

## Topic 5: Cluster-Robust Standard Errors

by Sai Zhang

**Key points:** The validity of Two-Way Cluster-Robust (TWCR) standard errors

**Disclaimer:** This note is compiled by Sai Zhang.

### 5.1 One-Way Clustering

First, consider the case of one-way clustering. The linear model with one-way clustering

$$y_{ig} = \mathbf{x}_{ig}\boldsymbol{\beta} + u_{ig}$$

where  $i$  denotes the  $i$ th of the  $N$  individuals in the sample,  $j$  denotes the  $g$ th of the  $G$  clusters, assume that

- $\mathbb{E}[u_{ig} | \mathbf{x}_{ig}] = 0$
- error independence across clusters: for  $i \neq j$

$$\mathbb{E}[u_{ig}u_{jg'} | \mathbf{x}_{ig}, \mathbf{x}_{jg'}] = 0 \quad (5.1)$$

unless  $g = g'$ , that is, errors for individuals within the same cluster may be correlated.

Grouping observations by cluster, get

$$\mathbf{y}_g = \mathbf{X}_g\boldsymbol{\beta} + \mathbf{u}$$

where  $\mathbf{X}_g$  has dimension  $N_g \times K$  and  $\mathbf{y}_g$  has dimension  $N_g \times 1$ , with  $N_g$  observations in cluster  $g$ . Stacking over cluster, get the matrix form of the model

$$\mathbf{y} = \mathbf{X}\boldsymbol{\beta} + \mathbf{u}$$

with  $\mathbf{y}, \mathbf{u}$  being  $N \times 1$  vectors,  $\mathbf{X}$  being an  $N \times K$  matrix. OLS estimator gives

$$\hat{\boldsymbol{\beta}} = (\mathbf{X}'\mathbf{X})^{-1} \mathbf{X}'\mathbf{y} = \left( \sum_{g=1}^G \mathbf{X}_g' \mathbf{X}_g \right)^{-1} \sum_{g=1}^G \mathbf{X}_g' \mathbf{y}_g \quad (5.2)$$

then, by CLT, we have that  $\sqrt{G}(\hat{\boldsymbol{\beta}} - \boldsymbol{\beta}) \xrightarrow{d} \mathcal{N}(0, \boldsymbol{\Sigma})$  where the variance matrix of the limit normal distribution  $\boldsymbol{\Sigma}$  is

$$\left( \lim_{G \rightarrow \infty} \frac{1}{G} \sum_{g=1}^G \mathbb{E}[\mathbf{X}_g' \mathbf{X}_g] \right)^{-1} \left( \lim_{G \rightarrow \infty} \frac{1}{G} \sum_{g=1}^G \mathbb{E}[\mathbf{X}_g' \mathbf{u}_g' \mathbf{u}_g \mathbf{X}_g] \right) \times \left( \lim_{G \rightarrow \infty} \frac{1}{G} \sum_{g=1}^G \mathbb{E}[\mathbf{X}_g' \mathbf{X}_g] \right)^{-1} \quad (5.3)$$

If the primary source of clustering is due to group-level common shocks, a useful approximation is that for the  $j$ th regressor, the default OLS variance estimate based on  $s^2(\mathbf{X}'\mathbf{X})^{-1}$  should be inflated by  $\tau_j \approx 1 + \rho_{x_j}\rho_u(\bar{N}_g - 1)$ , where

- $s$  is the estimated standard deviation of the error

- $\rho_{x_j}$  is a measure of within-cluster correlation of  $x_j$
- $\rho_u$  is the within-cluster error correlation
- $\bar{N}_g$  is the average cluster size

It's easy to see the  $\tau_j$  can be large even with small  $\rho_u$  (Kloek, 1981; Scott and Holt, 1982; Moulton, 1990). If assume the model for the cluster error variance matrices  $\mathbf{\Omega}_g = \mathbb{V}[\mathbf{u}_g | \mathbf{X}_g] = \mathbb{E}[\mathbf{u}_g \mathbf{u}_g' | \mathbf{X}_g]$ , and there is a consistent estimate  $\hat{\mathbf{\Omega}}_g$  of  $\mathbf{\Omega}_g$ , we can estimate  $\mathbb{E}[\mathbf{X}_g' \mathbf{u}_g \mathbf{u}_g' \mathbf{X}_g] = \mathbb{E}[\mathbf{X}_g' \mathbf{\Omega}_g \mathbf{X}_g]$  via GLS.

**Cluster-robust variance matrix estimate** consider

$$\hat{\mathbb{V}}[\hat{\beta}] = (\mathbf{X}'\mathbf{X})^{-1} \left( \sum_{g=1}^G \mathbf{X}_g' \hat{\mathbf{u}}_g \hat{\mathbf{u}}_g' \mathbf{X}_g \right) (\mathbf{X}'\mathbf{X})^{-1} \quad (5.4)$$

where  $\hat{\mathbf{u}}_g = \mathbf{y}_g - \mathbf{X}_g \hat{\beta}$ . This estimate is consistent if

$$G^{-1} \sum_{g=1}^G \mathbf{X}_g' \hat{\mathbf{u}}_g \hat{\mathbf{u}}_g' \mathbf{X}_g - G^{-1} \sum_{g=1}^G \mathbb{E}[\mathbf{X}_g' \mathbf{u}_g \mathbf{u}_g' \mathbf{X}_g] \xrightarrow{P} \mathbf{0}$$

as  $G \rightarrow \infty$ . An informal presentation of Eq.(5.4) is to rewrite the central matrix as

$$\hat{\mathbf{B}} = \sum_{g=1}^G \mathbf{X}_g' \hat{\mathbf{u}}_g \hat{\mathbf{u}}_g' \mathbf{X}_g = \mathbf{X}' \begin{bmatrix} \hat{\mathbf{u}}_1 \hat{\mathbf{u}}_1' & \mathbf{0} & \cdots & \mathbf{0} \\ \mathbf{0} & \hat{\mathbf{u}}_2 \hat{\mathbf{u}}_2' & & \vdots \\ \vdots & & \ddots & \mathbf{0} \\ \mathbf{0} & \cdots & & \hat{\mathbf{u}}_G \hat{\mathbf{u}}_G' \end{bmatrix} \mathbf{X} = \mathbf{X}' (\hat{\mathbf{u}} \hat{\mathbf{u}}' \otimes \mathbf{S}^G) \mathbf{X} \quad (5.5)$$

where  $\otimes$  denotes element-wise multiplication. The  $(p, q)$ th element of this matrix is

$$\sum_{i=1}^N \sum_{j=1}^N x_{ia} x_{jb} \hat{u}_i \hat{u}_j \cdot \mathbf{1}(i, j \text{ in the same cluster})$$

with  $\hat{u}_i = y_i - \mathbf{x}_i' \hat{\beta}$ .

$\mathbf{S}^G$  is an  $N \times N$  indicator matrix with  $\mathbf{S}_{ij}^G = 1$  only if the  $i$ th and  $j$ th observation belong to the same cluster: it zeros out a large amount of  $\hat{\mathbf{u}} \hat{\mathbf{u}}'$  (asymptotically equivalently,  $\mathbf{u} \mathbf{u}'$ ), specifically, only  $\sum_{g=1}^G N_g^2$  out of  $N^2 = \left( \sum_{g=1}^G N_g \right)^2$  terms are not zero (sub-matrices on the diagonal). Asymptotically

- for fixed  $N_g$ ,  $\frac{1}{N^2} \sum_{g=1}^G N_g^2 \xrightarrow{G \rightarrow \infty} 0$
- for balanced clusters  $N_g = N/G$ ,  $\frac{1}{N^2} \sum_{g=1}^G N_g^2 = \frac{1}{G} \xrightarrow{G \rightarrow \infty} 0$

A strand of literature popularizes this method:

- Liang and Zeger (1986): in a generalized estimatin equations setting
- Arellano (1987): fixed effects estimator in linear panel models
- Hansen (2007): asymptotic theory for panel data where  $T \rightarrow \infty$  in addition to  $N \rightarrow \infty$  (or  $N_g \rightarrow \infty$  in addition to  $G \rightarrow \infty$  in the notation above).

## 5.2 Two-Way Clustering

Now, consider the case of two-way clustering,

$$y_{i,gh} = \mathbf{x}'_{i,gh} \boldsymbol{\beta} + u$$

where each observation may belong to **two** dimension of groups: group  $g \in \{1, \dots, G\}$  and  $h \in \{1, \dots, H\}$ , and for  $i \neq j$

$$\mathbb{E} [u_{i,gh} u_{j,g'h'} | \mathbf{x}_{i,gh}, \mathbf{j}, \mathbf{g}', \mathbf{h}'] = 0 \quad (5.6)$$

unless  $g = g'$  or  $h = h'$ , that is, errors for individuals within the same group (along either  $g$  or  $h$ ) may be correlated.

**Cluster-robust variance matrix estimate** extending the one-way clustering case, keep elements of  $\hat{\mathbf{u}}\hat{\mathbf{u}}'$  where the  $i$ th and  $j$ th observations share a cluster in **any** dimension, then similar to Eq.(5.5)

$$\hat{\mathbf{B}} = \mathbf{X}' \left( \hat{\mathbf{u}}\hat{\mathbf{u}}' \otimes \mathbf{S}^{GH} \right) \mathbf{X} \quad (5.7)$$

here  $\mathbf{S}^{GH}$  is an  $N \times N$  indicator matrix with  $S_{ij}^{GH} = 1$  only if the  $i$ th and  $j$ th observation share any cluster, the  $(p, q)$ th element of this matrix is

$$\sum_{i=1}^N \sum_{j=1}^N x_{ia} x_{jb} \hat{u}_i \hat{u}_j \cdot \mathbf{1}(i, j \text{ share any cluster})$$

$\hat{\mathbf{B}}$  can also be presented in one-way cluster-robust fashion:

$$\hat{\mathbf{B}} = \mathbf{X}' \left( \hat{\mathbf{u}}\hat{\mathbf{u}}' \otimes \mathbf{S}^{GH} \right) \mathbf{X} = \mathbf{X}' \left( \hat{\mathbf{u}}\hat{\mathbf{u}}' \otimes \mathbf{S}^G \right) \mathbf{X} + \mathbf{X}' \left( \hat{\mathbf{u}}\hat{\mathbf{u}}' \otimes \mathbf{S}^H \right) \mathbf{X} - \mathbf{X}' \left( \hat{\mathbf{u}}\hat{\mathbf{u}}' \otimes \mathbf{S}^{G \cap H} \right) \mathbf{X} \quad (5.8)$$

where  $\mathbf{G}^{GH} = \mathbf{G}^G + \mathbf{G}^H - \mathbf{G}^{G \cap H}$ , with

- $\mathbf{G}^G$ :  $G_{ij}^G = 1$  only if the  $i$ th and  $j$ th observation belong to the same cluster  $g \in \{1, 2, \dots, G\}$
- $\mathbf{G}^H$ :  $G_{ij}^H = 1$  only if the  $i$ th and  $j$ th observation belong to the same cluster  $h \in \{1, 2, \dots, H\}$
- $\mathbf{G}^{G \cap H}$ :  $G_{ij}^{G \cap H} = 1$  only if the  $i$ th and  $j$ th observation belong to **both** the same cluster  $g \in \{1, 2, \dots, G\}$  and the same cluster  $h \in \{1, 2, \dots, H\}$

then, similar to one-way clustering case,

$$\begin{aligned} \hat{\mathbf{V}}[\hat{\boldsymbol{\beta}}] &= (\mathbf{X}'\mathbf{X})^{-1} \mathbf{X}' \left( \hat{\mathbf{u}}\hat{\mathbf{u}}' \otimes \mathbf{S}^G \right) \mathbf{X} (\mathbf{X}'\mathbf{X})^{-1} \\ &\quad + (\mathbf{X}'\mathbf{X})^{-1} \mathbf{X}' \left( \hat{\mathbf{u}}\hat{\mathbf{u}}' \otimes \mathbf{S}^H \right) \mathbf{X} (\mathbf{X}'\mathbf{X})^{-1} \\ &\quad - (\mathbf{X}'\mathbf{X})^{-1} \mathbf{X}' \left( \hat{\mathbf{u}}\hat{\mathbf{u}}' \otimes \mathbf{S}^{G \cap H} \right) \mathbf{X} (\mathbf{X}'\mathbf{X})^{-1} \end{aligned} \quad (5.9)$$

that is,

$$\hat{\mathbf{V}}[\hat{\boldsymbol{\beta}}] = \hat{\mathbf{V}}^G[\hat{\boldsymbol{\beta}}] + \hat{\mathbf{V}}^H[\hat{\boldsymbol{\beta}}] - \hat{\mathbf{V}}^{G \cap H}[\hat{\boldsymbol{\beta}}] \quad (5.10)$$

each of Eq.(5.10) can be separately computed by OLS of  $\mathbf{y}$  on  $\mathbf{X}$ , with variance matrix estimates  $\hat{\mathbf{V}}$  based on

- clustering on  $g \in \{1, 2, \dots, G\}$
- clustering on  $h \in \{1, 2, \dots, H\}$
- clustering on  $(g, h) \in \{(1, 1), \dots, (G, H)\}$

**Practical considerations** It is required to know what *ways* will be potentially important for clustering, which can be tested via checking the dimension of correlations in the errors. There are several ways to test

- estimate sample covariances of  $\mathbf{X}'\hat{\mathbf{u}}$  within dimensions, test the null that the **average** of such covariances is 0: rejecting this null is sufficient (not necessary) to reject the null of no clustering (White, 1980)
- for **small samples**, Eq. (5.4) is biased downwards. This is corrected (in Stata) by replacing  $\hat{\mathbf{u}}_g$  with  $\sqrt{c}\hat{\mathbf{u}}_g$ , where  $c = \frac{G}{G-1} \frac{N-1}{N-K} \simeq \frac{G}{G-1}$ . For two-way clustering (Eq. 5.8), there are 2 ways of correction:
  - choose correction terms for each of the 3 components:

$$c_1 = \frac{G}{G-1} \frac{N-1}{N-K}, c_2 = \frac{H}{H-1} \frac{N-1}{N-K}, c_3 = \frac{I}{I-1} \frac{N-1}{N-K}$$

with  $I$  being the number of unique clusters determined by  $G \cap H$

- choose a constant terms for all components:

$$c = \frac{J}{J-1} \frac{N-1}{N-K}$$

with  $J = \min(G, H)$

- **Var-cov matrix not positive-semidefinite**:  $\hat{\mathbf{V}}[\hat{\boldsymbol{\beta}}]$  might have negative elements on the diagonal (Eq. 5.10), informly, this is more likely to arise when clustering is done over the same groups as the fixed effects. One way to address this issue is using *eigendecomposition* technique:

$$\hat{\mathbf{V}}[\hat{\boldsymbol{\beta}}] = \mathbf{U}\mathbf{\Lambda}\mathbf{U}'$$

where

- $\mathbf{U}$  containing the eigenvectors of  $\hat{\mathbf{V}}$
- $\mathbf{\Lambda} = \text{diag}[\lambda_1, \dots, \lambda_d]$  contains the eigenvalues of  $\hat{\mathbf{V}}$

then create  $\mathbf{\Lambda}^+ = \text{diag}[\lambda_1^+, \dots, \lambda_d^+]$  with  $\lambda_j^+ = \max(0, \lambda_j)$  and use  $\hat{\mathbf{V}}^+[\hat{\boldsymbol{\beta}}] = \mathbf{U}\mathbf{\Lambda}^+\mathbf{U}'$  as the estimate

## 5.3 Multiway Clustering

Cameron et al. (2011) extended the framework<sup>1</sup> to allow clustering in  $D$  dimensions, then we can do the following reframing

- $G_d$ : the number of clusters in dimension  $d \in \{1, 2, \dots, D\}$
- $D$ -vector  $\boldsymbol{\delta}_i = \boldsymbol{\delta}(i)$ , with function  $\boldsymbol{\delta} : \{1, 2, \dots, N\} \rightarrow \times_{d=1}^D \{1, 2, \dots, G_d\}$  lists the cluster membership in each dimension of each observation

then we have

$$\mathbf{1}[i, j \text{ shares a cluster}] = 1 \Leftrightarrow \delta_{id} = \delta_{jd}$$

for some  $d \in \{1, 2, \dots, D\}$ , where  $\delta_{id}$  denotes the  $d$ th element of  $\boldsymbol{\delta}_i$ . Also

- $D$ -vector  $\mathbf{r}$ : define the set

$$R \equiv \{\mathbf{r} : r_d \in \{0, 1\}, d = 1, 2, \dots, D, \mathbf{r} \neq \mathbf{0}\}$$

elements of the set  $R$  can be used to index all cases where 2 observations share a cluster in at least one dimension. Define the function

$$\mathbf{I}_r(i, j) \equiv \mathbf{1}[r_d \delta_{id} = r_d \delta_{jd}, \forall d]$$

<sup>1</sup>Also proposed by Thompson (2011).

which indicates whether observations  $i$  and  $j$  have identical cluster membership for **all** dimensions  $d$  s.t.  $r_d = 1$ . Then we have a *aggregate* identifier

$$\mathbf{I}(i, j) = 1 \Leftrightarrow \mathbf{I}_r(i, j) = 1 \text{ for some } \mathbf{r} \in R$$

i.e., 2 observations share **at least** one dimension.

The define the  $2^D - 1$  matrices

$$\tilde{\mathbf{B}}_r \equiv \sum_{i=1}^N \sum_{j=1}^N \mathbf{x}_i \mathbf{x}_j' \hat{u}_i \hat{u}_j \mathbf{I}_r(i, j) \quad (5.11)$$

with  $\mathbf{r} \in R$ .

**Var-cov matrix estimator** consider, similarly, an estimator

$$\hat{\mathbb{V}}[\hat{\beta}] = (\mathbf{X}'\mathbf{X})^{-1} \tilde{\mathbf{B}} (\mathbf{X}'\mathbf{X})^{-1} \equiv (\mathbf{X}'\mathbf{X})^{-1} \left( \sum_{\|\mathbf{r}\|=k, \mathbf{r} \in R} (-1)^{k+1} \tilde{\mathbf{B}}_r \right) (\mathbf{X}'\mathbf{X})^{-1} \quad (5.12)$$

where cases of clustering on an odd number of dimensions are added, those of clustering on an even number of dimensions are subtracted. Consider the case of  $D = 3$ ,

$$(\tilde{\mathbf{B}}_{(1,0,0)} + \tilde{\mathbf{B}}_{(0,1,0)} + \tilde{\mathbf{B}}_{(0,0,1)}) - (\tilde{\mathbf{B}}_{(1,1,0)} + \tilde{\mathbf{B}}_{(1,0,1)} + \tilde{\mathbf{B}}_{(0,1,1)}) + \tilde{\mathbf{B}}_{(1,1,1)}$$

$\tilde{\mathbf{B}}$  is identical to  $\hat{\mathbf{B}}$  defined analogically as in Eq.(5.8), since

- no observation pair with  $\mathbf{I}(i, j) = 0$ : this is immediate, since  $\mathbf{I}(i, j) = 0 \Leftrightarrow \mathbf{I}_r(i, j) = 0, \forall \mathbf{r}$
- the covariance term corresponding to each observation pair with  $\mathbf{I}(i, j) = 1$  is included **exactly once** in  $\tilde{\mathbf{B}}$ : by inclusion-exclusion principle for set cardinality

$$\mathbf{I}(i, j) \Rightarrow \sum_{\|\mathbf{r}\|=k, \mathbf{r} \in R} (-1)^{k+1} \mathbf{I}_r(i, j) = 1$$

**Curse of dimensionality** this could arise in a setting with **many dimensions** of clustering, and in which one or more dimensions have **few** clusters<sup>2</sup>. **Cameron et al. (2011)** suggested an ad-hoc rule of thumb for approximating sufficient numbers of clusters.

### 5.3.1 Non-linear Estimators

**m-Estimators** Consider an  $m$ -estimator that solves

$$\sum_{i=1}^N \mathbf{h}_i(\hat{\theta}) = \mathbf{0}$$

under standard assumptions,  $\hat{\theta}$  is asymptotically normal with estimated variance matrix

$$\hat{\mathbb{V}}[\hat{\theta}] = \hat{\mathbf{A}}^{-1} \hat{\mathbf{B}} \hat{\mathbf{A}}'^{-1} \quad (5.13)$$

where  $\hat{\mathbf{A}} = \sum_i \frac{\partial \mathbf{h}_i}{\partial \theta'} \Big|_{\hat{\theta}}$  and  $\hat{\mathbf{B}}$  is an estimate of  $\mathbb{V}[\sum_i \mathbf{h}_i]$ .

<sup>2</sup>The square design (each dimension has the same number of clusters) with orthogonal dimensions has the **least** independence of observations.

- **one-way clustering**  $\hat{\mathbf{B}} = \sum_{g=1}^G \hat{\mathbf{h}}_g \hat{\mathbf{h}}_g'$  where  $\hat{\mathbf{h}}_g = \sum_{i=1}^{N_g} \hat{\mathbf{h}}_{ig}$ , clustering may not lead to parameter inconsistency, depending on whether  $\mathbb{E}[\mathbf{h}_i(\boldsymbol{\theta})] = \mathbf{0}$  with clustering
  - **population-averaged approach**: assume  $\mathbb{E}[y_{ig} | \mathbf{x}_{ig}] = \Phi(\mathbf{x}_{ig}'\boldsymbol{\beta})$
  - **random effects approach**: let  $y_{ig} = 1$  if  $y_{ig}^* > 0$  where  $y_{ig}^* = \mathbf{x}_{ig}'\boldsymbol{\beta} + \epsilon_g + \epsilon_{ig}$ , where
    - \* idiosyncratic error  $\epsilon_{ig} \sim \mathcal{N}(0, 1)$
    - \* cluster-specific error  $\epsilon_g \sim \mathcal{N}(0, \sigma_g^2)$
 then we have the alternative moment condition

$$\mathbb{E}[y_{ig} | \mathbf{x}_{ig}] = \Phi\left(\frac{\mathbf{x}_{ig}'\boldsymbol{\beta}}{\sqrt{1 + \sigma_g^2}}\right)$$

- **multiway clustering** replacing  $\hat{\mathbf{u}}_i \mathbf{x}_i$  in Eq.(5.11) with  $\hat{\mathbf{h}}_i$ , then we have

$$\hat{\mathbb{V}}[\hat{\boldsymbol{\theta}}] = \hat{\mathbf{A}}^{-1} \hat{\mathbf{B}} \hat{\mathbf{A}}^{-1}$$

where

$$\hat{\mathbf{A}} = \sum_i \frac{\partial \mathbf{h}_i}{\partial \boldsymbol{\theta}'} \bigg|_{\hat{\boldsymbol{\theta}}} \quad \hat{\mathbf{B}} = \sum_{\|\mathbf{r}\|=k, \mathbf{r} \in R} (-1)^{k+1} \tilde{\mathbf{B}}_{\mathbf{r}} \quad \tilde{\mathbf{B}}_{\mathbf{r}} \equiv \sum_{i=1}^N \sum_{j=1}^N \hat{\mathbf{h}}_i \hat{\mathbf{h}}_j' \mathbb{I}_{\mathbf{r}}(i, j)$$

with  $\mathbf{r} \in R^3$ .

**GMM estimation** Consider an example of over-identified models: linear two stage least squares with more instruments than endogenous regressors, we have

$$\hat{\boldsymbol{\theta}} = \arg \min_{\boldsymbol{\theta}} Q(\boldsymbol{\theta}) = \arg \min_{\boldsymbol{\theta}} \left( \sum_{i=1}^N \mathbf{h}_i(\boldsymbol{\theta}) \right)' \mathbf{W} \left( \sum_{i=1}^N \mathbf{h}_i(\boldsymbol{\theta}) \right)$$

where  $\mathbf{W}$  is a symmetric positive definite weighting matrix. Under standard regularity conditions,  $\hat{\boldsymbol{\theta}}$  is asymptotically normal, with estimated variance matrix

$$\hat{\mathbb{V}}[\hat{\boldsymbol{\theta}}] = (\hat{\mathbf{A}}' \mathbf{W} \hat{\mathbf{A}})^{-1} \hat{\mathbf{A}}' \mathbf{W} \hat{\mathbf{B}} \mathbf{W} \hat{\mathbf{A}} (\hat{\mathbf{A}}' \mathbf{W} \hat{\mathbf{A}})^{-1}$$

again,  $\hat{\mathbf{A}} = \sum_i \frac{\partial \mathbf{h}_i}{\partial \boldsymbol{\theta}'} \bigg|_{\hat{\boldsymbol{\theta}}}$ , and  $\hat{\mathbf{B}}$  is an estimate of  $\mathbb{V}[\sum_i \mathbf{h}_i]$ .

## 5.4 Menzel (2021): Asymptotic Gaussianity

One key of TWCR inference is the asymptotic Gaussianity, [Menzel \(2021\)](#) pointed out the potential non-Gaussianity of the limit distribution. Still, consider a random array  $(Y_{it})$  indexed by two dimensions by  $i = 1, \dots, N$  and  $t = 1, \dots, T$ . Clusters are sampled independently at random from an infinite population, but otherwise **unrestricted** in dependence within each row  $\mathbf{Y}_i := (Y_{i1} \dots, Y_{iT})$  and within each column  $\mathbf{Y}_{\cdot t} := (Y_{1t}, \dots, Y_{Nt})$ .

<sup>3</sup>This multiway clustering can be implemented using several one-way clustered bootstraps. Each of the one-way cluster robust matrices is estimated by a pairs cluster bootstrap that resamples with replacement from the appropriate cluster dimension. They are then combined as if they had been estimated analytically ([Cameron et al., 2011](#)).

### 5.4.1 Distribution of Sample Average

First, consider

$$\bar{Y}_{NT} := \frac{1}{NT} \sum_{i=1}^N \sum_{t=1}^T Y_{it}$$

and approximate the asymptotic distribution regardless of whether, or what type of, cluster-dependence is present.

**3 scenarios** of the array  $(Y_{it})$

- **no cluster-dependence**:  $(Y_{it})$  are mutually independent, CLT at a rate of  $(NT)^{-1/2}$  applies (under regularity conditions)
- **correlation within clusters**: the convergence rate of  $(Y_{it})$  is determined by the number of relevant clusters
- **non-separable models of heterogeneity (dependence with clusters, even uncorrelated)**<sup>4</sup>: The asymptotic behavior is non-standard

Consider 2 examples:

- **Additive factor model**

$$Y_{it} = \mu + \alpha_i + \gamma_t + \epsilon_{it}$$

where  $\mu$  is a constant, and  $\alpha_i, \gamma_t, \epsilon_{it}$  are zero-mean i.i.d. random variables for  $i = 1, \dots, N$  and  $t = 1, \dots, T$  with bounded second moments, and  $N = T$ . Based on a standard central limit theory, we have

- in the non-degenerate case with  $\text{Var}(\alpha_i) > 0$  or  $\gamma_t > 0$ , the sample distribution

$$\sqrt{N} \left( \bar{Y}_{NT} - \mathbb{E}[Y_{it}] \right) \xrightarrow{d} \mathcal{N}(0, \text{Var}(\alpha_i) + \text{Var}(\gamma_t))$$

- in the degenerate case of no clustering with  $\text{Var}(\alpha_i) = \text{Var}(\gamma_t) = 0$ , the sample distribution

$$\sqrt{NT} \left( \bar{Y}_{NT} - \mathbb{E}[Y_{it}] \right) \xrightarrow{d} \mathcal{N}(0, \text{Var}(\epsilon_{it}))$$

if marginal distributions of  $\alpha_i, \gamma_t, \epsilon_{it}$  are known, we can simulate from the joint distribution of  $(Y_{it})$  by sampling the individual components at random, a bootstrap procedure would be consistent. If **unknown**, consider estimators

$$\begin{aligned} \hat{\alpha}_i &:= \frac{1}{T} \sum_{t=1}^T (Y_{it} - \bar{Y}_{NT}) = \alpha_i + \frac{1}{T} \sum_{t=1}^T (\epsilon_{it} - \bar{\epsilon}_{NT}) \\ \hat{\gamma}_t &:= \frac{1}{N} \sum_{i=1}^N (Y_{it} - \bar{Y}_{NT}) = \gamma_t + \frac{1}{N} \sum_{i=1}^N (\epsilon_{it} - \bar{\epsilon}_{NT}) \\ \hat{\epsilon}_{it} &:= Y_{it} - \bar{Y}_{NT} - \hat{\alpha}_i - \hat{\gamma}_t \end{aligned}$$

then use these empirical distributions for estimation and form a bootstrap sample

$$Y_{it}^* := \bar{Y}_{NT} + \alpha_i^* + \gamma_t^* + \epsilon_{it}^*$$

<sup>4</sup>This is specific to clustering in 2 or more dimensions.

by drawing from these estimators and obtain  $\bar{Y}_{NT}^* := \frac{1}{NT} \sum_{i=1}^N \sum_{t=1}^T Y_{it}^*$ , and verify the conditional variances of the bootstrap distribution given the sample:

$$\begin{aligned} \frac{1}{N} \sum_{i=1}^N \left( \hat{\alpha}_i - \frac{1}{N} \sum_{j=1}^N \hat{\alpha}_j \right)^2 - \left[ \text{Var}(\alpha_i) + \frac{\text{Var}(\epsilon_{it})}{T} \right] &\xrightarrow{p} 0 \\ \frac{1}{T} \sum_{t=1}^T \left( \hat{\gamma}_t - \frac{1}{T} \sum_{s=1}^T \hat{\gamma}_s \right)^2 - \left[ \text{Var}(\gamma_t) + \frac{\text{Var}(\epsilon_{it})}{N} \right] &\xrightarrow{p} 0 \end{aligned}$$

then the bootstrap distribution is

– in the non-degenerate case,

$$\sqrt{N} \left( \bar{Y}_{NT}^* - \bar{Y}_{NT} \right) \xrightarrow{d} \mathcal{N} \left( 0, \text{Var}(\alpha_i) + \text{Var}(\gamma_t) \right)$$

the estimation error  $\hat{\alpha}_i$  does **NOT** affect the asymptotic variance.

– in the degenerate case,

$$\sqrt{NT} \left( \bar{Y}_{NT}^* - \bar{Y}_{NT} \right) \xrightarrow{d} \mathcal{N} \left( 0, 3\text{Var}(\epsilon_{it}) \right)$$

asymptotically overestimates the variance of the sampling distribution, leading to inconsistency of this naive bootstrapping procedure.

- **Non-Gaussian limit distribution**

$$Y_{it} = \alpha_i \gamma_t + \epsilon_{it}$$

where  $\alpha_i, \gamma_t, \epsilon_{it}$  are independently distributed with  $\mathbb{E}[\epsilon_{it}] = 0$ ,  $\text{Var}(\alpha_i) = \sigma_\alpha^2$ ,  $\text{Var}(\gamma_t) = \sigma_\gamma^2$ ,  $\text{Var}(\epsilon_{it}) = \sigma_\epsilon^2$ .

If  $\mathbb{E}[\alpha_i] = \mathbb{E}[\gamma_t] = 0$ , then CLT and Continuous Mapping Theorem (CMT) imply

$$\begin{aligned} \sqrt{NT} \cdot \bar{Y}_{NT} &= \frac{1}{\sqrt{NT}} \sum_{i=1}^N \sum_{t=1}^T (\alpha_i \gamma_t + \epsilon_{it}) \\ &= \left( \frac{1}{\sqrt{N}} \sum_{i=1}^N \alpha_i \right) \left( \frac{1}{\sqrt{T}} \sum_{t=1}^T \gamma_t \right) + \frac{1}{\sqrt{NT}} \sum_{i=1}^N \sum_{t=1}^T \epsilon_{it} \\ &\xrightarrow{d} \sigma_\alpha \sigma_\gamma Z_1 Z_2 + \sigma_\epsilon Z_3 \end{aligned}$$

then even without correlation within clusters, non-separable heterogeneity can still generate dependence in 2<sup>nd</sup> or higher moments in the limiting distribution<sup>5</sup>.

## 5.4.2 Menzel (2021)'s Bootstrap procedure

### 5.4.2.1 Notation

For the array  $(Y_{it})_{i,t}$ , denote

- $\mathbb{P}$ : joint distribution of  $(Y_{it})_{i,t}$

---

<sup>5</sup>2 major issues arise:

- The limiting distribution needs **not** be Gaussian: plug-in asymptotic inference based on the normal distribution is invalid
- It only comes from two-or-more-dimension cluster dependence, not single-dimension cluster dependence.



- $\mathbb{P}_{NT}$ : drifting DGP indexed by  $N, T$
- $\mathbb{P}_{NT}^*$ : bootstrap distribution for  $(Y_{it}^*)$  given the realizations  $(Y_{it} : i = 1, \dots, N; t = 1, \dots, T)$
- respective distributions  $\mathbb{E}, \mathbb{E}_{NT}, \mathbb{E}_{NT}^*$

### 5.4.2.2 Inference: Sample Mean

First, consider the assumption of *separate exchangeability*

#### Assumption 5.4.1: Separate Exchangeability

- A **separately exchangeable** array is an infinite array  $(Y_{it})_{i,t}$  such that for any integers  $\tilde{N}, \tilde{T}$  and permutations  $\pi_1 : \{1, \dots, \tilde{N}\} \rightarrow \{1, \dots, \tilde{N}\}$  and  $\pi_2 : \{1, \dots, \tilde{T}\} \rightarrow \{1, \dots, \tilde{T}\}$ , we have

$$(Y_{\pi_1(i), \pi_2(t)})_{i,t} \stackrel{d}{=} (Y_{it})_{i,t}$$

such an array is called **dissociated** if for any  $N_0, T_0 \geq 1$ ,  $(Y_{it})_{i=1, t=1}^{i=N_0, t=T_0}$  is independent of  $(Y_{it})_{i>N_0, t>T_0}$ .

- For dyadic data, consider the alternative assumption **jointly exchangeable** arrays  $(Y_{ij})_{i,j}$  satisfying

$$(Y_{\pi(i), \pi(j)})_{i,j} \stackrel{d}{=} (Y_{ij})_{i,j}$$

for any permutation  $\pi$  on  $\{1, \dots, \tilde{N}\}$ , in addition,  $(Y_{ij})_{i,j=1}^{N_0}$  is independent of  $(Y_{ij})_{i,j>N_0}$

This assumption can be interpreted as rows (and columns) corresponding to units that are drawn independently from a common population, where we then observe the joint outcome for every row-column pair, consider the requirements in the following applications

- **DiD/matched data**: the units corresponding to either dimension of the sample to represent independent draws from a common, infinite population
- **non-exhaustively matched data**: only observe joint outcomes for a possibly self-selected subset of unit pairs, sample selection should be (jointly or separately) exchangeable
- **U-/V-statistics**: the kernel  $Y_{i_1, \dots, i_D} := h(X_{i_1}, \dots, X_{i_D})$  evaluated at i.i.d. observations  $X_1, \dots, X_N$  forms a dissociated, jointly exchangeable array
- **Network**: unlabeled<sup>6</sup> data implies finite exchangeability, the sampled graph has joint (*infinite*) exchangeability if it is a subgraph of an infinite graph

Directly from Assumption 5.4.1, any dissociated separately exchangeable array can be represented as

$$Y_{it} = f(\alpha_i, \gamma_t, \epsilon_{it})$$

for some function  $f(\cdot)$  where  $\alpha_1, \dots, \alpha_N, \gamma_1, \dots, \gamma_T, \epsilon_{11}, \dots, \epsilon_{NT}$  are mutually independent, uniformly distributed random variables.

**Projection** now, decompose the array  $(Y_{it})_{i,t}$  as

$$Y_{it} = b + a_i + g_t + w_{it}$$

$$\mathbb{E}[w_{it} \mid a_i, g_t] = 0$$

<sup>6</sup>Unlabeled: model identifiers do not carry any significance for the statistical model.

where  $a_i$  and  $g_t$  are mean-zero and mutually independent, s.t. the joint distribution of  $Y_{it}$  can then be expanded as

$$\begin{aligned} Y_{it} &= \mathbb{E}[Y_{it}] + (\mathbb{E}[Y_{it} | \alpha_i] - \mathbb{E}[Y_{it}]) + (\mathbb{E}[Y_{it} | \gamma_t] - \mathbb{E}[Y_{it}]) \\ &\quad + (\mathbb{E}[Y_{it} | \alpha_i, \gamma_t] - \mathbb{E}[Y_{it} | \alpha_i] - \mathbb{E}[Y_{it} | \gamma_t] + \mathbb{E}[Y_{it}]) + (Y_{it} - \mathbb{E}[Y_{it} | \alpha_i, \gamma_t]) \\ &=: b + a_i + g_t + v_{it} + e_{it} \end{aligned}$$

with

- $e_{it} = Y_{it} - \mathbb{E}[Y_{it} | \alpha_i, \gamma_t]$
- $a_i = \mathbb{E}[Y_{it} | \alpha_i] - \mathbb{E}[Y_{it}]$ ,  $g_t = \mathbb{E}[Y_{it} | \gamma_t] - \mathbb{E}[Y_{it}]$
- $v_{it} = \mathbb{E}[Y_{it} | \alpha_i, \gamma_t] - \mathbb{E}[Y_{it} | \alpha_i] - \mathbb{E}[Y_{it} | \gamma_t] + \mathbb{E}[Y_{it}]$
- $b = \mathbb{E}[Y_{it}]$

here,

- temporal and cross-sectional units were drawn independently:  $a_1, \dots, a_N$  and  $g_1, \dots, g_T$  are independent of each other.
- by construction,  $\mathbb{E}[e_{it} | a_i, g_t, v_{it}] = 0$ ,  $\mathbb{E}[v_{it} | a_i] = \mathbb{E}[v_{it} | g_t] = 0$
- $e_{it}$ ,  $(a_i, g_t)$  and  $v_{it}$  are **uncorrelated**

then, rewrite the sample mean as

$$\begin{aligned} \hat{Y}_{NT} &= b + \bar{a}_N + \bar{g}_T + \bar{v}_{NT} + \bar{e}_{NT} \\ &=: b + \frac{1}{N} \sum_{i=1}^N a_i + \frac{1}{T} \sum_{t=1}^T g_t + \frac{1}{NT} \sum_{t=1}^T \sum_{i=1}^N v_{it} + \frac{1}{NT} \sum_{t=1}^T \sum_{i=1}^N e_{it} \end{aligned}$$

and the unconditional variances of the projections with

$$\sigma_a^2 := \text{Var}(a_i) \quad \sigma_g^2 := \text{Var}(g_t) \quad \sigma_v^2 := \text{Var}(v_{it}) \quad \sigma_e^2 := \text{Var}(e_{it})$$

let  $w_{it} := v_{it} + e_{it}$ , and denote its variance by  $\sigma_w^2 = \text{Var}(w_{it})$ . Then, assume integrability

#### Assumption 5.4.2: Integrability

Let  $Y_{it} = f(\alpha_i, \gamma_t, \epsilon_{it})$ , where  $\alpha_i, \gamma_t, \epsilon_{it}$  are random arrays with elements i.i.d. drawn from  $[0, 1]$  uniform distribution, assume

- $a_i/\sigma_a, g_t/\sigma_g, v_{it}/\sigma_v, e_{it}/\sigma_e$  are well-defined and have bounded moments up to the order  $4 + \delta$  for some  $\delta > 0$ , whenever the respective variances  $\sigma_a^2, \sigma_g^2, \sigma_v^2, \sigma_e^2$  are non-zero.
- $\sigma_a^2 + \sigma_g^2 > 0$ , or  $\sigma_v^2 + \sigma_e^2 > 0$

**Low-rank approximation** Consider the row/column projection

$$\bar{v}_{NT} \equiv \frac{1}{NT} \sum_{t=1}^T \sum_{i=1}^N (\mathbb{E}[Y_{it} | \alpha_i, \gamma_t] - \mathbb{E}[Y_{it} | \alpha_i] - \mathbb{E}[Y_{it} | \gamma_t] + \mathbb{E}[Y_{it}]) =: \frac{1}{NT} \sum_{t=1}^T \sum_{i=1}^N v(\alpha_i, \gamma_t)$$

as a generalized U-statistic with a kernel  $v(\alpha, \gamma)$  evaluated at the samples  $\alpha_1, \dots, \alpha_N$  and  $\gamma_1, \dots, \gamma_T$ . There are 2 major issues w.r.t. characterizing the distribution of  $\bar{Y}_{NT}$

- the presence of the projection error  $e_{it}$
- the factors  $\alpha_i, \gamma_t$  are not observable

Define,

$$v(\alpha, \gamma) := \mathbb{E}[Y_{it} \mid \alpha_i = \alpha, \gamma_t = \gamma] - \mathbb{E}[Y_{it} \mid \alpha_i = \alpha] - \mathbb{E}[Y_{it} \mid \gamma_t = \gamma] + \mathbb{E}[Y_{it}]$$

under Assumption 5.4.2, we have compact integral operators

$$S(u)(g) = \int v(a, g)u(a)F_\alpha(da) \quad S^*(u)(a) = \int v(a, g)u(g)F_\gamma(dg)$$

where  $F_\alpha, F_\gamma$  are the marginal distributions corresponding to the joint  $F_{\alpha\gamma}$  of  $\alpha_i, \gamma_t$ . Then the low-rank approximation is

$$v(\alpha, \gamma) = \sum_{k=1}^{\infty} c_k \phi_k(\alpha) \psi_k(\gamma) \quad (5.14)$$

under the  $L_2(F_{\alpha\gamma})$  norm on the space of smooth functions of  $(\alpha, \gamma) \in [0, 1]^2$ . Here

- $(c_k)_{k \geq 1}$ : a sequence of singular values,  $\lim |c_k| \rightarrow 0$
- $(\phi_k(\cdot))_{k \geq 1}$  and  $(\psi_k(\cdot))_{k \geq 1}$ : orthonormal bases for  $L_2([0, 1], F_\alpha)$  and  $L_2([0, 1], F_\gamma)$ :
  - By construction:

$$\mathbb{E}[v(a, \gamma_t)] = \mathbb{E}[v(\alpha_i, g)] = 0, \forall a, g \in [0, 1] \Rightarrow \mathbb{E}[\phi_k(\alpha_i)] = \mathbb{E}[\psi_k(\gamma_t)] = 0, \forall k = 1, 2, \dots$$

- the basis functions are orthonormal and  $\alpha_i$  and  $\gamma_t$  are independent, then  $\forall K < \infty$

$$\text{Cov}[(\phi_1(\alpha_i), \psi_1(\gamma_t), \dots, \phi_K(\alpha_i), \psi_K(\gamma_t))]$$

is the  $2K$ -dimensional identity matrix

- $(\phi_1(\alpha_i), \dots, \phi_K(\alpha_i))$  can be correlated with  $a_i$ :  $\sigma_{ak} := \text{Cov}(a_i, \phi_k(\alpha_i))$
- $(\psi_1(\gamma_t), \dots, \psi_K(\gamma_t))$  can be correlated with  $g_t$ :  $\sigma_{gk} := \text{Cov}(g_t, \psi_k(\gamma_t))$

with this representation of Eq.(5.14), we have<sup>7</sup>

$$\frac{1}{NT} \sum_{i=1}^N \sum_{t=1}^T v(\alpha_i, \gamma_t) = \sum_{k=1}^{\infty} c_k \left( \frac{1}{N} \sum_{i=1}^N \phi_k(\alpha_i) \right) \left( \frac{1}{T} \sum_{t=1}^T \psi_k(\gamma_t) \right)$$

and the second-order projection term can also be represented as a function of **countably many** sample averages of **i.i.d. mean-zero** random variables.

#### Assumption 5.4.3: Eigenfucntions and coefficients in the spectral representation (5.14)

The function  $v(\alpha, \gamma) := \mathbb{E}[Y_{it} \mid \alpha_i = \alpha, \gamma_t = \gamma] - \mathbb{E}[Y_{it} \mid \alpha_i = \alpha] - \mathbb{E}[Y_{it} \mid \gamma_t = \gamma] + \mathbb{E}[Y_{it}]$  admits a spectral representation

$$v(\alpha, \gamma) = \sum_{k=1}^{\infty} c_k \phi_k(\alpha) \psi_k(\gamma)$$

under the  $L_2(F_{\alpha\gamma})$  norm. And

- the singular values are uniformly bounded by a square summable null sequence  $\bar{c}_k$ :  $c_k \leq \bar{c}_k, \forall k = 1, 2, \dots$ , where  $\sum_{k=1}^{\infty} \bar{c}_k^2 < \infty$

<sup>7</sup>The limiting distribution of this term is not Gaussian, but can be represented as a linear combination of independent chi-squared random variables. This type of distributions is known as Wiener/Gaussian chaos.

- $\forall k = 1, 2, \dots$ , the first 3 moments of the eigenfunctions  $\phi_k(\alpha_i)$  and  $\psi_k(\gamma_t)$  are bounded by a constant  $B > 0$

To summarize the two assumptions

- Assumption 5.4.1 guarantees the pointwise consistency of the bootstrap
- Assumption 5.4.3 gives the uniform consistency of the bootstrap: it imposes common bounds on moments and singular values and restricts the set of joint distribution  $F$  to a **uniformity** class<sup>8</sup>.

### 5.4.2.3 Bootstrap procedure

For the sample mean  $\bar{Y}_{NT} - \mathbb{E}[Y_{it}]$ , the limiting distribution depends on the scale parameters:

- If observations are independent across rows and columns:  $\sqrt{NT} \left( \bar{Y}_{NT} - \mathbb{E}[Y_{it}] \right) \xrightarrow{d} \mathcal{N}(0, \sigma_e^2)$
- If  $N = T$ , within-cluster covariances are bounded from 0 in **at least one dimension**:  $\sqrt{N} \left( \bar{Y}_{NT} - \mathbb{E}[Y_{it}] \right) \xrightarrow{d} \mathcal{N}(0, \sigma_a^2 + \sigma_g^2)$

The bootstrap procedure should then be adaptive for both degenerate and non-degenerate cases. For the expansion

$$\begin{aligned} Y_{it} &= \mathbb{E}[Y_{it}] + (\mathbb{E}[Y_{it} | \alpha_i] - \mathbb{E}[Y_{it}]) + (\mathbb{E}[Y_{it} | \gamma_t] - \mathbb{E}[Y_{it}]) \\ &\quad + (\mathbb{E}[Y_{it} | \alpha_i, \gamma_t] - \mathbb{E}[Y_{it} | \alpha_i] - \mathbb{E}[Y_{it} | \gamma_t] + \mathbb{E}[Y_{it}]) + (Y_{it} - \mathbb{E}[Y_{it} | \alpha_i, \gamma_t]) \\ &=: b + a_i + g_t + v_{it} + e_{it} \end{aligned} \quad (5.15)$$

the sample analogs are:

$$\hat{a}_i := \frac{1}{T} \sum_{t=1}^T Y_{it} - \bar{Y}_{NT} \quad \hat{g}_t := \frac{1}{N} \sum_{i=1}^N Y_{it} - \bar{Y}_{NT} \quad \hat{w}_{it} := Y_{it} - \hat{a}_i - \hat{g}_t - \bar{Y}_{NT}$$

**Evaluating bootstrap performance** it is crucial at what rates these estimators are consistent depending on the extent of clustering in the true DGP. The variance of the projection terms are:

$$\text{Var}(\hat{a}_i) = \sigma_a^2 + \frac{\sigma_w^2}{T} \quad \text{Var}(\hat{g}_t) = \sigma_g^2 + \frac{\sigma_w^2}{N}$$

s.t. the **convolution error** depending on  $\sigma_w^2$  dominates in the degenerate case. Therefore, to correct for the contribution of the row/column averages of  $w_{it}$ , consider the scalar for the distribution of  $\hat{a}_i, \hat{g}_t$  by

$$\lambda_a = \frac{T\sigma_a^2}{T\sigma_a^2 + \sigma_w^2} \quad \lambda_g = \frac{N\sigma_g^2}{N\sigma_g^2 + \sigma_w^2}$$

---

<sup>8</sup>Here, the sequence  $c := (\tilde{c})_{k \geq 0}$  controls the magnitude of the error from a finite-dimensional approximation to  $v(\alpha, \gamma)$ .

**Component variance estimator** let

$$\begin{aligned}\hat{s}_a^2 &:= \frac{1}{N-1} \sum_{i=1}^N \left( \hat{a}_i - \bar{Y}_{NT} \right)^2 \\ \hat{s}_g^2 &:= \frac{1}{T-1} \sum_{t=1}^T \left( \hat{g}_t - \bar{Y}_{NT} \right)^2 \\ \hat{s}_w^2 &:= \frac{1}{NT - N - T} \sum_{i=1}^N \sum_{t=1}^T \left( Y_{it} - \hat{a}_i - \hat{g}_t - \bar{Y}_{NT} \right)^2\end{aligned}$$

then form the estimators as

$$\hat{\sigma}_a^2 = \max \left\{ 0, \hat{s}_a^2 - \frac{1}{T} \hat{s}_w^2 \right\} \quad \hat{\sigma}_g^2 = \max \left\{ 0, \hat{s}_g^2 - \frac{1}{N} \hat{s}_w^2 \right\} \quad \hat{\sigma}_w^2 := \hat{s}_w^2 \quad (5.16)$$

the rates of convergence for these estimators are given in the following lemma:

**Lemma 5.4.4: Stochastic Order of Variance Estimators**

Under Assumption 5.4.1,

$$\begin{aligned}\hat{\sigma}_a^2 - \sigma_a^2 &= O_p \left( \frac{1}{\sqrt{N}} \left( \sigma_a + \frac{\sigma_e}{\sqrt{T}} \right)^2 + \frac{\sigma_v^2}{T} \right) \\ \hat{\sigma}_g^2 - \sigma_g^2 &= O_p \left( \frac{1}{\sqrt{T}} \left( \sigma_g + \frac{\sigma_e}{\sqrt{N}} \right)^2 + \frac{\sigma_v^2}{N} \right) \\ \hat{\sigma}_w^2 - \sigma_w^2 &= O_p \left( \frac{\sigma_e^2}{\sqrt{NT}} + \left( \frac{1}{N} + \frac{1}{T} \right) \sigma_v^2 \right)\end{aligned}$$

and there exist **no estimators** for  $\sigma_a^2, \sigma_g^2, \sigma_w^2$  that converge at rates faster than these rates. Specifically,  $\sigma_a^2$  can **NOT** be estimated at a rate faster than  $T^{-1}$  even when  $\sigma_a^2 = 0^a$ .

<sup>a</sup>See the appendix of Menzel (2021) for the proof.

Hence, a bootstrap procedure can use a consistent pre-test for the presence of cluster dependence in the **first moment**, with the model selectors

$$\hat{D}_a(\kappa) := \mathbf{1} \{ T \hat{\sigma}_a^2 \geq \kappa \} \quad \hat{D}_g(\kappa) := \mathbf{1} \{ N \hat{\sigma}_g^2 \geq \kappa \}$$

$\forall \kappa \geq 0$ . And for some  $\kappa_a, \kappa_g$ , let

$$\hat{\lambda}_a := \frac{\hat{D}_a(\kappa_a) T \hat{\sigma}_a^2}{\hat{D}_a(\kappa_a) T \hat{\sigma}_a^2 + \hat{\sigma}_w^2} \quad \hat{\lambda}_g := \frac{\hat{D}_g(\kappa_g) T \hat{\sigma}_g^2}{\hat{D}_g(\kappa_g) N \hat{\sigma}_g^2 + \hat{\sigma}_w^2}$$

and estimate the asymptotic variance of the sample mean as

$$\hat{S}_{NT,sel}^2 := \hat{D}_a(\kappa_a) T \hat{\sigma}_a^2 + \hat{D}_g(\kappa_g) N \hat{\sigma}_g^2 + \hat{\sigma}_w^2 \quad (5.17)$$

**Bootstrap procedures** Menzel (2021) proposed the following resampling algorithm to estimate the sampling distribution for exhaustive sampling with cluster dependence in two dimensions

**Algorithm 5.4.5: Resampling Algorithm**

(a) For the  $b$ -th bootstrap iteration, draw

$$a_{i,b}^* := \hat{a}_{k_b^*(i)} \quad \mathcal{S}_{t,b}^* := \hat{\mathcal{S}}_{s_b^*(t)}$$

where  $k_b^*(i)$  and  $s_b^*(t)$  are i.i.d. draws from the discrete uniform distribution on the index sets  $\{1, \dots, N\}$  and  $\{1, \dots, T\}$  respectively

(b) Generate

$$w_{it,b}^* := \omega_{1i,b} \omega_{2t,b} \hat{w}_{k_b^*(i)s_b^*(t)}$$

where  $\omega_{1i,b}, \omega_{2t,b}$  are i.i.d. random variables with  $\mathbb{E}[\omega] = 0, \mathbb{E}[\omega^2] = \mathbb{E}[\omega^3] = 1^a$

(c) Generate a bootstrap sample of draws

$$Y_{it,b}^* = \bar{Y}_{NT} + \sqrt{\hat{\lambda}_a} a_{i,b}^* + \sqrt{\hat{\lambda}_g} \mathcal{S}_{t,b}^* + w_{it,b}^*$$

and get the bootstrapped statistic

$$\bar{Y}_{NT,b}^* := \frac{1}{NT} \sum_{i=1}^N \sum_{t=1}^T Y_{it,b}^*$$

(d) Repeat this procedure, get a sample of  $B$  replications and approximate the conditional distribution of  $\bar{Y}_{NT}^*$  given the sample with the empirical distribution over the bootstrap draws  $\bar{Y}_{NT,1}^*, \dots, \bar{Y}_{NT,B}^*$

<sup>a</sup>Typical choices of  $\omega_{1i,b}, \omega_{2t,b}$  are the Gamma distribution (with shape = 4, scale = 1/2).

For the **pivotal bootstrap**, the last step uses instead the empirical distribution of the studentized bootstrap draws to approximate the distribution of

$$\sqrt{NT} \left( \bar{Y}_{NT}^* - \bar{Y}_{NT} \right) / \hat{S}_{NT,sel}^*$$

where  $\hat{S}_{NT,sel}^*$  is the bootstrap analog of the variance estimator  $\hat{S}_{NT,sel}$ .

**Definition 5.4.6: Bootstrap Procedures**

Consider 3 versions of the bootstrap procedure based on 5.4.5:

- **BS-N** (bootstrap *without* model selection): apply steps (a) - (d), and set  $\kappa_a = \kappa_g = 0$
- **BS-S** (bootstrap *with* model selection): apply steps (a) - (d), and set  $\kappa_a, \kappa_g$  according to increasing sequences  $\kappa_g, \kappa_a \rightarrow \infty$  s.t.  $\kappa_a/T \rightarrow 0$  and  $\kappa_g/N \rightarrow 0$
- **BS-C** (*conservative* bootstrap): addition to the settings of **BS-S**, set

$$\hat{\lambda}_a := \frac{\hat{q}_a}{\hat{q}_a + \hat{\sigma}_w^2} \frac{\hat{q}_a}{T \hat{\sigma}_a^2} \quad \hat{\lambda}_g := \frac{\hat{q}_g}{\hat{q}_g + \hat{\sigma}_w^2} \frac{\hat{q}_g}{N \hat{\sigma}_g^2}$$

where

$$\hat{q}_a := \max \{ T \hat{\sigma}_a^2, \kappa_a \} \quad \hat{q}_g := \max \{ N \hat{\sigma}_g^2, \kappa_g \}$$

**Consistency of the bootstrap procedures**

- **BS-N** (bootstrap *with* model selection): **pointwise consistent** in  $\sigma_a^2, \sigma_g^2, \sigma_w^2$
- **BS-S** (bootstrap *without* model selection): **uniformly consistent** if the limiting distribution is Gaussian
- **BS-C** (*conservative* bootstrap): **consistent** in the nondegenerate case  $\sigma_a^2 + \sigma_g^2 > 0$ , but asymptotically **conservative** for the degenerate cases

To establish the consistency, define the **adaptive rate**  $r_{NT}$  as<sup>9</sup>

$$r_{NT}^{-2} := N^{-1}\sigma_a^2 + T^{-1}\sigma_g^2 + (NT)^{-1}\sigma_w^2 \equiv \text{Var}(\bar{Y}_{NT})$$

then consider the limiting distribution with the respective limits of normalized sequences:

$$\begin{aligned} q_{a,NT} &:= r_{NT}^2 N^{-1} \sigma_a^2 & q_{g,NT} &:= r_{NT}^2 T^{-1} \sigma_g^2 & q_{e,NT} &:= r_{NT}^2 (NT)^{-1} \sigma_e^2 & q_{v,NT} &:= r_{NT}^2 (NT)^{-1} \sigma_v^2 \\ q_{ak,NT} &:= r_{NT}^2 N^{-1} \sigma_{ak} & q_{gk,NT} &:= r_{NT}^2 T^{-1} \sigma_{gk} \end{aligned} \quad (5.18)$$

for  $k = 1, 2, \dots$ . Let  $\varrho_{NT} := r_{NT} (NT)^{-1/2}$ , then

$$q_{a,NT} + q_{g,NT} + q_{e,NT} + q_{v,NT} = 1$$

stacking the sequences as the vector

$$\mathbf{q}_{NT} := (q_{e,NT}, q_{a,NT}, q_{g,NT}, q_{a1,NT}, q_{g1,NT}, q_{a2,NT}, q_{g2,NT}, \dots)$$

and the singular values for the spectral decomposition (5.14):

$$\begin{aligned} \mathbf{c}_{NT} &:= (c_{1,NT}, c_{2,NT}, \dots) \in l^2 & \text{for } \mathbb{E}_{NT} [Y_{it} \mid \alpha_i, \gamma_t] \\ \mathbf{c} &:= (c_1, c_2, \dots) \in l^2 & \text{for } \mathbb{E} [Y_{it} \mid \alpha_i, \gamma_t] \end{aligned}$$

then for convergent sequences  $\mathbf{q}_{NT}, \mathbf{c}_{NT}, \mathbf{c}$ , denote the limits

$$\begin{aligned} q_a &:= \lim_{N,T} q_{a,NT} & q_g &:= \lim_{N,T} q_{g,NT} & q_e &:= \lim_{N,T} q_{e,NT} & q_v &:= \lim_{N,T} q_{v,NT} \\ \mathbf{q} &:= \lim_{N,T} \mathbf{q}_{NT} & \mathbf{c} &:= \lim_{N,T} \mathbf{c}_{NT} & \varrho &:= \lim_{N,T} \varrho_{NT} \end{aligned}$$

for any fixed values of  $\mathbf{q}, \mathbf{c}, \varrho \in [0, 1]$ , define

$$\mathcal{L}_0(\mathbf{q}, \mathbf{c}, \varrho) := \left( \sqrt{q_e} Z^e + \sqrt{q_a} Z^a + \sqrt{q_g} Z^g \right) + \varrho V \quad (5.19)$$

where

$$V := \sum_{k=1}^{\infty} c_k Z_k^\psi Z_k^\phi$$

and  $Z^e, Z_k^\psi, Z_k^\phi$  are i.i.d. standard normal random variables,  $Z^a, Z_g$  are standard normal random variables with

$$\text{Cov}(Z^a, Z_k^\phi) = \frac{q_{ak}}{\sqrt{q_a}} \quad \text{Cov}(Z^g, Z_k^\psi) = \frac{q_{gk}}{\sqrt{q_g}} \quad \text{Cov}(Z^a, Z^g) = \text{Cov}(Z^a, Z_k^\psi) = \text{Cov}(Z^g, Z_k^\psi) = 0$$

Then, the CLT for sampling distribution is established as

<sup>9</sup>Following Eq. (5.15),  $\text{Var}(\bar{Y}_{NT}) = \text{Var}(b + \bar{a}_N + \bar{g}_T + \bar{v}_{NT} + \bar{e}_{NT})$ .

**Theorem 5.4.7: CLT for Sampling Distribution**

Under Assumption 5.4.2,

(a) along *any* convergent sequence  $\mathbf{q}_{NT} \rightarrow \mathbf{q}$  and fixed  $\mathbf{c} = (c_1, c_2, \dots)$ , we have

$$\left\| \mathbb{P} \left( r_{NT} \left( \bar{Y}_{NT} - \mathbb{E}[Y_{it}] \right) \right) - \mathcal{L}_0(\mathbf{q}, \mathbf{c}, \varrho) \right\|_{\infty} \rightarrow 0$$

where  $\varrho := \lim_{N,T} \varrho_{NT}$ , and  $\|\cdot\|_{inf ty}$  denotes the Kolmogorov metric; the limiting distribution  $\mathcal{L}_0(\mathbf{q}, \mathbf{c}, \varrho)$  is continuous<sup>a</sup>.

(b) if in addition, Assumption 5.4.3 holds, (a) is robust under drifting sequences  $\mathbf{c}_{NT} \rightarrow \mathbf{c}^b$

<sup>a</sup>The convergence is pointwise w.r.t. the conditional mean function  $\mathbb{E}[Y_{it} | \alpha_i = \alpha, \gamma_t = \gamma]$

<sup>b</sup>The convergence is uniform within the class of distributions satisfying Assumption 5.4.3

**Estimating the asymptotic distribution** Lemma 5.4.4 establishes the consistency of the estimation for the components variances  $\sigma_a^2, \sigma_g^2, \sigma_w^2$ , but are they **fast** enough?

**Proposition 5.4.8: Estimability of Asymptotic Distribution**

Let  $\hat{\mathcal{L}}_{NT}$  denote an arbitrary estimator for  $\mathcal{L}_0$  based on an array of size  $N, T$  from the unknown distribution, then  $\exists \delta > 0$  s.t.

$$\liminf_{N,T \rightarrow \infty} \sup_{f \in \mathcal{F}} \mathbb{P}_{f,NT} \left( \left\| \hat{\mathcal{L}}_{NT} - \mathcal{L}_0(\mathbf{q}_{NT}(f), \mathbf{c}_{NT}(f), \varrho_{NT}(f)) \right\|_{\infty} > \delta \right) > 0$$

where

- $\mathcal{F}$ : the class of functions  $f(\alpha, \gamma, \epsilon)$  corresponding to distributions of  $Y_{it}$  satisfying Assump. 5.4.2 and 5.4.3, for i.i.d. uniform  $\alpha_i, \gamma_t, \epsilon_{it}$ <sup>a</sup>
- $\mathbb{P}_{f,NT}(\cdot)$ : probabilities for events w.r.t. an array of size  $N, T$ , generated according to  $f$
- $\mathbf{q}_{NT}(f) := (q_{e,NT}(f), q_{a,NT}(f), \dots)$ : the vector of normalized variances from Eq. 5.18

<sup>a</sup>From the Aldous-Hoover representation

Proposition 5.4.8 states that there exists **no estimator**<sup>10</sup> for the asymptotic distribution that achieves consistency uniformly over the space of distributions satisfying Assumption 5.4.2 and 5.4.3:

- Under Theorem 5.4.7, the sample mean  $\bar{Y}_{NT}$  converges to a continuous limiting distribution  $\mathcal{L}_0(\mathbf{q}, \mathbf{c}, \varrho)$  along sequences  $f_{NT} \in \mathcal{F}$  with proper limits for  $\mathbf{q}_{NT}, \mathbf{c}_{NT}$

<sup>10</sup>Consider the counterexample for this impossibility: for the model

$$Y_{it} = \alpha_i \gamma_t$$

where  $\alpha_i, \gamma_t$  are mutually independent with i.i.d. factors  $\alpha_i \sim \mathcal{N}(0, 1), \gamma_t \sim \mathcal{N}(\mu_\gamma, 1)$ . This model satisfies Assump. 5.4.2, hence Thm. 5.4.7 gives convergence results. However, for this model

$$\begin{aligned} a_i &:= \mathbb{E}[Y_{it} | \alpha_i] = \alpha_i \mu_\gamma & g_t &:= \mathbb{E}[Y_{it} | \gamma_t] = \gamma_t \mathbb{E}[\alpha_i] \equiv 0 \\ v_{it} &= \alpha_i(\gamma_t - \mu_\gamma) & \sigma_a^2 &= \mu_\gamma^2 & \sigma_v^2 &= 1 \end{aligned}$$

here,  $\mu_\gamma$  can **not** be estimated from the original data at a rate faster than  $T^{-1/2}$ , the fastest possible rate at which  $\mu_\gamma$  can be estimated from observing  $\gamma_1, \dots, \gamma_T$  directly. Therefore, no test can consistently distinguish the model  $\mu_\gamma = 0$  (asymptotic variance  $\sigma_v^2$ ) from a drifting sequence  $\tilde{\mu}_{T,\gamma} := T^{-1/2} m_\gamma$  (asymptotic variance  $m_\gamma^2 + \sigma_v^2$ ).



**Bootstrap Consistency** Consider the bootstrap analog of  $\hat{S}_{NT,sel}$  in Eq. 5.17

$$\hat{S}_{NT,sel}^{2*} := \hat{D}_a(\kappa_a)T\hat{\sigma}_a^{2*} + \hat{D}_g(\kappa_g)N\hat{\sigma}_g^{2*} + \hat{\sigma}_w^{2*}$$

where  $\hat{D}_a(\kappa_a), \hat{D}_g(\kappa_g)$  are fixed at the sample values,  $\kappa_a, \kappa_g$  are chosen according to whether the bootstrap is **with** or **without** model selection. Consider 2 versions based on the studentized sample mean:

- **non-pivotal** bootstrap: approximating the distribution of **the sample mean**  $r_{NT}(\bar{Y}_{NT} - \mathbb{E}[Y_{it}])$  with the bootstrap distribution  $r_{NT}(\bar{Y}_{NT}^* - \bar{Y}_{NT})$
- **pivotal** bootstrap: approximating the distribution of the **studentized sample mean**  $\frac{(NT)^{1/2}}{\hat{S}_{NT,sel}}(\bar{Y}_{NT} - \mathbb{E}[Y_{it}])$  with the bootstrap distribution  $\frac{(NT)^{1/2}}{\hat{S}_{NT,sel}^*}(\bar{Y}_{NT}^* - \bar{Y}_{NT})$

And we can establish the consistency

#### Theorem 5.4.9: Bootstrap Consistency

Under Assumption 5.4.2,

- (a) the bootstrap **with model selection** satisfies

$$\left\| \mathbb{P}_{NT}^* \left( r_{NT}(\bar{Y}_{NT}^* - \bar{Y}_{NT}) \right) - \mathbb{P}_{NT} \left( r_{NT}(\bar{Y}_{NT} - \mathbb{E}[\bar{Y}_{it}]) \right) \right\|_{\infty} \xrightarrow{\text{a.s.}} 0 \quad (5.20)$$

and its pivotal analog

$$\left\| \mathbb{P}_{NT}^* \left( \sqrt{NT} \frac{\bar{Y}_{NT}^* - \bar{Y}_{NT}}{\hat{S}_{NT,sel}^*} \right) - \mathbb{P}_{NT} \left( \sqrt{NT} \frac{\bar{Y}_{NT} - \mathbb{E}[Y_{it}]}{\hat{S}_{NT,sel}} \right) \right\|_{\infty} \xrightarrow{\text{a.s.}} 0 \quad (5.21)$$

**pointwise** for any fixed  $\sigma_a^2, \sigma_g^2, \sigma_e^2, \sigma_v^2$

- (b) the bootstrap **without model selection** satisfies Eq.5.20 and Eq.5.21 **uniformly** if  $q_v = 0$   
 (c) the **conservative** bootstrap satisfies

$$\left\| \mathbb{P}_{NT}^* \left( r_{NT}(\bar{Y}_{NT}^* - \bar{Y}_{NT}) \right) - \mathcal{L}_0(\bar{\mathbf{q}}, \mathbf{c}, \rho) \right\|_{\infty} \xrightarrow{\text{P}} 0 \quad (5.22)$$

and its pivotal analog

$$\left\| \mathbb{P}_{NT}^* \left( \sqrt{NT} \frac{\bar{Y}_{NT}^* - \bar{Y}_{NT}}{\hat{S}_{NT,sel}^*} \right) - \mathcal{L}_0(\bar{\mathbf{q}}, \mathbf{c}, \rho) \right\|_{\infty} \xrightarrow{\text{P}} 0 \quad (5.23)$$

uniformly over the **entire parameter space**, where  $\bar{\mathbf{q}} = (q_c, \bar{q}_a, \bar{q}_g, 0, 0, \dots)$ , with  $\bar{q}_a := \max\{\kappa_a/T, q_a\}$  and  $\bar{q}_g := \max\{\kappa_g/T, q_g\}$ , which increases as  $N, T \rightarrow \infty$ .

Theorem 5.4.7 gives that

- **bootstrap with model selection**: pointwise valid asymptotically
- **bootstrap without model selection**: valid uniformly w.r.t. clustering in means, but **inconsistent** if  $q_v > 0$
- **conservative bootstrap**: uniformly valid without any qualifications. In degenerate cases ( $q_e + q_v > 0$ ), the scale of the estimated asymptotic distribution **diverges** at a rate  $\kappa_a/T + \kappa_g/N$

Notice that  $\mathcal{L}_0(\bar{\mathbf{q}}, \mathbf{c}, \rho)$  in Thm. 5.4.9 is a mean-preserving spread of  $\mathcal{L}_0(\mathbf{q}, \mathbf{c}, \rho)$  in Thm. 5.4.7, hence estimates of percentiles from the conservative bootstrap are **biased outwards** away from 0, leading to asymptotic conservative CIs.

**Refinements** Using standard results on Edgeworth expansions, get

**Proposition 5.4.10: Refinements**

Under Assumption 5.4.2 for any  $0 < \delta < \infty$ , and the distributions of  $a_i$  and  $g_t$  satisfy Cramer's condition<sup>a</sup>

$$\limsup_{\|t\| \rightarrow \infty} |\mathbb{E} [\exp(it' \mathbf{X})]| < 1$$

then if  $\sigma_a^2 + \sigma_g^2 \geq C$  for some  $C > 0$ , we have

$$\left\| \mathbb{P}_{NT}^* \left( \sqrt{NT} \frac{\bar{Y}_{NT}^* - \bar{Y}_{NT}}{\hat{S}_{NT,sel}^*} - \mathbb{P}_{NT} \left( \sqrt{NT} \frac{\bar{Y}_{NT} - \mathbb{E}[Y_{it}]}{\hat{S}_{NT,sel}} \right) \right) \right\|_{\infty} = O_p \left( r_{NT}^{-2} \vee (NT)^{-1/2} \right)$$

for all three versions of the bootstrap.

<sup>a</sup>Cramer's condition states that  $\mathbf{X}$  has a non-degenerate, absolutely continuous component.

#### 5.4.2.4 Inference in Regression Models

Consider the linear projection model

$$y_{it} = \mathbf{x}_{it}' \boldsymbol{\beta} + u_{it} \quad (5.24)$$

with the dependent variable  $y_{it}$  and the vector of  $k$  regressors  $\mathbf{x}_{it} \in \mathbb{R}^k$ . Consider LS estimator

$$\hat{\boldsymbol{\beta}}_{LS} := (\mathbf{X}'\mathbf{X})^{-1} \mathbf{X}'\mathbf{y} = \boldsymbol{\beta} + (\mathbf{X}'\mathbf{X})^{-1} \left( \frac{1}{NT} \sum_{i=1}^N \sum_{t=1}^T \mathbf{x}_{it} u_{it} \right)$$

assume  $(\mathbf{x}_{it} u_{it})_{i,t}$  constitute a dissociated, separately exchangeable array, then we can have the Aldous-Hoover representation

$$F_{it} := \mathbf{x}_{it} u_{it} = f(\alpha_i, \gamma_t, \epsilon_{it})$$

then denote

$$\begin{aligned} \mathbf{a}_i &:= \mathbb{E}[\mathbf{x}_{it} u_{it} \mid \alpha_i] & \mathbf{g}_t &:= \mathbb{E}[\mathbf{x}_{it} u_{it} \mid \gamma_t] \\ \mathbf{v}_{it} &:= \mathbb{E}[\mathbf{x}_{it} u_{it} \mid \alpha_i, \gamma_t] - \mathbf{a}_i - \mathbf{g}_t & \mathbf{e}_{it} &:= \mathbf{x}_{it} u_{it} - \mathbb{E}[\mathbf{x}_{it} u_{it} \mid \alpha_i, \gamma_t] \\ \mathbf{w}_{it} &:= \mathbf{x}_{it} u_{it} - \mathbf{a}_i - \mathbf{g}_t = \mathbf{v}_{it} + \mathbf{e}_{it} \end{aligned}$$

and the unconditional component variances as  $\sigma_{al}^2, \sigma_{gl}^2, \sigma_{vl}^2, \sigma_{el}^2, \sigma_{wl}^2 = \sigma_{vl}^2 + \sigma_{el}^2$ . The empirical analog of this decomposition is given by

$$\begin{aligned} \hat{\mathbf{a}}_i &:= \frac{1}{T} \sum_{t=1}^T \mathbf{x}_{it} \hat{u}_{it} & \hat{\mathbf{g}}_t &:= \frac{1}{N} \sum_{i=1}^N \mathbf{x}_{it} \hat{u}_{it} \\ \hat{\mathbf{w}}_{it} &:= \mathbf{x}_{it} \hat{u}_{it} - \hat{\mathbf{a}}_i - \hat{\mathbf{g}}_t \end{aligned}$$

for each  $l = 1, \dots, k$ , then construct

**Bootstrap procedure for regression**

- for the  $b$ th bootstrap iteration, draw  $\mathbf{a}_{i,b}^* := \hat{\mathbf{a}}_{k_b^*(i)}^*$  and  $\mathbf{g}_{t,b}^* := \hat{\mathbf{g}}_{s_b^*(t)}^*$  where  $k_b^*(i)$  and  $s_b^*(t)$  are i.i.d. draws from the discrete uniform distribution on the index sets  $\{1, \dots, N\}$  and  $\{1, \dots, T\}$ , respectively
- generate  $\mathbf{w}_{it,b}^* := \omega_{1i,b} \omega_{2t,b} \hat{\mathbf{w}}_{k_b^*(i)s_b^*(t)}^*$ , where  $\omega_{1i,b}$ ,  $\omega_{2t,b}$  are i.i.d. random variables with  $\mathbb{E}[\omega] = 0$ ,  $\mathbb{E}[\omega^2] = \mathbb{E}[\omega^3] = 1$
- simulate values of  $\mathbf{z}_{it,b}^* = (z_{it1,b}^*, \dots, z_{itk,b}^*)'$ , where the  $l$ th component is given by

$$z_{itl,b}^* := \sqrt{\hat{\lambda}_{al}} a_{il,b}^* + \sqrt{\hat{\lambda}_{gl}} g_{tl,b}^* + w_{itl,b}^*$$

where the scalars are  $\hat{\lambda}_{al} := \frac{\hat{D}_{al}(\kappa_a) T \hat{\sigma}_{al}^2}{\hat{D}_{al}(\kappa_a) T \hat{\sigma}_{al}^2 + \hat{\sigma}_{wl}^2}$  and  $\hat{\lambda}_{gl} := \frac{\hat{D}_{gl}(\kappa_g) N \hat{\sigma}_{gl}^2}{\hat{D}_{gl}(\kappa_g) N \hat{\sigma}_{gl}^2 + \hat{\sigma}_{wl}^2}$

- then compute

$$\hat{\beta}_{LS,b}^* = \hat{\beta}_{LS} + (\mathbf{X}'\mathbf{X})^{-1} \left( \frac{1}{NT} \sum_{i=1}^N \sum_{t=1}^T \mathbf{z}_{it,b}^* \right)$$

for each bootstrap sample.

Next, we can approximate the asymptotic distribution of  $r_{NT}(\hat{\beta}_{LS} - \beta)$  with the simulated distribution of  $r_{NT}(\hat{\beta}_{LS,b}^* - \hat{\beta}_{LS})$ .

**Assumption 5.4.11: Regression**

Assume the model in Eq.(5.24) with  $\mathbf{x}_{it}u_{it} = f(\alpha_i, \gamma_t, \epsilon_{it})$  and  $\alpha_i, \gamma_t, \epsilon_{it}$  are i.i.d. uniform on  $[0, 1]$ . And

- $\mathbf{X}$  has full column rank
- $\forall l = 1, \dots, k$  and some  $\delta > 0$ , the  $(4 + \delta)$ th absolute moments of  $x_{itl}$  are bounded, and the  $(4 + \delta)$ th conditional moments of each component  $\frac{a_{il}}{\sqrt{\text{Var}(a_{il}|\mathbf{X})}}$ ,  $\frac{g_{tl}}{\sqrt{\text{Var}(g_{tl}|\mathbf{X})}}$ ,  $\frac{v_{itl}}{\sqrt{\text{Var}(v_{itl}|\mathbf{X})}}$  and  $\frac{e_{itl}}{\sqrt{\text{Var}(e_{itl}|\mathbf{X})}}$  given  $\mathbf{X}$  are bounded whenever the conditional variance of either component is strictly positive.
- unconditional variance:  $\text{Var}(a_{il}) + \text{Var}(g_{tl}) > 0$  or  $\text{Var}(w_{itl}) > 0$  for each  $l = 1, \dots, k$ .
- For each component of  $\mathbf{z}_{it} = \mathbf{x}_{it}u_{it}$ , there exists a spectral representation satisfying Assumption 5.4.3

then analogous to the sample mean inference, we have

**Proposition 5.4.12: Regression Inference**

Under Assumption 5.4.11, then

- $\hat{\beta}_{LS}$  is consistent at the  $r_{NT}$  rate
- The bootstrap with model selection satisfies Eq.(5.20) and (5.21) pointwise as  $\sigma_{al}^2, \sigma_{gl}^2, \sigma_{el}^2, \sigma_{vl}^2$  are held fixed for all  $l = 1, \dots, k$
- The bootstrap without mode selection satisfies Eq.(5.20) and (5.21) uniformly if  $q_{vl} = 0$  for all  $l = 1, \dots, k$
- The conservative bootstrap satisfies Eq.(5.22) and (5.23) uniformly over the entire parameter space

**Asymptotic Gaussian of the LS estimator** for conditional asymptotic normality of bilinear forms  $V_k := \mathbf{Z}'_{1k} \mathbf{X} \mathbf{Z}_{2k}$  of random vectors  $\mathbf{Z}_{1k}, \mathbf{Z}_{2k}$  given the matrix  $\mathbf{X}$ . Under the conditions of this paper,  $V_k$  is asymptotically Gaussian if  $\check{\mathbf{x}}_{it}, \check{\mathbf{x}}_{js}$  are mean-independent for any  $(j, s) \neq (i, t)$ <sup>11</sup>.

## 5.5 Latest Development

### 5.5.1 LLN and CLT for Exchangeable Arrays

Davezie et al. (2021) establish uniform LLN and CLT to show consistency and asymptotic normality of **nonlinear** estimators under weak regularity conditions.

#### 5.5.1.1 Set up

**Notations** For any  $A \subset \mathbb{R}, B \subset \mathbb{R}^k$  for some  $k \geq 2$ , then let

$$A^+ = A \cap (0, \infty)$$

$$\overline{B} = \left\{ b = (b_1, \dots, b_k) \in B : \forall (i, j) \in \{1, \dots, k\}^2, i \neq j, b_i \neq b_j \right\}$$

and let

- $\mathbb{I}_k = \overline{\mathbb{N}^{+k}}$  denote the set of  $k$ -tuples of  $\mathbb{N}^+$  **without** repetition
- for any  $n \in \mathbb{N}^+$ , let  $\mathbb{I}_{n,k} = \overline{\{1, \dots, n\}^k}$
- for any  $\mathbf{i} = (i_1, \dots, i_k), \mathbf{j} = (j_1, \dots, j_k)$  in  $\mathbb{N}^k$ , let  $\mathbf{i} \odot \mathbf{j} = (i_1 j_1, \dots, i_k j_k)$ , and denote the distinct elements of  $\mathbf{i}$  as  $\{\mathbf{i}\}$
- for any  $r \in \{1, \dots, k\}$ , let

$$\mathcal{E}_r = \left\{ (e_1, \dots, e_k) \in \{0, 1\}^k : \sum_{j=1}^k e_j = r \right\}$$

- for any  $A \subset \mathbb{N}^+$ , let  $\mathfrak{S}(A)$  denote the set of permutations on  $A$ , then for any  $\mathbf{i} = (i_1, \dots, i_k) \in \mathbb{N}^{+k}$  and  $\pi \in \mathfrak{S}(\mathbb{N}^+)$ , let  $\pi(\mathbf{i}) = (\pi(i_1), \dots, \pi(i_k))$

**Polyadic data** For random variables  $Y_{\mathbf{i}}$  indexed by  $\mathbf{i} \in \mathbb{I}_k$ <sup>12</sup>, it's assumed that the random variables are generated according to a **jointly exchangeable** and **dissociated** array:

#### Assumption 5.5.1: Jointly Exchangeable and Dissociated Arrays

For any  $\pi \in \mathfrak{S}(\mathbb{N}^+)$ ,

$$(Y_{\mathbf{i}})_{\mathbf{i} \in \mathbb{I}_k} \stackrel{d}{=} (Y_{\pi(\mathbf{i})})_{\mathbf{i} \in \mathbb{I}_k}$$

and for any disjoint subsets of  $\mathbb{N}^+, A, B$ , with  $\min(|A|, |B|) \geq k$ ,  $(Y_{\mathbf{i}})_{\mathbf{i} \in \overline{A^k}}$  is **independent** of  $(Y_{\mathbf{i}})_{\mathbf{i} \in \overline{B^k}}$

The assumption implies that

<sup>11</sup>For difference-in-differences designs with a regressor  $x_{it1} := \mathbf{1}\{t \geq T_i\}$  for unit-specific intervention date  $T_i$ , or when  $\mathbf{x}_{it} := \mathbf{x}(\xi_i, \zeta_t)$  are a non-additive function of row- and column-level attributes  $\xi_i$  and  $\zeta_t$ , respectively, these conditions need not hold in general.

<sup>12</sup>Some examples are:  $Y_{i_1, i_2}$  corresponds to export flows from country  $i_1$  to  $i_2$ , or whether there is a link between node  $i_1$  and  $i_2$  in a network.  $\{i_1, \dots, i_k\}$  can also correspond to the different dimensions of clustering.

- **jointly exchangeability**: the joint distribution of the data remains identical under any possible permutation of labels, i.e., labeling conveys no information
- **dissociation**: the variables are independent if they have **no unit** in common, that is  $Y_{i_1, i_2}$  must be independent of  $Y_{j_1, j_2}$  if  $\{i_1, i_2\} \cap \{j_1, j_2\} = \emptyset$

the dependence structure under such assumptions are

#### Lemma 5.5.2: Key Dependence Structure

Assumption 5.5.1 holds **if and only if** there exists i.i.d. variables  $(U_J)_{J \subset \mathbb{N}^+, 1 \leq |J| \leq k}$  and a measurable function  $\tau$  s.t. almost surely

$$Y_{\mathbf{i}} = \tau \left( \left( U_{\{\mathbf{i} \odot \mathbf{e}\}^+} \right)_{\mathbf{e} \in \bigcup_{r=1}^k \mathcal{E}_r} \right), \forall \mathbf{i} \in \mathbb{I}_k$$

this result is referred to as the AHK representation (Aldous 1981, Hoover 1979, Kallenberg 1989)<sup>a</sup>.

<sup>a</sup>Consider dyadic data ( $k = 2$ ), then for every  $i_1 < i_2$  (the ranking is precise),  $Y_{i_1, i_2} = \tau(U_{i_1}, U_{i_2}, U_{\{i_1, i_2\}})$ , that is, the outcome  $Y$  depends of factors specific to  $i_1$  and  $i_2$ , and factors relating both.

#### 5.5.1.2 Uniform LLN and CLT

Let  $\mathcal{F}$  denote a class of real-valued functions admitting a first moment w.r.t. the distribution  $P$ , let  $Pf$  denote the corresponding moment  $\mathbb{E}[f(Y_1)]$ , with  $\mathbf{1}$  as the  $k$ -tuple  $(1, \dots, k)$ . Assume that

#### Assumption 5.5.3: Measurability Assumption

$\exists$  a countable subclass  $\mathcal{G} \subset \mathcal{F}$  s.t. elements of  $\mathcal{F}$  are pointwise limits of sequences of elements of  $\mathcal{G}$

Consider

$$\mathbb{P}_n f = \frac{(n-k)!}{n!} \sum_{\mathbf{i} \in \mathbb{I}_{n,k}} f(Y_{\mathbf{i}})$$

$$\mathbb{G}_n f = \sqrt{n} (\mathbb{P}_n f - P f)$$

and the restrictions on  $\mathcal{F}$ : for any  $\eta > 0$  and any seminorm  $\|\cdot\|$  on a space containing  $\mathcal{F}$ , let

- $N(\eta, \mathcal{F}, \|\cdot\|)$ : the minimal number of  $\|\cdot\|$ -closed balls of radius  $\eta$  with centers in  $\mathcal{F}$  needed to cover  $\mathcal{F}$
- $N_{[]}(\eta, \mathcal{F}, \|\cdot\|)$ : the minimal number of  $\eta$ -brackets needed cover  $\mathcal{F}$ , where an  $\eta$ -bracket for  $f \in \mathcal{F}$  is a pair of functions  $(l, u)$  s.t.  $l \leq f \leq u$  and  $\|u - l\| < \eta$

Davezies et al. (2021) considered the seminorms  $\|f\|_{\mu, r} = \left( \int |f|^r d\mu \right)^{1/r}$  for any  $r \geq 1$  and probability measure (cdf)  $\mu$ . An envelope of  $\mathcal{F}$  is measurable function  $F$  satisfying  $F(u) \geq \sup_{f \in \mathcal{F}} |f(u)|$ , satisfying

#### Assumption 5.5.4: Assumptions of $\mathcal{F}$

**A** The class  $\mathcal{F}$

- (i) either admits an envelope  $F$  with  $PF < \infty$  and  $\forall \eta > 0$ ,

$$\sup_{Q \in \mathcal{Q}} N \left( \eta \|F\|_{Q,1}, \mathcal{F}, \|\cdot\|_{Q,1} \right) < \infty$$

(ii) or satisfies  $N_{[]}(\eta, \mathcal{F}, \|\cdot\|_{L_1(P)}) < \infty$  for all  $\eta > 0$

**B** and it

(i) **uniform entropy integral**: either admits an envelope  $F$  with  $PF^2 < \infty$  and

$$\int_0^\infty \sup_{Q \in \mathcal{Q}} \sqrt{\log N(\eta \|F\|_{Q,2}, \mathcal{F}, \|\cdot\|_{Q,2})} d\eta < \infty$$

(ii) **bracketing entropy integral**: or satisfies  $\int_0^\infty \sqrt{\log N_{[]}(\eta, \mathcal{F}, \|\cdot\|_{L_2(P)})} d\eta < \infty$

Assumption 5.5.4 are same as the conditions imposed on i.i.d. data for uniform LLNs and CLTs. Under these assumptions, Davezies et al. (2021) established the uniform LLNs and CLTs as

#### Theorem 5.5.5: Uniform LLNs and CLTs

Under Assumption 5.5.1 and 5.5.3,

- if (A) of Assumption 5.5.4 holds,  $\sup_{f \in \mathcal{F}} |\mathbb{P}_n f - P f| \xrightarrow{\text{a.s.}} 0$  and in  $L^1$
- if (B) of Assumption 5.5.4 holds,  $\mathbb{G}_n$  converges weakly in  $l^\infty(\mathcal{F})$  to a centered Gaussian process  $\mathbb{G}$  on  $\mathcal{F}$  as  $n \rightarrow \infty$ , the covariance kernel  $K$  of  $\mathbb{G}$  satisfies

$$K(f_1, f_2) = \frac{1}{(k-1)!^2} \sum_{(\pi, \pi') \in \mathfrak{S}(\{1\}) \times \mathfrak{S}(\{1'\})} \text{Cov}(f_1(Y_{\pi(1)}), f_2(Y_{\pi'(1')}))$$

Here, (A) of Assumption 5.5.4 is stronger than necessary to obtain the uniform LLNs. To establish the exact characterization, consider the norms:

$$\begin{aligned} \|f\|_{1,1} &= \frac{1}{n} \sum_{i_1=1}^n \left| \frac{1}{n-1} \sum_{i_2 \neq i_1} f(Y_{i_1, i_2}) + f(Y_{i_2, i_1}) \right| \\ \|f\|_{1,2} &= \frac{1}{n(n-1)} \sum_{1 \leq i_1 < i_2 \leq n} |\mathbb{E}[f(Y_{i_1, i_2}) + f(Y_{i_2, i_1})] \mid U_{\{i_1, i_2\}}| \end{aligned}$$

and the exact characterization is established as

#### Proposition 5.5.6: Exact Characterization of Uniform LLNs

Under Assumption 5.5.1 and 5.5.3, and  $\mathcal{F}$  admits an envelope  $F$  with  $PF < \infty$ , then

$$\sup_{f \in \mathcal{F}} |\mathbb{P}_n f - P f| \xrightarrow{\text{a.s.}} 0$$

**if and only if** both  $\log N(\epsilon, \mathcal{F}, \|\cdot\|_{1,2}) / n^2$  and  $\log N(\epsilon, \mathcal{F}, \|\cdot\|_{1,1}) / n$  tend to 0 in outer probability.

and 2 aspects of dissociated, exchangeable arrays are emphasized:

- **i.i.d. variations**: through the random entropy term related to  $\|\cdot\|_{1,2}$ , which only involves  $(U_{\{i_1, i_2\}})_{i \in \mathbb{I}_{n,2}}$
- **U-statistic**: through the random entropy term related to  $\|\cdot\|_{1,1}$ , up to negligible terms,  $\|f\|_{1,1}$  only depends on  $(U_{i_1})_{1 \leq i_1 \leq n}$

### 5.5.1.3 Convergence of the bootstrap process

Davezies et al. (2021), extending the pigeonhole bootstrap (McCullagh, 2000; Owen, 2007), established the following bootstrap process:

- 1  $n$  units are sampled independently in  $\{1, \dots, n\}$  with replacement and equal probability,  $W_i$  denotes the number of times unit  $i$  is sampled.
- 2 the  $k$ -tuple  $\mathbf{i} = (i_1, \dots, i_k) \in \mathbb{I}_{n,k}$  is then selected  $W_{\mathbf{i}} = \prod_{j=1}^k W_{i_j}$  times in the bootstrap sample

then consider  $\mathbb{P}_n^*$  and  $\mathbb{G}_n^*$  defined on  $\mathcal{F}$  by

$$\mathbb{P}_n^* f = \frac{(n-k)!}{n!} \sum_{\mathbf{i} \in \mathbb{I}_{n,k}} W_{\mathbf{i}} f(Y_{\mathbf{i}})$$

$$\mathbb{G}_n^* f = \sqrt{n} (\mathbb{P}_n^* f - \mathbb{P}_n f)$$

the validity of the bootstrap is then established as:

#### Theorem 5.5.7: Bootstrapping Validity

Under Assumption 5.5.1 and 5.5.3, if (B-i) of Assumption 5.5.4 also holds, the process  $\mathbb{G}_n^*$  converges weakly in  $l^\infty(\mathcal{F})$  to  $\mathbb{G}$ , conditional on  $(Y_{\mathbf{i}})_{\mathbf{i} \in \mathbb{I}_k}$  and outer almost surely.

The proof of this theorem boils down to proving

$$\sup_{h \in BL_1} \left| \mathbb{E} \left( h(\mathbb{G}_n^*) \mid (Y_{\mathbf{i}})_{\mathbf{i} \in \mathbb{I}_k} \right) - \mathbb{E}(h(\mathbb{G})) \right| \xrightarrow{\text{a.s. outer}} 0$$

where  $BL_1$  is the set of bounded and Lipschitz functions from  $l^\infty(\mathcal{F})$  to  $[0, 1]$ . With the standard bootstrap for i.i.d. data

$$\mathbb{E} [\mathbb{P}_n^*(f) \mid (Y_{\mathbf{i}})_{\mathbf{i} \in \mathbb{I}_k}] = \frac{1}{n^k} \sum_{\mathbf{i} \in \mathbb{I}_{n,k}} f(Y_{\mathbf{i}}) \xrightarrow{n \rightarrow \infty} \mathbb{P}_n f$$

hence, Davezies et al. (2021) established the a.s. conditional convergence of  $\sqrt{n} \left( \mathbb{P}_n^* f - \frac{1}{n^k} \sum_{\mathbf{i} \in \mathbb{I}_{n,k}} f(Y_{\mathbf{i}}) \right)$ .

### 5.5.1.4 Nonlinear estimators

Davezies et al. (2021) considered 2 classes of estimators: **Z-estimators** and **smooth functionals of the empirical cdf**:

**Z-estimators** Let

- $\Theta$  denote a normed space, endowed with norm  $\|\cdot\|_\Theta$
- $(\psi_{\theta,h})_{(\theta,h) \in \Theta \times \mathcal{H}}$  denote a class of real, measurable functions
- $\Psi(\theta)(h) = P\psi_{\theta,h}$ ,  $\Psi_n(\theta)(h) = \mathbb{P}_n \psi_{\theta,h}$ ,  $\Psi_n^*(\theta)(h) = \mathbb{P}_n^* \psi_{\theta,h}$
- for any real function  $g$  on  $\mathcal{H}$ ,  $\|g\|_{\mathcal{H}} = \sup_{h \in \mathcal{H}} |g(h)|$

The parameter of interest  $\theta_0$ , satisfying  $\Psi(\theta_0) = 0$ , is estimated by  $\hat{\theta} = \arg \min_{\theta \in \Theta} \|\Psi_n(\theta)\|_{\mathcal{H}}$ , define the bootstrap counterpart of  $\hat{\theta}$  as

$$\hat{\theta}^* = \arg \min_{\theta \in \Theta} \|\Psi_n^*(\theta)\|_{\mathcal{H}}$$

then we have the convergence

### Theorem 5.5.8: Convergence of Z-estimators Bootstrap

Under Assumption 5.5.1, if also

- 1  $\|\Psi(\theta_m)\|_{\mathcal{H}} \rightarrow 0 \Rightarrow \|\theta_m - \theta_0\|_{\Theta} \rightarrow 0, \forall (\theta_m)_{m \in \mathbb{N}} \in \Theta$
  - 2  $\{\psi_{\theta,h} : (\theta, h) \in \Theta \times \mathcal{H}\}$  satisfies Assumption 5.5.3 and (A) of 5.5.4, with  $PF < \infty$
  - 3  $\exists \delta > 0$  s.t.  $\{\psi_{\theta,h} : \|\theta - \theta_0\|_{\Theta} < \delta, h \in \mathcal{H}\}$  satisfies Assumption 5.5.3 and (B) of 5.5.4, with  $PF_{\delta}^2 < \infty$
  - 4  $\lim_{\theta \rightarrow \theta_0} \sup_{h \in \mathcal{H}} P(\psi_{\theta,h} - \psi_{\theta_0,h})^2 = 0$
  - 5  $\forall \eta > 0, \|\Psi_n(\hat{\theta})\|_{\mathcal{H}} = o_p(n^{-1/2})$  and  $P\left(\|\sqrt{n}\Psi_n^*(\hat{\theta}^*)\|_{\mathcal{H}} > \eta \mid (Y_i)_{i \in \mathbb{I}_k}\right) = o_p(1)$
  - 6  $\theta \mapsto \Psi(\theta)$  is Frechet-differentiable at  $\theta_0$ , with continuously invertible derivative  $\Psi_{\theta_0}$
- Then the convergence can be established as
- $\sqrt{n}(\hat{\theta} - \theta_0)$  converges in distribution to a centered Gaussian process  $\mathbb{G}$
  - conditional on  $(Y_i)_{i \in \mathbb{I}_k}$ ,  $\sqrt{n}(\hat{\theta}^* - \hat{\theta}) \xrightarrow{d} \mathbb{G}$  almost surely

**Smooth functionals of  $F_Y$**  For the cdf of  $Y_i$ , suppose that  $\mathcal{Y} \subset \mathbb{R}^p$  for some  $p \in \mathbb{N}^+$  and  $\theta_0 = g(F_Y)$ , where  $g$  is Hadamard differentiable<sup>13</sup>. Estimate  $\theta_0$  with  $\hat{\theta} = g(\hat{F}_Y)$ , where  $\hat{F}_Y$  denotes the empirical cdf of  $(Y_i)_{i \in \mathbb{I}_{n,k}}$ , and let  $\hat{\theta}^*$  denote the bootstrap counterpart of  $\hat{\theta}$ . Davezies et al. (2021) established the convergence results as

### Theorem 5.5.9: Convergence of Smooth Functionals of the Empirical CDF

Suppose that  $g$  is Hadamard differentiable at  $F_Y$  tangentially to a set  $\mathbb{D}_0$ , with derivative equal to  $g'_{F_Y}$ . Under Assumption 5.5.1,

- $\sqrt{n}(\hat{F}_Y - F_Y)$  converges weakly, as a process indexed by  $y$ , to a Gaussian process  $\mathbb{G}$  with kernel  $K$  satisfying

$$K(y_1, y_2) = \frac{1}{(k-1)!^2} \sum_{(\pi, \pi') \in \mathfrak{S}(\{1\}) \times \mathfrak{S}(\{1'\})} \text{Cov}\left(\mathbf{1}_{\{Y_{\pi(1)} \leq y_1\}}, \mathbf{1}_{\{Y_{\pi'(1')} \leq y_2\}}\right)$$

- If  $\mathbb{G} \in \mathbb{D}_0^a$  with probability 1,

$$\sqrt{n}(\hat{\theta} - \theta_0) \xrightarrow{d} \mathcal{N}\left(0, \mathbb{V}\left(g'_{F_Y}(\mathbb{G})\right)\right)$$

conditional on  $(Y_i)_{i \in \mathbb{I}_k}$ ,  $\sqrt{n}(\hat{\theta}^* - \hat{\theta}) \xrightarrow{d} \mathcal{N}\left(0, \mathbb{V}\left(g'_{F_Y}(\mathbb{G})\right)\right)$ , almost surely.

<sup>a</sup>In practice,  $\mathbb{D}_0$  often corresponds to the set of functions that are continuous everywhere or at a certain point  $y_0$ .

#### 5.5.1.5 Extensions

Davezies et al. (2021) also considered several extensions of the main results:

- **Degenerate cases:** consider the simple  $k = 2$  situations where  $K(f, f) = 0, \forall f \in \mathcal{F}$ . Generally, when  $K(f, f) = 0$ , the rate of convergence of  $\mathbb{P}_n f - P f$  is  $n^{-1}$  rather than  $n^{-1/2}$ , and the asymptotic distribution

<sup>13</sup>No linearity assumed under Hadamard differentiability.



is not necessarily normal.  $\forall (i_1, i_2) \in \mathbb{I}_2$ , let  $Y_{i_1, i_2} = \tau(U_{i_1}, U_{i_2}, U_{\{i_1, i_2\}})$  be the Aldous-Hoover-Kallenberg representation. WLoG, assume  $U$  to be uniform on  $[0, 1]$ .

Under a more stringent version of (B-i) Assumption 5.5.4, that is,  $\mathcal{F}$  admits an envelope  $F$  with  $PF^2 < \infty$  and

$$\int_0^\infty \sup_{Q \in \mathcal{Q}} \log N\left(\eta \|F\|_{Q,2}, \mathcal{F}, \|\cdot\|_{Q,2}\right) d\eta < \infty$$

Davezies et al. (2021) showed that for  $k = 2$ ,  $K(f, f) = 0$  for all  $f \in \mathcal{F}$ ,  $\sqrt{n}G_n$  converges **weakly** in  $l^\infty(\mathcal{F})$  to  $\mathbb{G}^d$ . However, the proposed bootstrap process does **not** generally converge to  $\mathbb{G}^d$ .

- **An alternative bootstrap process:** Davezies et al. (2021) established that under Assumption 5.5.1 and 5.5.3,  $PF^2 < \infty$  for all  $f \in \mathcal{F}$  and  $\mathcal{F}$  admits an envelope  $F$  s.t.  $PF^{1+\delta} < \infty$  for some  $\delta > 0$ , then if conditional on  $(Y_i)_{i \in \mathbb{I}_k}$ , the process  $\mathbb{G}_n^*$  outer almost surely converges weakly in  $l^\infty(\mathcal{F})$  to a centered Gaussian process  $\mathbb{G}$ , the process  $\mathbb{G}_n$  also converges weakly in  $l^\infty(\mathcal{F})$  to  $\mathbb{G}$ . Combined with Theorem 5.5.7, we have

$$\mathbb{G}_n \rightarrow \mathbb{G} \Leftrightarrow \mathbb{G}_n^* \xrightarrow{\text{a.s.-outer}} \mathbb{G}$$

consider an alternative, the multiplier bootstrap process adapted to jointly exchangeable arrays. Let  $(\xi_i)_{i=1}^n$  be a sequence of i.i.d. random variables, independent from the original data  $(Y_i)_{i \in \mathbb{I}_{n,2}}$ , then the process

$$\mathbb{G}_n^{m*} : f \mapsto \frac{1}{\sqrt{n}} \sum_{i_1=1}^n \xi_{i_1} \left( \frac{1}{n-1} \sum_{1 \leq i_2 \neq i_1 \leq n} [f(Y_{i_1, i_2}) + f(Y_{i_2, i_1})] - 2\mathbb{P}_n f \right)$$

also outer almost surely converges weakly in  $l^\infty(\mathcal{F})$  to  $\mathbb{G}$ , just as the proposed process  $\mathbb{G}_n^*$ .

- **Separately exchangeable arrays** For the case of separately exchangeable arrays (the  $n$  units stem from  $k$  different populations, or multiway clustering as in Menzel (2021)), we must assume a stronger version of Assumption 5.5.1

**Assumption 5.5.10: Stronger assumptions for separately exchangeable arrays**

Consider random variables  $Y_i$  where  $\mathbf{i} = (i_1, \dots, i_k) \in \mathbb{N}^{+k}$ , implying that repetitions are allowed. Then assume that for any  $(\pi_1, \dots, \pi_k) \in \mathfrak{S}(\mathbb{N}^+)^k$ ,

$$(Y_i)_{i \in \mathbb{N}^{+k}} \stackrel{d}{=} (Y_{\pi_1(i_1), \dots, \pi_k(i_k)})_{i \in \mathbb{N}^{+k}}$$

and for any  $A, B$ , disjoint subsets of  $\mathbb{N}^+$ ,  $(Y_i)_{i \in A^k}$  is independent of  $(Y_i)_{i \in B^k}$ .

Under this stronger assumption, we have equality in distribution even for  $\pi_1 = \dots = \pi_k$ . Here, denote

- $\mathbf{1} = (1, \dots, 1)$
- $\mathbf{n} = (n_1, \dots, n_k)$ , where  $n_j \geq 1$  denotes the number of units observed in population (or cluster)  $j$ , in general,  $n_j \neq n_{j'}$  for  $j \neq j'$
- sample at hand:  $(Y_i)_{1 \leq i \leq \mathbf{n}}$ , where  $\mathbf{i} \geq \mathbf{i}'$  means that  $i_j \geq i'_j, \forall j = 1, \dots, k$ .
- $\underline{n} = \min(n_1, \dots, n_k)$

then the empirical measure and empirical process for separately exchangeable arrays are

$$\mathbb{P}_n f = \frac{1}{\prod_{j=1}^k n_j} \sum_{1 \leq \mathbf{i} \leq \mathbf{n}} f(Y_i) \quad \mathbb{G}_n f = \sqrt{\underline{n}} (\mathbb{P}_n f - P f)$$

Consider the **pigeonhole bootstrap** process (McCullagh, 2000), which is close the bootstrap in Theorem 5.5.7, except that weights are now independent from one coordinate to another:

- 1 For each  $j \in \{1, \dots, k\}$ ,  $n_j$  elements are sampled with replacement and equal probability in the set  $\{1, \dots, n_j\}$ . For each  $i_j$ , let  $W_{i_j}^j$  denote the number of times  $i_j$  is selected
  - 2  $k$ -tuple  $\mathbf{i} = (i_1, \dots, i_k)$  is then selected  $W_{\mathbf{i}} = \prod_{j=1}^k W_{i_j}^j$  times in the bootstrap sample
- the bootstrap process  $\mathbb{G}_{\mathbf{n}}^*$  is then defined on  $\mathcal{F}$  by

$$\mathbb{G}_{\mathbf{n}}^* f = \sqrt{\underline{n}} \left( \frac{1}{\prod_{j=1}^k n_j} \sum_{1 \leq \mathbf{i} \leq \mathbf{n}} (W_{\mathbf{i}} - 1) f(Y_{\mathbf{i}}) \right)$$

as with multisample U-statistics, assume an index  $m \in \mathbb{N}^+$  and increasing functions  $g_1, \dots, g_k$  s.t. for all  $j$ ,  $n_j = g_j(m) \xrightarrow{m \rightarrow \infty} \infty$ , and w.l.o.g.,  $\forall m \in \mathbb{N}^+, \exists j$  s.t.  $g_j(m+1) > g_j(m)$ , then

**Theorem 5.5.11: Bootstrap Convergence: Separately Exchangeable Arrays**

Under Assumption 5.5.3 and 5.5.10 and for every  $j = 1, \dots, k$ ,  $\exists \lambda_j \geq 0$  s.t.  $\underline{n}/n_j \rightarrow \lambda_j \geq 0$ , then

- 1 If (A) of Assumption 5.5.4 holds,  $\sup_{f \in \mathcal{F}} |\mathbb{P}_{\mathbf{n}} f - P f| \xrightarrow{\text{a.s.}} 0$  and in  $L^1$
- 2 If (B-i) of Assumption 5.5.4 holds, the process  $\mathbb{G}_n$  converges weakly in  $l^\infty(\mathcal{F})$  to a centered Gaussian process  $\mathbb{G}_\lambda$  on  $\mathcal{F}$  as  $n \rightarrow \infty$ , and the covariance kernel  $K_\lambda$  of  $\mathbb{G}_\lambda$  satisfies

$$K_\lambda(f_1, f_2) = \sum_{j=1}^k \lambda_j \text{Cov}(f_1(Y_{\mathbf{1}}), f_2(Y_{\mathbf{2}_j}))$$

where  $\mathbf{2}_j$  is the  $k$ -tuple with 2 in each entry but 1 in entry  $j$ .

- 3 If (B-i) of Assumption 5.5.4 holds,  $\mathbb{G}_n^* \rightarrow \mathbb{G}_\lambda$  weakly, conditional on  $(Y_{\mathbf{i}})_{\mathbf{i} \in \mathbb{N}^{+k}}$  and outer almost surely

Here, the case where  $\lambda_j = 0$  for some  $j$  corresponds to **strongly unbalanced** designs with different rates of convergence to  $\infty$  along the different dimensions of the array. In such case, only the dimensions with the slowest rate of convergence contribute to the asymptotic distribution.

## 5.5.2 CLT for the estimator with heterogeneous clusters

Yap (2023) provides general conditions such that the plug-in mean estimator is asymptotically normal, and the Cameron et al. (2011) variance estimator is consistent even when clusters are heterogeneous. The conditions mimic one-way clustering conditions, assuming that two observations are independent when they do not share any cluster, and **not** assuming separate exchangeability.

Consider for vectors  $\{\mathbf{W}_i\}_{i=1}^n$ , where  $\mathbf{W}_i := (W_{i1}, \dots, W_{iK})' \in \mathbb{R}^K$  and  $i$  is the unit of observation, for the population of size  $n$ . The goal is to establish a central limit theorem (CLT) for a weighted sum of the random vector,  $\sum_i \omega_i \mathbf{W}_i$  where  $\omega_i$  are non-stochastic scalar weights as  $n \rightarrow \infty$ .

**Notation** Consider 2 clustering dimension  $G$  and  $H$ , then

- $g(i), h(i)$ : the cluster where observations  $i$  belongs on the  $G$  and  $H$  dimensions, respectively
- **partition**: For  $C \in \{G, H\}$ , let  $\mathcal{N}_C^C$  denote the set of observations in cluster  $c$  on dimension  $C$
- $N_C^C = |\mathcal{N}_C^C|$ : the cluster size for  $C \in \{G, H\}$ , and  $N_{gh} := |\mathcal{N}_g^G \cap \mathcal{N}_h^H|$

**Assumptions** Several assumptions are imposed to establish the main result

**Assumption 5.5.12: Assumptions of Yap (2023)**

- 1 **dependence structure:**  $\mathbf{W}_i \perp \mathbf{W}_j$  if  $g(i) \neq g(j)$  and  $h(i) \neq h(j)$
- 2 **additional assumptions:** For  $C \in \{G, H\}$  and  $k \in \{1, 2, \dots, K\}$ ,  $\exists K_0 < \infty$  s.t.
  - $\forall i, \mathbb{E}[W_{ik}^4] \leq K_0$ : bounded fourth moment, stronger than the one-way clustering condition.
  - $\frac{1}{\lambda_n} \max_c \left( \sum_{i \in \mathcal{N}_c^c} |\omega_i| \right)^2 \rightarrow 0$ : the cluster with the largest weight to have relatively **small** contribution to the total variance, s.t. removing one cluster does not change the variance substantively, allowing the ratio of the size of any 2 clusters to diverge.
  - $\frac{1}{\lambda_n} \sum_c \sum_{i,j \in \mathcal{N}_c^c} A_{ij} |\omega_i \omega_j| \leq K_0$ : ruling out the purely interactive model in Menzel (2021)

The key feature of Assumption 5.5.12 (1) is that it is agnostic about the dependent structure when  $\mathbf{W}_i$  and  $\mathbf{W}_j$  share at least one cluster. The DGP can be arbitrarily heterogeneous across different clusters. Consider a 0 – 1 indicator:

$$A_{ij} := \mathbf{1}[\mathbf{W}_i \not\perp \mathbf{W}_j]$$

then we have  $A_{ij} = A_{ji}$  and  $A_{ii} = 1$ . Yap (2023) established the following results

**Theorem 5.5.13: Asymptotic Normality of Yap (2023)**

Under Assumption 5.5.12,

$$Q_n^{-1/2} \sum_{i=1}^n \omega_i (\mathbf{W}_i - \mathbb{E}[\mathbf{W}_i]) \xrightarrow{d} \mathcal{N}(0, \mathbf{I}_K)$$

further

- If  $\mathbb{E}[\mathbf{W}_i] = 0, \forall i$ , then  $Q_n^{-1} \hat{Q}_n \xrightarrow{P} \mathbf{I}_K$ , where  $\hat{Q}_n := \sum_i \sum_{j \in \mathcal{N}_i} \omega_i \omega_j \mathbf{W}_i \mathbf{W}_j'$
- If  $\mathbb{E}[\mathbf{W}_i] = \mu, \forall i^a$  and  $\frac{1}{\lambda_n} \sum_c \sum_{i,j \in \mathcal{N}_c^c} |\omega_i \omega_j| \leq K_0$  for some  $K_0 < \infty$ , then

$$\bar{\mathbf{W}} \xrightarrow{P} \mu \qquad Q_n^{-1} \hat{Q}_n \xrightarrow{P} \mathbf{I}_K$$

for  $\hat{\mathbf{W}} = (\sum_i \omega_i \mathbf{W}_i) / (\sum_j \omega_j)$  and  $\hat{Q}_n := \sum_i \sum_{j \in \mathcal{N}_i} \omega_i \omega_j (\mathbf{W}_i - \bar{\mathbf{W}}) (\mathbf{W}_j - \bar{\mathbf{W}})'$

one-way clustering is a special case of these results when one dimension is weakly nested within the other.

<sup>a</sup>Here, the variance estimator need not be consistent, or even conservative. See Remark 2 of Yap (2023) for details.

### 5.5.3 One-way robust clustering with large clusters

Sasaki and Wang (2022) established robust clustering for the cases wher the distribution of cluster sizes follows a power law with exponent less than 2, the conventional clustering method fails<sup>14</sup>. Consider again the linear model

$$Y_{gi} = \mathbf{X}_{gi}' \boldsymbol{\theta} + U_{gi} \qquad \mathbb{E}[\mathbf{U}_g \mid \mathbf{X}_g] = 0$$

<sup>14</sup>That is,  $\sup_g N_g/N \rightarrow 0$ . An example is that California consists of 10% of total sample among the 51 US states.

with a clustered sample  $\left\{ \left( Y_{gi}, \mathbf{X}'_{gi} \right)' \right\}_{i=1}^{N_g}$ . The OLS estimator gives

$$\hat{\theta} = \left( \sum_{g=1}^G \sum_{i=1}^{N_g} \mathbf{X}_{gi} \mathbf{X}'_{gi} \right)^{-1} \left( \sum_{g=1}^G \sum_{i=1}^{N_g} \mathbf{X}_{gi} Y_{gi} \right) = \left( \sum_{g=1}^G \mathbf{X}'_g \mathbf{X}_g \right)^{-1} \left( \sum_{g=1}^G \mathbf{X}'_g \mathbf{Y}_g \right)$$

and the commonly cluster-robust variance estimators take the form of

$$\hat{V}_{\hat{\theta}}^{\text{CR}} = a_n \left( \sum_{g=1}^G \mathbf{X}'_g \mathbf{X}_g \right)^{-1} \left( \sum_{g=1}^G \hat{\mathbf{S}}_g \hat{\mathbf{S}}'_g \right) \left( \sum_{g=1}^G \mathbf{X}'_g \mathbf{X}_g \right)^{-1}$$

where  $a_n \rightarrow 1$  is a suitable finite-sample adjustment, and  $\hat{\mathbf{S}}_g = \sum_i^{N_g} = 1 \mathbf{X}_{gi} \hat{U}_{gi}$  with  $\hat{U}_{gi} = Y_{gi} - \mathbf{X}'_{gi} \hat{\theta}$ . Conventionally, the adjustment is

$$a_n = \left( \frac{\sum_{g=1}^G N_g - 1}{\sum_{g=1}^G N_g - p} \right) \left( \frac{G}{G-1} \right)$$

with  $p$  denoting the dimension of  $\mathbf{X}_{gi}$ . Asymptotically,

$$\sqrt{G} (\hat{\theta} - \theta) = \left( \underbrace{\frac{1}{G} \sum_{g=1}^G \sum_{i=1}^{N_g} \mathbf{X}_{gi} \mathbf{X}'_{gi}}_{\xrightarrow{P} \mathbf{Q}} \right)^{-1} \left( \underbrace{\frac{1}{\sqrt{G}} \sum_{g=1}^G \sum_{i=1}^{N_g} \mathbf{X}_{gi} U_{gi}}_{\xrightarrow{d} \mathcal{N}(0, \mathbf{V})} \right) \xrightarrow{d} \mathcal{N}(\mathbf{0}, \mathbf{Q}^{-1} \mathbf{V} \mathbf{Q}^{-1})$$

where  $\mathbf{Q} = \mathbb{E} \left[ \sum_{i=1}^{N_g} \mathbf{X}_{gi} \mathbf{X}'_{gi} \right]$ ,  $\mathbf{V} = \text{Var} \left[ \sum_{i=1}^{N_g} \mathbf{X}_{gi} U_{gi} \right]$  and  $\sum_{i=1}^{N_g} \mathbf{X}_{gi} U_{gi}$  have finite second moments.

**Main results** for a given  $k \in \{1, \dots, p\}$ , let  $\sum_g$  and  $Z_{gi}$  be the  $k$ -th coordinate of  $\sum_{i=1}^{N_g} \mathbf{X}_{gi} U_{gi}$  and the  $k$ -th coordinate of  $\mathbf{X}_{gi} U_{gi}$  respectively. Consider the following property, **regularly varying (RV)**, of a distribution function  $F$ ,

$$\frac{1 - F(xt)}{1 - F(t)} \xrightarrow{t \rightarrow \infty} x^{-\alpha}, \quad \forall x > 0$$

for some constant  $\alpha > 0$ , which is referred to as the **tail exponent**, measuring the tail heaviness of  $F$ . Then, let

- $F$  denote the marginal distribution of  $Z_{gi}$
- $C^n$  denote, for each  $n \geq 2$ , the copula s.t.

$$\mathbb{P}(Z_{g1} \leq z_1, \dots, Z_{gn} \leq z_n) = C^n(F(z_1), \dots, F(z_n))$$

Sasaki and Wang (2022) make the following assumptions

**Assumption 5.5.14: Assumptions on the Distribution of  $\{Z_{gi}\}_{g,i}$** 

- $\{Z_{gi}\}_{g,i}$  is identically distributed with CDF  $F$ , which is **RV** at infinity with  $\alpha > 1^a$
- The copula density  $c(u_1, \dots, u_n) = \partial^n C(u_1, \dots, u_n) / \partial u_1 \dots \partial u_n$  exists and is **uniformly bounded**<sup>b</sup>
- $N_g$  is independent of  $(Z_{g1}, Z_{g2}, \dots)$  and its distribution  $H$  is **RV** at infinity with  $\beta > 1^c$

<sup>a</sup> $Z_{gi}$  should have a finite mean, and a regularly varying tail (satisfied by many common heavy-tailed distributions including Pareto, Student-t, Cauchy, F). The tail condition characterize the moment conditions as

$$\mathbb{E}[|Z_{gi}|^r] = \infty \forall r < \alpha$$

$$\mathbb{E}[|Z_{gi}|^r] < \infty \forall r > \alpha$$

<sup>b</sup>Allowing dependence among  $Z_{g1}, \dots, Z_{gN_g}$  within each cluster  $g$

<sup>c</sup>The distribution of the cluster size  $N_g$  is also regularly varying. One way to think about this is considering  $N_g$  as the integer part of some continuous random variable with a regularly varying tail.

Under Assumption 5.5.14, **Sasaki and Wang (2022)** establish that

**Theorem 5.5.15: When Conventional Robust Clustering Fails**

If  $\beta < \alpha$ , then  $\forall z > 0$ , as  $t \rightarrow \infty$

$$\frac{\mathbb{P}(\sum_g > zt)}{\mathbb{P}(\sum_g > t)} = z^{-\beta}$$

here, the tail of the distribution of  $N_g$  dominates that of  $Z_{gi}$ , the tail heaviness of the summation  $\sum_g = \sum_{i=1}^{N_g} Z_{gi}$  is dictated by that of  $N_g$ . Therefore, even if  $Z_{gi}$  has a finite  $r$ -th moment for  $r < \alpha$ , the  $r$ -th moment of  $\sum_g$  might still be infinite if  $r > \beta$ . Hence, the conventional robust clustering may fail to exist even if  $Z_{gi}$  has a bounded second moment. **Sasaki and Wang (2022)** illustrated this problem with state clustering of PSID and some recent studies published on Econometrica. The way they choose to show that  $\beta < 2$  is by using the Hill plot (**Drees et al., 2000**).

**Proposed alternative estimators** **Sasaki and Wang (2022)** proposed a simple fix as

$$\hat{\theta}^{\text{WCR}} = \left( \sum_{g=1}^G N_g^{-1} \sum_{i=1}^{N_g} \mathbf{x}_g i \mathbf{x}'_{gi} \right)^{-1} \left( \sum_{g=1}^G N_g^{-1} \sum_{i=1}^{N_g} \mathbf{x}_{gi} Y_{gi} \right) = \left( \sum_{g=1}^G N_g^{-1} \mathbf{X}' \mathbf{X} \right)^{-1} \left( \sum_{g=1}^G N_g^{-1} \mathbf{X}'_g \mathbf{y}_g \right)$$

$$\hat{V}_{\hat{\theta}}^{\text{WCR}} = a_n \left( \sum_{g=1}^G N_g^{-1} \mathbf{X}'_g \mathbf{X}_g \right)^{-1} \left( \sum_{g=1}^G N_g^{-2} \hat{\mathbf{s}}_g \hat{\mathbf{s}}'_g \right) \left( \sum_{g=1}^G N_g^{-1} \mathbf{X}'_g \mathbf{X}_g \right)^{-1}$$

where  $a_n \rightarrow 1$ . Again, for a given  $k \in \{1, \dots, p\}$ , let  $\sum_g$  and  $Z_{gi}$  be the  $k$ -th coordinate of  $\sum_{i=1}^{N_g} \mathbf{x}_{gi} U_{gi}$  and the  $k$ -th coordinate of  $\mathbf{x}_{gi} U_{gi}$  respectively, and then similar to Thm.5.5.15,

**Theorem 5.5.16: Sasaki and Wang (2022)'s Modification**

Under Assumption 5.5.14,  $\forall z > 0$ , as  $t \rightarrow \infty$ ,  $\frac{\mathbb{P}(\tilde{\sum}_g > zt)}{\mathbb{P}(\tilde{\sum}_g > t)} = z^{-\beta}$  where  $\tilde{\sum}_g = \sum_g / N_g = \sum_{i=1}^{N_g} Z_{gi}$

Thm.5.5.16 established that  $\tilde{\sum}_g$  has the same tail as the original score  $Z_{gi}$ . It requires the second moments of

the score  $\mathbf{X}_{gi}U_{gi}$  is **bounded** regardless of whether the second moment of  $N_g$  is finite or not.

**Theorem 5.5.17: CLT of Sasaki and Wang (2022)'s Clustering**

Under Assumption 5.5.14, and in addition, assume that  $(N_g, \mathbf{X}_g, \mathbf{X}_g)$  is i.i.d. across  $g^a$ , and  $\mathbb{E} \left[ N_g^{-1} \mathbf{X}_g' \mathbf{X}_g \right]$  is non-singular<sup>b</sup>. With  $\alpha > 2$ , then

$$\sqrt{G} \left( \hat{\boldsymbol{\theta}}^{\text{WCR}} - \boldsymbol{\theta} \right) \xrightarrow{P} \mathcal{N} \left( \mathbf{0}, \mathbf{V}^{\text{WCR}} \right)$$

as  $G \rightarrow \infty$ , where

$$\mathbf{V}^{\text{WCR}} = \left( \mathbb{E} \left[ N_g^{-1} \mathbf{X}_g' \mathbf{X}_g \right] \right)^{-1} \left( \mathbb{E} \left[ N_g^{-2} \mathbf{X}_g' \mathbf{X}_g \mathbf{U}_g^2 \right] \right) \left( \mathbb{E} \left[ N_g^{-1} \mathbf{X}_g' \mathbf{X}_g \right] \right)^{-1}$$

and  $G \hat{\mathbf{V}}^{\text{WCR}} \xrightarrow{P} \mathbf{V}_{\hat{\boldsymbol{\theta}}}^{\text{WCR}}$  as  $G \rightarrow \infty$ .

<sup>a</sup>Requiring the i.i.d. sampling **across** clusters, while allowing for arbitrary dependence within each cluster.

<sup>b</sup>Ruling out multi-collinearity.

Thm. 5.5.13 gives that the proposed clustering approach based on  $\hat{\boldsymbol{\theta}}^{\text{WCR}}$  works as far as  $\alpha > 2$  is true, regardless of whether  $\beta < 2$  is true or not.

### 5.5.4 Two-Way Robust Clustering: Justification

Chiang and Sasaki (2023) developed a CLT for means of two-way clustered triangular arrays under mild conditions, which provides a theoretical justification for the asymptotic gaussianity of the statistics commonly occurs under two-way clustering.

Consider

$$D_{it} = f_{NT}(\alpha_i, \gamma_t, \epsilon_{it})$$

normalize it, w.l.o.g., to  $\mathbb{E}[D_{it}] = 0$ , define

$$\hat{\theta}_{NT} = \frac{1}{NT} \sum_{i=1}^N \sum_{t=1}^T D_{it}$$

the goal is to show the **asymptotic Gaussianity** of  $\hat{\theta}_{NT}$ . The Hoeffding-type decomposition gives

$$\hat{\theta}_{NT} = \underbrace{\sum_{i=1}^N a_i}_{:=L_{NT}} + \underbrace{\sum_{t=1}^T b_t}_{:=W_{NT}} + \underbrace{\sum_{i=1}^N \sum_{t=1}^T w_{it} + \sum_{i=1}^N \sum_{t=1}^T r_{it}}_{:=R_{NT}}$$

where

$$\begin{aligned} a_i &= \frac{1}{N} \mathbb{E}[D_{it} \mid \alpha_i] & b_t &= \frac{1}{T} \mathbb{E}[D_{it} \mid \gamma_t] \\ w_{it} &= \frac{1}{NT} (\mathbb{E}[D_{it} \mid \alpha_i, \gamma_t] - \mathbb{E}[D_{it} \mid \alpha_i] - \mathbb{E}[D_{it} \mid \gamma_t]) & r_{it} &= \frac{1}{NT} (D_{it} - \mathbb{E}[D_{it} \mid \alpha_i, \gamma_t]) \end{aligned}$$

straightforwardly,

$$\begin{aligned}\mathbb{E}[a_i] &= \mathbb{E}[b_t] = \mathbb{E}[w_{it}] = \mathbb{E}[r_{it}] = 0 \\ \mathbb{E}[w_{it} | \alpha_i] &= \mathbb{E}[w_{it} | \gamma_t] = 0 \\ \mathbb{E}[a_i w_{it}] &= \mathbb{E}[a_i r_{it}] = \mathbb{E}[\gamma_t w_{it}] = \mathbb{E}[\gamma_t r_{it}] = 0\end{aligned}$$

here,  $W_{NT} = \sum_{i=1}^N \sum_{t=1}^T w_{it}$  is the potentially non-gaussian part, as discussed by [Menzel \(2021\)](#). [Chiang and Sasaki \(2023\)](#) argue that if the DGP is treated as a triangular array,  $W_{NT}$  is often asymptotically gaussian. For a simple, generic DGP

$$D_{it} = \alpha_{i0} + \gamma_{t0} + \sum_{j=1}^J \lambda_j \alpha_{ij} \gamma_{tj} + \epsilon_{ij}$$

where  $\{\alpha_{ij}\}_{j=0}^J$  are  $i$ -specific latent factors,  $\{\gamma_{tj}\}_{j=0}^J$  are  $t$ -specific latent factors,  $\epsilon_{it}$  is an idiosyncratic component. Suppose that the factor loading  $\lambda_j$  are non-zero, then [Chiang and Sasaki \(2023\)](#) provide a summary on when to use the TWCR standard error for  $\hat{\theta} = (NT)^{-1} \sum_{i=1}^N \sum_{t=1}^T D_{it}$ : Except for the extreme

Table 5.1: [Chiang and Sasaki \(2023\)](#)'s Summary on TWCR SE Validity

Small $J$	$\alpha_{i0}$	$\gamma_{t0}$	$\epsilon_{it}$	TWCR SE valid?
Yes	Degenerate	Degenerate	Degenerate	<b>No</b>
	Other cases			<b>Yes</b>

case where **the number of factors  $J$  is small**, latent factors on both dimension  $i$  and  $t$  are degenerate, the idiosyncratic component  $\epsilon_{it}$  is also degenerate.

## 5.6 Clustering in Experiments

### 5.6.1 When to cluster

[Abadie et al. \(2023\)](#) highlight 3 common misconceptions on clustering adjustments:

- clustering when there is a nonzero correlation between residuals for units within the same cluster
- when clustering makes a difference, one should cluster
- either fully adjust for clustering, or not at all

and they propose a new design-based clustering framework to address the overuse and overconservatism of the commonly used clustering.

They propose a framework with **3 sources** of sampling variations:

- 1 variation across samples in which units are observed in each cluster
- 2 variation in which clusters are observed
- 3 variation in the treatment assignment across units

Standard framework for clustering focuses on the first 2 sources of uncertainty. How much the 3 sources of variations matter depends on **the sampling process**, **the assignment process**, and **the heterogeneity** in the treatment effects across clusters.

### 5.6.1.1 Sampling process

Consider a **sequence of populations** indexed by  $k$ , and

- the  $k$ -th population has  $n_k$  units, indexed by  $i = 1, \dots, n_k$
- the population is partitioned into  $m_k$  clusters, and  $m_{k,i} \in \{1, \dots, m_k\}$  denote the cluster to which unit  $i$  of population  $k$  belongs
- number of units in cluster  $m$  of population  $k$  is  $n_{k,m} \geq 1$
- 2 potential outcomes: treated  $y_{k,i}(1)$ , control  $y_{k,i}(0)$

Thus, the population is characterized by triples  $(m_{k,i}, y_{k,i}(0), y_{k,i}(1))$ , for units  $1, \dots, n_k$  and clusters  $1, \dots, m_k$ . Then,

- **population ATE**

$$\tau_k = \frac{1}{n_k} \sum_{i=1}^{n_k} (y_{k,i}(1) - y_{k,i}(0))$$

- **population ATE by cluster**

$$\tau_{k,m} = \frac{1}{n_{k,m}} \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} (y_{k,i}(1) - y_{k,i}(0))$$

naturally. we have

$$\tau_k = \sum_{m=1}^{m_k} \frac{n_{k,m}}{n_k} \tau_{k,m}$$

For unit  $i$  in population  $k$ , there are 2 components of the stochastic nature:

- **Sampling process**: a random variable  $R_{k,i} = 1$  if unit  $i$  belongs to the sample, 0 if not. The sampling process is independent of the potential outcomes and the assignments, and consists of 2 stages
  - clusters are sampled with cluster sampling probability  $q_k \in (0, 1]$ 
    - \*  $q_k = 1$ : sample all clusters, i.e. **random sampling**
    - \*  $q_k < 1$ : **clustered sampling**
    - \*  $q_k \rightarrow 0$ : only a small fraction of the clusters sampled
  - units are sampled from the subpopulation consisting of all the sampled clusters, with probability  $p_k \in (0, 1]$ 
    - \*  $p_k = 1$ : sample all units in the population
    - \*  $p_k \rightarrow 0$ : only a small sample of units sampled
- **Assignment process** the stochastic treatment indicator  $W_{k,i} \in \{0, 1\}$  is determined by a 2-stage process as well:
  - for cluster  $m$  in population  $k$ , **randomly** draw an assignment probability  $A_{k,m} \in [0, 1]$  from a distribution with  $\mu_k$  (bounded away from 0 and 1 uniformly in  $k$ ), variance  $\sigma_k^2$ , **independently** for each cluster
    - \*  $\sigma_k^2 = 0$ :  $A_{k,m}$  is the same across all clusters, i.e., **random assignment**
    - \*  $\sigma_k^2 > 0$ : assignment probabilities depend on clustering, and
      - **clustered assignment**,  $\sigma_k^2 = \mu_k(1 - \mu_k)$ : **no within-cluster** variation in  $W_{k,i}$ <sup>15</sup>
      - **partially clustered assignment**,  $0 < \sigma_k^2 < \mu_k(1 - \mu_k)$ : assignment depends on cluster but not all units in the same cluster necessarily have the same value of  $W_{k,i}$
  - each unit in cluster  $m$  is assigned to the treatment independently, with cluster-specific probability  $A_{k,m}$

<sup>15</sup>This is the upper bound of  $\sigma_k^2$ , attained when  $A_{k,m}$  can only take the values 0 or 1



### 5.6.1.2 LS estimator and variance

Let

$$N_{k,1} = \sum_{i=1}^{n_k} R_{k,i} W_{k,i} \quad N_{k,0} = \sum_{i=1}^{n_k} R_{k,i} (1 - W_{k,i})$$

be the number of **treated** and **untreated** units in the sample (both randomly variables), total sample size is then  $N_k = N_{k,1} + N_{k,0}$ , then for the regression,

$$Y_{k,i} = \alpha + \tau_k W_{k,i} + \epsilon_{k,i}$$

the OLS estimator of  $\beta$  is equal to the difference in means<sup>16</sup>:

$$\hat{\tau}_k = \frac{1}{N_{k,1} \vee 1} \sum_{i=1}^{n_k} R_{k,i} W_{k,i} Y_{k,i} - \frac{1}{N_{k,0} \vee 1} \sum_{i=1}^{n_k} R_{k,i} (1 - W_{k,i}) Y_{k,i}$$

Abadie et al. (2023) assume that

- $m_k q_k \rightarrow \infty$ : the expected number of sampled clusters goes to infinity
- $\liminf_{k \rightarrow \infty} p_k \min_m n_{k,m} > 0$ : average number of observations sampled per cluster (conditional on sampled) does not go to 0
- $\limsup_{k \rightarrow \infty} \frac{\max_m n_{k,m}}{\min_m n_{k,m}} < \infty$ : the imbalance between the number of units across clusters is bounded

**Large  $k$  distribution of  $\hat{\tau}_k$**  Let

$$\alpha_k = \frac{1}{n_k} \sum_{i=1}^{n_k} y_{k,i}(0) \quad u_{k,i}(1) = y_{k,i} - (\alpha_k + \tau_k) \quad u_{k,i}(0) = y_{k,i}(0) - \alpha_k$$

then

$$\frac{\sqrt{N_k} (\hat{\tau}_k - \tau_k)}{\sqrt{v_k}} \xrightarrow{d} \mathcal{N}(0, 1)$$

where

$$\begin{aligned} v_k = & \frac{1}{n_k} \sum_{i=1}^{n_k} \left( \frac{u_{k,i}^2(1)}{\mu_k} + \frac{u_{k,i}^2(0)}{1 - \mu_k} \right) \\ & - p_k \frac{1}{n_k} \sum_{i=1}^{n_k} (u_{k,i}(1) - u_{k,i}(0))^2 - p_k \sigma_k^2 \frac{1}{n_k} \sum_{i=1}^{n_k} \left( \frac{u_{k,i}(1)}{\mu_k} + \frac{u_{k,i}(0)}{1 - \mu_k} \right)^2 \\ & + p_k (1 - q_k) \frac{1}{n_k} \sum_{m=1}^{m_k} \left( \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} (u_{k,i}(1) - u_{k,i}(0)) \right)^2 \\ & + p_k \sigma_k^2 \frac{1}{n_k} \sum_{m=1}^{m_k} \left( \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} \left( \frac{u_{k,i}(1)}{\mu_k} + \frac{u_{k,i}(0)}{1 - \mu_k} \right) \right)^2 \end{aligned}$$

Under the following cases:

---

<sup>16</sup>  $\leftarrow a \vee b = \max\{a, b\}$

- **random sampling** ( $q_k = 1$ ) and **random assignment** ( $\sigma_k^2 = 0$ ): then simply  $v_k$  as

$$v_k = \underbrace{\frac{1}{n_k} \sum_{i=1}^{n_k} \left( \frac{u_{k,i}^2(1)}{\mu_k} + \frac{u_{k,i}^2(0)}{1-\mu_k} \right)}_{\text{robust variance estimator}} - \underbrace{p_k \frac{1}{n_k} \sum_{i=1}^{n_k} (u_{k,i}(1) - u_{k,i}(0))^2}_{\text{finite-sample correction}}$$

here, the finite-sample correction vanishes if there is either **no heterogeneity** in the treatment effects ( $u_{k,i}(1) - u_{k,i}(0) = y_{k,i}(1) - y_{k,i}(0) - \tau_k = 0$ ), or the sample is **small** ( $p_k \simeq 0$ )

- **clustered sampling** ( $q_k < 1$ ), the component

$$p_k (1 - q_k) \frac{1}{n_k} \sum_{m=1}^{m_k} \left( \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} (u_{k,i}(1) - u_{k,i}(0)) \right)^2 = p_k (1 - q_k) \frac{1}{n_k} \sum_{m=1}^{m_k} n_{k,m}^2 (\tau_{k,m} - \tau_k)^2$$

this term vanishes when there is no heterogeneity in the average treatment effect across clusters. Information on whether to include this term to adjust for clustered sampling ( $q_k < 1$ ) must come from **outside the sample**.

- **clustered assignment** ( $\sigma_k^2 > 0$ ), then 2 more terms added

$$-p_k \sigma_k^2 \frac{1}{n_k} \sum_{i=1}^{n_k} \left( \frac{u_{k,i}(1)}{\mu_k} + \frac{u_{k,i}(0)}{1-\mu_k} \right)^2 + p_k \sigma_k^2 \frac{1}{n_k} \sum_{m=1}^{m_k} \left( \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} \left( \frac{u_{k,i}(1)}{\mu_k} + \frac{u_{k,i}(0)}{1-\mu_k} \right) \right)^2$$

the sign of it depends on the amount of variation in potential outcomes that can be explained by the clusters. In contrast to the lack of sample information about the need to adjust for clustered sampling, the sample is potentially informative about the need to account for clustered assignment.

The 5 terms of  $v_k$  might be of **different** orders:

- $\frac{1}{n_k} \sum_{i=1}^{n_k} \left( \frac{u_{k,i}^2(1)}{\mu_k} + \frac{u_{k,i}^2(0)}{1-\mu_k} \right)$ : average of bounded terms, of **order**  $O(1)$
- $p_k \frac{1}{n_k} \sum_{i=1}^{n_k} (u_{k,i}(1) - u_{k,i}(0))^2 + p_k \sigma_k^2 \frac{1}{n_k} \sum_{i=1}^{n_k} \left( \frac{u_{k,i}(1)}{\mu_k} + \frac{u_{k,i}(0)}{1-\mu_k} \right)^2$ : **at most** of the same order as the first term. If  $p_k \simeq 0$  (the sample is small relative to the population of sampled clusters), dominated by the first term
- the order of the last 2 terms depends on asymptotic of cluster sizes

$$p_k (1 - q_k) \frac{1}{n_k} \sum_{m=1}^{m_k} \left( \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} (u_{k,i}(1) - u_{k,i}(0)) \right)^2 + p_k \sigma_k^2 \frac{1}{n_k} \sum_{m=1}^{m_k} \left( \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} \left( \frac{u_{k,i}(1)}{\mu_k} + \frac{u_{k,i}(0)}{1-\mu_k} \right) \right)^2$$

- if cluster sizes are bounded as  $k$  increases: they are also order  $O(1)$
- if cluster sizes increase with  $k$ : they can be of higher order and dominate the variance

it also depends on  $p_k$ , clustering in sampling, clustering in assignment, heterogeneity in potential outcomes.

**Robust and cluster robust variance estimators** Let

$$\hat{U}_{k,i} = Y_{k,i} - \hat{\alpha}_k - \hat{\tau}_k W_{k,i}$$

be the residuals from the regression of  $Y_{k,i}$  on a constant and  $W_{k,i}$ , then the 2 common estimators of the variance of  $\sqrt{N_k}(\hat{\tau}_k - \tau_k)$  are

- **robust variance estimator:**

$$\hat{V}_k^{\text{robust}} = \frac{1}{\bar{W}_k^2 (1 - \bar{W}_k)^2} \left\{ \frac{1}{N_k} \sum_{i=1}^{n_k} R_{k,i} \hat{U}_{k,i}^2 (W_{k,i} - \bar{W}_k)^2 \right\}$$

where  $\bar{W}_k = \frac{1}{N_k} \sum_{i=1}^{n_k} R_{k,i} W_{k,i}$ , let  $v_k^{\text{robust}} = \frac{1}{n_k} \sum_{i=1}^{n_k} \left( \frac{u_{k,i}^2(1)}{\mu_k} + \frac{u_{k,i}^2(0)}{1-\mu_k} \right)$ , then under some regularity conditions,

$$\frac{\hat{V}_k^{\text{robust}}}{v_k} = \frac{v_k^{\text{robust}}}{v_k} + o_p(1)$$

notice that  $v_k^{\text{robust}} - v_k$  can be positive or negative in general, so the robust variance estimator can be invalid in large samples.

- **cluster variance estimator:**

$$\hat{V}_k^{\text{cluster}} = \frac{1}{\bar{W}_k^2 (1 - \bar{W}_k)^2} \times \left\{ \frac{1}{N_k} \sum_{m=1}^{m_k} \left( \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} R_{k,i} \hat{U}_{k,i} (W_{k,i} - \bar{W}_k) \right)^2 \right\}$$

define

$$\begin{aligned} v_k^{\text{cluster}} &= \frac{1}{n_k} \sum_{i=1}^{n_k} \left( \frac{u_{k,i}^2(1)}{\mu_k} + \frac{u_{k,i}^2(0)}{1-\mu_k} \right) \\ &\quad - p_k \frac{1}{n_k} \sum_{i=1}^{n_k} (u_{k,i}(1) - u_{k,i}(0))^2 - p_k \sigma_k^2 \frac{1}{n_k} \sum_{i=1}^{n_k} \left( \frac{u_{k,i}(1)}{\mu_k} + \frac{u_{k,i}(0)}{1-\mu_k} \right)^2 \\ &\quad + p_k \frac{1}{n_k} \sum_{m=1}^{m_k} \left( \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} (u_{k,i}(1) - u_{k,i}(0)) \right)^2 \\ &\quad + p_k \sigma_k^2 \frac{1}{n_k} \sum_{m=1}^{m_k} \left( \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} \left( \frac{u_{k,i}(1)}{\mu_k} + \frac{u_{k,i}(0)}{1-\mu_k} \right) \right)^2 \end{aligned}$$

then  $\hat{V}_k^{\text{cluster}}$  is close to  $v_k^{\text{cluster}}$  s.t.

$$\frac{\hat{V}_k^{\text{cluster}}}{v_k} = \frac{v_k^{\text{cluster}}}{v_k} + o_p(1)$$

and  $v_k^{\text{cluster}} - v_k$  is always **nonnegative**. For large  $k$ , the cluster variance can be conservative.

**Comparison** Comparing  $v_k^{\text{robust}}$  and  $v_k$ , we have

$$\begin{aligned} v_k^{\text{robust}} - v_k &= p_k \frac{1}{n_k} \sum_{i=1}^{n_k} (u_{k,i}(1) - u_{k,i}(0))^2 - p_k (1 - q_k) \frac{1}{n_k} \sum_{m=1}^{m_k} \left( \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} (u_{k,i}(1) - u_{k,i}(0)) \right)^2 \\ &\quad + p_k \sigma_k^2 \frac{1}{n_k} \sum_{i=1}^{n_k} \left( \frac{u_{k,i}(1)}{\mu_k} + \frac{u_{k,i}(0)}{1-\mu_k} \right)^2 - p_k \sigma_k^2 \frac{1}{n_k} \sum_{m=1}^{m_k} \left( \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} \left( \frac{u_{k,i}(1)}{\mu_k} + \frac{u_{k,i}(0)}{1-\mu_k} \right) \right)^2 \end{aligned}$$

which consists of 2 terms: the first term

$$p_k \frac{1}{n_k} \left[ \sum_{i=1}^{n_k} (u_{k,i}(1) - u_{k,i}(0))^2 - (1 - q_k) \sum_{m=1}^{m_k} n_{k,m}^2 (\tau_{k,m} - \tau_k)^2 \right]$$

- = 0 with **homogeneous treatment effects**:  $u_{k,i}(1) - u_{k,i}(0) = 0$ , for  $i = 1, \dots, n_k$  and  $\tau_{k,m} - \tau_k = 0$  for all  $m = 1, \dots, m_k$
- > 0 with **all clusters sampled**,  $q_k = 1$ , and **heterogeneous treatment effects**
- **possibly < 0**:  $q_k < 1$  (only a fraction of clusters are sampled), and  $n_{k,m}^2$  large enough.

and the second term

$$\begin{aligned}
 & p_k \sigma_k^2 \frac{1}{n_k} \sum_{i=1}^{n_k} \left( \frac{u_{k,i}(1)}{\mu_k} + \frac{u_{k,i}(0)}{1 - \mu_k} \right)^2 - p_k \sigma_k^2 \frac{1}{n_k} \sum_{m=1}^{m_k} \left( \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} \left( \frac{u_{k,i}(1)}{\mu_k} + \frac{u_{k,i}(0)}{1 - \mu_k} \right) \right)^2 \\
 &= p_k \sigma_k^2 \sum_{m=1}^{m_k} \frac{n_{k,m}}{n_k} \left[ \frac{1}{n_{k,m}} \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} \left( \frac{u_{k,i}(1)}{\mu_k} + \frac{u_{k,i}(0)}{1 - \mu_k} \right)^2 \right. \\
 &\quad \left. - n_{k,m} \left( \frac{1}{n_{k,m}} \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} \left( \frac{u_{k,i}(1)}{\mu_k} + \frac{u_{k,i}(0)}{1 - \mu_k} \right) \right)^2 \right]
 \end{aligned}$$

- = 0 if there is no clustered assignment:  $\sigma_k^2 = 0$
- close to 0 if the heterogeneity in potential outcomes is small:  $u_{k,i}(1), u_{k,i}(0)$  close to 0
- > 0 if there is heterogeneity in potential outcomes, but average potential outcomes are **nearly constant** across clusters
- < 0 if the clusters explain enough heterogeneity in potential outcomes, could be very large if  $n_{k,m}$  is large

Compare  $v_k^{\text{cluster}}$  and  $v_k$ , we have

$$\begin{aligned}
 v_k^{\text{cluster}} - v_k &= p_k q_k \frac{1}{n_k} \sum_{m=1}^{m_k} \left( \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} (u_{k,i}(1) - u_{k,i}(0)) \right)^2 \\
 &= \left( \frac{p_k n_k}{m_k} \right) q_k \left\{ \frac{1}{m_k} \sum_{m=1}^{m_k} \left( \frac{n_{k,m} m_k}{n_k} \right)^2 (\tau_{k,m} - \tau_k)^2 \right\}
 \end{aligned}$$

hence,

- $v_k^{\text{cluster}} - v_k \geq 0$ : cluster standard errors are (very) conservative in general
- $v_k^{\text{cluster}} - v_k \simeq 0$  when  $q_k$  is small (the expected fraction of clusters in the sample), or when the average treatment effect is nearly **constant** between clusters.

### 5.6.1.3 New variance estimators

**Abadie et al. (2023)** proposed 2 estimators of the variance of  $\hat{\tau}_k$ , one analytic based on a correction to  $\hat{V}_k^{\text{cluster}}$ , one based on resampling.

**Random Sampling:**  $q_k = 1$  When all clusters are observed ( $q_k = 1$ ), but allowing for general  $p_k$ , let

$$U_{k,i} = W_{k,i} u_{k,i}(1) + (1 - W_{k,i}) u_{k,i}(0)$$

first, approximate the normalized error of  $\hat{\tau}_k$  by a normalized sample average over clusters

$$\frac{\sqrt{N_k} (\hat{\tau}_k - \tau_k)}{\sqrt{v_k}} = \frac{1}{\sqrt{n_k p_k v_k} \mu_k (1 - \mu_k)} \sum_{m=1}^{m_k} C_{k,m} + o_p(1)$$

where  $C_{k,m} = \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} R_{k,i} (W_{k,i} - \mu_k) U_{k,i}$  are independent across clusters. Then For the term  $C_{k,m}$ , its expectation  $\mathbb{E}[C_{k,m}]$  is **not** 0 in general for each cluster separately, but equals to 0 when sum over **all** clusters:

$$\mathbb{E}[C_{k,m}] = n_{k,m} p_k \mu_k (1 - \mu_k) (\tau_{k,m} - \tau_k) \quad \sum_{m=1}^{m_k} \mathbb{E}[C_{k,m}] = p_k \mu_k (1 - \mu_k) \sum_{m=1}^{m_k} n_{k,m} (\tau_{k,m} - \tau_k) = 0$$

Now, since we have

$$\frac{\hat{V}_k^{\text{cluster}}}{v_k} = \frac{1}{n_k p_k v_k} \left( \frac{1}{\mu_k (1 - \mu_k)} \right)^2 \sum_{m=1}^{m_k} C_{k,m}^2 + o_p(1)$$

and  $\text{var}(C_{k,m}) \leq \mathbb{E}[C_{k,m}^2]$ , the expectation of  $\frac{1}{n_k p_k v_k} \left( \frac{1}{\mu_k (1 - \mu_k)} \right)^2 \sum_{m=1}^{m_k} C_{k,m}^2$  is  $\hat{V}_k^{\text{cluster}}$  is bigger than the variance of  $\frac{\sqrt{N_k}(\hat{\tau}_k - \tau_k)}{\sqrt{v_k}}$ , leading to  $\hat{V}_k^{\text{cluster}}$  being conservative.

Plug in  $\sum_{m=1}^{m_k} \mathbb{E}[C_{k,m}] = 0$  back into normalized  $\hat{\tau}_k$ , get

$$\begin{aligned} \frac{\sqrt{N_k}(\hat{\tau}_k - \tau_k)}{\sqrt{v_k}} &= \frac{1}{\sqrt{n_k p_k v_k} \mu_k (1 - \mu_k)} \sum_{m=1}^{m_k} C_{k,m} + o_p(1) \\ &= \frac{1}{\sqrt{n_k p_k v_k} \mu_k (1 - \mu_k)} \sum_{m=1}^{m_k} (C_{k,m} - \mathbb{E}[C_{k,m}]) + o_p(1) \\ &= \frac{1}{\sqrt{n_k p_k v_k} \mu_k (1 - \mu_k)} \sum_{m=1}^{m_k} (C_{k,m,1} + C_{k,m,2}) + o_p(1) \end{aligned}$$

where

$$\begin{aligned} C_{k,m,1} &= \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} (R_{k,i} - p_k) (\tau_{k,m} - \tau_k) \mu_k (1 - \mu_k) \\ C_{k,m,2} &= \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} R_{k,i} [(W_{k,i} - \mu_k) U_{k,i} - (\tau_{k,m} - \tau_k) \mu_k (1 - \mu_k)] \end{aligned}$$

$C_{k,m,1}$  and  $C_{k,m,2}$  are **mean 0** and **uncorrelated**, and **uncorrelated across clusters**, and

- variance of  $\frac{\sum_{m=1}^{m_k} C_{k,m,1}}{\sqrt{n_k p_k} \mu_k (1 - \mu_k)}$  is

$$(1 - p_k) \sum_{m=1}^{m_k} \frac{n_{k,m}}{n_k} (\tau_{k,m} - \tau_k)^2$$

- a direct estimator of the variance of  $\sum_{m=1}^{m_k} C_{k,m,2}$  is

$$\sum_{m=1}^{m_k} \left( \sum_{i=1}^{n_k} \mathbf{1}\{m_{k,i} = m\} R_{k,i} \left( (W_{k,i} - \bar{W}_k) \hat{U}_{k,i} - (\hat{\tau}_{k,m} - \hat{\tau}_k) \bar{W}_k (1 - \bar{W}_k) \right) \right)^2$$

where  $\hat{\tau}_{k,m}$  is the estimated treatment effect (sample average) in cluster  $m$ . This estimator could be biased from the correlation between the estimation errors of its components (due to  $\bar{W}_k$ ), to address this, [Abadie et al. \(2023\)](#)

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