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A note on a matrix inequality for generalized means

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Abstract

In Statistics, generalized means of positive random variables are often considered. As is well-known, the generalized mean of order α, $\{E(v^\alpha)\}^{\frac{1}{\alpha}}$, is smaller (greater) than the ordinary mean if $\alpha < 1$ ($\alpha > 1$). This result can be generalized to a corresponding inequality involving matrix random variables of a specific type. In the special case when $\alpha = -1$, we have a matrix inequality that has applications in various fields of Statistics. Two such applications are presented.

1. Introduction

Consider a positive random variable v. Its generalized mean of order α, $\alpha \neq 0$, is given by

$$\mu_\alpha = \{E(v^\alpha)\}^{\frac{1}{\alpha}}.$$

For $\alpha = 1$ we have the ordinary mean, for $\alpha = 2$ the quadratic mean, and for $\alpha = -1$ the harmonic mean. The definition of μ_α can be extended to the case $\alpha = 0$:

$$\mu_0 = \exp(E(\ln v)),$$

which is the geometric mean of v.

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Suppose v is not restricted to a single point, we then have the well-known inequality,

$$\mu_\alpha < \mu < \mu_\beta,$$

whenever $\alpha < 1 < \beta$, assuming that $\mu := Ev = \mu_1$ is finite, e.g. [1, Section 2.6].

This inequality can be generalized to the case where v is multiplied by a matrix of the form ww', w being a random vector with $Eww' = I$.

2. The matrix inequality

The following lemma generalizes the concept of a generalized mean of order α.

Lemma. *Suppose that x, y are random variables with $x \geqslant 0$ and $y > 0$ a.s., $Ex = 1$, $E(xy) < \infty$, and for any $d > 0$, $P(y \neq d, x \neq 0) > 0$. Then, for any $\alpha < 1$ and $\beta > 1$,*

$$\{E(xy^\alpha)\}^{\frac{1}{\alpha}} < E(xy) < \{E(xy^\beta)\}^{\frac{1}{\beta}}, \tag{1}$$

where, by definition,

$$\{E(xy^\alpha)\}^{\frac{1}{\alpha}}\Big|_{\alpha=0} := \mathrm{e}^{E(x \ln y)}.$$

Proof. Let $F(x, y)$ be the joint d.f. of x and y, then for each $c \neq 0$,

$$E(xy^c) = \int_{\mathbb{R}} y^c \, \mathrm{d}G(y), \tag{2}$$

with

$$G(y) = \int_{\mathbb{R}} x \, \mathrm{d}F(x, y), \quad y \in \mathbb{R}.$$

The function $G(y)$ is a probability d.f. since,

$$G(+\infty) = Ex = 1.$$

Denote the integral on the right hand side of (2) by $E_x y^c$. Note that the distribution $G(y)$ is not concentrated in a single point $y = d$ because this would imply $P(y \neq d, x \neq 0) = 0$ contrary to the assumption. For $\alpha \neq 0$ the inequality (1) can be written in the form

$$(E_x y^\alpha)^{\frac{1}{\alpha}} < E_x y < (E_x y^\beta)^{\frac{1}{\beta}}$$

and follows from Jensen's inequality, because $E_x y = E(xy) < \infty$. For $\alpha = 0$ we have to show that

$$\mathrm{e}^{E_x \ln y} < E_x y, \tag{3}$$

where

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$$E_x \ln y := \int_{\mathbb{R}} \ln y \, \mathrm{d}G(y).$$

But this also follows from Jensen's inequality, which completes the proof. □

Proposition. *Let v be a positive random variable with a distribution which has no atoms and w be a random (column) vector in $\mathbb{R}^m$, with $Eww' = I_m$. Let α and β be real numbers as previously stipulated and assume $E(\|w\|^2 v) < \infty$. Then, for $\alpha \neq 0$,*

$$\operatorname{tr}\{E(ww'v^\alpha)\}^{\frac{1}{\alpha}} < \operatorname{tr} E(ww'v) < \operatorname{tr}\{E(ww'v^\beta)\}^{\frac{1}{\beta}}. \quad (4)$$

This inequality holds true also for $\alpha = 0$, where by definition

$$\{E(ww'v^\alpha)\}^{\frac{1}{\alpha}}\Big|_{\alpha=0} := \mathrm{e}^{E(ww' \ln v)}.$$

Proof. Assume $\alpha \neq 0$ and <1. First note that, by the lemma, $E(\|w\|^2 v^\alpha) < \infty$ (set $\frac{1}{m}\|w\|^2 = x$ and $v = y$). Now, let (λ, φ) be an eigenvalue and normalized eigenvector pair of $E(ww'v^\alpha)$. Then

$$\begin{aligned} \varphi'\{E(ww'v^\alpha)\}^{\frac{1}{\alpha}}\varphi &= \lambda^{\frac{1}{\alpha}} = \{\varphi' E(ww'v^\alpha)\varphi\}^{\frac{1}{\alpha}} \\ &= \{E(\varphi'w)^2 v^\alpha\}^{\frac{1}{\alpha}} = \{E(xv^\alpha)\}^{\frac{1}{\alpha}}, \end{aligned}$$

where $x = (\varphi'w)^2$, and $Ex = \varphi'(Eww')\varphi = \varphi'\varphi = 1$. We can now show that for each $d > 0$, $P(v \neq d, x \neq 0) > 0$. For suppose $P(v \neq d \text{ and } x \neq 0) = 0$ or equivalently $P(v = d \text{ or } x = 0) = 1$, then $P(v = d) > 0$ because $P(v = d) = 0$ would imply $P(x = 0) = 1$, which is impossible since $Ex = 1$. But $P(v = d) > 0$ contradicts the assumption that the distribution of v has no atoms. From the lemma we have,

$$\{E(xv^\alpha)\}^{\frac{1}{\alpha}} < E(xv) = \varphi' E(ww'v)\varphi$$

and, therefore,

$$\varphi'\{E(ww'v^\alpha)\}^{\frac{1}{\alpha}}\varphi < \varphi' E(ww'v)\varphi.$$

Summing over all φ belonging to an eigenbasis of $E(ww'v^\alpha)$, we obtain the left hand side of (4). The right hand side is established in a similar way.

For the case $\alpha = 0$ we start with (λ, φ) being an eigenvalue and eigenvector pair of $E(ww' \ln v)$. By arguments analogous to the above, we find that,

$$\varphi' \mathrm{e}^{E(ww' \ln v)}\varphi < \varphi' E(ww'v)\varphi,$$

from which again the left hand side of (4) follows for $\alpha = 0$. □

For an earlier version of the proof see [2].

Remark 1. If the distribution of v has atoms, then (4) holds with nonstrict inequality.

Remark 2. For $\alpha = -1$, we have the stronger proposition,

$$\{E(ww'v^{-1})\}^{-1} \leqslant E(ww'v)$$

in the Loewner sense. For a proof see [4].

3. Statistical applications

We have two applications in the theory of measurement error models of the proposition and its stronger version for $\alpha = -1$, see Remark 2.

3.1. Poisson model

Shklyar and Schneeweiss [4] consider the log-linear Poisson model. It is given by a Poisson distributed random variable y with mean parameter λ, where $\log \lambda$ is a linear function of a random vector ξ: $\log \lambda = \beta_0 + \beta_1'\xi$ with an unknown parameter vector $\beta = (\beta_0, \beta_1')'$. Assume that $\xi \sim N(\mu_\xi, \Sigma_\xi)$ with known mean vector μ_ξ and covariance matrix Σ_ξ. Assume further that ξ is unobservable. Instead a random vector x is observed, which is related to ξ by the equation $x = \xi + \delta$, $\delta \sim N(0, \Sigma_\delta)$, δ being the vector of measurement errors with known covariance matrix Σ_δ. δ is assumed to be independent of ξ and y.

In this model, it is possible to evaluate the conditional mean and variance of y given x as functions of x and β:

$$E(y \mid x) = m(x, \beta)$$

$$V(y \mid x) = v(x, \beta).$$

They can by used to construct unbiased estimating functions $\psi(y, x, \beta)$ with the property that $E[\psi(y, x, \beta)] = 0$. In particular the following two estimating functions will be considered:

$$\psi_1(y, x, \beta) = [y - m(x, \beta)]v^{-1}(x, \beta)m_\beta(x, \beta),$$

$$\psi_2(y, x, \beta) = [y - m(x, \beta)]m^{-1}(x, \beta)m_\beta(x, \beta),$$

where $m_\beta(x, \beta) = \frac{\partial}{\partial \beta}m(x, \beta)$. Given an i.i.d. sample (x_i, y_i), $i = 1, \ldots, n$, consistent estimators $\widehat{\beta_1}$ and $\widehat{\beta_2}$ of β are found by solving the equations

$$\sum_{i=1}^{n} \psi_j(y_i, x_i, \widehat{\beta_j}) = 0, \quad j = 1, 2,$$

respectively.

The asymptotic covariance matrices of $\widehat{\beta_1}$ and $\widehat{\beta_2}$ are, respectively, given by

$$\Sigma_1 = \left[E\left(v^{-1}m_\beta m_\beta'\right)\right]^{-1},$$

$$\Sigma_2 = \left[E\left(m^{-1}m_\beta m_\beta'\right)\right]^{-1} E\left(vm^{-2}m_\beta m_\beta'\right)\left[E\left(m^{-1}m_\beta m_\beta'\right)\right]^{-1},$$

with v, m, m_β being obvious notational abbreviations.

The problem of comparing Σ_1 with Σ_2 can be reduced algebraically to the situation of the proposition with $\alpha = -1$ and thus to the case mentioned in Remark 2. It is thus seen that

$$\Sigma_1 \leqslant \Sigma_2$$

in the Loewner sense, i.e., $\widehat{\beta_1}$ is asymptotically more efficient than $\widehat{\beta_2}$.

This result also follows from a general theorem by Heyde (1997). However, Shklyar and Schneeweiss [4] can prove, for the Poisson model, the stronger result that $\Sigma_1 < \Sigma_2$ if $\Sigma_\delta\beta_1 \neq 0$.

3.2. Polynomial model

Kukush and Schneeweiss [2] and Kukush et al. [3] consider the polynomial model with measurement errors:

$$y = \beta_0 + \beta_1\xi + \cdots + \beta_k\xi^k + \epsilon,$$

$$x = \xi + \delta,$$

with the same assumptions on ξ and δ (scalar variables this time) as in the previous example and with $\epsilon \sim N(0, \sigma_\epsilon^2)$, ϵ being independent of ξ and δ.

One can again set up an estimating function like ψ_1 with a corresponding estimator of β and an asymptotic covariance matrix, again denoted by $\widehat{\beta_1}$ and Σ_1, respectively.

One can also estimate β via a completely different estimating function:

$$\psi_3(y, x, \beta) = H\beta - h,$$

where h is a vector with components $h_r = yt_r(x)$ and H a matrix with elements $H_{rs} = t_{r+s}(x)$, $r = 0, \ldots, k$, $s = 0, \ldots, k$, and $t_r(x)$ is a polynomial in x of degree r such that $E[t_r(x) \mid \xi] = \xi^r$. The corresponding estimator $\widehat{\beta_3}$ is consistent and has an asymptotic covariance matrix denoted by Σ_3.

It is difficult to compare the relative efficiencies of $\widehat{\beta_1}$ and $\widehat{\beta_3}$. However, in border line cases, when σ_δ^2 and σ_ϵ^2 both become small, a comparison is possible. Let $\sigma_\delta^2/\sigma_\epsilon^2 = \varkappa^2$ and let σ_δ^2 and σ_ϵ^2 go to zero such that $\varkappa > 0$ remains fixed, then,

$$\Sigma_1 = \sigma_\epsilon^2\left[E\left(v^{-1}\zeta\zeta'\right)\right]^{-1} + \mathrm{O}\left(\sigma_\epsilon^4\right),$$

$$\Sigma_3 = \sigma_\epsilon^2[E(\zeta\zeta')]^{-1}E(v\zeta\zeta')[E(\zeta\zeta')]^{-1} + \mathrm{O}\left(\sigma_\epsilon^4\right),$$

where $\zeta = (1, \xi, \ldots, \xi^k)'$ and $v = 1 + \varkappa^2\left(\frac{d\zeta'}{d\xi}\beta\right)^2$. Applying the proposition, one can see that

$$\lim_{\sigma_\epsilon \to 0} \sigma_\epsilon^{-2} \operatorname{tr} \Sigma_1 < \lim_{\sigma_\epsilon \to 0} \sigma_\epsilon^{-2} \operatorname{tr} \Sigma_2$$

and, by Remark 2, that

$$\lim_{\sigma_\epsilon \to 0} \sigma_\epsilon^{-2} \Sigma_1 \leqslant \lim_{\sigma_\epsilon \to 0} \sigma_\epsilon^{-2} \Sigma_2$$

in the Loewner sense.

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