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A generalized likelihood ratio test for multivariate analysis of variance in high dimension

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Abstract: This paper considers in the high dimensional setting a canonical testing problem, namely testing the equality of multiple mean vectors of normal distribution. Motivated by Roy's union-intersection principal, we propose a generalized likelihood ratio test. The critical value is determined by permutation method. We introduce an algorthm for permuting procedure, whose complexity does not depend on data dimension. The limiting distribution of the test statistic is derived in two different setting: non-spiked covariance and spiked covariance. Theoretical results and simulation studies show that the test is particularly powerful under spiked covariance.

Key words and phrases: Balanced incomplete block design, Hadamard matrix, nearly balanced incomplete block design, orthogonal array.

1. Introduction Suppose there are k ($k \geq 2$) groups of p dimensional data. Within the ith group ($1 \leq i \leq k$), we have observations $\{X_{ij}\}_{j=1}^{n_i}$ which are independent and identically distributed (i.i.d.) as $N_p(\mu_i, \Sigma)$, the

p dimensional normal distribution with mean vector μ_i and variance matrix Σ. We would like to test the hypotheses

$$H : \mu_1 = \mu_2 = \cdots = \mu_k \quad \text{v.s.} \quad K : \mu_i \neq \mu_j \text{ for some } i \neq j. \tag{1.1}$$

This testing problem is known as one-way multivariate analysis of variance (MANOVA) and has been well studied in the classical setting, i.e., p is small compared to n, where $n = \sum_{i=1}^{k} n_i$ is the total sample size.

Let $F = \sum_{i=1}^{k} n_i(\bar{\mathbf{X}}_i - \bar{\mathbf{X}})(\bar{\mathbf{X}}_i - \bar{\mathbf{X}})^T$ be the sum-of-squares between groups and $G = \sum_{i=1}^{k} \sum_{j=1}^{n_i} (X_{ij} - \bar{\mathbf{X}}_i)(X_{ij} - \bar{\mathbf{X}}_i)^T$ be the sum-of-squares within groups, where $\bar{\mathbf{X}}_i = n_i^{-1} \sum_{j=1}^{n_i} X_{ij}$ is the sample mean of group i and $\bar{\mathbf{X}} = n^{-1} \sum_{i=1}^{k} \sum_{j=1}^{n_i} X_{ij}$ is the pooled sample mean. There are four classical test statistics for hypothesis (1.1), which are all based on the eigenvalues of FG^{-1}.

Wilks' Lambda:	$\lvert G + F\rvert / \lvert G\rvert$
Pillai trace:	$\mathrm{tr}[F(G + F)^{-1}]$
Hotelling-Lawley trace:	$\mathrm{tr}[FG^{-1}]$
Roy's maximum root:	$\lambda_{\max}(FG^{-1})$

In some modern scientific applications, people would like to test hypothesis (1.1) in high dimensional setting, i.e., p is greater than n. See, for example, Tsai and Chen (2009). However, when $p \geq n$, the four classical test statistics can not be defined. Researchers have done extensive work to study the testing problem (1.1) in high dimensional setting. So far, most tests in the literature are designed for two sample case, i.e. $k = 2$. See, for example, Bai and Saranadasa (1996); Chen and Qin (2010); Srivastava (2009); Feng et al. (2016); Tony et al. (2013). For multiple sample case, Schott (2007) modified Hotelling-Lawley trace and proposed the test statistic

$$T_{SC} = \frac{1}{\sqrt{n-1}}\Big(\frac{1}{k-1}\,\mathrm{tr}\left(F\right) - \frac{1}{n-k}\,\mathrm{tr}\left(G\right)\Big).$$

In another work, Cai and Xia (2014) proposed a test statistic

$$T_{CX} = \max_{1\le i\le p} \sum_{1\le j<l\le k} \frac{n_j n_l}{n_j + n_l} \frac{(\Omega(\bar{X}_j - \bar{X}_l))_i^2}{\omega_{ii}},$$

Where $\Omega = (\omega)_{ij} = \Sigma^{-1}$ is the precision matrix. When Ω is unknown, they substitute it by an estimator $\hat{\Omega}$. Stitistics T_{SC} and T_{CX} are the representatives of two popular methodologies for high dimensional tests. T_{SC} is a so-called sum-of-squares type statistic as it is based on an estimation of squared Euclidean norm $\sum_{i=1}^k n_i\|\mu_i - \bar{\mu}\|^2$, where $\bar{\mu} = n^{-1}\sum_{i=1}^k n_i\mu_i$. T_{CX} is an extreme value type statistic.

Note that both sum-of-squares type statistic and extreme value type statistic are not based on likelihood function. While the likelihood ration test (LRT), i.e., Wilks' Lambda, is not defined if $p > n - k$, It remains a problem how to construct likelihood-based tests in high dimensional setting.

In a recent work, Zhao and Xu (2016) proposed a generalized likelihood ratio test in the context of one-sample mean vector test. Inspired by Roy's union-intersection principle (Roy (1953)), they wrote the null hypothesis as the intersection of a class of component hypotheses. For each component hypotheses, the likelihood ratio test is constructed. They use a least favorable argument to construct test statistic based on component tests. Their simulation results showed that their test has good power performance, especially when the variables are dependent.

Following Zhao and Xu (2016)'s methodology, we proposed a generalized likelihood ratio test for hypothesis (1.1). To understand the power behavior of the new test, we derive the asymptotic distribution of the new statistic under two different settings. In first setting, we assume the eigenvalues of Σ are bounded. It's a common assumption in high dimensional statistics. In fact, most existing tests for hypothesis (1.1) imposed conditions which prevent from large leading eigenvalues of Σ. However, when the correlations between variables are determined by a small number of factors, Σ is spiked in the sense that a few leading eigenvalues are much larger than the others. See, for example Cai et al. (2013) and Shen et al. (2013). We then derive the asymptotic distribution of the test statistic under spiked covariance. From the theoretical results we give, it can be seen that the new test is particularly powerful under spiked covariance. We also conduct a simulation study to examine the numerical performance of the test.

The rest of the paper is organized as follows.

2. Methodology

The Roy's union intersection principle plays an important role in the methodology of Zhao and Xu (2016). In this section, first we shall revisit Roy's union intersection principle. It turns out that some existing tests for (1.1) can be derived by Roy's union intersection principle. Then we will deal with the problem in high dimensions and propose a new test statistic.

2.1 Roy's union intersection principle

Roy's union intersection principle is proposed by Roy (1953). The idea of Roy's union intersection principle is to reduce testing problem to a class of pseudo-univariate problems and the null hypothesis is accepted if all component null hypotheses are aceepted. More precisely, Roy's union intersection principle can be decomposed into 3 main steps:

1. Decompose the hypothesis H and K into component hypohteses

$$H = \bigcap_{\gamma\in\Gamma} H_\gamma \quad \text{v.s.} \quad K = \bigcup_{\gamma\in\Gamma} K_\gamma,$$

   where Γ is an index set.

2. Construct a test for H_γ against H_γ.

3. Accept H if all component tests accept the null hypotheses. Or equivalently, reject H if any component test reject the null hypothesis.

Roy's union intersection principle is a powerful tool when H or K are complicated. In fact, the Roy's maximum root test statistic is derived in Roy (1953) as an example of Roy's union intersection principle.

For convenience, now we introduce some notations. Let

$$\mathbf{Z} = (X_{11}, \ldots, X_{1n_1}, \ldots, X_{k1}, \ldots, X_{kn_k})$$

be the pooled sample matrix. Define

$$J = \begin{pmatrix} \frac{1}{\sqrt{n_1}}\mathbf{1}_{n_1} & \mathbf{0} & \mathbf{0} \\ \mathbf{0} & \frac{1}{\sqrt{n_2}}\mathbf{1}_{n_2} & \mathbf{0} \\ \vdots & \vdots & \vdots \\ \mathbf{0} & \mathbf{0} & \frac{1}{\sqrt{n_k}}\mathbf{1}_{n_k} \end{pmatrix}.$$

Then the matrices $I_n - JJ^T$, $JJ^T - \frac{1}{n}\mathbf{1}_n\mathbf{1}_n^T$ and $\frac{1}{n}\mathbf{1}_n\mathbf{1}_n^T$ are three $n \times n$ projection matrices which are pairwise orthogonal with rank $n-k$, $k-1$ and 1 respectively. Let $\tilde{J}$ be an $n \times (n-k)$ matrix satisfying $\tilde{J}\tilde{J}^T = I - JJ^T$. Note that $I_k - \frac{1}{n}J^T\mathbf{1}_n\mathbf{1}_n^T J$ is a $k \times k$ projection matrix with rank $k-1$. Let C be a $k \times (k-1)$ matrix satisfying $CC^T = I_k - \frac{1}{n}J^T\mathbf{1}_n\mathbf{1}_n^T J$. Then we have

$$G = Z(I_n - JJ^T)Z^T = Z\tilde{J}\tilde{J}^T Z^T,$$

and

$$F = Z(JJ^T - \frac{1}{n}\mathbf{1}_n\mathbf{1}_n^T)Z^T = ZJ(I_k - \frac{1}{n}J^T\mathbf{1}_n\mathbf{1}_n^T J)J^T Z^T = ZJCC^T J^T Z^T.$$

Define $\Xi = (\sqrt{n_1}\mu_1, \ldots, \sqrt{n_k}\mu_k)$ and the null hypothesis H is equivalent to $\Xi C = O_{p\times(k-1)}$.

Note that there is a one-to-one mappting between the data $\mathbf{Z}$ and the set of data $\{a^T\mathbf{Z} \mid a \in \mathbb{R}^p, a^T a = 1\}$. This transformation naturally induces the decomposition

$$H = \bigcap_{a\in\mathbb{R}^p, a^T a=1} H_a \quad \text{and} \quad K = \bigcup_{a\in\mathbb{R}^p, a^T a=1} K_a,$$

where

$$H_a : a^T\Xi C = O_{1\times(k-1)} \quad \text{and} \quad K_a : a^T\Xi C \neq O_{1\times(k-1)}.$$

Note that the likelihood function of $a^T\mathbf{Z}$ is

$$f_a(a^T\mathbf{Z}; \mu_1, \ldots, \mu_k, \Sigma) = (2\pi)^{-n/2}|a^T\Sigma a|^{-n/2} \exp\Big(-\frac{1}{2a^T\Sigma a}\sum_{i=1}^{k}\sum_{j=1}^{n_i}(a^T X_{ij} - a^T\mu_i)^2\Big).$$

The likelihood ratio test statistic for H_a against $K_a$is

$$\mathrm{LRT}_a = \frac{\sup_{\mu_1,\ldots,\mu_k,\Sigma} f_a(a^T\mathbf{Z}; \mu_1, \ldots, \mu_k, \Sigma)}{\sup_{\mu,\Sigma} f_a(a^T\mathbf{Z}; \mu, \ldots, \mu, \Sigma)} = \Big(1 + \frac{a^T F a}{a^T G a}\Big)^{n/2}.$$

By Roy's union intersection principle, H is rejected when $\max_{a^T a=1} \mathrm{LRT}_a$ is large. If $p \leq n-k$, G is invertible and $\max_{a^T a=1} \mathrm{LRT}_a = (1+\lambda_{\max}(FG^{-1}))^{n/2}$, which is an increase function of Roy's maximum root test statistic.

Another classical test statistic, Hotelling-Lawley trace, can also be derived by Roy's union intersection principle. This is shown by Mudholkar et al. (1974). In that paper, they consider the transformed data $\{M^T\mathbf{Z} \mid M is (k-1)\times p \text{ matrix}\}$ and the decomposition of hypotheses:

$$H = \bigcap_M H_M \quad \text{and} \quad K = \bigcup_M K_M,$$

where

$$H_M : \text{tr}(M\Xi C) = 0 \quad \text{and} \quad K_M : \text{tr}(M\Xi C) > 0.$$

Note that $\mathrm{E}\, Z = \Xi J^T$, hence the uniformly minimum variance unbiased estimator of $\text{tr}(M\Xi C)$ is $\text{tr}(MZJC)$. It can be seen that $\text{tr}\big(MZJC\big) \sim N\big(\text{tr}(M\Xi C), \text{tr}(M\Sigma M^T)\big)$.

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$$\text{tr}\big(MZJC\big) = \text{tr}\big(CMZJ\big)$$

$ZJ = (\sqrt{n_1}\bar{\mathbf{X}}_1, \ldots, \sqrt{n_k}\bar{\mathbf{X}}_k)$. Note that $CM\sqrt{n_i}\bar{\mathbf{X}}_i \sim N_{k-1}(\sqrt{n_i}CM\mu_i, CM\Sigma M^TC^T)$. Hence we have that

$$\text{tr}\big(CMZJ\big) \sim N(\text{tr}(CM\Xi), \text{tr}(CM\Sigma M^TC^T)) \sim N(\text{tr}(M\Xi C), \text{tr}(M\Sigma M^T)).$$

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Hence it's natural to use one side t type statistic

$$T_M = \frac{\text{tr}\big(MZJC\big)}{\sqrt{\text{tr}(MGM^T)}}$$

to test H_M against K_M.

By Cauchy inequality $\max_B \operatorname{tr}(AB^T)/\operatorname{tr}^{1/2}(BB^T) = \operatorname{tr}^{1/2}(AA^T)$, we have

$$\max_M T_M = \max_M \frac{\operatorname{tr}\left(MG^{1/2}G^{-1/2}ZJC\right)}{\sqrt{\operatorname{tr}(MG^{1/2}(MG^{1/2})^T)}} = \operatorname{tr}^{1/2}((ZJC)^T G^{-1} ZJC)$$
$$= \operatorname{tr}^{1/2}(ZJC(ZJC)^T G^{-1}) = \operatorname{tr}^{1/2}(FG^{-1}).$$

==================== In fact, T_{CX} can also be derived by Roy's union intersection principle. Suppose Ω is known,. ====================

Note that T_{CX} is designed for high dimensional data, which implies that Roy's union intersection principle may be useful in high dimensional setting. In fact, Roy's union intersection principle can be used to decompose high dimensional problem into univariate problems which are easy to deal with.

2.2 The new test statistic

We are interested in the case when $p > n-k$. In this setting, $\max_{a^Ta=1} \mathrm{LRT}_a = +\infty$ and Roy's maximum root test is not defined. In another viewpoint, union intersection principal finds an direction a along which the evidence against null hypothesis is maximized. Such an a is data dependent. In the classical setting, the evidence of direction a is LRT_a. In the current context, there are a class of a such that LRT_a achieve the infinity, the largest evidence in classical sense. We need to further choose a single a

from $\{a \,|\, \text{LRT}_a = +\infty \text{ and } a^T a = 1\}$. From the expression of LRT_a, we would like to make the largest discrepancy between $a^T F a$ and $a^T G a$. Note that if $\text{LRT}_\alpha = +\infty$, then $a^T G a = 0$. Hence it's natural to choose a as

$$a^* = \underset{a^T a = 1, a^T G a = 0}{\arg\max} \; a^T F a.$$

Since $a^{*T} G a^* = 0$, we propose the following test statistic for H:

$$T = a^{*T} F a^* = \max_{a^T a = 1, a^T G a = 0} a^T F a.$$

When T is large enough, we reject H. The above strategy is first proposed by Zhao and Xu (2016) in the context of testing one sample mean vector.

Next we derive the explicit forms of the test statistic. Let $J = \text{diag}(n_1^{-1/2}\mathbf{1}_{n_1}, \ldots, n_k^{-1/2}\mathbf{1}_{n_k})$. Then the matrices $I_n - JJ^T$, $JJ^T - \frac{1}{n}\mathbf{1}_n\mathbf{1}_n^T$ and $\frac{1}{n}\mathbf{1}_n\mathbf{1}_n^T$ are three $n \times n$ projection matrices which are pairwise orthogonal with rank $n-k$, $k-1$ and 1. Let $\tilde{J}$ be a $n \times (n-k)$ matrix satisfying $\tilde{J}\tilde{J}^T = I - JJ^T$. Note that $I_k - \frac{1}{n}J^T\mathbf{1}_n\mathbf{1}_n^T J$ is a $k \times k$ projection matrix with rank $k-1$. Let C be a $k \times (k-1)$ matrix satisfying $CC^T = I_k - \frac{1}{n}J^T\mathbf{1}_n\mathbf{1}_n^T J$. Then

$$G = Z(I_n - JJ^T)Z^T = Z\tilde{J}\tilde{J}^T Z^T.$$

and

$$F = \sum_{i=1}^k n_i(\bar{\mathbf{X}}_i - \bar{\mathbf{X}})(\bar{\mathbf{X}}_i - \bar{\mathbf{X}})^T = Z(JJ^T - \frac{1}{n}\mathbf{1}_n\mathbf{1}_n^T)Z^T = ZJ(I_k - \frac{1}{n}J^T\mathbf{1}_n\mathbf{1}_n^T J)J^T Z^T = ZJCC^T J^T Z^T.$$

Let $Z\tilde{J} = U_{Z\tilde{J}} D_{Z\tilde{J}} V_{Z\tilde{J}}^T$ be the singular value decomposition of $Z\tilde{J}$, where $U_{Z\tilde{J}}$ and $V_{Z\tilde{J}}$ are $p \times (n-k)$ and $(n-k) \times (n-k)$ column orthogonal matrices respectively, $D_{Z\tilde{J}}$ is $(n-k) \times (n-k)$ diagonal matrix. Let $H_{Z\tilde{J}} = U_{Z\tilde{J}} U_{Z\tilde{J}}^T$ be the projection on the column space of A. Then by Proposition 1,

$$T(Z) = \lambda_{\max}\big(ZJCC^T J^T Z^T (I_p - H_{Z\tilde{J}})\big) = \lambda_{\max}\big(C^T J^T Z^T (I_p - H_{Z\tilde{J}}) ZJC\big). \tag{2.2}$$

Next we introduce another form of T. By the relationship

$$\begin{pmatrix} J^T Z^T Z J & J^T Z^T Z \tilde{J} \\ \tilde{J}^T Z^T Z J & \tilde{J}^T Z^T Z \tilde{J} \end{pmatrix}^{-1} = \left(\begin{pmatrix} J^T \\ \tilde{J}^T \end{pmatrix} Z^T Z \begin{pmatrix} J & \tilde{J} \end{pmatrix} \right)^{-1} = \begin{pmatrix} J^T (Z^T Z)^{-1} J & J^T (Z^T Z)^{-1} \tilde{J} \\ \tilde{J}^T (Z^T Z)^{-1} J & \tilde{J}^T (Z^T Z)^{-1} \tilde{J} \end{pmatrix}$$

and matrix inverse formula, we have that

$$\big(J^T (Z^T Z)^{-1} J\big)^{-1} = J^T Z^T Z J - J^T Z^T Z \tilde{J} (\tilde{J}^T Z^T Z \tilde{J})^{-1} \tilde{J}^T Z^T Z J = J^T Z^T (I_p - H_{Z\tilde{J}}) Z J.$$

Thus,

$$T(Z) = \lambda_{\max}\Big(C^T \big(J^T (Z^T Z)^{-1} J\big)^{-1} C\Big). \tag{2.3}$$

While the form (2.2) is used for theoretical analysis, the form (2.3) is well suited for computation, as we shall see.

2.3 Permutation method

Permutation method is a powerful tool to determine the critical value of a test statistic. The test procedure resulting from permutation method is exact as long as the null distribution of observations are exchangeable. See, for example, Romano (1990). The major down-side to permutation method is that it can be computationally intensive. Fortunately, for our statistic, there is a fast implementation of the permutation method. Using expression (2.3), a permuted statistic can be written as

$$T(Z\Gamma) = \lambda_{\max}\Big(C^T \big(J^T \Gamma^T (Z^T Z)^{-1} \Gamma J\big)^{-1} C\Big), \tag{2.4}$$

where Γ is an $n \times n$ permutation matrix. Note that $(Z^T Z)^{-1}$, the most time-consuming component, can be calculated aforehand. The permutation procedure for our statistic can be summarized as:

1. Calculate $T(Z)$ according to (2.3), hold intermediate result $(Z^T Z)^{-1}$.
2. For a large M, independently generate M random permutation matrix $\Gamma_1, \ldots, \Gamma_M$ and calculate $T(Z\Gamma_1), \ldots, T(Z\Gamma_M)$ according to (2.4).
3. Calculate the p-value by $\tilde{p} = (M+1)^{-1}\big[1+\sum_{i=1}^{M} I\{T(Z\Gamma_i) \geq T(Z)\}\big]$. Reject the null hypothesis if $\tilde{p} \leq \alpha$.

Here M is the permutation times. It can be shown that for any integer $M > 0$, the resulting test controls the Type I error. More precisely, we have $\Pr(\tilde{p} \leq u) \leq u$ for all $0 \leq u \leq 1$. Moreover, as M tends to ∞, $\lim_{M\to\infty} \Pr(\tilde{p} \leq u) = u$. See, for example, E. L. Lehmann (2005), Chapter 15.

It can be seen that the time complexities of step (I) and step (II) are $O(n^2p+n^3)$ and $O(n^2M)$, respectively. In large sample or high dimensional setting, $M/(p+n)$ is small. In this case, the permutation procedure has negligible effect on total time complexity.

3. Theory

In this section, we investigate the asymptotic behavior of our test statistic when p is much larger than n. More precisely, we shall assume $p/n \to \infty$. In high dimensional setting, it is a common phenomenon that the asymptotic distribution of statistic relies on the covariance structure. See, for example, Ma et al. (2015) and Rui Wang's paper. We shall investigate the asymptotics of our statistic under two different covariance structures: non-spiked covariance and spiked covariance.

Let W_{k-1} be a $(k-1)\times(k-1)$ symmetric random matrix whose entries above the main diagonal are i.i.d. $N(0,1)$ and the entries on the diagonal are i.i.d. $N(0,2)$.

Let $\Sigma = U\Lambda U^T$ be the eigenvalue decomposition of Σ, where $\Lambda = \mathrm{diag}(\lambda_1,\ldots,\lambda_p)$.

Theorem 1. *Suppose $p/n \to \infty$, $c \le \lambda_p(\Sigma) \le \cdots \le \lambda_1(\Sigma) \le C$ and*

$$\operatorname{tr}\left(\Sigma - \frac{1}{p}(\operatorname{tr}\Sigma) I_p\right)^2 = o\left(\frac{p}{n}\right).$$

Under local alternative $p^{-1}\|\Xi C\|_F^2 \to 0$, we have

$$\frac{T(Z) - \frac{p-n+k}{p}\operatorname{tr}(\Sigma)}{\sqrt{\operatorname{tr}(\Sigma^2)}} \sim \lambda_{\max}\left(W_{k-1} + \frac{1}{\sqrt{\operatorname{tr}(\Sigma^2)}} C^T \Xi^T (I_p - H_{Z\tilde{J}})\Xi C\right) + o_P(1).$$

The spiked covariance model assumes that a few eigenvalues of Σ are significantly larger than the others. This model is a standard model in many problems and takes factor model as a special case. See, for example,.

Assumption 1. *Let r be a fixed integer. Suppose $\lambda_r n/p \to \infty$ and $C \ge \lambda_{r+1} \ge \ldots \ge \lambda_p \ge c$, where c and C are absolute constant.*

Let $U = (U_1, U_2)$ where U_1 is $p \times r$ and U_2 is $p \times (p-r)$. Let $\Lambda_1 = \operatorname{diag}(\lambda_1, \ldots, \lambda_r)$ and $\Lambda_2 = \operatorname{diag}(\lambda_{r+1}, \ldots, \lambda_p)$. Then $\Sigma = U_1\Lambda_1 U_1^T + U_2 \Lambda_2 U_2^T$.

Theorem 2. *Under Assumption* (1)*, suppose $p/n \to \infty$, $\frac{\lambda_1^2 p}{\lambda_r^2 n^2} \to 0$ and*

$$\operatorname{tr}\left(\Lambda_2 - \frac{1}{p-r}(\operatorname{tr}\Lambda_2) I_{p-r}\right)^2 = o\left(\frac{p}{n}\right).$$

Then under local alternative

$$\frac{1}{\sqrt{p}}\|\Xi C\|_F^2 = O(1),$$

we have

$$\frac{T(Z) - \frac{p-r-n+k}{p-r}\operatorname{tr}(\Lambda_2)}{\sqrt{\operatorname{tr}(\Lambda_2^2)}} \sim \lambda_{\max}\left(W_{k-1} + \frac{1}{\sqrt{\operatorname{tr}(\Lambda_2^2)}} C^T \Xi^T (I_p - H_{Z\tilde{J}})\Xi C\right) + o_P(1).$$

4. Simulation Results

In this section, we evaluate the numerical performance of the new test. For comparison, we also carried out simulation for the test of Tony Cai and Yin Xia and the test of Schott. These tests are denoted respectively by NEW, CX and SC.

In the simulations, we set $k = 3$. Note that the new test is invariant under orthogonal transformation. Without loss of generality, we only consider diagonal Σ. We set $\Sigma = \mathrm{diag}(p, 1, \ldots, 1)$. Define signal-to-noise ratio (SNR) as

$$\mathrm{SNR} = \frac{\|\mu_f\|_F^2}{\sqrt{\sum_{i=2}^{p} \lambda_i(\Sigma)^2}}.$$

We use SNR to characterize the signal strength. We consider two alternative hypotheses: the non-sparse alternative and the sparse alternative. In the non-sparse case, we set $\mu_1 = \kappa 1_p$, $\mu_2 = -\kappa 1_p$ and $\mu_3 = 0_p$, where κ is selected to make the SNR equal to the given value. In the sparse case, we set $\mu_1 = \kappa(1_{p/5}^T, 0_{4p/5}^T)^T$, $\mu_2 = \kappa(0_{p/5}^T, 1_{p/5}^T, 0_{3p/5}^T)^T$ and $\mu_3 = 0_p$. Again, κ is selected to make the SNR equal to the given value.

Table 1: Empirical powers of tests under non-sparse alternative with $\alpha = 0.05$, $k = 3$, $n_1 = n_2 = n_3 = 10$. Based on 1000 replications.

SNR	$p = 50$			$p = 75$			$p = 100$		
	CX	SC	NEW	CX	SC	NEW	CX	SC	NEW
0	0.035	0.048	0.052	0.057	0.052	0.057	0.053	0.048	0.045
1	0.060	0.049	0.096	0.081	0.050	0.092	0.063	0.062	0.085
2	0.100	0.058	0.140	0.073	0.045	0.169	0.086	0.055	0.171
3	0.145	0.066	0.234	0.119	0.070	0.266	0.117	0.056	0.307
4	0.126	0.064	0.317	0.121	0.059	0.380	0.122	0.061	0.402
5	0.179	0.072	0.392	0.178	0.068	0.541	0.141	0.071	0.579
6	0.198	0.070	0.513	0.189	0.071	0.639	0.143	0.066	0.717
7	0.249	0.085	0.629	0.227	0.084	0.777	0.206	0.073	0.822
8	0.268	0.092	0.685	0.252	0.084	0.822	0.217	0.078	0.894
9	0.324	0.100	0.786	0.256	0.090	0.911	0.246	0.074	0.949
10	0.342	0.115	0.828	0.303	0.097	0.937	0.270	0.075	0.973

Table 2: Empirical powers of tests under non-sparse alternative with $\alpha = 0.05$, $k = 3$, $n_1 = n_2 = n_3 = 25$. Based on 1000 replications.

SNR	$p = 100$			$p = 150$			$p = 200$		
	CX	SC	NEW	CX	SC	NEW	CX	SC	NEW
0	0.050	0.043	0.050	0.056	0.066	0.048	0.062	0.045	0.054
1	0.069	0.048	0.063	0.046	0.052	0.091	0.068	0.048	0.095
2	0.097	0.046	0.131	0.086	0.053	0.164	0.068	0.057	0.173
3	0.113	0.061	0.200	0.117	0.057	0.270	0.101	0.045	0.313
4	0.135	0.053	0.247	0.130	0.054	0.402	0.118	0.066	0.485
5	0.158	0.065	0.357	0.134	0.066	0.526	0.134	0.073	0.616
6	0.198	0.081	0.433	0.161	0.052	0.668	0.138	0.067	0.765
7	0.217	0.068	0.514	0.191	0.067	0.759	0.174	0.068	0.862
8	0.229	0.063	0.582	0.223	0.075	0.853	0.187	0.060	0.927
9	0.264	0.094	0.680	0.218	0.080	0.918	0.227	0.067	0.966
10	0.298	0.091	0.758	0.245	0.076	0.934	0.228	0.052	0.982

Table 3: Empirical powers of tests under sparse alternative with $\alpha = 0.05$, $k = 3$, $n_1 = n_2 = n_3 = 10$. Based on 1000 replications.

SNR	$p = 50$			$p = 75$			$p = 100$		
	CX	SC	NEW	CX	SC	NEW	CX	SC	NEW
0	0.063	0.056	0.052	0.048	0.049	0.048	0.057	0.047	0.042
1	0.087	0.058	0.071	0.069	0.044	0.096	0.076	0.051	0.080
2	0.091	0.066	0.116	0.113	0.037	0.133	0.080	0.058	0.139
3	0.155	0.065	0.177	0.131	0.062	0.228	0.113	0.058	0.218
4	0.184	0.065	0.246	0.174	0.076	0.308	0.144	0.061	0.310
5	0.225	0.081	0.337	0.214	0.075	0.386	0.176	0.083	0.417
6	0.270	0.088	0.425	0.266	0.085	0.507	0.228	0.071	0.508
7	0.364	0.080	0.501	0.307	0.078	0.571	0.302	0.087	0.629
8	0.405	0.105	0.549	0.381	0.080	0.698	0.362	0.089	0.721
9	0.470	0.121	0.634	0.408	0.078	0.774	0.391	0.070	0.797
10	0.547	0.128	0.702	0.484	0.109	0.819	0.415	0.088	0.877

Table 4: Empirical powers of tests under sparse alternative with $\alpha = 0.05$, $k = 3$, $n_1 = n_2 = n_3 = 25$. Based on 1000 replications.

SNR	$p = 100$			$p = 150$			$p = 200$		
	CX	SC	NEW	CX	SC	NEW	CX	SC	NEW
0	0.048	0.045	0.046	0.053	0.046	0.043	0.051	0.034	0.046
1	0.079	0.055	0.082	0.066	0.063	0.079	0.063	0.059	0.100
2	0.097	0.054	0.119	0.088	0.055	0.138	0.085	0.055	0.160
3	0.133	0.069	0.167	0.113	0.066	0.223	0.114	0.054	0.235
4	0.149	0.062	0.212	0.126	0.084	0.298	0.132	0.057	0.344
5	0.204	0.060	0.281	0.169	0.066	0.427	0.154	0.057	0.469
6	0.252	0.060	0.352	0.227	0.070	0.548	0.195	0.072	0.641
7	0.310	0.072	0.429	0.252	0.059	0.614	0.220	0.061	0.711
8	0.372	0.088	0.529	0.314	0.085	0.719	0.297	0.060	0.800
9	0.427	0.083	0.547	0.362	0.085	0.794	0.300	0.057	0.881
10	0.449	0.093	0.619	0.396	0.072	0.853	0.340	0.076	0.911

Table 5: Empirical powers of tests under non-sparse alternative with $\alpha = 0.05$, $k = 3$, $n_1 = n_2 = n_3 = 25$. The diagonal elements of Σ are generated from sort(Unif(1,100)). Based on 1000 replications.

SNR	$p = 100$			$p = 150$			$p = 200$		
	CX	SC	NEW	CX	SC	NEW	CX	SC	NEW
0	0.063	0.054	0.058	0.052	0.040	0.042	0.045	0.049	0.070
1	0.141	0.120	0.115	0.126	0.120	0.112	0.103	0.110	0.102
2	0.181	0.209	0.169	0.330	0.260	0.210	0.200	0.227	0.201
3	0.692	0.367	0.244	0.759	0.385	0.341	0.468	0.413	0.394
4	0.753	0.539	0.420	0.744	0.573	0.515	0.516	0.554	0.561
5	0.828	0.690	0.509	0.871	0.697	0.693	0.556	0.724	0.727
6	0.809	0.812	0.622	0.822	0.824	0.766	0.959	0.838	0.859
7	1.000	0.882	0.780	0.979	0.916	0.903	0.990	0.923	0.947
8	0.993	0.955	0.789	1.000	0.965	0.954	0.999	0.972	0.971
9	1.000	0.979	0.911	0.999	0.981	0.979	0.964	0.986	0.987
10	1.000	0.991	0.877	0.989	0.996	0.988	0.996	0.996	0.997

Table 6: Empirical powers of tests under sparse alternative with $\alpha = 0.05$, $k = 3$, $n_1 = n_2 = n_3 = 25$. The diagonal elements of Σ are generated from sort(Unif(1,100)). Based on 1000 replications.

SNR	$p = 100$			$p = 150$			$p = 200$		
	CX	SC	NEW	CX	SC	NEW	CX	SC	NEW
0	0.052	0.055	0.047	0.055	0.057	0.053	0.044	0.055	0.057
1	0.068	0.124	0.065	0.070	0.130	0.085	0.049	0.116	0.087
2	0.085	0.233	0.112	0.076	0.239	0.149	0.067	0.241	0.161
3	0.110	0.388	0.161	0.090	0.408	0.215	0.097	0.417	0.227
4	0.120	0.530	0.184	0.112	0.552	0.282	0.103	0.556	0.309
5	0.167	0.708	0.238	0.142	0.699	0.387	0.140	0.687	0.394
6	0.196	0.807	0.261	0.168	0.820	0.472	0.162	0.823	0.547
7	0.217	0.875	0.318	0.177	0.892	0.505	0.173	0.896	0.646
8	0.234	0.935	0.378	0.220	0.951	0.625	0.195	0.948	0.749
9	0.312	0.965	0.407	0.222	0.970	0.672	0.224	0.979	0.809
10	0.334	0.976	0.505	0.292	0.987	0.773	0.254	0.989	0.881

5. Conclusion remarks

Appendix

Proposition 1. *Suppose A is a $p \times r$ matrix with rank r and B is a $p \times p$ non-zero semi-definite matrix. Denote by $A = U_A D_A V_A^T$ the singular value decomposition of A, where U_A and V_A are $p \times r$ and $r \times r$ column orthogonal matrix, D_A is a $r \times r$ diagonal matrix. Let $H_A = U_A U_A^T$ be the projection on the column space of A. Then*

$$\max_{a^T a=1, a^T A A^T a=0} a^T B a = \lambda_{\max}\big(B(I_p - H_A)\big). \tag{5.5}$$

Proof. Note that $a^T A A^T a = 0$ is equivalent to $H_A a = 0$ which in turn is equivalent to $a = (I_p - H_A)a$. Then

$$\max_{a^T a=1, a^T A A^T a=0} a^T B a = \max_{a^T a=1, H_A a=0} a^T (I_p - H_A) B (I_p - H_A) a, \tag{5.6}$$

which is obviously no greater than $\lambda_{\max}\big((I - H_A)B(I - H_A)\big)$. To prove that they are equal, without loss of generality, we can assume $\lambda_{\max}\big((I - H_A)B(I - H_A)\big) > 0$. Let α_1 be one eigenvector corresponding to the largest eigenvalue of $(I - H_A)B(I - H_A)$. Since $(I - H_A)B(I - H_A)H_A = (I - H_A)B(H_A - H_A) = O_{p\times p}$ and H_A is symmetric, the rows of H_A are eigenvetors of $(I - H_A)B(I - H_A)$ corresponding to eigenvalue 0. It follows that $H_A \alpha_1 = 0$. Therefore, α_1 satisfies the constraint of (5.6) and (5.6) is no less than $\lambda_{\max}\big((I-H_A)B(I-H_A)\big)$. The conclusion now follows by noting that $\lambda_{\max}\big((I-H_A)B(I-H_A)\big)=\lambda_{\max}\big(B(I-H_A)\big)$.

$\square$

Proof of the main results It can be seen that ZJC is independent of $Z\tilde{J}$. Since $\mathrm{E}(Z\tilde{J})=O_{p\times(n-k)}$, we can write $Z\tilde{J}=U\Lambda^{1/2}G_1$, where G_1 is a $p\times(n-k)$ matrix with i.i.d. $N(0,1)$ entries. We write $ZJC=\mu_f+U\Lambda^{1/2}G_2$, where G_2 is a $p\times(k-1)$ matrix with i.i.d. $N(0,1)$ entries.

Then

$$\begin{aligned} C^TJ^TZ^T(I_p-H_{Z\tilde{J}})ZJC=&G_2^T\Lambda^{1/2}U^T(I_P-H_{Z\tilde{J}})U\Lambda^{1/2}G_2+\mu_f^T(I_p-H_{Z\tilde{J}})\mu_f+\\ &\mu_f^T(I_p-H_{Z\tilde{J}})U\Lambda^{1/2}G_2+G_2^T\Lambda^{1/2}U^T(I_P-H_{Z\tilde{J}})\mu_f. \end{aligned} \tag{5.7}$$

The first term of (5.7) can be represented as

$$G_2^T\Lambda^{1/2}U^T(I_p-H_{Z\tilde{J}})U\Lambda^{1/2}G_2=\sum_{i=1}^{p}\lambda_i(\Lambda^{1/2}U^T(I_p-H_{Z\tilde{J}})U\Lambda^{1/2})\xi_i\xi_i^T, \tag{5.8}$$

where $\xi_i\overset{i.i.d.}{\sim}N(0,I_{k-1})$.

Proof of Theorem 1. First we deal with the first term of (5.7). Note that for $i=1,\dots,p$, we have

$$\lambda_i(\Lambda^{1/2}U^T(I_p-H_{Z\tilde{J}})U\Lambda^{1/2})\le\lambda_i(\Lambda). \tag{5.9}$$

Note that $H_{Z\tilde{J}}$ has rank $n-k$. For $i=1,\dots,p-n+k$, by Weyl's inequality, we have

$$\lambda_i(\Lambda^{1/2}U^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2}) \geq \lambda_{i+n-k}(\Lambda). \tag{5.10}$$

Then we have

$$\frac{\lambda_1^2(\Lambda^{1/2}U^T(I_p - H_{Z\tilde{Z}})U\Lambda^{1/2})}{\sum_{i=1}^p \lambda_i^2(\Lambda^{1/2}U^T(I_p - H_{Z\tilde{Z}})U\Lambda^{1/2})} \leq \frac{C}{c(p-n+k)} \to 0.$$

Apply Lyapunov central limit theorem conditioning on $Z\tilde{J}$, we have

$$\Big(\sum_{i=1}^p \lambda_i^2(\Lambda^{1/2}U^T(I_p - H_{Z\tilde{Z}})U\Lambda^{1/2})\Big)^{-1/2}$$
$$\Big(G_2^T\Lambda^{1/2}U^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2}G_2 - \sum_{i=1}^p \lambda_i(\Lambda^{1/2}U^T(I_p - H_{Z\tilde{Z}})U\Lambda^{1/2})I_{k-1}\Big) \xrightarrow{\mathcal{L}} W_{k-1}.$$

Also by (5.9) and (5.10), we have

$$\sum_{i=n-k+1}^p \lambda_i^2 \leq \operatorname{tr}\left[(\Lambda^{1/2}U^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2})^2\right] \leq \operatorname{tr}(\Lambda^2).$$

Hence we have

$$\operatorname{tr}\left[(\Lambda^{1/2}U^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2})^2\right] = \operatorname{tr}(\Lambda^2) + O_P(n) = \big(1 + O_P(\frac{n}{p})\big)\operatorname{tr}(\Lambda^2).$$

Note that

$$\operatorname{tr}(\Lambda^{1/2}U^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2}) = \operatorname{tr}(\Lambda) - \operatorname{tr}(H_{Z\tilde{J}}U\Lambda U^T).$$

and

$$\big|\operatorname{tr}(H_{Z\tilde{J}}U\Lambda U^T) - \frac{n-k}{p}\operatorname{tr}(\Lambda)\big| = \big|\operatorname{tr}\Big(H_{Z\tilde{J}}U\big(\Lambda - \frac{1}{p}(\operatorname{tr}\Lambda)I_p\big)U^T\Big)\big|$$
$$\leq \sqrt{\operatorname{tr}\left(H_{Z\tilde{J}}^2\right)}\sqrt{\operatorname{tr}\Big(\Lambda - \frac{1}{p}\big(\operatorname{tr}\Lambda\big)I_p\Big)^2} = \sqrt{(n-k)\operatorname{tr}\Big(\Lambda - \frac{1}{p}\big(\operatorname{tr}\Lambda\big)I_p\Big)^2} = o(\sqrt{p}).$$

Hence

$$\operatorname{tr}(\Lambda^{1/2}U^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2}) = \frac{p-n+k}{p}\operatorname{tr}(\Lambda) + o(\sqrt{p}).$$

It follows that

$$\begin{aligned}
&\Big(\sum_{i=1}^{p}\lambda_i^2(\Lambda^{1/2}U^T(I_p - H_{Z\tilde{Z}})U\Lambda^{1/2})\Big)^{-1/2} \\
&\Big(G_2^T\Lambda^{1/2}U^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2}G_2 - \sum_{i=1}^{p}\lambda_i(\Lambda^{1/2}U^T(I_p - H_{Z\tilde{Z}})U\Lambda^{1/2})I_{k-1}\Big) \\
=&\Big((1+O_P(\frac{n}{p}))\operatorname{tr}(\Lambda^2)\Big)^{-1/2}\Big(G_2^T\Lambda^{1/2}U^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2}G_2 - \big(\frac{p-n+k}{p}\operatorname{tr}(\Lambda) + O_P(\sqrt{p})\big)I_{k-1}\Big)
\end{aligned}$$

By Slutsky's theorem, we have that

$$\frac{1}{\sqrt{\operatorname{tr}(\Lambda_2^2)}}\Big(G_2^T\Lambda^{1/2}U^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2}G_2 - \frac{p-n+k}{p}\operatorname{tr}(\Lambda)I_{k-1}\Big) \xrightarrow{\mathcal{L}} W_{k-1}$$

Note that

$$\begin{aligned}
&\operatorname{E}\big[\|C^T\Xi^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2}G_2\|_F^2\big] \\
=&(k-1)\operatorname{E}\big[\operatorname{tr}\big(C^T\Xi^T(I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})\Xi C\big)\big] \\
\le&(k-1)\operatorname{E}\big[\lambda_1\big((I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})\big)\big]\|\Xi C\|_F^2 \\
\le&(k-1)\lambda_1(\Lambda)\|\Xi C\|_F^2 \le (k-1)C\|\Xi C\|_F^2 = o(p),
\end{aligned}$$

we have

$$\frac{1}{\sqrt{\operatorname{tr}(\Lambda_2^2)}}\Big(C^TJ^TZ^T(I_p - H_{Z\tilde{J}})ZJC - \frac{p-n+k}{p}\operatorname{tr}(\Sigma)I_{k-1} - C^T\Xi^T(I_p - H_{Z\tilde{J}})\Xi C\Big) \xrightarrow{\mathcal{L}} W_{k-1}.$$

Equivalently, we have

$$\begin{aligned}
&\frac{1}{\sqrt{\operatorname{tr}(\Lambda_2^2)}}\Big(C^TJ^TZ^T(I_p - H_{Z\tilde{J}})ZJC - \frac{p-n+k}{p}\operatorname{tr}(\Sigma)I_{k-1}\Big) \\
\sim&\frac{1}{\sqrt{\operatorname{tr}(\Lambda_2^2)}}C^T\Xi^T(I_p - H_{Z\tilde{J}})\Xi C + W_{k-1} + o_P(1).
\end{aligned}$$

Then the conclusion follows by taking the maximum eigenvalue. $\square$

The following lemma gives the asympototics of $\lambda_i(\tilde{J}^T Z^T Z \tilde{J})$, $i = 1, \ldots, r$.

Lemma 1. *Under the Assumptions of Theorem 2, we have* $\lambda_i(\tilde{J}^T Z^T Z \tilde{J}) = \lambda_i n(1 + o_P(1))$, $i = 1, \ldots, r$.

Proof. Note that $\tilde{J}^T Z^T Z \tilde{J} = G_1^T \Lambda G_1 = V_{Z\tilde{J}} D_{Z\tilde{J}}^2 V_{Z\tilde{J}}^T$, and $G_1^T \Lambda G_1 = G_{1[1:r,]}^T \Lambda_1 G_{1[1:r,]} + G_{1[(r+1):p,]}^T \Lambda_2 G_{1[(r+1):p,]}$. We have

$$V_{Z\tilde{J}} D_{Z\tilde{J}}^2 V_{Z\tilde{J}}^T = G_{1[1:r,]}^T \Lambda_1 G_{1[1:r,]} + G_{1[(r+1):p,]}^T \Lambda_2 G_{1[(r+1):p,]}.$$

For $i = 1, \ldots, r$,

$$\begin{aligned} &\lambda_i(G_{1[1:r,]}^T \Lambda_1 G_{1[1:r,]}) \geq \lambda_i(G_{1[1:r,]}^T \operatorname{diag}(\lambda_i I_i, O_{(r-i)\times(r-i)}) G_{1[1:r,]}) \\ =&\lambda_i \lambda_i(G_{1[1:i,]} G_{1[1:i,]}^T) = \lambda_i n(1 + o_P(1)), \end{aligned} \tag{5.11}$$

where the last equality holds since $n^{-1} G_{1[1:i,]} G_{1[1:i,]}^T \xrightarrow{P} I_i$ by law of large numbers. On the other hand, for $i = 1, \ldots, r$,

$$\begin{aligned} &\lambda_i(G_{1[1:r,]}^T \Lambda_1 G_{1[1:r,]}) \\ =&\lambda_i\Big(G_{1[1:r,]}^T \big(\operatorname{diag}(\lambda_1, \ldots, \lambda_{i-1}, O_{(r-i+1)\times(r-i+1)}) + \operatorname{diag}(O_{(i-1)\times(i-1)}, \lambda_i, \ldots, \lambda_r)\big) G_{1[1:r,]}\Big) \\ \leq&\lambda_1(G_{1[1:r,]}^T \operatorname{diag}(O_{(i-1)\times(i-1)}, \lambda_i, \ldots, \lambda_r) G_{1[1:r,]}) \leq \lambda_1(G_{1[1:r,]}^T \operatorname{diag}(O_{(i-1)\times(i-1)}, \lambda_i I_{r-i+1}) G_{1[1:r,]}) \\ =&\lambda_i \lambda_1(G_{1[i:r,]} G_{1[i:r,]}^T) = \lambda_i n(1 + o_P(1)) \end{aligned} \tag{5.12}$$

where the first inequality holds by Weyl's inequality. It follows from (5.11) and (5.12) that $\lambda_i(G_{1[1:r,]}^T \Lambda_1 G_{1[1:r,]}) = \lambda_i n(1 + o_P(1))$ for $i = 1, \ldots, r$.

Note that $\lambda_{\max}(G_{1[(r+1):p,]}^T \Lambda_2 G_{1[(r+1):p,]}) \leq C\lambda_{\max}(G_{1[(r+1):p,]}^T G_{1[(r+1):p,]}) = O_P(p)$ by Bai-Yin's law. By assumption $\lambda_r n/p \to \infty$, we can deduce that $D_{Z\tilde{J}[i,i]}^2 = \lambda_i(G_1^T \Lambda G_1) = \lambda_i n(1 + o_P(1))$, $i = 1, \ldots, r$.

□

The next lemma gives the asymptotics of $U_{Z\tilde{J}[,1:r]}$.

Lemma 2. *Under the Assumptions of Theorem 2, we have*

$$\lambda_{\max}(I_r - U_1^T U_{Z\tilde{J}[,1:r]} U_{Z\tilde{J}[,1:r]}^T U_1) = O_P(\frac{p}{\lambda_r n}).$$

Proof. Note that $U\Lambda^{1/2} G_1 G_1^T \Lambda^{1/2} U^T = U_{Z\tilde{J}} D_{Z\tilde{J}}^2 U_{Z\tilde{J}}^T$, we have $G_1 G_1^T = \Lambda^{-1/2} U^T U_{Z\tilde{J}} D_{Z\tilde{J}}^2 U_{Z\tilde{J}}^T U \Lambda^{-1/2}$. Thus,

$$\begin{aligned} G_{1[(r+1):p,]} G_{1[(r+1):p,]}^T &= \Lambda_2^{-1/2} U_{[,(r+1):p]}^T U_{Z\tilde{J}} D_{Z\tilde{J}}^2 U_{Z\tilde{J}}^T U_{[,(r+1):p]} \Lambda_2^{-1/2} \\ &\geq \Lambda_2^{-1/2} U_{[,(r+1):p]}^T U_{Z\tilde{J}[,1:r]} D_{Z\tilde{J}[1:r,1:r]}^2 U_{Z\tilde{J}[,1:r]}^T U_{[,(r+1):p]} \Lambda_2^{-1/2} \\ &\geq D_{Z\tilde{J}[r,r]}^2 \Lambda_2^{-1/2} U_{[,(r+1):p]}^T U_{Z\tilde{J}[,1:r]} U_{Z\tilde{J}[,1:r]}^T U_{[,(r+1):p]} \Lambda_2^{-1/2}. \end{aligned}$$

It follows that

$$\lambda_{\max}(U_{[,(r+1):p]}^T U_{Z\tilde{J}[,1:r]} U_{Z\tilde{J}[,1:r]}^T U_{[,(r+1):p]}) \leq \frac{C}{D_{Z\tilde{J}[r,r]}^2} \lambda_1(G_{1[(r+1):p,]} G_{1[(r+1):p,]}^T) = O_P(\frac{p}{\lambda_r n}),$$

where the last equality follows by Lemma 1 and Weyl's inequality.

The conclusion follows by the following simple relationship

$$
\begin{aligned}
&\lambda_{\max}(U^T_{[,(r+1):p]}U_{Z\tilde{J}[,1:r]}U^T_{Z\tilde{J}[,1:r]}U_{[,(r+1):p]}) = \lambda_{\max}(U^T_{Z\tilde{J}[,1:r]}U_{[,(r+1):p]}U^T_{[,(r+1):p]}U_{Z\tilde{J}[,1:r]})\\
=&\lambda_{\max}(U^T_{Z\tilde{J}[,1:r]}(I_p - U_1U_1^T)U_{Z\tilde{J}[,1:r]}) = \lambda_{\max}(I_r - U^T_{Z\tilde{J}[,1:r]}U_1U_1^TU_{Z\tilde{J}[,1:r]})\\
=&1 - \lambda_{\min}(U^T_{Z\tilde{J}[,1:r]}U_1U_1^TU_{Z\tilde{J}[,1:r]}) = 1 - \lambda_{\min}(U_1^TU_{Z\tilde{J}[,1:r]}U^T_{Z\tilde{J}[,1:r]}U_1)\\
=&\lambda_{\max}(I_r - U_1^TU_{Z\tilde{J}[,1:r]}U^T_{Z\tilde{J}[,1:r]}U_1).
\end{aligned}
$$

□

Proof of Theorem 2. As in the proof of Theorem 1, for $i = r+1, \ldots, p$, we have that

$$\lambda_i(\Lambda^{1/2}U^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2}) \le \lambda_i(\Lambda). \tag{5.13}$$

And for $i = 1, \ldots, p-n+k$, we have

$$\lambda_i(\Lambda^{1/2}U^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2}) \ge \lambda_{i+n-k}(\Lambda). \tag{5.14}$$

Next, we need to give an upper bound for $\lambda_i(\Lambda^{1/2}U^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2})$, $i = 1, \ldots, r$. Note that the positive eigenvalues of $\Lambda^{1/2}U^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2}$ are equal to the eigenvalues of $(I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})$. Write $(I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})$ as the sum of two terms

$$
\begin{aligned}
&(I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})\\
=&(I_p - H_{Z\tilde{J}})U_1\Lambda_1U_1^T(I_p - H_{Z\tilde{J}}) + (I_p - H_{Z\tilde{J}})U_2\Lambda_2U_2^T(I_p - H_{Z\tilde{J}}) \stackrel{def}{=} R_1 + R_2.
\end{aligned}
$$

Note that

$$\lambda_{\max}\big(R_1\big) = \lambda_{\max}\big(\Lambda_1^{1/2}U_1^T(I_p - H_{Z\tilde{J}})U_1\Lambda_1^{1/2}\big) \le \lambda_{\max}\big(\Lambda_1^{1/2}U_1^T(I_p - U_{Z\tilde{J}[,1:r]}U_{Z\tilde{J}[,1:r]}^T)U_1\Lambda_1^{1/2}\big)$$

$$\le \lambda_1\lambda_{\max}\big(U_1^T(I_p - U_{Z\tilde{J}[,1:r]}U_{Z\tilde{J}[,1:r]}^T)U_1\big) = \lambda_1\lambda_{\max}\big(I_r - U_1^T U_{Z\tilde{J}[,1:r]}U_{Z\tilde{J}[,1:r]}^T U_1\big) = O_P(\frac{\lambda_1 p}{\lambda_r n}).$$

The last equality follows by Lemma 2.

Thus, for $i = 1, \ldots, r$, we have

$$\lambda_i\big((I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})\big) = \lambda_i(R_1 + R_2) \le \lambda_1(R_1 + R_2) \le \lambda_1(R_1) + \lambda_1(R_2) = O_P(\frac{\lambda_1 p}{\lambda_r n}) + C.$$

As a consequence of these bounds, we have

$$\sum_{i=n-k+1}^{p} \lambda_i^2 \le \operatorname{tr}\big((I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})\big)^2 \le r(O_P(\frac{\lambda_1 p}{\lambda_r n}) + C)^2 + \sum_{i=r+1}^{p} \lambda_i^2,$$

or

$$\Big|\operatorname{tr}\big((I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})\big)^2 - \sum_{i=r+1}^{p} \lambda_i^2\Big| \le r(O_P(\frac{\lambda_1 p}{\lambda_r n}) + C)^2 + O(n). \tag{5.15}$$

Note that

$$\operatorname{tr}(R_2) = \operatorname{tr}(\Lambda_2) - \operatorname{tr}(H_{Z\tilde{J}}U_2\Lambda_2U_2^T).$$

and

$$\Big|\operatorname{tr}(H_{Z\tilde{J}}U_2\Lambda_2U_2^T) - \frac{n-k}{p-r}\operatorname{tr}(\Lambda_2)\Big| = \Big|\operatorname{tr}\Big(H_{Z\tilde{J}}U\big(\Lambda_2 - \frac{1}{p-r}(\operatorname{tr}\Lambda_2)I_{p-r}\big)U^T\Big)\Big|$$

$$\le\sqrt{\operatorname{tr}\big(H_{Z\tilde{J}}^2\big)}\sqrt{\operatorname{tr}\Big(\Lambda_2 - \frac{1}{p-r}\big(\operatorname{tr}\Lambda_2\big)I_{p-r}\Big)^2} = \sqrt{(n-k)\operatorname{tr}\Big(\Lambda_2 - \frac{1}{p-r}\big(\operatorname{tr}\Lambda_2\big)I_{p-r}\Big)^2} = o(\sqrt{p}).$$

Hence

$$\operatorname{tr}(R_2) = \frac{p-r-n+k}{p-r}\operatorname{tr}(\Lambda_2) + o(\sqrt{p}).$$

Then

$$\left| \operatorname{tr}[(R_1+R_2)] - \frac{p-r-n+k}{p-r}\operatorname{tr}(\Lambda_2)\right| \le rO_P\Big(\frac{\lambda_1 p}{\lambda_r n}\Big) + o(\sqrt{p}). \qquad (5.16)$$

Equation (5.15) and (5.16), conbined with the assumptions, yield

$$\operatorname{tr}\big((I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})\big)^2 = (1+o_P(1))\operatorname{tr}(\Lambda_2),$$

and

$$\operatorname{tr}\big((I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})\big) = \frac{p-r-n+k}{p-r}\operatorname{tr}(\Lambda_2) + o_P(\sqrt{p}).$$

Now we have the Lyapunov condition

$$\frac{\lambda_1\Big(\big((I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})\big)^2\Big)}{\operatorname{tr}\Big(\big((I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})\big)^2\Big)} = \frac{\big(O_P(\frac{\lambda_1 p}{\lambda_r n}) + C\big)^2}{(1+o_P(1))\operatorname{tr}(\Lambda_2)} \xrightarrow{P} 0.$$

Apply Lyapunov central limit theorem conditioning on $H_{Z\tilde{J}}$, we have

$$\Big(\operatorname{tr}\Big(\big((I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})\big)^2\Big)\Big)^{-1/2}$$
$$\big(G_2^T\Lambda^{1/2}U^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2}G_2 - \operatorname{tr}\big((I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})\big)I_{k-1}\big) \xrightarrow{\mathcal{L}} W_{k-1},$$

where W_{k-1} is a $(k-1)\times(k-1)$ symmetric random matrix whose entries above the main diagonal are i.i.d. $N(0,1)$ and the entries on the diagonal are i.i.d. $N(0,2)$. By Slutsky's theorem, we have

$$\frac{1}{\sqrt{\operatorname{tr}(\Lambda_2^2)}}\big(G_2^T\Lambda^{1/2}U^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2}G_2 - \tfrac{p-r-n+k}{p-r}\operatorname{tr}(\Lambda_2)I_{k-1}\big) \xrightarrow{\mathcal{L}} W_{k-1}.$$

As for the cross term of (5.7), we have

$$
\begin{aligned}
&\mathrm{E}[\|C^T\Xi^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2}G_2\|_F^2|Z\tilde{J}] \\
=&(k-1)\operatorname{tr}(C^T\Xi^T(I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})\Xi C) \\
\le&(k-1)\lambda_1\big((I_p - H_{Z\tilde{J}})U\Lambda U^T(I_p - H_{Z\tilde{J}})\big)\|\Xi C\|_F^2 \\
=&(k-1)O_P(\frac{\lambda_1 p}{\lambda_r n})\|\Xi C\|_F^2 \\
=&(k-1)O_P(\frac{\lambda_1\sqrt{p}}{\lambda_r n})\sqrt{p}\|\Xi C\|_F^2 = o_P(p)
\end{aligned}
$$

The last equality holds when we assume $\frac{1}{\sqrt{p}}\|\Xi C\|_F^2 = O(1)$. Hence $\|C^T\Xi^T(I_p - H_{Z\tilde{J}})U\Lambda^{1/2}G_2\|_F^2 = o_P(p)$, and we have

$$
\frac{1}{\sqrt{\operatorname{tr}(\Lambda_2^2)}}\big(C^TJ^TZ^T(I_p - H_{Z\tilde{J}})ZJC - \tfrac{p-r-n+k}{p-r}\operatorname{tr}(\Lambda_2)I_{k-1} - C^T\Xi^T(I_p - H_{Z\tilde{J}})\Xi C\big) \xrightarrow{\mathcal{L}} W_{k-1}.
$$

Equivalently, we have

$$
\begin{aligned}
&\frac{1}{\sqrt{\operatorname{tr}(\Lambda_2^2)}}\Big(C^TJ^TZ^T(I_p - H_{Z\tilde{J}})ZJC - \frac{p-r-n+k}{p-r}\operatorname{tr}(\Lambda_2)I_{k-1}\Big) \\
\sim&\frac{1}{\sqrt{\operatorname{tr}(\Lambda_2^2)}}C^T\Xi^T(I_p - H_{Z\tilde{J}})\Xi C + W_{k-1} + o_P(1).
\end{aligned}
$$

Then the conclusion follows by taking the maximum eigenvalue. □

Supplementary Materials

Contain the brief description of the online supplementary materials.

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