\Big|G_g = 1\right] \\
&= \mathbb{E}\left[G_g\right]^{-1}\mathbb{E}\left[G_g \mathbb{E}\left[\frac{(1 - D_t)}{1 - P\left(D_t = 1|X\right)}(Y_t - Y_{g-1})\Big|X\right]\right] \\
&= \mathbb{E}\left[G_g\right]^{-1}\mathbb{E}\left[\mathbb{E}\left[G_g|X\right]\mathbb{E}\left[\frac{(1 - D_t)}{1 - P\left(D_t = 1|X\right)}(Y_t - Y_{g-1})\Big|X\right]\right] \\
&= \mathbb{E}\left[G_g\right]^{-1}\mathbb{E}\left[\mathbb{E}\left[\frac{P\left(G_g = 1|X\right)\left(1 - D_t\right)}{1 - P\left(D_t = 1|X\right)}(Y_t - Y_{g-1})\Big|X\right]\right] \\
&= \mathbb{E}\left[G_g\right]^{-1}\mathbb{E}\left[\frac{P\left(G_g = 1|X\right)\left(1 - D_t\right)}{1 - P\left(D_t = 1|X\right)}(Y_t - Y_{g-1})\right] \\
&= \frac{\mathbb{E}\left[\dfrac{P\left(G_g = 1|X\right)\left(1 - D_t\right)}{1 - P\left(D_t = 1|X\right)}(Y_t - Y_{g-1})\right]}{\mathbb{E}\left[\dfrac{P\left(G_g = 1|X\right)\left(1 - D_t\right)}{1 - P\left(D_t = 1|X\right)}\right]}, \quad (\text{C.5})
\end{aligned}
$$

where (C.5) follows from

$$\mathbb{E}\left[\frac{P\left(G_g = 1|X\right)\left(1 - D_t\right)}{1 - P\left(D_t = 1|X\right)}\right] = \mathbb{E}\left[\mathbb{E}\left[\frac{P\left(G_g = 1|X\right)\left(1 - D_t\right)}{1 - P\left(D_t = 1|X\right)}|X\right]\right]$$

$$= \mathbb{E}\left[\frac{P\left(G_g = 1|X\right)}{1 - P\left(D_t = 1|X\right)}\mathbb{E}\left[\left(1 - D_t\right)|X\right]\right]$$

$$= \mathbb{E}\left[\frac{P\left(G_g = 1|X\right)}{1 - P\left(D_t = 1|X\right)}\left(1 - P\left(D_t = 1|X\right)\right)\right]$$

$$= \mathbb{E}\left[P\left(G_g = 1|X\right)\right]$$

$$= \mathbb{E}\left[\mathbb{E}\left[G_g|X\right]\right]$$

$$= \mathbb{E}\left[G_g\right].$$

The proof is completed by combining (C.4) and (C.5). □

Once we have establish nonparametric identification of $ATT(g,t)$, we can follow a similar two-step estimation strategy as described in Section 3 . More precisely, under Assumptions C.1 - C.4 and for $2 \le g \le t \le \mathcal{T}$, one can estimate $ATT(g,t)$ by

$$\widehat{ATT}_{n.yet}(g,t) = \mathbb{E}_n\left[\left(\frac{G_g}{\mathbb{E}_n\left[G_g\right]} - \frac{\frac{\hat{p}_{G_g}\left(X\right)\left(1 - D_t\right)}{1 - \hat{p}_{D_t}\left(X\right)}}{\mathbb{E}_n\left[\frac{\hat{p}_{G_g}\left(X\right)\left(1 - D_t\right)}{1 - \hat{p}_{D_t}\left(X\right)}\right]}\right)\left(Y_t - Y_{g-1}\right)\right],$$

where $\hat{p}_{G_g}\left(X\right)$ is an estimate of $P\left(G_g = 1|X\right)$, and $\hat{p}_{D_t}\left(X\right)$ is an estimate of $P\left(D_t = 1|X\right)$. In contrast to the case analyzed in the main text, here we need to estimate two propensity scores. These can be estimated separately, using binary choice models (e.g. logit), or jointly, using multinomial choice models (e.g. multinomial logit).

Following similar steps as in Theorems 2 and 3, one can show that under suitable regularity conditions akin to those in Assumption 5, $\widehat{ATT}_{n.yet}(g,t)$ is consistent and asymptotically normal, and that one can use a multiplier bootstrap similar to the one described in Algorithm 1 to conduct asymptotically valid inference. Nonetheless, it is worth mentioning that the asymptotic linear representation of $\widehat{ATT}_{n.yet}(g,t)$ will be different from that of $\widehat{ATT}(g,t)$, because the former is based on of two different propensity scores whereas the later is based only on one. A detailed and formal derivation of the aforementioned results is beyond the scope of this article.

# Appendix D: Additional Results for the Case without Covariates

## Panel Data

The case where the DID assumption holds without conditioning on covariates is of particular interest. In this appendix, we briefly consider whether or not it is possible to obtain $ATT(g,t)$ using a regression approach like the two period - two group case. A natural starting point is the model

$$Y_{igt} = \alpha_t + c_g + \gamma_{gt}G_{igt} + u_{igt}$$

where $\alpha_t$ is a vector of time period fixed effects (we normalize $\alpha_1$ to be equal to zero and $\gamma_{g1}$ to be equal to 1), $c_g$ is time invariant unobserved heterogeneity that can be distributed differently across groups, and $G_{igt}$ is is a dummy variable indicating whether or not individual $i$ is a member group $g$ and the time period is $t$. Differencing the model across time periods results in

$$\Delta Y_{igt} = \tilde{\alpha}_t + \gamma_{gt}G_{igt} + \Delta u_{igt},$$

where $\tilde{\alpha}_t = \alpha_t - \alpha_{t-1}$. Notice that this is a fully saturated model in group and time effects. It is straightforward to show that

$$\gamma_{gt} = E[\Delta Y_t | G_g = 1] - E[\Delta Y_t | C = 1].$$

When $g = t$, this is exactly the DID estimator. Under the augmented unconditional version of the parallel trends assumption, $\gamma_{gt}$ should be equal to 0 for all $g > t$, and it is straightforward to test this using output from standard regression software (e.g. Wald test). For $t > g$, the long difference estimate of $ATT(g,t)$ can be constructed by

$$
\begin{aligned}
ATT(g,t) &= E[Y_t - Y_{g-1} | G_g = 1] - E[Y_t - Y_{g-1} | C = 1] \\
&= \sum_{s=g}^{t} \left( E[\Delta Y_s | G_g = 1] - E[\Delta Y_s | C = 1] \right) \\
&= \sum_{s=g}^{t} \gamma_{gs}
\end{aligned}
$$

This implies that, under the (augmented) unconditional parallel trends assumption, $ATT(g,t)$ can be recovered using a regression approach. However, combining the estimates of the parameters in this way does not seem to offer much convenience relative to simply computing the estimates directly using the main approach suggested in the paper. Thus, unlike the 2-period case, it does not appear that there is as exact of a mapping from a regression coefficient to a group-time average

treatment effect.

## Common Approaches to Pre-Testing in the Unconditional Case

Finally in this section, we consider the most common approach to pre-testing the augmented unconditional version of the parallel trends assumption, that is, to run the following regression (see Autor et al. (2007) and Angrist and Pischke (2008)).

$$Y_{it} = \alpha_t + \theta_g + \beta_0 D_{it} + \sum_{j=1}^{q} \beta_j \Delta D_{it,t+j} + u_{it} \tag{D.1}$$

where $D_{it}$ is a dummy variable for whether or not individual $i$ is treated in period $t$ (notice that this is not whether they are *first treated* in period $t$ but whether or not they are treated at all; it is a post-treatment dummy variable), $\Delta D_{it,t+j}$ is a $j$ period lead for individual $i$ who is first treated in period $t + j$. For example, when $t = 2$, $\Delta D_{i2,4} = 1$ (for $j = 2$) for individuals who are first treated in period 4, which indicates that the group of individuals first treated in period 4 will be treated 2 periods from period $t$.

Then, one can pre-test the unconditional parallel trends assumption by testing if $\beta_j = 0$ for $j = 1, \ldots, q$. Under the Unconditional DID Assumption, each $\beta_j$ will be 0. One advantage of this approach is that it allows simple graphs of pre-treatment trends. However, it is possible for this approach to miss departures from the unconditional parallel trends assumption that our test would not miss.

Consider the case with four periods and three groups – the control group, a group first treated in period 4, and a group first treated in period 3. Also, consider the case with $q = 1$. It is easy to show that $\beta_1 = \mathbb{E}[\Delta Y_3 | G_4 = 1] - \mathbb{E}[\Delta Y_3 | C = 1]$ and $\beta_1 = \mathbb{E}[\Delta Y_2 | G_3 = 1] - \mathbb{E}[\Delta Y_1 | C = 1]$ so that the estimate of $\beta_1$ will be a weighted average of these two pre-trends. Thus, the unconditional augmented parallel trends assumption could be violated in ways that offset each other leading to $\beta_1$ being equal to 0. Even more importantly, the weights associate with the regression coefficient $\beta_1$ may not be convex; see Propositions 3 and 7 in Abraham and Sun (2018) for detailed arguments. As a consequence, tests for pre-trends based on (D.1) may not be reliable under treatment effect heterogeneity. Our approach described in Remark 7 in the main text, on the other hand, does not suffer from this potential drawback.

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