

SIMULTANEOUS ESTIMATION OF SCALE MATRICES IN TWO-SAMPLE PROBLEM UNDER ELLIPTICALLY CONTOURED DISTRIBUTIONS

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Abstract

Two-sample problems of estimating $p \times p$ scale matrices are investigated under elliptically contoured distributions. Two loss functions are employed; one is sum of Stein's loss functions of one-sample problem of estimating a normal covariance matrix and the other is a quadratic loss function for $\Sigma_2 \Sigma_1^{-1}$, where Σ_1 and Σ_2 are $p \times p$ scale matrices of elliptically contoured distributions. It is shown that improvement of the estimators obtained under the normality assumption remains robust under elliptically contoured distributions. A Monte Carlo study is also conducted to evaluate the risk performances of the improved estimators under three elliptically contoured distributions.

Keywords: covariance matrix, Stein's loss, Stein-Haff identity, two-sample problem

1 Introduction

Since the pioneer works of Stein (1956) and James and Stein (1961), there has been a great deal of effort to construct improved estimators for a covariance matrix of a multivariate normal distribution. The literature includes Haff (1980, 1982, 1991) and Dey and Srinivasan (1985). Two sample analogue of estimating covariance matrices has been also considered by several authors such as Muirhead and Veerathaworn (1985) and Loh (1991a, 1991b). On the other hand Kubokawa and Srivastava (1999) showed that improvement of minimax estimators for a covariance matrix obtained under the normality assumption remains robust under elliptically contoured distribution models. In this paper, following the set-up considered by Loh (1991a, 1991b), we examine two-sample problems of estimating scale matrices of elliptically contoured distributions.

The precise set-up of the problems is as follows: Let \mathbf{Y}_1 and \mathbf{Y}_2 be $N_1 \times p$ and $N_2 \times p$ random matrices which take multivariate linear models of the form

$$\mathbf{Y}_1 = \mathbf{C}_1 \boldsymbol{\beta}_1 + \boldsymbol{\epsilon}_1 \quad \text{and} \quad \mathbf{Y}_2 = \mathbf{C}_2 \boldsymbol{\beta}_2 + \boldsymbol{\epsilon}_2. \quad (1)$$

Here $\boldsymbol{\epsilon}_i$ ($i = 1, 2$) are $N_i \times p$ random matrices, \mathbf{C}_i are known $N_i \times m$ matrices with full rank, and $\boldsymbol{\beta}_i$ are unknown $m \times p$ matrices. We suppose that $N_i \geq m + p$. We also assume that the error matrices $\boldsymbol{\epsilon}_1$ and $\boldsymbol{\epsilon}_2$ are marginally distributed as elliptically contoured distributions. But we assume the error matrices to have two forms of the joint density functions: First, two error matrices $\boldsymbol{\epsilon}_1$ and $\boldsymbol{\epsilon}_2$ are independently distributed and have the joint density function

$$\prod_{i=1}^2 |\boldsymbol{\Sigma}_i|^{-N_i/2} g_i(\text{tr}(\boldsymbol{\Sigma}_i^{-1} \boldsymbol{\epsilon}'_i \boldsymbol{\epsilon}_i)), \quad (2)$$

where $\boldsymbol{\Sigma}_i$ ($i = 1, 2$) are $p \times p$ unknown positive definite matrices and g_i are nonnegative real-valued functions: Secondly two error matrices $\boldsymbol{\epsilon}_1$ and $\boldsymbol{\epsilon}_2$ are uncorrelatedly distributed and have the joint density function

$$|\boldsymbol{\Sigma}_1|^{-N_1/2} |\boldsymbol{\Sigma}_2|^{-N_2/2} g(\text{tr}(\boldsymbol{\Sigma}_1^{-1} \boldsymbol{\epsilon}'_1 \boldsymbol{\epsilon}_1 + \boldsymbol{\Sigma}_2^{-1} \boldsymbol{\epsilon}'_2 \boldsymbol{\epsilon}_2)), \quad (3)$$

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where g is nonnegative real-valued function. Here $|\mathbf{P}|$, $\text{tr}(\mathbf{P})$ and \mathbf{P}' stand for the determinant, the trace and the transpose of a square matrix \mathbf{P} , respectively.

Following the approaches due to Loh (1991a, 1991b), we consider two sets of estimation problems as follows.

(i) Under the model (1) with the assumption (2), the problem of estimating $(\boldsymbol{\Sigma}_1, \boldsymbol{\Sigma}_2)$ with unknown parameters $(\boldsymbol{\beta}_1, \boldsymbol{\beta}_2)$ is considered under a loss function

$$L_1(\widehat{\boldsymbol{\Sigma}}_1, \widehat{\boldsymbol{\Sigma}}_2, \boldsymbol{\Sigma}_1, \boldsymbol{\Sigma}_2) = \sum_{i=1}^2 \{\text{tr}(\widehat{\boldsymbol{\Sigma}}_i \boldsymbol{\Sigma}_i^{-1}) - \log |\widehat{\boldsymbol{\Sigma}}_i \boldsymbol{\Sigma}_i^{-1}| - p\}, \quad (4)$$

where $\widehat{\boldsymbol{\Sigma}}_i$, $i = 1, 2$, are estimators of $\boldsymbol{\Sigma}_i$, respectively. This loss function is a natural extension of Stein's loss function in the one-sample case. We denote by $R_1(\widehat{\boldsymbol{\Sigma}}_1, \widehat{\boldsymbol{\Sigma}}_2, \boldsymbol{\Sigma}_1, \boldsymbol{\Sigma}_2)$ the corresponding risk function, the expected loss function L_1 , where the expectation is taken with respect to the joint distributions of \mathbf{Y}_1 and \mathbf{Y}_2 .

(ii) Under the model (1) with the assumption (3), the problem of estimating $\boldsymbol{\zeta} = \boldsymbol{\Sigma}_2 \boldsymbol{\Sigma}_1^{-1}$ with unknown parameters $(\boldsymbol{\beta}_1, \boldsymbol{\beta}_2)$ is considered under a loss function

$$L_2(\widehat{\boldsymbol{\zeta}}, \boldsymbol{\zeta}) = \text{tr}\{\boldsymbol{\Sigma}_2^{-1}(\widehat{\boldsymbol{\zeta}} - \boldsymbol{\zeta})\mathbf{S}_1(\widehat{\boldsymbol{\zeta}} - \boldsymbol{\zeta})'\}/\text{tr}(\boldsymbol{\zeta}), \quad (5)$$

where $\widehat{\boldsymbol{\zeta}}$ is an estimator of $\boldsymbol{\zeta}$ and $\mathbf{S}_1 = \mathbf{Y}'_1(\mathbf{I}_{N_1} - \mathbf{C}_1(\mathbf{C}'_1\mathbf{C}_1)^{-1}\mathbf{C}'_1)\mathbf{Y}_1$. Again we denote by $R_2(\widehat{\boldsymbol{\zeta}}, \boldsymbol{\zeta})$ the corresponding risk function. This estimation problem is related to estimation of the common mean of two multivariate distributions. See a possible motivation for Loh (1991b). Furthermore, the eigenvalues of $\boldsymbol{\zeta}$ are important, for example, in the problem of testing the null hypotheses $\boldsymbol{\Sigma}_1 = \boldsymbol{\Sigma}_2$ against the alternative hypotheses $\boldsymbol{\Sigma}_1 \neq \boldsymbol{\Sigma}_2$. For estimating these eigenvalues, see Muirhead and Verathaworn (1985), Muirhead (1987), and DasGupta (1989).

This paper is organized in the following way. In Section 2, we treat the problem (i). We adapt the extended Stein and Haff identity due to Kubokawa and Srivastava (1999) for two sample set-up (which is stated in Section 4) and give a result that, with respect to the loss function (4), an alternative estimator improves upon the James-Stein estimator $(\mathbf{T}_1\mathbf{D}_1\mathbf{T}'_1, \mathbf{T}_2\mathbf{D}_2\mathbf{T}'_2)$ under the elliptically contoured error. Here \mathbf{T}_i , $i = 1, 2$, is the lower triangular matrix with positive diagonal elements such that $\mathbf{S}_i = \mathbf{T}_i\mathbf{T}'_i$ and \mathbf{D}_i is diagonal matrix with the j -th diagonal element $1/(N_i - m + p + 1 - 2j)$, $j = 1, 2, \dots, p$, where $\mathbf{S}_i = \mathbf{Y}'_i(\mathbf{I}_{N_i} - \mathbf{C}_i(\mathbf{C}'_i\mathbf{C}_i)^{-1}\mathbf{C}'_i)\mathbf{Y}_i$, $i = 1, 2$. Simulation study is conducted to evaluate risk performances of alternative estimators under the multivariate normal distribution, the matrix-variate t -distribution, and the matrix-variate Kotz-type distribution. Since these matrix-variate distributions except the normal distribution are not independent sampling, we also conduct simulation study based on independently and identically sampling model from the multivariate t -distribution and the Kotz-type distribution, respectively. Finding in this Section is that the estimators obtained under the error distribution (2) (i.e., which is different from independently and identically sampling model) perform well under independently and identically sampling from the elliptically contoured error models. In Section 3, we treat the problem (ii). In this problem, we treat the joint density function (3) only since we fail to obtain the suitable integration-by-parts formula under the joint density (2) to get improved estimators. We first obtain the best estimator among the constant multiple of $\mathbf{S}_2\mathbf{S}_1^{-1}$. Next we consider several types of improvement over the best constant multiple of $\mathbf{S}_2\mathbf{S}_1^{-1}$ and conduct simulation study in the much same way as that of Section 2. In Section 4 we sketch the proofs of the results obtained in Sections 2 and 3. In Section 5, we give the concluding remark on the result we obtained in this paper.

2 Simultaneous estimation of $(\boldsymbol{\Sigma}_1, \boldsymbol{\Sigma}_2)$

To consider the estimation problem of $(\boldsymbol{\Sigma}_1, \boldsymbol{\Sigma}_2)$, we shall employ the loss function (4) and evaluate performance of estimators of $(\widehat{\boldsymbol{\Sigma}}_1, \widehat{\boldsymbol{\Sigma}}_2)$ by means of their risk function R_1 , i.e., $\mathbf{E}[L_1(\widehat{\boldsymbol{\Sigma}}_1, \widehat{\boldsymbol{\Sigma}}_2, \boldsymbol{\Sigma}_1, \boldsymbol{\Sigma}_2)]$, where the expectation is taken with respect to the joint distributions of two error distributions (2).

2.1 Class of estimators

First we introduce estimators obtained from one-sample problem of estimating a normal covariance matrix. We define the usual estimator

$$(\widehat{\boldsymbol{\Sigma}}_1^{US}, \widehat{\boldsymbol{\Sigma}}_2^{US}) = (\mathbf{S}_1/n_1, \mathbf{S}_2/n_2), \quad (6)$$

where $n_i = N_i - m$, $i = 1, 2$. Also we put the James-Stein estimator

$$(\widehat{\boldsymbol{\Sigma}}_1^{JS}, \widehat{\boldsymbol{\Sigma}}_2^{JS}) = (\mathbf{T}_1 \mathbf{D}_1 \mathbf{T}_1', \mathbf{T}_2 \mathbf{D}_2 \mathbf{T}_2'), \quad (7)$$

where \mathbf{T}_i , $i = 1, 2$, is the lower triangular matrix with positive diagonal elements such that $\mathbf{S}_i = \mathbf{T}_i \mathbf{T}_i'$ and \mathbf{D}_i is diagonal matrix with the j -th diagonal element $1/(n_i + p + 1 - 2j)$, $j = 1, 2, \dots, p$. Note that the James-Stein estimator (7) is invariant under the group of transformations given by

$$\boldsymbol{\Sigma}_i \rightarrow \mathbf{P}_i \boldsymbol{\Sigma}_i \mathbf{P}_i', \quad \mathbf{S}_i \rightarrow \mathbf{P}_i \mathbf{S}_i \mathbf{P}_i', \quad i = 1, 2,$$

where \mathbf{P}_i is any $p \times p$ lower triangular matrix with positive diagonal elements. From the argument from James and Stein (1961), we can see that the estimator (7) has smaller risk than that of the estimator (6). In the same fashion as in Kubokawa and Srivastava (1999), we can see that improvement of the estimator (7) over the usual estimator (6) remains robust for all possible functions g_1 and g_2 in (2). To improve upon the estimator (7) by using both \mathbf{S}_1 and \mathbf{S}_2 simultaneously, we adapt the argument due to Loh (1991) and consider a class of invariant estimator under the group of the transformations

$$\boldsymbol{\Sigma}_i \rightarrow \mathbf{Q} \boldsymbol{\Sigma}_i \mathbf{Q}', \quad \mathbf{S}_i \rightarrow \mathbf{Q} \mathbf{S}_i \mathbf{Q}', \quad i = 1, 2, \quad (8)$$

where \mathbf{Q} is any $p \times p$ nonsingular matrix. From Loh (1991a), we can see that an invariant estimator under the above group transformations has the form

$$(\widehat{\boldsymbol{\Sigma}}_1^{EQ}, \widehat{\boldsymbol{\Sigma}}_2^{EQ}) = (\mathbf{B}^{-1} \boldsymbol{\Psi}(\mathbf{F}) \mathbf{B}'^{-1}, \mathbf{B}^{-1} \boldsymbol{\Phi}(\mathbf{F}) \mathbf{B}'^{-1}). \quad (9)$$

Here we assume that \mathbf{B} is a nonsingular matrix such that $\mathbf{B}(\mathbf{S}_1 + \mathbf{S}_2) \mathbf{B}' = \mathbf{I}_p$, $\mathbf{B} \mathbf{S}_2 \mathbf{B}' = \mathbf{F}$, and $\mathbf{F} = \mathbf{diag}(f_1, f_2, \dots, f_p)$ with $f_1 \geq f_2 \geq \dots \geq f_p > 0$, and that $\boldsymbol{\Psi}(\mathbf{F}) = \mathbf{diag}(\psi_1(\mathbf{F}), \psi_2(\mathbf{F}), \dots, \psi_p(\mathbf{F}))$ and $\boldsymbol{\Phi}(\mathbf{F}) = \mathbf{diag}(\phi_1(\mathbf{F}), \phi_2(\mathbf{F}), \dots, \phi_p(\mathbf{F}))$ are diagonal matrices whose elements are functions of \mathbf{F} . In the sequel of the paper, we abbreviate $\boldsymbol{\Psi}$, $\boldsymbol{\Phi}$, ψ_i , ϕ_i ($i = 1, 2, \dots, p$) for $\boldsymbol{\Psi}(\mathbf{F})$, $\boldsymbol{\Phi}(\mathbf{F})$, $\psi_i(\mathbf{F})$, $\phi_i(\mathbf{F})$, respectively.

2.2 An improvement upon the James-Stein estimator

The form of the invariant estimators (9) suggests us to utilize the eigenvalues of $\mathbf{S}_2(\mathbf{S}_1 + \mathbf{S}_2)^{-1}$ for the simultaneous estimation of $\boldsymbol{\Sigma}_1$ and $\boldsymbol{\Sigma}_2$. Furthermore, it is well-known that the eigenvalues of $\mathbf{S}_2(\mathbf{S}_1 + \mathbf{S}_2)^{-1}$ are more spread out than those of $\mathbf{E}[\mathbf{S}_2(\mathbf{S}_1 + \mathbf{S}_2)^{-1}]$. In the light of this fact, Loh (1991a) introduced an alternative estimator as a special case of $(\widehat{\boldsymbol{\Sigma}}_1^{EQ}, \widehat{\boldsymbol{\Sigma}}_2^{EQ})$, which is given by

$$(\widehat{\boldsymbol{\Sigma}}_1^{LO}, \widehat{\boldsymbol{\Sigma}}_2^{LO}) = (\mathbf{B}^{-1} \boldsymbol{\Psi}^{LO} \mathbf{B}'^{-1}, \mathbf{B}^{-1} \boldsymbol{\Phi}^{LO} \mathbf{B}'^{-1}), \quad (10)$$

where $\boldsymbol{\Psi}^{LO} = \mathbf{diag}(\psi_1^{LO}, \dots, \psi_p^{LO})$ and $\boldsymbol{\Phi}^{LO} = \mathbf{diag}(\phi_1^{LO}, \dots, \phi_p^{LO})$ with the j -th diagonal elements $\psi_j^{LO} = (1 - f_j)/(n_1 - p - 1 + 2j)$ and $\phi_j^{LO} = f_j/(n_2 + p + 1 - 2j)$, $j = 1, 2, \dots, p$, respectively. He showed that his estimator improves upon the James-Stein estimator $(\widehat{\boldsymbol{\Sigma}}_1^{JS}, \widehat{\boldsymbol{\Sigma}}_2^{JS})$ under the assumption where the error matrices in (1) follow the multivariate normal distributions independently, that is, he obtained the constant multiples in ψ_i^{LO} and ϕ_i^{LO} such that the estimator (10) is better than the James-Stein estimator (7) in terms of the loss function (4) when \mathbf{Y}_1 and \mathbf{Y}_2 follow independently as the multivariate normal distributions. However, it is not obvious whether this improvement remains robust under the assumption where two error matrices follow the elliptically contoured error distributions independently. The following theorem answers this question affirmatively.

Theorem 2.1 *The estimator $(\widehat{\boldsymbol{\Sigma}}_1^{LO}, \widehat{\boldsymbol{\Sigma}}_2^{LO})$ is better than $(\widehat{\boldsymbol{\Sigma}}_1^{JS}, \widehat{\boldsymbol{\Sigma}}_2^{JS})$ for arbitrary g_1 and g_2 in (2) in terms of the loss function (4).*

Our approach to prove Theorem 1 is called “unbiased risk estimate methods” and it is summarized in the following: We compute the unbiased estimate of the risk of equivariant estimators with respect to certain expectation via Kubokawa and Srivastava’s integration-by-parts method for the elliptically contoured distributions, which results in sufficient conditions for equivariant estimators to improve upon the usual estimator. Then we determine promising alternative estimators from the sufficient conditions. We put detail proof of Theorem 1 and necessary technical lemmas in Section 4.

Remark : When the error (ϵ_1, ϵ_2) have the joint density (3), we can see the dominance result similar to that in Theorem above.

2.3 Numerical studies

Furthermore we introduce the Dey-Srinivasan estimator

$$(\widehat{\Sigma}_1^{DS}, \widehat{\Sigma}_2^{DS}) = (\mathbf{H}_1 \mathbf{D}_1 \mathbf{K}_1 \mathbf{H}_1', \mathbf{H}_2 \mathbf{D}_2 \mathbf{K}_2 \mathbf{H}_2'), \quad (11)$$

where \mathbf{H}_i , $i = 1, 2$, is a $p \times p$ orthogonal matrix such that $\mathbf{S}_i = \mathbf{H}_i \mathbf{K}_i \mathbf{H}_i'$ and $\mathbf{K}_i = \mathbf{diag}(k_{i1}, k_{i2}, \dots, k_{ip})$ with $k_{i1} \geq k_{i2} \geq \dots \geq k_{ip} > 0$.

Note that the Dey-Srinivasan estimator (11) is invariant under the group of transformations given by

$$\Sigma_i \rightarrow \mathbf{O}_i \Sigma_i \mathbf{O}_i', \quad \mathbf{S}_i \rightarrow \mathbf{O}_i \mathbf{S}_i \mathbf{O}_i', \quad i = 1, 2,$$

where \mathbf{O}_i is any $p \times p$ orthogonal matrix. In the same way as in Kubokawa and Srivastava (1999), we can see that the estimator (11) improves upon the estimator (7) for the two-sample set-up. However, it is difficult to compare $(\widehat{\Sigma}_1^{LO}, \widehat{\Sigma}_2^{LO})$ with $(\widehat{\Sigma}_1^{DS}, \widehat{\Sigma}_2^{DS})$ analytically. Therefore, to compare the risk performances of these estimators, we carry out Monte Carlo simulations.

Our simulations are based on 10,000 independent replications. We consider three-type of error distributions which are given in the following.

1. The matrix-variate normal distribution: The joint density function (ϵ_1, ϵ_2) is given by

$$\prod_{i=1}^2 c_{i1} |\Sigma_i|^{-N_i/2} \exp[-(1/2) \text{tr}(\Sigma_i^{-1} \epsilon_i' \epsilon_i)],$$

where $c_{i1} = (2\pi)^{-N_i p/2}$.

2. The t -distribution: The joint density function (ϵ_1, ϵ_2) is given by

$$\prod_{i=1}^2 c_{i2} |\Sigma_i|^{-N_i/2} \{1 + (1/v_i) \text{tr}(\Sigma_i^{-1} \epsilon_i' \epsilon_i)\}^{-(v_i + N_i p)/2},$$

where $c_{i2} = \Gamma[\{v_i + N_i p\}/2] / \{(\pi v_i)^{N_i p/2} \Gamma[v_i/2]\}$, $v_i > 0$. Here we denote by $\Gamma(\cdot)$ the Gamma function.

3. The Kotz-type distribution: The joint density function (ϵ_1, ϵ_2) is given by

$$\prod_{i=1}^2 c_{i3} |\Sigma_i|^{-N_i/2} \{\text{tr}(\Sigma_i^{-1} \epsilon_i' \epsilon_i)\}^{u_i-1} \exp[-r_i \{\text{tr} \Sigma_i^{-1} \epsilon_i' \epsilon_i\}^{s_i}],$$

where $r_i > 0$, $s_i > 0$, $2u_i + N_i p > 2$, and

$$c_{i3} = \frac{s_i \Gamma[N_i p/2] r_i^{\{u_i + N_i p/2 - 1\}/s_i}}{\pi^{N_i p/2} \Gamma[\{u_i + N_i p/2 - 1\}/s_i]}.$$

For generating a random number of the Kotz-type distribution above, see Fang, Kotz, and Ng (1990) for example.

For Monte Carlo simulations, we took $N_1 = N_2 = 15$, $m = 1$, and $p = 3$ and we also put $v_1 = v_2 = 3$ for the t -distribution and $(u_i, r_i, s_i) = (5, 0.1, 2)$, $i = 1, 2$, for the Kotz-type distribution. We also suppose that $\beta_1 = \beta_2 = (0, 0, 0)'$ and that the parameter $\Sigma_2 \Sigma_1^{-1}$ is the diagonal matrix with typical elements. The estimated risks of these cases are given by Tables 1–3 and their estimated standard errors are in parentheses. Three distributions we used in the simulation study are characterized via the kurtosis parameter for the elliptically contoured distribution in the following way. The multivariate

normal distribution has zero kurtosis while the Kotz-type distribution with parameters $(u_1, r_1, s_1) = (5, 0.1, 2)$ has the negative kurtosis. The t -distribution with $v_1 = 3$ dose not have the kurtosis parameter.

In Tables 1–3, ‘ US ’, ‘ JS ’, ‘ DS ’, and ‘ LO ’ stand for the usual estimator $(\widehat{\Sigma}_1^{US}, \widehat{\Sigma}_2^{US})$, the James-Stein estimator $(\widehat{\Sigma}_1^{JS}, \widehat{\Sigma}_2^{JS})$, the Dey-Srinivasan estimator $(\widehat{\Sigma}_1^{DS}, \widehat{\Sigma}_2^{DS})$, and the Loh estimator $(\widehat{\Sigma}_1^{LO}, \widehat{\Sigma}_2^{LO})$, respectively and ‘ AI ’ stands for average of improvement in risk over $(\widehat{\Sigma}_1^{US}, \widehat{\Sigma}_2^{US})$.

We also carried out simulations when the rows of $\epsilon_i (i = 1, 2)$ have densities

$$|\Sigma_i|^{-N_i/2} h(\mathbf{e}'_{ij} \Sigma_i^{-1} \mathbf{e}_{ij}), \quad \text{for } j = 1, \dots, N_i,$$

where $\epsilon_i = (\mathbf{e}'_{i1}, \mathbf{e}'_{i2}, \dots, \mathbf{e}'_{iN_i})'$. That is, the rows of each error matrix ϵ_i are independently and identically distributed (i.i.d.) as an elliptically contoured distribution.

For Monte Carlo simulations, we suppose that the rows of ϵ_i follow the vector-valued t -distributions, i.e., the density function of the random vectors $\mathbf{e}_{ij} (i = 1, 2, j = 1, 2, \dots, N_i)$ are given by

$$c_{i4} |\Sigma_i|^{-1/2} (1 + \mathbf{e}'_{ij} \Sigma_i^{-1} \mathbf{e}_{ij} / v_i)^{-(v_i+p)/2}, \quad (12)$$

where $v_i > 0$ and $c_{i4} = \Gamma[(v_i + p)/2] / \{(\pi v_i)^{p/2} \Gamma[v_i/2]\}$, and we also suppose that the rows of ϵ_i follow the vector-valued Kotz-type distributions, i.e., the density functions of the random vectors $\mathbf{e}_{ij} (i = 1, 2, j = 1, 2, \dots, N_i)$ are given by

$$c_{i5} |\Sigma_i|^{-1/2} \{\mathbf{e}'_{ij} \Sigma_i^{-1} \mathbf{e}_{ij}\}^{u_i-1} \exp[-r_i \{\mathbf{e}'_{ij} \Sigma_i^{-1} \mathbf{e}_{ij}\}^{s_i}], \quad (13)$$

where $r_i > 0$, $s_i > 0$, $2u_i + p > 2$, and

$$c_{i5} = \frac{s_i \Gamma[p/2] r_i^{\{u_i+p/2-1\}/s_i}}{\pi^{p/2} \Gamma[\{u_i + p/2 - 1\}/s_i]}.$$

For simulations, we took $N_1 = N_2 = 15$, $m = 1$, and $p = 3$ and we put $v_1 = v_2 = 3$ for the t -distributions and $(u_i, r_i, s_i) = (5, 0.1, 2)$, $i = 1, 2$, for the Kotz-type distributions. The estimated risks of these cases are given by Tables 4 and 5.

Furthermore, we put $N_1 = N_2 = 15$ and carried out simulation for higher dimensions, i.e., $p = 5, 8, 10$ when the error matrices follow the normal distribution (Table 6). Since Table 1 indicates that the maximum risk reduction is observed when the eigenvalues of $\Sigma_2 \Sigma_1^{-1}$ are equal and that the minimum risk reduction is observed when the eigenvalues of $\Sigma_2 \Sigma_1^{-1}$ are spread out, we choose the type of eigenvalues of $\Sigma_2 \Sigma_1^{-1}$ for two typical cases. To see how the kurtosis parameter effects risk reduction, we also carried out simulation when the error matrices follow the Kotz-type distribution with zero-kurtosis (Tables 7 and 8). The parameters can be chosen in the following manner: For p -variate Kotz-type distribution with the density (13), its kurtosis is given by

$$\kappa_5 = \left(\Gamma\left(\frac{2(u_i + 2) + p - 2}{2s_i}\right) / \Gamma\left(\frac{2u_i + p - 2}{2s_i}\right) r_i^{2/s_i} p(p + 2) \right) - 1.$$

Its kurtosis is equal to zero if

$$\begin{cases} s_i = 1/2, \\ r_i = 4 \sqrt{\frac{(2u_i + p + 1)(2u_i + p)(2u_i + p - 1)(2u_i + p - 2)}{p(p + 2)}}, \\ u_i > (2 - p)/2. \end{cases}$$

We summarize our numerical results in Tables 1–8 as follows:

1. Tables 1–5 indicate that AI of LO is large when the eigenvalues of $\Sigma_2 \Sigma_1^{-1}$ are close together and that AI of LO is small when the eigenvalues of $\Sigma_2 \Sigma_1^{-1}$ are spread out:
2. DS is better than LO :
3. AIs of all alternative estimators are relatively small under non-normal error:

4. Tables 4 and 5 indicate that AIs of LO and DS are substantial under independently and identically sampling set-up from non-normal distribution, although we cannot prove improvement of alternative estimators over the usual estimator under this situation. Hence, these results suggest that the improvement under densities (2) remains robust even if the rows of errors are i.i.d.:
5. Table 6 indicates that the AI's increase with increasing dimension p when the sample size is fixed:
6. When the error follows the Kotz-type distribution with the kurtosis zero, the AI's are the same as those of the normal case. It seems that the AI's depend on the kurtosis rather than distribution type as long as the distributions of the error matrices belong to the elliptically contoured distributions.

3 Estimation of $\Sigma_2 \Sigma_1^{-1}$

In this section we consider the problem (ii) given in Section 1 under elliptically contoured error with density (3) and we treat the problem under the loss function

$$L_2(\hat{\zeta}; \zeta) = \text{tr}\{\Sigma_2^{-1}(\hat{\zeta} - \zeta)\mathbf{S}_1(\hat{\zeta} - \zeta)'\}/\text{tr}(\zeta), \quad (14)$$

as considered by Loh (1988). Recall that $\mathbf{S}_i = \mathbf{Y}'_i(\mathbf{I}_{N_i} - \mathbf{C}_i(\mathbf{C}'_i\mathbf{C}_i)^{-1}\mathbf{C}'_i)\mathbf{Y}_i$ for $i = 1, 2$. As pointed out in Loh (1991b), the problem (ii) is invariant under the group of transformations given by (8) and the estimators which is invariant under this transformation group has the form

$$\hat{\zeta} = \mathbf{A}^{-1}\Xi\mathbf{A},$$

where \mathbf{A} is a nonsingular matrix such that $\mathbf{A}\mathbf{S}_1\mathbf{A}' = \mathbf{I}_p$ and $\mathbf{A}\mathbf{S}_2\mathbf{A}' = \mathbf{L}$ with $\mathbf{L} = \text{diag}(l_1, \dots, l_p)$ ($l_1 \geq l_2 \geq \dots \geq l_p > 0$) and further $\Xi = \text{diag}(\xi_1, \dots, \xi_p)$ whose diagonal elements are functions of \mathbf{L} .

3.1 The best constant multiplier of $\mathbf{S}_2\mathbf{S}_1^{-1}$

Consider a class of estimators of the form $\hat{\zeta}^{US} = \alpha\mathbf{S}_2\mathbf{S}_1^{-1}$, where α is a constant. Then this estimator can be rewritten as

$$\hat{\zeta}^{US} = \mathbf{A}^{-1}\Xi^{US}\mathbf{A},$$

where Ξ^{US} is diagonal matrix whose the j -th diagonal element is $\xi_j^{US} = \alpha l_j$, $j = 1, 2, \dots, p$.

Theorem 3.1 *For any function g in (3), the best usual estimator of ζ under the loss function (14) is given by*

$$\hat{\zeta}^{BU} = \mathbf{A}^{-1}\Xi^{BU}\mathbf{A}, \quad (15)$$

where $\Xi = \text{diag}(\xi_1^{BU}, \dots, \xi_p^{BU})$ with $\xi_j^{BU} = [(n_1 - p - 1)/(n_2 + p + 1)]l_j$.

3.2 Improved Estimators

We next discuss an improvement on the estimator (15). It is expected that the eigenvalues of $\mathbf{S}_2\mathbf{S}_1^{-1}$ are more spread out than those of $\mathbf{E}[\mathbf{S}_2\mathbf{S}_1^{-1}]$. To reduce the biases of the estimators for eigenvalues, we consider

$$\Xi^{LO} = \text{diag}(\xi_1^{LO}, \dots, \xi_p^{LO}), \quad \xi_j^{LO} = (n_1 - p - 1)l_j/(n_2 + p + 3 - 2j), \quad (16)$$

for $j = 1, 2, \dots, p$. Then we have the following theorem:

Theorem 3.2 *Under the loss function (14), $\hat{\zeta}^{LO} = \mathbf{A}^{-1}\Xi^{LO}\mathbf{A}$ is better than $\hat{\zeta}^{BU}$ for any function g in (3).*

Further we consider an improved estimator on $\hat{\zeta}^{LO}$ as in Loh (1988). Define the Berger-type estimator as

$$\hat{\zeta}^{BE} = \mathbf{A}^{-1} \mathbf{\Xi}^{BE} \mathbf{A}, \quad (17)$$

where $\mathbf{\Xi}^{BE} = \mathbf{diag}(\xi_1^{BE}, \dots, \xi_p^{BE})$ with

$$\xi_j^{BE} = \xi_j^{LO} + \frac{c}{b+u}, \quad u = \sum_{j=1}^p \left(\frac{n_2 + p + 3 - 2j}{(n_1 - p - 1)l_j} \right)^2.$$

Here $c : R^+ \mapsto R$ is a differentiable function of u and b is a suitable positive constant.

Then we have the following theorem:

Theorem 3.3 *Assume that*

- (I) $p \geq 3$, $n_1 \geq p$ and $n_2 \geq p$;
- (II) $c(u) \geq 0$ and $c'(u) \geq 0$ for all $u \geq 0$;
- (III) $\sup_u c(u)/\sqrt{b} \leq 4(p^2 + p - 4)(n_2 - p + 3)/[\sqrt{p}(n_1 - p - 1)(n_2 - p + 7)]$.

Then, under the loss function (14), $\hat{\zeta}^{BE}$ is better than $\hat{\zeta}^{LO}$ for any function g in (3).

Since $l_1 \geq \dots \geq l_p$, the diagonal elements of $\mathbf{\Xi}$ should have the ordering property, i.e., $\xi_1 \geq \dots \geq \xi_p$. Hence, to improve on $\hat{\zeta}^{LO}$, we can consider the estimator, for example, with Stein's isotonic regression on the ξ_j^{LO} 's (see Lin and Perlman, 1985).

Further we can also consider the Stein-type estimator (Stein, 1977)

$$\hat{\zeta}^{ST} = \mathbf{A}^{-1} \mathbf{\Xi}^{ST} \mathbf{A}, \quad (18)$$

where $\mathbf{\Xi}^{ST} = \mathbf{diag}(\xi_1^{ST}, \dots, \xi_p^{ST})$ with

$$\xi_j^{ST} = (n_1 - p - 1)l_j / \left(n_2 + p + 1 + 2 \sum_{k \neq j} \frac{l_k}{l_j - l_k} \right), \quad j = 1, 2, \dots, p.$$

The derivation of this estimator is given by Loh (1991b).

By applying methods of Berger and of Stein, the Stein-Berger estimator is given by

$$\hat{\zeta}^{SB} = \mathbf{A}^{-1} \mathbf{\Xi}^{SB} \mathbf{A},$$

where $\mathbf{\Xi}^{SB} = \mathbf{diag}(\xi_1^{SB}, \dots, \xi_p^{SB})$ with

$$\xi_j^{SB} = \bar{\xi}_j^{ST} + \frac{c}{b+u} \quad (j = 1, 2, \dots, p), \quad u = \sum_{j=1}^p \left(\frac{n_2 + p + 3 - 2j}{(n_1 - p - 1)l_j} \right)^2,$$

and the $\bar{\xi}_j^{ST}$'s are constructed by Stein's isotonic regression on the ξ_j^{ST} 's. However, we cannot analytically compare $\hat{\zeta}^{BE}$ with $\hat{\zeta}^{SB}$ and hence, in the next subsection, we examine these risk performances by using a numerical study.

3.3 Numerical studies

We have carried out Monte Carlo simulations (10,000 runs) to observe the risk performances of several estimators in the previous subsection.

1. The matrix-variate normal distribution: The joint density function of $(\boldsymbol{\epsilon}_1, \boldsymbol{\epsilon}_2)$ is given by

$$\gamma_1 |\boldsymbol{\Sigma}_1|^{-N_1/2} |\boldsymbol{\Sigma}_2|^{-N_2/2} \exp[-(1/2) \text{tr}\{\boldsymbol{\Sigma}_1^{-1} \boldsymbol{\epsilon}_1' \boldsymbol{\epsilon}_1 + \boldsymbol{\Sigma}_2^{-1} \boldsymbol{\epsilon}_2' \boldsymbol{\epsilon}_2\}],$$

where $\gamma_1 = (2\pi)^{-(N_1+N_2)p/2}$.

2. The t -distribution: The joint density function of $(\boldsymbol{\epsilon}_1, \boldsymbol{\epsilon}_2)$ is given by

$$\gamma_2 |\boldsymbol{\Sigma}_1|^{-N_1/2} |\boldsymbol{\Sigma}_2|^{-N_2/2} [1 + (1/v) \text{tr}\{\boldsymbol{\Sigma}_1^{-1} \boldsymbol{\epsilon}'_1 \boldsymbol{\epsilon}_1 + \boldsymbol{\Sigma}_2^{-1} \boldsymbol{\epsilon}'_2 \boldsymbol{\epsilon}_2\}]^{-\{v+(N_1+N_2)p\}/2},$$

where $\gamma_2 = \Gamma[\{v + (N_1 + N_2)p\}/2] / \{(\pi v)^{(N_1+N_2)p/2} \Gamma[v/2]\}$, $v > 0$.

3. The Kotz-type distribution: The joint density function of $(\boldsymbol{\epsilon}_1, \boldsymbol{\epsilon}_2)$ is given by

$$\begin{aligned} & \gamma_3 |\boldsymbol{\Sigma}_1|^{-N_1/2} |\boldsymbol{\Sigma}_2|^{-N_2/2} [\text{tr}\{\boldsymbol{\Sigma}_1^{-1} \boldsymbol{\epsilon}'_1 \boldsymbol{\epsilon}_1 + \boldsymbol{\Sigma}_2^{-1} \boldsymbol{\epsilon}'_2 \boldsymbol{\epsilon}_2\}]^{u-1} \\ & \times \exp[-r \text{tr}\{\boldsymbol{\Sigma}_1^{-1} \boldsymbol{\epsilon}'_1 \boldsymbol{\epsilon}_1 + \boldsymbol{\Sigma}_2^{-1} \boldsymbol{\epsilon}'_2 \boldsymbol{\epsilon}_2\}^s], \end{aligned}$$

where $r > 0$, $s > 0$, $2u + (N_1 + N_2)p > 2$, and

$$\gamma_3 = \frac{s \Gamma[(N_1 + N_2)p/2] r^{\{u+(N_1+N_2)p/2-1\}/s}}{\pi^{(N_1+N_2)p/2} \Gamma[\{u + (N_1 + N_2)p/2 - 1\}/s]}.$$

The estimated risks with the error above are given in Tables 9–11. For Monte Carlo simulations, we took $N_1 = N_2 = 15$, $m = 1$, and $p = 3$ and we also put $v = 5$ for the t -distribution and $(u, r, s) = (5, 0.1, 2)$ for the Kotz-type distribution. We supposed that $\boldsymbol{\beta}_1 = \boldsymbol{\beta}_2 = (0, 0, 0)'$ and that the parameter $\boldsymbol{\Sigma}_2 \boldsymbol{\Sigma}_1^{-1}$ is the diagonal matrix with typical elements.

In tables, ‘ BU ’, ‘ LO ’, ‘ BE ’, and ‘ SB ’ stand for $\hat{\zeta}^{BU}$, $\hat{\zeta}^{LO}$, $\hat{\zeta}^{BE}$ respectively and $\hat{\zeta}^{SB}$, and ‘AI’ stands for the average of improving over risk of $\hat{\zeta}^{BU}$. For $\hat{\zeta}^{BE}$ and $\hat{\zeta}^{SB}$, we set $b = 100$ and

$$c = \frac{2(p^2 + p - 4)(n_2 - p + 3)\sqrt{b}}{\sqrt{p}(n_1 - p - 1)(n_2 - p + 7)}.$$

We also studied simulations when the rows of the error are independently and identically distributed as in the densities (12) and (13). We assume that the distributions of rows of the errors are (12) and are (13) in the previous section. This results are given in Tables 12 and 13. Furthermore we got the similar results as those in Tables 7 and 8.

We summarize our numerical results in Tables 9–16 as follows:

1. When the diagonal elements of $\boldsymbol{\Sigma}_2 \boldsymbol{\Sigma}_1^{-1}$ are close together, AIs of BE and SB are substantially large under normal error as well as under elliptically contoured error;
2. Among the alternative estimators, AI of SB is the largest when the diagonal elements of $\boldsymbol{\Sigma}_2 \boldsymbol{\Sigma}_1^{-1}$ are equal or when the smallest diagonal element of $\boldsymbol{\Sigma}_2 \boldsymbol{\Sigma}_1^{-1}$ is far from the others and the others are large.
3. When the largest diagonal element of $\boldsymbol{\Sigma}_2 \boldsymbol{\Sigma}_1^{-1}$ is far from the second, AIs are relatively small.
4. The results in Tables 12 and 13 suggest that the improvement by $\hat{\zeta}^{LO}$ and $\hat{\zeta}^{BE}$ under density (3) remains robust even if the rows of errors are i.i.d.
5. Tables 14–16 indicate that the AI’s decrease with increasing dimension p when the sample size is fixed.

4 Proofs of Theorems

In this section we show theorems and corollaries in Sections 2 and 3. To give the proofs, we state a canonical form of our problems and list useful lemmas. Listed lemmas consist of two ingredients. First, we adapted the integration-by-parts formula from Kubokawa and Srivastava (1999) for our problems. We introduce two types of integration-by-parts formulas which concern the joint density function (2) and the joint density function (3), respectively. Second, we quote lemmas on eigenstructure from Loh (1988).

4.1 Preliminaries

To derive a canonical form, write $(\mathbf{\Gamma}_i \mathbf{C}_i)' = ((\mathbf{C}'_i \mathbf{C}_i)^{1/2}, \mathbf{0})$, $i = 1, 2$, where $\mathbf{\Gamma}_i$ is an $N_i \times N_i$ orthogonal matrix. Also put $\boldsymbol{\theta}_i = (\mathbf{C}'_i \mathbf{C}_i)^{1/2} \boldsymbol{\beta}_i$ and $n_i = N_i - m$. Furthermore, write $(\mathbf{X}'_i, \mathbf{Z}'_i)' = \mathbf{\Gamma}_i \mathbf{Y}_i$, where \mathbf{X}_i and \mathbf{Z}_i are $m \times p$ and $n_i \times p$ matrices respectively. Then the densities (2) are rewritten as

$$|\boldsymbol{\Sigma}_i|^{-N_i/2} g_i[\text{tr}\{\boldsymbol{\Sigma}_i^{-1}(\mathbf{X}_i - \boldsymbol{\theta}_i)'(\mathbf{X}_i - \boldsymbol{\theta}_i)\} + \text{tr}(\boldsymbol{\Sigma}_i^{-1} \mathbf{Z}'_i \mathbf{Z}_i)] \quad (19)$$

for $i = 1, 2$.

Next we introduce notation for integration-by-parts formula with respect to the joint densities (2). Let

$$G_i(x) = \frac{1}{2} \int_x^{+\infty} g_i(t) dt$$

and let $\boldsymbol{\theta} = (\boldsymbol{\theta}_1, \boldsymbol{\theta}_2)$ and $\boldsymbol{\Sigma} = (\boldsymbol{\Sigma}_1, \boldsymbol{\Sigma}_2)$. For a function $U \equiv U(\mathbf{X}_1, \mathbf{X}_2, \mathbf{Z}_1, \mathbf{Z}_2)$, define

$$\mathbf{E}_{\boldsymbol{\theta}, \boldsymbol{\Sigma}}^{g_1 g_2} [U] = \int U \times \left\{ \prod_{i=1}^2 |\boldsymbol{\Sigma}_i|^{-N_i/2} g_i(w_i) \right\} d\mathbf{X}_1 d\mathbf{X}_2 d\mathbf{Z}_1 d\mathbf{Z}_2, \quad (20)$$

$$\mathbf{E}_{\boldsymbol{\theta}, \boldsymbol{\Sigma}}^{G_1 g_2} [U] = \int U \times |\boldsymbol{\Sigma}_1|^{-N_1/2} |\boldsymbol{\Sigma}_2|^{-N_2/2} G_1(w_1) g_2(w_2) d\mathbf{X}_1 d\mathbf{X}_2 d\mathbf{Z}_1 d\mathbf{Z}_2, \quad (21)$$

$$\mathbf{E}_{\boldsymbol{\theta}, \boldsymbol{\Sigma}}^{g_1 G_2} [U] = \int U \times |\boldsymbol{\Sigma}_1|^{-N_1/2} |\boldsymbol{\Sigma}_2|^{-N_2/2} g_1(w_1) G_2(w_2) d\mathbf{X}_1 d\mathbf{X}_2 d\mathbf{Z}_1 d\mathbf{Z}_2, \quad (22)$$

where $w_i = \text{tr}\{\boldsymbol{\Sigma}_i^{-1}(\mathbf{X}_i - \boldsymbol{\theta}_i)'(\mathbf{X}_i - \boldsymbol{\theta}_i)\} + \text{tr}(\boldsymbol{\Sigma}_i^{-1} \mathbf{Z}'_i \mathbf{Z}_i)$ ($i = 1, 2$).

Put $\mathbf{S}_i = \mathbf{Z}'_i \mathbf{Z}_i$ and recall that $\mathbf{S}_i = \mathbf{Y}'_i \{\mathbf{I}_{N_i} - \mathbf{C}_i (\mathbf{C}'_i \mathbf{C}_i)^{-1} \mathbf{C}'_i\} \mathbf{Y}_i$ for $i = 1, 2$ and let $\mathbf{H} \equiv \mathbf{H}(\mathbf{S}_1, \mathbf{S}_2) = (h_{ij})$ be a $p \times p$ matrix such that the (j, k) -element h_{jk} is a function of $\mathbf{S}_1 = (s_{1.jk})$ and $\mathbf{S}_2 = (s_{2.jk})$. For $i = 1, 2$, let

$$\{\mathbf{D}_{S_i} \mathbf{H}\}_{jk} = \sum_{a=1}^p d_{i.ja} h_{ak}, \quad (23)$$

where

$$d_{i.ja} = \frac{1}{2} (1 + \delta_{ja}) \frac{\partial}{\partial s_{i.ja}}$$

with $\delta_{ja} = 1$ for $j = a$ and $\delta_{ja} = 0$ for $j \neq a$. Also put $\mathbf{Z}_i = (\mathbf{z}'_{i1}, \dots, \mathbf{z}'_{in_i})'$ and $\mathbf{z}_{ij} = (z_{i.j1}, \dots, z_{i.jp})$ for $i = 1, 2$, and $j = 1, 2, \dots, n_i$. Hence we have $\mathbf{S}_i = \mathbf{Z}'_i \mathbf{Z}_i = \sum_{j=1}^{n_i} \mathbf{z}'_{ij} \mathbf{z}_{ij}$ for $i = 1, 2$.

Now we adapt the extended Stein-Haff identity due to Kubokawa and Srivastava (1999) for our problem. The difference between derivations of our identity and of that by Kubokawa and Srivastava (1999) is what expectation for the variables of integration are multiplied. Hence, we state the following formula without the proof.

Lemma 4.1 *Let*

$$\mathbf{H}_i \equiv \mathbf{H}_i \left(\sum_{j_1=1}^{n_1} \mathbf{z}'_{1j_1} \mathbf{z}_{1j_1}, \sum_{j_2=1}^{n_2} \mathbf{z}'_{2j_2} \mathbf{z}_{2j_2} \right), \quad i = 1, 2,$$

be a $p \times p$ matrix whose element is differentiable with respect to $z_{i.jk}$ ($j_i = 1, 2, \dots, n_i$, $k = 1, 2, \dots, p$). Furthermore, assume that

(a) $\mathbf{E}_{\boldsymbol{\theta}, \boldsymbol{\Sigma}}^{g_1 g_2} \left[\left| \text{tr} \mathbf{H}_i \boldsymbol{\Sigma}_i^{-1} \right| \right]$ ($i = 1, 2$) is finite;

(b) $\lim_{z_{i.jk} \rightarrow \pm\infty} |z_{i.jk}| \mathbf{H}_i \left(\sum_{j_1=1}^{n_1} \mathbf{z}'_{1j_1} \mathbf{z}_{1j_1}, \sum_{j_2=1}^{n_2} \mathbf{z}'_{2j_2} \mathbf{z}_{2j_2} \right) \left(\sum_{j_i=1}^{n_i} \mathbf{z}'_{1j_i} \mathbf{z}_{1j_i} \right)^{-1} G_i(z_{i.jk} + a) = 0$
for any real a .

Then we have

$$\begin{aligned} \mathbf{E}_{\boldsymbol{\theta}, \boldsymbol{\Sigma}}^{g_1 g_2} [\text{tr}(\mathbf{H}_1 \boldsymbol{\Sigma}_1^{-1})] + \mathbf{E}_{\boldsymbol{\theta}, \boldsymbol{\Sigma}}^{g_1 g_2} [\text{tr}(\mathbf{H}_2 \boldsymbol{\Sigma}_2^{-1})] &= \mathbf{E}_{\boldsymbol{\theta}, \boldsymbol{\Sigma}}^{G_1 g_2} [(n_1 - p - 1) \text{tr}(\mathbf{H}_1 \mathbf{S}_1^{-1}) + 2 \text{tr}(\mathbf{D}_{S_1} \mathbf{H}_1)] \\ &\quad + \mathbf{E}_{\boldsymbol{\theta}, \boldsymbol{\Sigma}}^{g_1 G_2} [(n_2 - p - 1) \text{tr}(\mathbf{H}_2 \mathbf{S}_2^{-1}) + 2 \text{tr}(\mathbf{D}_{S_2} \mathbf{H}_2)] \end{aligned}$$

for $\boldsymbol{\theta} = (\boldsymbol{\theta}_1, \boldsymbol{\theta}_2)$ and $\boldsymbol{\Sigma} = (\boldsymbol{\Sigma}_1, \boldsymbol{\Sigma}_2)$.

To derive the integration-by-parts formula with respect to the density (3), we make an orthogonal transformation \mathbf{Y}_1 and \mathbf{Y}_2 to rewrite the density (3) as

$$|\boldsymbol{\Sigma}_1|^{-N_1/2} |\boldsymbol{\Sigma}_2|^{-N_2/2} g \left\{ \sum_{i=1}^2 [\text{tr}\{\boldsymbol{\Sigma}_i^{-1}(\mathbf{X}_i - \boldsymbol{\theta}_i)'(\mathbf{X}_i - \boldsymbol{\theta}_i)\} + \text{tr}(\boldsymbol{\Sigma}_i^{-1} \mathbf{Z}'_i \mathbf{Z}_i)] \right\}, \quad (24)$$

where \mathbf{X}_i , \mathbf{Z}_i and $\boldsymbol{\theta}_i$ are defined in the same way to obtain (19). For a real-valued function U , denote

$$\begin{aligned} \mathbf{E}_{\boldsymbol{\theta}, \boldsymbol{\Sigma}}^g[U] &= \int U \times \left(\prod_{i=1}^2 |\boldsymbol{\Sigma}_i|^{-N_i/2} \right) g(w) d\mathbf{X}_1 d\mathbf{X}_2 d\mathbf{Z}_1 d\mathbf{Z}_2, \\ \mathbf{E}_{\boldsymbol{\theta}, \boldsymbol{\Sigma}}^G[U] &= \int U \times \left(\prod_{i=1}^2 |\boldsymbol{\Sigma}_i|^{-N_i/2} \right) G(w) d\mathbf{X}_1 d\mathbf{X}_2 d\mathbf{Z}_1 d\mathbf{Z}_2, \end{aligned}$$

where $w = \sum_{i=1}^2 [\text{tr}\{\boldsymbol{\Sigma}_i^{-1}(\mathbf{X}_i - \boldsymbol{\theta}_i)'(\mathbf{X}_i - \boldsymbol{\theta}_i)\} + \text{tr}(\boldsymbol{\Sigma}_i^{-1} \mathbf{Z}'_i \mathbf{Z}_i)]$, $G(x) = (1/2) \int_x^{+\infty} g(t) dt$, $\boldsymbol{\theta} = (\boldsymbol{\theta}_1, \boldsymbol{\theta}_2)$, and $\boldsymbol{\Sigma} = (\boldsymbol{\Sigma}_1, \boldsymbol{\Sigma}_2)$. From preliminaries as above, we get the integration-by-parts formula for the density (3):

Lemma 4.2 *Assume that*

$$\mathbf{H} \equiv \mathbf{H}(\mathbf{S}_1, \mathbf{S}_2) \quad (25)$$

is differentiable with respect to $z_{i,jk}$ ($i = 1, 2, j = 1, 2, \dots, n_i, k = 1, 2, \dots, p$) and that

- (a) $\mathbf{E}_{\boldsymbol{\theta}, \boldsymbol{\Sigma}}^g[|\text{tr}(\mathbf{H}\boldsymbol{\Sigma}_i^{-1})|]$ *is finite for $i = 1, 2$;*
- (b) $\lim_{z_{i,jk} \rightarrow \pm\infty} |z_{i,jk}| \mathbf{H}(\mathbf{S}_1, \mathbf{S}_2) (\mathbf{S}_i)^{-1} G(z_{i,jk} + a) = 0$ *where $\mathbf{S}_i = \sum_{j=1}^{n_i} \mathbf{z}'_{ij} \mathbf{z}_{ij}$, a is any real number and $i = 1, 2$.*

Then, for $i = 1, 2$,

$$\mathbf{E}_{\boldsymbol{\theta}, \boldsymbol{\Sigma}}^g[\text{tr}(\mathbf{H}\boldsymbol{\Sigma}_i^{-1})] = \mathbf{E}_{\boldsymbol{\theta}, \boldsymbol{\Sigma}}^G[(n_i - p - 1)\text{tr}(\mathbf{H}\mathbf{S}_i^{-1}) + 2\text{tr}(\mathbf{D}_{S_i} \mathbf{H})]. \quad (26)$$

Furthermore, we need the following lemmas to show main theorems and their corollaries.

Lemma 4.3 [Loh, 1991a] *Let \mathbf{S}_1 and \mathbf{S}_2 be $p \times p$ symmetric and positive-definite matrices. Also let \mathbf{B} be nonsingular matrix such that $\mathbf{B}(\mathbf{S}_1 + \mathbf{S}_2)\mathbf{B}' = \mathbf{I}_p$, $\mathbf{B}\mathbf{S}_2\mathbf{B}' = \mathbf{F}$ where $\mathbf{F} = \text{diag}(f_1, f_2, \dots, f_p)$ with $f_1 \geq f_2 \geq \dots \geq f_p$. Furthermore, let $\boldsymbol{\Psi} = \text{diag}(\psi_1, \psi_2, \dots, \psi_p)$ and $\boldsymbol{\Phi} = \text{diag}(\phi_1, \phi_2, \dots, \phi_p)$, where the ψ_j and the ϕ_j ($j = 1, 2, \dots, p$) are differentiable functions of \mathbf{F} . Then we have*

$$\begin{aligned} \text{tr}(\mathbf{D}_{S_1} \mathbf{B}^{-1} \boldsymbol{\Psi} \mathbf{B}'^{-1}) &= \sum_{j=1}^p \left[\psi_j + f_j \frac{\partial \psi_j}{\partial (1 - f_j)} + \psi_j \sum_{k \neq j} \frac{f_k}{f_k - f_j} \right], \\ \text{tr}(\mathbf{D}_{S_2} \mathbf{B}^{-1} \boldsymbol{\Phi} \mathbf{B}'^{-1}) &= \sum_{j=1}^p \left[\phi_j + (1 - f_j) \frac{\partial \phi_j}{\partial f_j} + \phi_j \sum_{k \neq j} \frac{1 - f_k}{f_j - f_k} \right], \end{aligned}$$

where \mathbf{D}_{S_1} and \mathbf{D}_{S_2} are defined as (23).

Lemma 4.4 [Loh, 1991b] *Let \mathbf{S}_1 and \mathbf{S}_2 be $p \times p$ symmetric and positive-definite matrices. Also let \mathbf{A} be nonsingular matrix such that $\mathbf{A}\mathbf{S}_1\mathbf{A}' = \mathbf{I}_p$, $\mathbf{A}\mathbf{S}_2\mathbf{A}' = \mathbf{L}$ where $\mathbf{L} = \text{diag}(l_1, l_2, \dots, l_p)$ with $l_1 \geq l_2 \geq \dots \geq l_p$. Let $\boldsymbol{\Xi} = \text{diag}(\xi_1, \xi_2, \dots, \xi_p)$, where the ξ_j ($j = 1, 2, \dots, p$) are differentiable functions of \mathbf{L} . Then we have*

$$\begin{aligned} \text{tr}(\mathbf{D}_{S_1} \mathbf{A}^{-1} \boldsymbol{\Xi} \mathbf{A}'^{-1}) &= \sum_{j=1}^p \left[\xi_j - l_j \frac{\partial \xi_j}{\partial l_j} + \xi_j \sum_{k \neq j} \frac{l_k}{l_k - l_j} \right], \\ \text{tr}(\mathbf{D}_{S_2} \mathbf{A}^{-1} \boldsymbol{\Xi}^2 \mathbf{A}'^{-1}) &= \sum_{j=1}^p \left[2\xi_j \frac{\partial \xi_j}{\partial l_j} + \xi_j^2 \sum_{k \neq j} \frac{1}{l_j - l_k} \right], \end{aligned}$$

where \mathbf{D}_{S_1} and \mathbf{D}_{S_2} are defined as (23).

4.2 Proofs of Theorem 2.1 in Section 2

To prove Theorem 2.1, we give sufficient conditions for the estimators of the form (9) to improve upon the James-Stein estimator (7).

Lemma 4.5 *Suppose that we wish to estimate $(\boldsymbol{\Sigma}_1, \boldsymbol{\Sigma}_2)$ simultaneously under the loss function (4). An invariant estimator $(\widehat{\boldsymbol{\Sigma}}_1^{EQ}, \widehat{\boldsymbol{\Sigma}}_2^{EQ})$ is better than the James-Stein estimator $(\widehat{\boldsymbol{\Sigma}}_1^{JS}, \widehat{\boldsymbol{\Sigma}}_2^{JS})$ for arbitrary g_1 and g_2 in (2) if the following three inequalities hold almost surely:*

$$(i) \quad p - (n_1 - p - 1) \sum_{j=1}^p \frac{\psi_j}{1 - f_j} - 2 \sum_{j=1}^p \left[\psi_j + f_j \frac{\partial \psi_j}{\partial (1 - f_j)} + \psi_j \sum_{k \neq j} \frac{f_k}{f_k - f_j} \right] \geq 0,$$

$$(ii) \quad p - (n_2 - p - 1) \sum_{j=1}^p \frac{\phi_j}{f_j} - 2 \sum_{j=1}^p \left[\phi_j + (1 - f_j) \frac{\partial \phi_j}{\partial f_j} + \phi_j \sum_{k \neq j} \frac{1 - f_k}{f_j - f_k} \right] \geq 0,$$

$$(iii) \quad \sum_{j=1}^p \left[-(\log d_{1j} + \log d_{2j}) + \log \frac{\psi_j}{1 - f_j} + \log \frac{\phi_j}{f_j} \right] \geq 0,$$

where $d_{1j} = 1/(n_1 + p + 1 - 2j)$ and $d_{2j} = 1/(n_2 + p + 1 - 2j)$, $j = 1, 2, \dots, p$.

Proof. From Lemmas 4.1 and 4.3, the risk of the estimator $(\widehat{\boldsymbol{\Sigma}}_1^{EQ}, \widehat{\boldsymbol{\Sigma}}_2^{EQ})$ can be expressed as

$$\begin{aligned} R_1(\widehat{\boldsymbol{\Sigma}}_1^{EQ}, \widehat{\boldsymbol{\Sigma}}_2^{EQ}) &= \mathbf{E}_{\theta, \Sigma}^{G_1 g_2} \left[(n_1 - p - 1) \sum_{j=1}^p \frac{\psi_j}{1 - f_j} + 2 \sum_{j=1}^p \left\{ \psi_j + f_j \frac{\partial \psi_j}{\partial (1 - f_j)} + \psi_j \sum_{k \neq j} \frac{f_k}{f_k - f_j} \right\} \right] \\ &\quad + \mathbf{E}_{\theta, \Sigma}^{g_1 G_2} \left[(n_2 - p - 1) \sum_{j=1}^p \frac{\phi_j}{f_j} + 2 \sum_{j=1}^p \left\{ \phi_j + (1 - f_j) \frac{\partial \phi_j}{\partial f_j} + \phi_j \sum_{k \neq j} \frac{1 - f_k}{f_j - f_k} \right\} \right] \\ &\quad + \mathbf{E}_{\theta, \Sigma}^{g_1 g_2} \left[- \sum_{j=1}^p \left\{ \log \frac{\psi_j}{1 - f_j} + \log \frac{\phi_j}{f_j} \right\} + \sum_{i=1}^2 \{ -\log |\mathbf{S}_i| + \log |\boldsymbol{\Sigma}_i| - p \} \right]. \end{aligned} \quad (27)$$

Similarly the risk of the James-Stein estimator $(\widehat{\boldsymbol{\Sigma}}_1^{JS}, \widehat{\boldsymbol{\Sigma}}_2^{JS})$ is given by

$$\begin{aligned} R_1(\widehat{\boldsymbol{\Sigma}}_1^{JS}, \widehat{\boldsymbol{\Sigma}}_2^{JS}) &= \mathbf{E}_{\theta, \Sigma}^{G_1 g_2} [p] + \mathbf{E}_{\theta, \Sigma}^{g_1 G_2} [p] \\ &\quad + \mathbf{E}_{\theta, \Sigma}^{g_1 g_2} \left[- \sum_{j=1}^p \left\{ \log d_{1j} + \log d_{2j} \right\} + \sum_{i=1}^2 \{ -\log |\mathbf{S}_i| + \log |\boldsymbol{\Sigma}_i| - p \} \right], \end{aligned} \quad (28)$$

where $d_{ij} = 1/(n_i + p + 1 - 2j)$. Hence, comparing the integrands with respect to each expectation of (20), (21), and (22) in the right-hand side of the equations (27) and (28), we complete the proof. \square

Now we see that the estimator (10) satisfies three conditions given in Lemma 4.5, which give the proof of Theorem 2.1.

Proof of THEOREM 2.1: From Lemma 4.5, it suffices to show that

$$p - (n_1 - p + 1) \sum_{j=1}^p d_{1j}^* - 2 \sum_{j=1}^p \left[(1 - f_j) d_{1j}^* \sum_{k \neq j} \frac{f_k}{f_k - f_j} \right] \geq 0, \quad (29)$$

$$p - (n_2 - p + 1) \sum_{j=1}^p d_{2j}^* - 2 \sum_{j=1}^p \left[f_j d_{2j}^* \sum_{k \neq j} \frac{1 - f_k}{f_j - f_k} \right] \geq 0, \quad (30)$$

where $d_{1j}^* = 1/(n_1 - p - 1 + 2j)$ and $d_{2j}^* = 1/(n_2 + p + 1 - 2j)$. We here note that the last term of the left-hand side in (29) is evaluated as

$$\begin{aligned} -\sum_{j=1}^p \left\{ (1-f_j)d_{1j}^* \sum_{k \neq j} \frac{f_k}{f_k - f_j} \right\} &= \sum_{j < k} \frac{f_k(1-f_j)d_{1j}^* - f_j(1-f_k)d_{1k}^*}{f_j - f_k} \\ &= \sum_{j < k} \frac{f_k(1-f_j)(d_{1j}^* - d_{1k}^*)}{f_j - f_k} - \sum_{j < k} d_{1k}^* \\ &\geq -\sum_{j < k} d_{1k}^*, \end{aligned}$$

where the last inequality in the above display is derived by the fact that $f_k < f_j < 1$ and $d_{1k}^* < d_{1j}^*$ for $j < k$. Since $\sum_{j < k} d_{1k}^* = \sum_{j=1}^p (j-1)d_{1j}^*$, we get

$$\begin{aligned} p - (n_1 - p + 1) \sum_{j=1}^p d_{1j}^* - 2 \sum_{j=1}^p \left\{ (1-f_j)d_{1j}^* \sum_{k \neq j} \frac{f_k}{f_j - f_k} \right\} \\ \geq p - \sum_{j=1}^p (n_1 - p - 1 + 2j)d_{1j}^* = 0. \end{aligned}$$

The proof of the inequality (30) can proceed similarly. Note that the last term of the left-hand side in (30) is evaluated as

$$-\sum_{j=1}^p \left[f_j d_{2j}^* \sum_{k \neq j} \frac{1-f_k}{f_j - f_k} \right] = \sum_{j > k} \frac{f_k(1-f_j)(d_{2k}^* - d_{2j}^*)}{f_j - f_k} - \sum_{j > k} d_{2k}^* \geq -\sum_{j=1}^p (p-j)d_{2j}^*.$$

Putting the above inequality into the left-hand side of (30), we get the desired result. \square

4.3 Proofs of Theorems 3.1, 3.2, and 3.3 in Section 3

First, we give the following lemma:

Lemma 4.6 *For estimation of $\zeta = \Sigma_2 \Sigma_1^{-1}$ in model (3), we consider an estimator of the form*

$$\hat{\zeta} = \mathbf{A}^{-1} \Xi \mathbf{A}, \quad (31)$$

where \mathbf{A} is a nonsingular matrix such that $\mathbf{A} \mathbf{S}_1 \mathbf{A}' = \mathbf{I}_p$ and $\mathbf{A} \mathbf{S}_2 \mathbf{A}' = \mathbf{L}$ with $\mathbf{L} = \text{diag}(l_1, l_2, \dots, l_p)$ with $l_1 \geq l_2 \geq \dots \geq l_p > 0$, and $\Xi = \text{diag}(\xi_1, \xi_2, \dots, \xi_p)$. Then, under the loss function (5), the risk of $\hat{\zeta}$ is given by

$$\begin{aligned} R_2(\hat{\zeta}; \zeta) &= \mathbf{E}_{\theta, \Sigma}^G \left[n_1 + \sum_j \left\{ \frac{n_2 - p - 1}{l_j} \xi_j^2 + 2\xi_j^2 \sum_{k \neq j} \frac{1}{l_j - l_k} + 4\xi_j \frac{\partial \xi_j}{\partial l_j} \right. \right. \\ &\quad \left. \left. - 2(n_1 - p + 1)\xi_j + 4\xi_j \sum_{k \neq j} \frac{l_k}{l_j - l_k} + 4l_j \frac{\partial \xi_j}{\partial l_j} \right\} / \text{tr} \zeta \right]. \end{aligned} \quad (32)$$

Proof. We can write the risk function as

$$R_2(\hat{\zeta}; \zeta) = \mathbf{E}_{\theta, \Sigma}^g [\text{tr} \{ \Sigma_1^{-1} \Sigma_2 \Sigma_1^{-1} \mathbf{S}_1 + \Sigma_2^{-1} \mathbf{A}^{-1} \Phi^2 \mathbf{A}'^{-1} - 2 \Sigma_1^{-1} \mathbf{A}^{-1} \Phi \mathbf{A}'^{-1} \} / \text{tr} \zeta].$$

Noting that $\mathbf{E}_{\theta, \Sigma}^g[\mathbf{S}_1] = \mathbf{E}_{\theta, \Sigma}^G[n_1 \Sigma_1]$ and using Lemmas 4.2 and 4.4, we get (32). \square

Proof of THEOREM 3.1: Substituting αl_j for ξ_j in (32), we get

$$R_2(\hat{\zeta}^{US}; \zeta) = \mathbf{E}_{\theta, \Sigma}^G [n_1 + \{(n_2 + p + 1)\alpha^2 - 2(n_1 - p - 1)\alpha\} \sum_{j=1}^p l_j / \text{tr} \zeta]. \quad (33)$$

Hence we can see that (33) is minimized at $\alpha = (n_1 - p - 1)/(n_2 + p + 1)$. \square

Proof of THEOREMS 3.2: The proof proceeds in similar way as in that of Theorem 3.5 in Loh (1988). We reproduce it for reader's convenience. Write $\tilde{d}_j = (n_1 - p - 1)/(n_2 + p + 3 - 2j)$, $j = 1, 2, \dots, p$. From (33), we have

$$(\text{tr}\zeta)\{R_2(\hat{\zeta}^{BU}; \zeta) - \mathbf{E}_{\theta, \Sigma}^G[n_1]\} = \mathbf{E}_{\theta, \Sigma}^G \left[-\frac{(n_1 - p - 1)^2}{n_2 + p + 1} \sum_{i=1}^p l_i \right].$$

From Lemma 4.6, we have

$$\begin{aligned} (\text{tr}\zeta)\{R_2(\hat{\zeta}^{LO}; \zeta) - \mathbf{E}_{\theta, \Sigma}^G[n_1]\} &= \mathbf{E}_{\theta, \Sigma}^G \left[\sum_{j=1}^p \left\{ (n_2 - p + 3)\tilde{d}_j^2 l_j + 2 \sum_{k < j} \frac{\tilde{d}_k^2 (l_j^2 - l_k^2)}{l_j - l_k} \right. \right. \\ &\quad \left. \left. + 2 \sum_{k < j} \frac{l_j^2 (\tilde{d}_j^2 - \tilde{d}_k^2)}{l_j - l_k} - 2(n_1 - p - 1)\tilde{d}_j l_j + 4 \sum_{k < j} \frac{l_j l_k (\tilde{d}_j - \tilde{d}_k)}{l_j - l_k} \right\} \right]. \end{aligned} \quad (34)$$

Noting that

$$2 \sum_{k < j} \frac{l_j^2 (\tilde{d}_j^2 - \tilde{d}_k^2)}{l_j - l_k} + 4 \sum_{k < j} \frac{l_j l_k (\tilde{d}_j - \tilde{d}_k)}{l_j - l_k} \leq 0$$

and that

$$\sum_{j=1}^p \sum_{k < j} \frac{\tilde{d}_k^2 (l_j^2 - l_k^2)}{l_j - l_k} = \sum_{j=1}^p \left\{ (p - j)\tilde{d}_j^2 + \sum_{k < j} \tilde{d}_k^2 \right\} l_j,$$

we can see that the right-hand side of (34) is less than

$$\begin{aligned} &\mathbf{E}_{\theta, \Sigma}^G \left[\sum_{j=1}^p \left\{ (n_2 + p + 3 - 2j)\tilde{d}_j^2 - 2(n_1 - p - 1)\tilde{d}_j + 2 \sum_{k < j} \tilde{d}_k^2 \right\} l_j \right] \\ &= \mathbf{E}_{\theta, \Sigma}^G \left[\sum_{j=1}^p \left\{ -\frac{(n_1 - p - 1)^2}{n_2 + p + 1} - 2(n_1 - p - 1)^2 \left(\frac{j - 1}{(n_2 + p + 1)(n_2 + p + 3 - 2j)} \right. \right. \right. \\ &\quad \left. \left. \left. - \sum_{k < j} \frac{1}{(n_2 + p + 3 - 2k)^2} \right) \right\} l_j \right]. \end{aligned}$$

Furthermore, from mathematical induction on j , we can see that

$$\frac{j - 1}{(n_2 + p + 1)(n_2 + p + 3 - 2j)} \geq \sum_{k < j} \frac{1}{(n_2 + p + 3 - 2k)^2}.$$

From this inequality, we finally get

$$(\text{tr}\zeta)\{R_2(\hat{\zeta}^{LO}; \zeta) - \mathbf{E}_{\theta, \Sigma}^G[n_1]\} \leq (\text{tr}\zeta)\{R_2(\hat{\zeta}^{BU}; \zeta) - \mathbf{E}_{\theta, \Sigma}^G[n_1]\},$$

which completes the proof. \square

Proof of THEOREM 3.3: The proof proceed in the same way as in that of Theorem 3.5 in Loh (1988). However, we reproduce it for reader's convenience when c is a positive constant which satisfies Assumption (III). Put $\alpha_j = c/\{\tilde{d}_j l_j (b + u)\}$, $j = 1, 2, \dots, p$. Hence $\xi_j^{BE} = \tilde{d}_j l_j (1 + \alpha_j)$. Also note that

$$\frac{\partial \xi_j^{BE}}{\partial l_j} = \tilde{d}_j (1 + \alpha_j) + \tilde{d}_j l_j \frac{\partial \alpha_j}{\partial l_j} = \tilde{d}_j (1 + \alpha_j) + \frac{c}{l_j (b + u)} \left[\frac{2}{\tilde{d}_j^2 l_j^2 (b + u)} - 1 \right].$$

From tedious calculation, we have

$$\begin{aligned}
& (\text{tr}\zeta)\{R_2(\hat{\zeta}^{BE}; \zeta) - R_2(\hat{\zeta}^{LO}; \zeta)\} \\
&= \mathbf{E}_{\theta, \Sigma}^G \sum_{j=1}^p \left[(n_2 - p - 1) \tilde{d}_j^2 l_j (2\alpha_j + \alpha_j^2) + 4 \sum_{k < j} \frac{\tilde{d}_j^2 l_j^2 \alpha_j - \tilde{d}_k^2 l_k^2 \alpha_k}{l_j - l_k} + 4 \tilde{d}_j^2 l_j (2\alpha_j + \alpha_j^2) \right. \\
&\quad + 4 \tilde{d}_j^2 l_j^2 (1 + \alpha_j) \frac{\partial \alpha_j}{\partial l_j} - 2(n_1 - p + 1) \tilde{d}_j l_j \alpha_j + 4 \sum_{k < j} \frac{\tilde{d}_j l_j \alpha_j l_k - \tilde{d}_k l_k \alpha_k l_j}{l_j - l_k} \\
&\quad \left. + 4 l_j \tilde{d}_j \alpha_j + 4 \tilde{d}_j l_j^2 \frac{\partial \alpha_j}{\partial l_j} \right] \\
&= \mathbf{E}_{\theta, \Sigma}^G \sum_{j=1}^p \left[(n_2 - p + 3) \tilde{d}_j^2 l_j \alpha_j^2 + 4(\tilde{d}_j^2 + \tilde{d}_j) l_j^2 \frac{\partial \alpha_j}{\partial l_j} + 4 \tilde{d}_j^2 l_j^2 \alpha_j \frac{\partial \alpha_j}{\partial l_j} - 4 l_j \alpha_j \tilde{d}_j^2 (p - j) \right. \\
&\quad \left. + 4 \sum_{k < j} \frac{\tilde{d}_j^2 l_j^2 \alpha_j - \tilde{d}_k^2 l_k^2 \alpha_k}{l_j - l_k} + 4 \sum_{k < j} \frac{\tilde{d}_j l_j \alpha_j l_k - \tilde{d}_k l_k \alpha_k l_j}{l_j - l_k} \right]. \tag{35}
\end{aligned}$$

Now we observe that

$$\sum_{j=1}^p \left(\sum_{k < j} \left\{ \frac{\tilde{d}_j^2 l_j^2 \alpha_j - \tilde{d}_k^2 l_k^2 \alpha_k}{l_j - l_k} + \frac{\tilde{d}_j l_j \alpha_j l_k - \tilde{d}_k l_k \alpha_k l_j}{l_j - l_k} \right\} - l_j \alpha_j \tilde{d}_j^2 (p - j) \right) \leq \frac{c(p - p^2)}{2(b + u)}. \tag{36}$$

Furthermore, noting that $\tilde{d}_1 < \tilde{d}_2 < \dots < \tilde{d}_p$ and that $\sum_{j=1}^p 1/\{\tilde{d}_j^2 l_j^2 (b + u)\} = u/(b + u) \leq 1$, we have

$$\begin{aligned}
\sum_{j=1}^p (\tilde{d}_j^2 + \tilde{d}_j) l_j^2 \frac{\partial \alpha_j}{\partial l_j} &= \frac{c}{b + u} \sum_{j=1}^p \left[-1 + \frac{2}{\tilde{d}_j^2 l_j^2 (b + u)} - \tilde{d}_j + \frac{2\tilde{d}_j}{\tilde{d}_j^2 l_j^2 (b + u)} \right] \\
&\leq -\frac{c}{b + u} \left[p - 2 - \tilde{d}_p + \sum_{j=1}^{p-1} \tilde{d}_j \right] \leq -\frac{c}{b + u} (p - 2). \tag{37}
\end{aligned}$$

The last inequality follows from $\sum_{j=1}^{p-1} \tilde{d}_j - \tilde{d}_p \geq 0$. Since $l_1 > l_2 > \dots > l_p$, we have

$$4 \sum_{j=1}^p \tilde{d}_j^2 l_j^2 \alpha_j \frac{\partial \alpha_j}{\partial l_j} = \frac{4c^2}{(b + u)^2} \sum_{j=1}^p \left[\frac{2}{\tilde{d}_j^2 l_j^3 (b + u)} - \frac{1}{l_j} \right] \leq \frac{4c^2}{(b + u)^2} \left[\frac{2}{l_p} - \sum_{j=1}^p \frac{1}{l_j} \right],$$

which gives

$$\begin{aligned}
\sum_{j=1}^p \{ (n_2 - p + 3) \tilde{d}_j^2 l_j \alpha_j^2 + 4 \tilde{d}_j^2 l_j^2 \alpha_j (\partial \alpha_j / \partial l_j) \} &\leq \frac{c^2}{(b + u)^2} \left[\frac{8}{l_p} + (n_2 - p - 1) \sum_{j=1}^p \frac{1}{l_j} \right] \\
&\leq (n_2 - p + 7) \frac{c^2 \tilde{d}_p}{(b + u)^2} \sum_{j=1}^p \frac{1}{\tilde{d}_j l_j} \leq \frac{(n_1 - p - 1)(n_2 - p + 7)c^2 \sqrt{p}}{2(n_2 - p + 3)(b + u)\sqrt{b}}. \tag{38}
\end{aligned}$$

The last inequality follows from the inequality

$$\max_{y_j > 0} \frac{\sum_{j=1}^p y_j}{b + \sum_{j=1}^p y_j^2} \leq \frac{\sqrt{p}}{2\sqrt{b}}.$$

Finally putting (36)–(38) into the right-hand side of (35), we have

$$\begin{aligned}
& \text{tr}\zeta\{R_2(\hat{\zeta}^{BE}; \zeta) - R_2(\hat{\zeta}^{LO}; \zeta)\} \\
&\leq \mathbf{E}_{\theta, \Sigma}^G \left[\frac{c}{b + u} \left\{ \frac{(n_1 - p - 1)(n_2 - p + 7)c\sqrt{p}}{2(n_2 - p + 3)\sqrt{b}} - 2(p^2 + p - 4) \right\} \right] \leq 0.
\end{aligned}$$

The last inequality follows from Assumption (III) of Theorem 4. \square

5 Concluding Remarks

Out of findings through our simulation study the following is most interesting. From our simulation study, improved estimators obtained under the normal assumption give some savings in risk when the observation is taken from i.i.d. sampling of the elliptically contoured distributions although robust improvement is shown theoretically under non-i.i.d. sampling set-up.

We used three types of distributions with different kurtosis parameters. Our simulation indicates that the AI's depend on the kurtosis rather than distribution type as long as the distributions of the error matrices belong to the elliptically contoured distributions.

Robustness of improvement is shown as long as the errors have the elliptically contoured distributions. One of the anonymous referees raised an interesting question whether shrinkage method we used in this paper still works under a family of non-normal distributions which are more broad than the elliptically contoured distributions. However, the method underlying in this paper heavily depends on the form of density functions of the elliptically contoured distributions so it is difficult to apply the methods we used in this paper to wider class of distributions. Although it is worthy of being investigated, it is beyond the scope of this paper.

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Table 1. Estimated risks for estimation of (Σ_1, Σ_2) under normal distributions
 (Estimated standard errors are in parentheses; Kurtosis is zero)

$\Sigma_2 \Sigma_1^{-1}$	<i>US</i>	<i>JS</i>	AI	<i>DS</i>	AI	<i>LO</i>	AI
diag (1, 1, 1)	0.9356 (0.0038)	0.8938 (0.0037)	4.47%	0.6649 (0.0033)	28.9%	0.7517 (0.0035)	19.7%
diag (10, 1, 1)	0.9356 (0.0038)	0.8938 (0.0037)	4.47%	0.7068 (0.0034)	24.5%	0.8324 (0.0035)	11.0%
diag (100, 1, 1)	0.9356 (0.0038)	0.8938 (0.0037)	4.47%	0.7100 (0.0034)	24.1%	0.8468 (0.0036)	9.48%
diag (1000, 1, 1)	0.9356 (0.0038)	0.8938 (0.0037)	4.47%	0.7099 (0.0034)	24.1%	0.8480 (0.0036)	9.36%
diag (10, 5, 1)	0.9356 (0.0038)	0.8938 (0.0037)	4.47%	0.7397 (0.0036)	20.9%	0.8331 (0.0035)	11.0%
diag (10, 10, 1)	0.9356 (0.0038)	0.8938 (0.0037)	4.47%	0.7748 (0.0042)	17.2%	0.8325 (0.0035)	11.0%
diag (100, 10, 1)	0.9356 (0.0038)	0.8938 (0.0037)	4.47%	0.7523 (0.0035)	19.6%	0.8812 (0.0036)	5.81%
diag (100, 100, 1)	0.9356 (0.0038)	0.8938 (0.0037)	4.47%	0.8046 (0.0076)	14.0%	0.8470 (0.0036)	9.47%
diag (1000, 10, 1)	0.9356 (0.0038)	0.8938 (0.0037)	4.47%	0.7510 (0.0035)	19.7%	0.8879 (0.0036)	5.10%
diag (1000, 100, 1)	0.9356 (0.0038)	0.8938 (0.0037)	4.47%	0.7694 (0.0035)	17.8%	0.8878 (0.0036)	5.11%
diag (1000, 1000, 1)	0.9356 (0.0038)	0.8938 (0.0037)	4.47%	0.7941 (0.0129)	15.1%	0.8482 (0.0036)	9.34%

Table 2. Estimated risks for estimation of (Σ_1, Σ_2) under *t*-distributions
 (Estimated standard errors are in parentheses; Kurtosis is not defined)

$\Sigma_2 \Sigma_1^{-1}$	<i>US</i>	<i>JS</i>	AI	<i>DS</i>	AI	<i>LO</i>	AI
diag (1, 1, 1)	10.116 (0.4976)	10.078 (0.5042)	0.38%	9.4165 (0.4812)	6.92%	9.6676 (0.4888)	4.44%
diag (10, 1, 1)	10.116 (0.4976)	10.078 (0.5042)	0.38%	9.5315 (0.4838)	5.78%	9.8952 (0.4937)	2.19%
diag (100, 1, 1)	10.116 (0.4976)	10.078 (0.5042)	0.38%	9.5425 (0.4843)	5.67%	9.9337 (0.4938)	1.81%
diag (1000, 1, 1)	10.116 (0.4976)	10.078 (0.5042)	0.38%	9.5423 (0.4843)	5.67%	9.9370 (0.4937)	1.77%
diag (10, 5, 1)	10.116 (0.4976)	10.078 (0.5042)	0.38%	9.6207 (0.4876)	4.90%	9.9112 (0.4973)	2.03%
diag (10, 10, 1)	10.116 (0.4976)	10.078 (0.5042)	0.38%	9.7393 (0.4886)	3.73%	9.9116 (0.4961)	2.02%
diag (100, 10, 1)	10.116 (0.4976)	10.078 (0.5042)	0.38%	9.6664 (0.4887)	4.45%	10.040 (0.4980)	0.76%
diag (100, 100, 1)	10.116 (0.4976)	10.078 (0.5042)	0.38%	9.8365 (0.4921)	2.77%	9.9501 (0.4961)	1.64%
diag (1000, 10, 1)	10.116 (0.4976)	10.078 (0.5042)	0.38%	9.6613 (0.4887)	4.50%	10.058 (0.4980)	0.58%
diag (1000, 100, 1)	10.116 (0.4976)	10.078 (0.5042)	0.38%	9.7189 (0.4895)	3.93%	10.058 (0.4980)	0.58%
diag (1000, 1000, 1)	10.116 (0.4976)	10.078 (0.5042)	0.38%	9.8939 (0.5104)	2.20%	9.9531 (0.4960)	1.61%

Table 3. Estimated risks for estimation of (Σ_1, Σ_2) under Kotz-type distributions

(Estimated standard errors are in parentheses; Kurtosis is -0.9374)

$\Sigma_2 \Sigma_1^{-1}$	<i>US</i>	<i>JS</i>	AI	<i>DS</i>	AI	<i>LO</i>	AI
diag (1, 1, 1)	4.6140 (0.0060)	4.5725 (0.0059)	0.90%	4.5145 (0.0059)	2.16%	4.5363 (0.0059)	1.68%
diag (10, 1, 1)	4.6140 (0.0060)	4.5725 (0.0059)	0.90%	4.5250 (0.0059)	1.93%	4.5572 (0.0059)	1.23%
diag (100, 1, 1)	4.6140 (0.0060)	4.5725 (0.0059)	0.90%	4.5257 (0.0059)	1.91%	4.5609 (0.0059)	1.15%
diag (1000, 1, 1)	4.6140 (0.0060)	4.5725 (0.0059)	0.90%	4.5257 (0.0059)	1.91%	4.5612 (0.0059)	1.14%
diag (10, 5, 1)	4.6140 (0.0060)	4.5725 (0.0059)	0.90%	4.5330 (0.0059)	1.76%	4.5572 (0.0059)	1.23%
diag (10, 10, 1)	4.6140 (0.0060)	4.5725 (0.0059)	0.90%	4.5435 (0.0059)	1.53%	4.5570 (0.0059)	1.24%
diag (100, 10, 1)	4.6140 (0.0060)	4.5725 (0.0059)	0.90%	4.5366 (0.0059)	1.68%	4.5696 (0.0059)	0.96%
diag (100, 100, 1)	4.6140 (0.0060)	4.5725 (0.0059)	0.90%	4.5516 (0.0062)	1.35%	4.5607 (0.0059)	1.16%
diag (1000, 10, 1)	4.6140 (0.0060)	4.5725 (0.0059)	0.90%	4.5365 (0.0059)	1.68%	4.5713 (0.0059)	0.93%
diag (1000, 100, 1)	4.6140 (0.0060)	4.5725 (0.0059)	0.90%	4.5408 (0.0059)	1.59%	4.5713 (0.0059)	0.93%
diag (1000, 1000, 1)	4.6140 (0.0060)	4.5725 (0.0059)	0.90%	4.5503 (0.0068)	1.38%	4.5610 (0.0059)	1.15%

Table 4. Estimated risks for estimation of (Σ_1, Σ_2) under *t*-distributions (i.i.d.)

(Estimated standard errors are in parentheses; Kurtosis is not defined)

$\Sigma_2 \Sigma_1^{-1}$	<i>US</i>	<i>JS</i>	AI	<i>DS</i>	AI	<i>LO</i>	AI
diag (1, 1, 1)	8.3045 (0.2110)	7.7993 (0.1879)	6.08%	7.0405 (0.1845)	15.2%	7.2989 (0.1849)	12.1%
diag (10, 1, 1)	8.3045 (0.2110)	7.7993 (0.1879)	6.08%	7.1560 (0.1847)	13.8%	7.5063 (0.1853)	9.61%
diag (100, 1, 1)	8.3045 (0.2110)	7.7993 (0.1879)	6.08%	7.2055 (0.1853)	13.2%	7.6093 (0.1861)	8.37%
diag (1000, 1, 1)	8.3045 (0.2110)	7.7993 (0.1879)	6.08%	7.2099 (0.1855)	13.2%	7.6232 (0.1865)	8.20%
diag (10, 5, 1)	8.3045 (0.2110)	7.7993 (0.1879)	6.08%	7.1335 (0.1841)	14.1%	7.4849 (0.1853)	9.87%
diag (10, 10, 1)	8.3045 (0.2110)	7.7993 (0.1879)	6.08%	7.2077 (0.1838)	13.2%	7.5056 (0.1853)	9.62%
diag (100, 10, 1)	8.3045 (0.2110)	7.7993 (0.1879)	6.08%	7.2281 (0.1842)	13.0%	7.6732 (0.1859)	7.60%
diag (100, 100, 1)	8.3045 (0.2110)	7.7993 (0.1879)	6.08%	7.3523 (0.1852)	11.5%	7.6110 (0.1862)	8.35%
diag (1000, 10, 1)	8.3045 (0.2110)	7.7993 (0.1879)	6.08%	7.2351 (0.1846)	12.9%	7.7301 (0.1865)	6.92%
diag (1000, 100, 1)	8.3045 (0.2110)	7.7993 (0.1879)	6.08%	7.3381 (0.1852)	11.6%	7.7286 (0.1865)	6.93%
diag (1000, 1000, 1)	8.3045 (0.2110)	7.7993 (0.1879)	6.08%	7.3524 (0.1867)	11.5%	7.6243 (0.1864)	8.19%

Table 5. Estimated risks for estimation of (Σ_1, Σ_2) under Kotz-type distributions (i.i.d.)

(Estimated standard errors are in parentheses; Kurtosis is 0.8333)

$\Sigma_2 \Sigma_1^{-1}$	<i>US</i>	<i>JS</i>	<i>AI</i>	<i>DS</i>	<i>AI</i>	<i>LO</i>	<i>AI</i>
diag (1, 1, 1)	1.5235 (0.0037)	1.5228 (0.0037)	0.04%	1.1793 (0.0030)	22.6%	1.3040 (0.0033)	14.4%
diag (10, 1, 1)	1.5235 (0.0037)	1.5228 (0.0037)	0.04%	1.2461 (0.0032)	18.2%	1.4343 (0.0035)	5.85%
diag (100, 1, 1)	1.5235 (0.0037)	1.5228 (0.0037)	0.04%	1.2489 (0.0032)	18.0%	1.4503 (0.0036)	4.80%
diag (1000, 1, 1)	1.5235 (0.0037)	1.5228 (0.0037)	0.04%	1.2487 (0.0032)	18.0%	1.4517 (0.0036)	4.71%
diag (10, 5, 1)	1.5235 (0.0037)	1.5228 (0.0037)	0.04%	1.3037 (0.0038)	14.4%	1.4425 (0.0034)	5.32%
diag (10, 10, 1)	1.5235 (0.0037)	1.5228 (0.0037)	0.04%	1.3743 (0.0057)	9.79%	1.4337 (0.0035)	5.89%
diag (100, 10, 1)	1.5235 (0.0037)	1.5228 (0.0037)	0.04%	1.3174 (0.0034)	13.5%	1.5069 (0.0037)	1.09%
diag (100, 100, 1)	1.5235 (0.0037)	1.5228 (0.0037)	0.04%	1.4237 (0.0136)	6.55%	1.4498 (0.0036)	4.84%
diag (1000, 10, 1)	1.5235 (0.0037)	1.5228 (0.0037)	0.04%	1.3166 (0.0034)	13.6%	1.5141 (0.0037)	0.62%
diag (1000, 100, 1)	1.5235 (0.0037)	1.5228 (0.0037)	0.04%	1.3374 (0.0034)	12.2%	1.5141 (0.0037)	0.62%
diag (1000, 1000, 1)	1.5235 (0.0037)	1.5228 (0.0037)	0.04%	1.4000 (0.0233)	8.10%	1.4512 (0.0036)	4.74%

Table 6. Estimated risks for estimation of (Σ_1, Σ_2) under normal distributions

(Estimated standard errors are in parentheses; Kurtosis is zero)

$\Sigma_2 \Sigma_1^{-1}$	<i>UB</i>	<i>ST</i>	<i>AI</i>	<i>DS</i>	<i>AI</i>	<i>LO</i>	<i>AI</i>
$(N_1, N_2, p) = (15, 15, 5)$							
diag (1, 1, 1, 1, 1)	2.4750 (0.0065)	2.2659 (0.0060)	8.45%	1.4628 (0.0053)	40.9%	1.7783 (0.0056)	28.2%
diag ($10^4, 10^3, 10^2, 10, 1$)	2.4750 (0.0065)	2.2659 (0.0060)	8.45%	1.8577 (0.0059)	24.9%	2.2359 (0.0058)	9.66%
$(N_1, N_2, p) = (15, 15, 8)$							
diag (1, 1, 1, 1, 1, 1, 1, 1)	6.6374 (0.0113)	5.6895 (0.0099)	14.3%	3.3701 (0.0094)	49.2%	4.3066 (0.0097)	35.1%
diag ($10^7, 10^6, 10^5, 10^4, 10^3, 10^2, 10, 1$)	6.6374 (0.0113)	5.6895 (0.0099)	14.3%	4.5516 (0.0102)	31.4%	5.6254 (0.0097)	15.3%
$(N_1, N_2, p) = (15, 15, 10)$							
diag (1, 1, 1, 1, 1, 1, 1, 1, 1, 1)	11.210 (0.0159)	9.1777 (0.0134)	18.1%	5.4054 (0.0135)	51.8%	6.9257 (0.0134)	38.2%
diag ($10^9, 10^8, 10^7, 10^6, 10^5, 10^4, 10^3, 10^2, 10, 1$)	11.210 (0.0159)	9.1777 (0.0134)	18.1%	7.3720 (0.0145)	34.2%	9.0567 (0.0130)	19.2%

Table 7. Estimated risks for estimation of (Σ_1, Σ_2) under Kotz-type distributions

(Estimated standard errors are in parentheses; Kurtosis is zero)

$\Sigma_2 \Sigma_1^{-1}$	UB	ST	AI	DS	AI	LO	AI
$(N_1, N_2, p) = (15, 15, 3), (u, r, s) = (5, 8.0348, 0.5)$							
diag (1, 1, 1)	1.0262 (0.0042)	0.9840 (0.0041)	4.11%	0.7578 (0.0038)	26.2%	0.8439 (0.0039)	17.8%
diag (100, 10, 1)	1.0262 (0.0042)	0.9840 (0.0041)	4.11%	0.8436 (0.0039)	17.8%	0.9724 (0.0040)	5.24%
$(N_1, N_2, p) = (15, 15, 5), (u, r, s) = (5, 9.6924, 0.5)$							
diag (1, 1, 1, 1, 1)	2.6084 (0.0068)	2.3937 (0.0063)	8.23%	1.5982 (0.0058)	38.7%	1.9108 (0.0060)	26.7%
diag ($10^4, 10^3, 10^2, 10, 1$)	2.6084 (0.0068)	2.3937 (0.0063)	8.23%	1.9902 (0.0064)	23.7%	2.3623 (0.0062)	9.43%
$(N_1, N_2, p) = (15, 15, 8), (u, r, s) = (5, 11.7725, 0.5)$							
diag (1, 1, 1, 1, 1, 1, 1, 1)	6.7849 (0.0115)	5.8292 (0.0101)	14.1%	3.5290 (0.0099)	48.0%	4.4610 (0.0100)	34.3%
diag ($10^7, 10^6, 10^5, 10^4, 10^3, 10^2, 10, 1$)	6.7849 (0.0115)	5.8292 (0.0101)	14.1%	4.7058 (0.0107)	30.6%	5.7636 (0.0099)	15.1%
$(N_1, N_2, p) = (15, 15, 10), (u, r, s) = (5, 12.9797, 0.5)$							
diag (1, 1, 1, 1, 1, 1, 1, 1, 1, 1)	11.341 (0.0161)	9.3073 (0.0136)	17.9%	5.5591 (0.0140)	51.0%	7.0676 (0.0138)	37.7%
diag ($10^9, 10^8, 10^7, 10^6, 10^5, 10^4, 10^3, 10^2, 10, 1$)	11.341 (0.0161)	9.3073 (0.0136)	17.9%	7.5121 (0.0148)	33.8%	9.1873 (0.0133)	19.0%

Table 8. Estimated risks for estimation of (Σ_1, Σ_2) under Kotz-type distributions (i.i.d.)

(Estimated standard errors are in parentheses; Kurtosis is zero)

$\Sigma_2 \Sigma_1^{-1}$	UB	ST	AI	DS	AI	LO	AI
$(N_1, N_2, p) = (15, 15, 3), (u_1, r_1, s_1) = (5, 6.3261, 0.5), (u_2, r_2, s_2) = (5, 6.3261, 0.5)$							
diag (1, 1, 1)	0.7769 (0.0032)	0.7482 (0.0031)	3.70%	0.5080 (0.0026)	34.6%	0.5964 (0.0028)	23.2%
diag (100, 10, 1)	0.7769 (0.0032)	0.7482 (0.0031)	3.70%	0.6024 (0.0028)	22.5%	0.7383 (0.0031)	4.97%
$(N_1, N_2, p) = (15, 15, 5), (u_1, r_1, s_1) = (5, 5.9437, 0.5), (u_2, r_2, s_2) = (5, 5.9437, 0.5)$							
diag (1, 1, 1, 1, 1)	2.3326 (0.0062)	2.1413 (0.0057)	8.20%	1.3246 (0.0049)	43.2%	1.6438 (0.0053)	29.5%
diag ($10^4, 10^3, 10^2, 10, 1$)	2.3326 (0.0062)	2.1413 (0.0057)	8.20%	1.7259 (0.0056)	26.0%	2.1140 (0.0055)	9.37%
$(N_1, N_2, p) = (15, 15, 8), (u_1, r_1, s_1) = (5, 5.8395, 0.5), (u_2, r_2, s_2) = (5, 5.8395, 0.5)$							
diag (1, 1, 1, 1, 1, 1, 1, 1)	6.6518 (0.0114)	5.6927 (0.0099)	14.4%	3.3851 (0.0093)	49.1%	4.3190 (0.0096)	35.1%
diag ($10^7, 10^6, 10^5, 10^4, 10^3, 10^2, 10, 1$)	6.6518 (0.0114)	5.6927 (0.0099)	14.4%	4.5602 (0.0101)	31.4%	5.6293 (0.0097)	15.4%
$(N_1, N_2, p) = (15, 15, 10), (u_1, r_1, s_1) = (5, 5.8820, 0.5), (u_2, r_2, s_2) = (5, 5.8820, 0.5)$							
diag (1, 1, 1, 1, 1, 1, 1, 1, 1, 1)	11.334 (0.0158)	9.2751 (0.0133)	18.2%	5.5416 (0.0134)	51.1%	7.0546 (0.0134)	37.8%
diag ($10^9, 10^8, 10^7, 10^6, 10^5, 10^4, 10^3, 10^2, 10, 1$)	11.334 (0.0158)	9.2751 (0.0133)	18.2%	7.4992 (0.0146)	33.8%	9.1526 (0.0129)	19.3%

Table 9. Estimated risks for estimation of $\Sigma_2 \Sigma_1^{-1}$ under normal distributions

(Estimated standard errors are in parentheses; Kurtosis is zero)

$\Sigma_2 \Sigma_1^{-1}$	<i>BU</i>	<i>LO</i>	<i>AI</i>	<i>BE</i>	<i>AI</i>	<i>SB</i>	<i>AI</i>
diag (1, 1, 1)	6.192 (0.033)	5.391 (0.033)	12.9%	4.822 (0.032)	22.1%	4.424 (0.032)	28.6%
diag (10, 1, 1)	6.186 (0.042)	6.008 (0.042)	2.87%	5.853 (0.041)	5.38%	5.931 (0.042)	4.12%
diag (100, 1, 1)	6.178 (0.047)	6.160 (0.047)	0.30%	6.142 (0.047)	0.60%	6.151 (0.047)	0.44%
diag (1000, 1, 1)	6.176 (0.047)	6.174 (0.047)	0.03%	6.172 (0.047)	0.06%	6.173 (0.047)	0.05%
diag (10, 5, 1)	6.198 (0.037)	5.833 (0.037)	5.89%	5.705 (0.036)	7.96%	5.671 (0.036)	8.67%
diag (10, 10, 1)	6.205 (0.037)	5.790 (0.036)	6.69%	5.691 (0.036)	8.28%	5.517 (0.036)	11.1%
diag (100, 10, 1)	6.184 (0.044)	6.124 (0.044)	0.97%	6.106 (0.044)	1.27%	6.167 (0.044)	0.27%
diag (100, 100, 1)	6.205 (0.038)	5.845 (0.037)	5.80%	5.835 (0.037)	5.97%	5.564 (0.037)	10.3%
diag (1000, 10, 1)	6.177 (0.047)	6.171 (0.047)	0.09%	6.169 (0.047)	0.12%	6.176 (0.047)	0.02%
diag (1000, 100, 1)	6.183 (0.044)	6.134 (0.044)	0.79%	6.132 (0.044)	0.82%	6.178 (0.044)	0.07%
diag (1000, 1000, 1)	6.204 (0.038)	5.849 (0.037)	5.73%	5.848 (0.037)	5.75%	5.566 (0.037)	10.3%

Table 10. Estimated risks for estimation of $\Sigma_2 \Sigma_1^{-1}$ under *t*-distributions

(Estimated standard errors are in parentheses; Kurtosis is 7.3333)

$\Sigma_2 \Sigma_1^{-1}$	<i>BU</i>	<i>LO</i>	<i>AI</i>	<i>BE</i>	<i>AI</i>	<i>SB</i>	<i>AI</i>
diag (1, 1, 1)	10.26 (0.145)	8.954 (0.131)	12.8%	8.033 (0.122)	21.7%	7.359 (0.114)	28.3%
diag (10, 1, 1)	10.14 (0.146)	9.860 (0.144)	2.77%	9.608 (0.141)	5.25%	9.745 (0.145)	3.91%
diag (100, 1, 1)	10.10 (0.152)	10.07 (0.152)	0.29%	10.04 (0.152)	0.58%	10.05 (0.152)	0.44%
diag (1000, 1, 1)	10.09 (0.153)	10.09 (0.153)	0.03%	10.08 (0.153)	0.06%	10.08 (0.153)	0.04%
diag (10, 5, 1)	10.20 (0.143)	9.609 (0.137)	5.83%	9.401 (0.135)	7.86%	9.341 (0.134)	8.46%
diag (10, 10, 1)	10.24 (0.145)	9.568 (0.139)	6.56%	9.408 (0.137)	8.11%	9.123 (0.135)	10.9%
diag (100, 10, 1)	10.12 (0.148)	10.03 (0.147)	0.91%	9.999 (0.147)	1.21%	10.11 (0.148)	0.16%
diag (100, 100, 1)	10.22 (0.145)	9.633 (0.139)	5.71%	9.616 (0.139)	5.87%	9.169 (0.135)	10.3%
diag (1000, 10, 1)	10.09 (0.153)	10.08 (0.152)	0.09%	10.08 (0.152)	0.12%	10.09 (0.153)	0.02%
diag (1000, 100, 1)	10.11 (0.148)	10.04 (0.147)	0.74%	10.04 (0.147)	0.77%	10.12 (0.149)	-0.02%
diag (1000, 1000, 1)	10.21 (0.145)	9.631 (0.139)	5.65%	9.630 (0.139)	5.66%	9.165 (0.135)	10.2%

Table 11. Estimated risks for estimation of $\Sigma_2 \Sigma_1^{-1}$ under Kotz-type distributions(Estimated standard errors are in parentheses; Kurtosis is -0.9704)

$\Sigma_2 \Sigma_1^{-1}$	<i>BU</i>	<i>LO</i>	AI	<i>BE</i>	AI	<i>SB</i>	AI
diag (1, 1, 1)	1.072 (0.006)	0.933 (0.006)	12.9%	0.835 (0.006)	22.1%	0.766 (0.006)	28.5%
diag (10, 1, 1)	1.074 (0.007)	1.043 (0.007)	2.87%	1.017 (0.007)	5.37%	1.030 (0.007)	4.14%
diag (100, 1, 1)	1.074 (0.008)	1.071 (0.008)	0.30%	1.068 (0.008)	0.59%	1.070 (0.008)	0.44%
diag (1000, 1, 1)	1.074 (0.008)	1.074 (0.008)	0.03%	1.074 (0.008)	0.06%	1.074 (0.008)	0.05%
diag (10, 5, 1)	1.073 (0.006)	1.010 (0.006)	5.89%	0.988 (0.006)	7.95%	0.982 (0.006)	8.47%
diag (10, 10, 1)	1.072 (0.006)	1.000 (0.006)	6.70%	0.982 (0.006)	8.28%	0.953 (0.006)	11.1%
diag (100, 10, 1)	1.074 (0.008)	1.064 (0.007)	0.96%	1.061 (0.007)	1.25%	1.072 (0.008)	0.25%
diag (100, 100, 1)	1.072 (0.006)	1.009 (0.006)	5.81%	1.008 (0.006)	5.97%	0.961 (0.006)	10.3%
diag (1000, 10, 1)	1.074 (0.008)	1.073 (0.008)	0.09%	1.073 (0.008)	0.12%	1.074 (0.008)	0.01%
diag (1000, 100, 1)	1.074 (0.008)	1.066 (0.008)	0.78%	1.065 (0.008)	0.81%	1.074 (0.008)	0.06%
diag (1000, 1000, 1)	1.072 (0.006)	1.010 (0.006)	5.73%	1.010 (0.006)	5.75%	0.961 (0.006)	10.3%

Table 12. Estimated risks for estimation of $\Sigma_2 \Sigma_1^{-1}$ under t -distributions (i.i.d.)

(Estimated standard errors are in parentheses; Kurtosis is 7.3333)

$\Sigma_2 \Sigma_1^{-1}$	<i>BU</i>	<i>LO</i>	AI	<i>BE</i>	AI	<i>SB</i>	AI
diag (1, 1, 1)	20.73 (0.883)	19.65 (0.883)	5.21%	18.92 (0.888)	8.72%	17.89 (0.876)	13.7%
diag (10, 1, 1)	20.48 (0.963)	20.28 (0.964)	0.97%	20.06 (0.965)	2.03%	20.01 (0.960)	2.26%
diag (100, 1, 1)	20.57 (1.140)	20.55 (1.140)	0.06%	20.53 (1.140)	0.18%	20.52 (1.140)	0.20%
diag (1000, 1, 1)	20.63 (1.193)	20.63 (1.193)	0.00%	20.63 (1.193)	0.02%	20.62 (1.193)	0.02%
diag (10, 5, 1)	20.14 (0.825)	19.67 (0.825)	2.35%	19.47 (0.826)	3.31%	19.08 (0.816)	5.25%
diag (10, 10, 1)	20.02 (0.792)	19.49 (0.793)	2.63%	19.34 (0.793)	3.38%	18.78 (0.783)	6.22%
diag (100, 10, 1)	20.33 (1.008)	20.29 (1.008)	0.20%	20.26 (1.008)	0.34%	20.27 (1.006)	0.27%
diag (100, 100, 1)	19.97 (0.808)	19.52 (0.809)	2.23%	19.50 (0.809)	2.31%	18.85 (0.799)	5.59%
diag (1000, 10, 1)	20.55 (1.149)	20.55 (1.149)	0.00%	20.55 (1.149)	0.01%	20.55 (1.149)	0.01%
diag (1000, 100, 1)	20.32 (1.014)	20.29 (1.014)	0.14%	20.29 (1.014)	0.15%	20.29 (1.012)	0.16%
diag (1000, 1000, 1)	19.98 (0.810)	19.53 (0.810)	2.21%	19.53 (0.810)	2.21%	18.87 (0.801)	5.56%

Table 13. Estimated risks for estimation of $\Sigma_2 \Sigma_1^{-1}$ under Kotz-type distributions (i.i.d.)

(Estimated standard errors are in parentheses; Kurtosis is 0.8333)

$\Sigma_2 \Sigma_1^{-1}$	<i>BU</i>	<i>LO</i>	AI	<i>BE</i>	AI	<i>SB</i>	AI
diag (1, 1, 1)	7.815 (0.029)	6.333 (0.029)	19.0%	5.297 (0.028)	32.2%	4.703 (0.029)	39.8%
diag (10, 1, 1)	7.761 (0.042)	7.424 (0.041)	4.35%	7.150 (0.041)	7.88%	7.328 (0.043)	5.58%
diag (100, 1, 1)	7.752 (0.048)	7.714 (0.048)	0.49%	7.682 (0.048)	0.90%	7.703 (0.049)	0.63%
diag (1000, 1, 1)	7.753 (0.049)	7.749 (0.049)	0.05%	7.746 (0.049)	0.09%	7.748 (0.049)	0.06%
diag (10, 5, 1)	7.782 (0.035)	7.122 (0.034)	8.49%	6.903 (0.034)	11.3%	7.001 (0.033)	10.0%
diag (10, 10, 1)	7.799 (0.034)	7.022 (0.033)	9.96%	6.854 (0.033)	12.1%	6.629 (0.034)	15.0%
diag (100, 10, 1)	7.756 (0.045)	7.633 (0.045)	1.59%	7.602 (0.045)	2.00%	7.731 (0.046)	0.33%
diag (100, 100, 1)	7.796 (0.035)	7.115 (0.035)	8.73%	7.097 (0.035)	8.96%	6.684 (0.035)	14.3%
diag (1000, 10, 1)	7.752 (0.049)	7.740 (0.049)	0.16%	7.736 (0.049)	0.21%	7.750 (0.049)	0.03%
diag (1000, 100, 1)	7.756 (0.045)	7.654 (0.045)	1.32%	7.651 (0.045)	1.36%	7.751 (0.046)	0.07%
diag (1000, 1000, 1)	7.795 (0.035)	7.123 (0.035)	8.62%	7.122 (0.035)	8.64%	6.689 (0.035)	14.2%

Table 14. Estimated risks for estimation of $\Sigma_2 \Sigma_1^{-1}$ under normal distributions

(Estimated standard errors are in parentheses; Kurtosis is zero)

$\Sigma_2 \Sigma_1^{-1}$	<i>BU</i>	<i>LO</i>	AI	<i>BE</i>	AI	<i>SB</i>	AI
$(N_1, N_2, p) = (15, 15, 5)$							
diag (1, 1, 1, 1, 1)	8.402 (0.027)	7.118 (0.027)	15.3%	5.557 (0.027)	33.9%	5.512 (0.027)	34.4%
diag (10 ⁴ , 10 ³ , 10 ² , 10, 1)	8.427 (0.045)	8.381 (0.045)	0.54%	8.380 (0.045)	0.56%	8.420 (0.045)	0.09%
$(N_1, N_2, p) = (15, 15, 8)$							
diag (1, 1, 1, 1, 1, 1, 1, 1)	10.97 (0.023)	9.592 (0.023)	12.5%	7.548 (0.027)	31.2%	8.371 (0.024)	23.7%
diag (10 ⁷ , 10 ⁶ , 10 ⁵ , 10 ⁴ , 10 ³ , 10 ² , 10, 1)	11.01 (0.046)	10.98 (0.046)	0.27%	10.98 (0.046)	0.27%	11.00 (0.046)	0.08%
$(N_1, N_2, p) = (15, 15, 10)$							
diag (1, 1, 1, 1, 1, 1, 1, 1, 1, 1)	12.32 (0.022)	11.28 (0.022)	8.42%	10.16 (0.025)	17.6%	10.79 (0.022)	12.4%
diag (10 ⁹ , 10 ⁸ , 10 ⁷ , 10 ⁶ , 10 ⁵ , 10 ⁴ , 10 ³ , 10 ² , 10, 1)	12.29 (0.053)	12.27 (0.053)	0.17%	12.27 (0.053)	0.17%	12.28 (0.053)	0.08%

Table 15. Estimated risks for estimation of $\Sigma_2 \Sigma_1^{-1}$ under Kotz-type distributions

(Estimated standard errors are in parentheses; Kurtosis is zero)

$\Sigma_2 \Sigma_1^{-1}$	BU	LO	AI	BE	AI	SB	AI
$(N_1, N_2, p) = (15, 15, 3), (u, r, s) = (5, 10.4301, 0.5)$							
diag (1, 1, 1)	6.179 (0.038)	5.386 (0.038)	12.8%	4.818 (0.037)	22.0%	4.426 (0.037)	28.4%
diag (100, 10, 1)	6.136 (0.045)	6.078 (0.045)	0.95%	6.059 (0.044)	1.25%	6.123 (0.045)	0.22%
$(N_1, N_2, p) = (15, 15, 5), (u, r, s) = (5, 12.9797, 0.5)$							
diag (1, 1, 1, 1, 1)	8.375 (0.029)	7.100 (0.028)	15.2%	5.553 (0.028)	33.7%	5.506 (0.028)	34.3%
diag ($10^4, 10^3, 10^2, 10, 1$)	8.363 (0.047)	8.317 (0.046)	0.55%	8.315 (0.046)	0.57%	8.353 (0.047)	0.11%
$(N_1, N_2, p) = (15, 15, 8), (u, r, s) = (5, 16.0716, 0.5)$							
diag (1, 1, 1, 1, 1, 1, 1, 1, 1)	10.94 (0.025)	9.577 (0.025)	12.5%	7.548 (0.028)	31.0%	8.366 (0.026)	23.5%
diag ($10^7, 10^6, 10^5, 10^4, 10^3, 10^2, 10, 1$)	10.90 (0.047)	10.87 (0.047)	0.26%	10.87 (0.047)	0.26%	10.89 (0.047)	0.07%
$(N_1, N_2, p) = (15, 15, 10), (u, r, s) = (5, 17.8392, 0.5)$							
diag (1, 1, 1, 1, 1, 1, 1, 1, 1, 1)	12.31 (0.023)	11.27 (0.023)	8.41%	10.15 (0.026)	17.5%	10.78 (0.023)	12.4%
diag ($10^9, 10^8, 10^7, 10^6, 10^5, 10^4, 10^3, 10^2, 10, 1$)	12.32 (0.050)	12.30 (0.050)	0.17%	12.30 (0.050)	0.17%	12.31 (0.050)	0.08%

Table 16. Estimated risks for estimation of $\Sigma_2 \Sigma_1^{-1}$ under Kotz-type distributions (i.i.d.)

(Estimated standard errors are in parentheses; Kurtosis is zero)

$\Sigma_2 \Sigma_1^{-1}$	BU	LO	AI	BE	AI	SB	AI
$(N_1, N_2, p) = (15, 15, 3), (u_1, r_1, s_1) = (5, 6.3261, 0.5), (u_2, r_2, s_2) = (5, 6.3261, 0.5)$							
diag (1, 1, 1)	5.917 (0.028)	4.988 (0.027)	15.7%	4.332 (0.026)	26.8%	3.922 (0.027)	33.7%
diag (100, 10, 1)	5.941 (0.038)	5.867 (0.038)	1.25%	5.846 (0.037)	1.60%	5.923 (0.038)	0.30%
$(N_1, N_2, p) = (15, 15, 5), (u_1, r_1, s_1) = (5, 5.9437, 0.5), (u_2, r_2, s_2) = (5, 5.9437, 0.5)$							
diag (1, 1, 1, 1, 1)	8.348 (0.026)	6.993 (0.026)	16.2%	5.336 (0.025)	36.1%	5.310 (0.026)	36.4%
diag ($10^4, 10^3, 10^2, 10, 1$)	8.320 (0.042)	8.270 (0.042)	0.60%	8.269 (0.042)	0.62%	8.314 (0.043)	0.07%
$(N_1, N_2, p) = (15, 15, 8), (u_1, r_1, s_1) = (5, 5.8395, 0.5), (u_2, r_2, s_2) = (5, 5.8395, 0.5)$							
diag (1, 1, 1, 1, 1, 1, 1, 1, 1)	10.90 (0.022)	9.536 (0.022)	12.6%	7.500 (0.026)	31.2%	8.321 (0.024)	23.7%
diag ($10^7, 10^6, 10^5, 10^4, 10^3, 10^2, 10, 1$)	10.95 (0.045)	10.92 (0.045)	0.27%	10.92 (0.045)	0.27%	10.94 (0.045)	0.08%
$(N_1, N_2, p) = (15, 15, 10), (u_1, r_1, s_1) = (5, 5.8820, 0.5), (u_2, r_2, s_2) = (5, 5.8820, 0.5)$							
diag (1, 1, 1, 1, 1, 1, 1, 1, 1, 1)	12.17 (0.022)	11.15 (0.022)	8.37%	10.06 (0.025)	17.3%	10.67 (0.022)	12.3%
diag ($10^9, 10^8, 10^7, 10^6, 10^5, 10^4, 10^3, 10^2, 10, 1$)	12.17 (0.047)	12.15 (0.047)	0.17%	12.15 (0.047)	0.17%	12.16 (0.047)	0.08%