A Bivariate Gamma Distribution Whose Marginals are Finite Mixtures of Gamma Distributions

Maryam Rafiei\textsuperscript{1}, Anis Iranmanesh\textsuperscript{1,}* , Daya K. Nagar \textsuperscript{2}

\textsuperscript{1}Department of Mathematics and Statistics, Mashhad Branch, Islamic Azad University, Mashhad, Iran
\textsuperscript{2}Instituto de Matemáticas, Universidad de Antioquia, Medellín, Colombia

Abstract
In this article a new bivariate distribution, whose both the marginals are finite mixtures of gamma distributions, has been defined. Several of its properties such as moments, correlation coefficients, measure of skewness, moment generating function, Renyi and Shannon entropies have been derived. Simulation study has been conducted to evaluate the performance of maximum likelihood method.

Keywords Bivariate Distribution, Beta Distribution, Entropy, Information Matrix, Gamma Distribution, Simulation

AMS 2010 subject classifications 62H10, 62E15

DOI: 10.19139/soic-2310-5070-1001

1. Introduction

The univariate gamma distribution is one of the most commonly used statistical distributions to analyze skewed data in many disciplines and has been studied extensively in scientific literature. The chi-square distribution, which is of utmost importance in statistical inference, is a special case of gamma distribution. Probability distributions such as exponential and Erlang are also special cases of the gamma distribution. Several univariate generalizations and variants of gamma distribution have also been developed and applied in various areas.

The univariate gamma distribution has been generalized to the bivariate case in many different ways and many forms of bivariate gamma distribution are available. Several techniques to generate bivariate distributions have also been proposed in the scientific literature, e.g., see Balakrishnan and Lai [1], Mardia [12], and Zhang and Singh [30].

Bivariate gamma distributions have found useful applications in many areas. They have been used for representing joint probabilistic properties of multivariate hydrological events such as floods and storms or in the modeling of rainfall at two nearby rain gauges, data obtained from rainmaking experiments, the dependence between annual streamflow and aerial precipitation, wind gust modeling (Smith and Adelfang [25], Smith, Adelfang, and Tubbs [26]), and the dependence between rainfall and runoff (see Nadarajah and Gupta [16], Nadarajah [14, 15] and references therein). For an interesting review of bivariate gamma distributions for hydrological application, the reader is referred to Yue, Ouarda, and Bobée [29] and Zhang and Singh [30].

Nadarajah [13] has listed a number of bivariate gamma distributions such as McKay’s bivariate gamma distribution, Dussauchoy and Berland’s bivariate gamma distribution, Cherians bivariate gamma distribution, Arnold and Strauss’ bivariate gamma distribution, Becker and Roux’s bivariate gamma distribution, and Smith and Adelfang’s bivariate gamma distribution.

\textsuperscript{*}Correspondence to: Anis Iranmanesh (Email: iranmanesh0030@mshdiau.ac.ir), Department of Mathematics and Statistics, Mashhad Branch, Islamic Azad University, Mashhad, Iran.
Chatelain and Touneret [3] proposed a family of bivariate gamma distributions whose marginals have different shape parameters and indicated its usefulness in detecting changes in two synthetic radar aperture (SAR) images acquired by different sensors and having different numbers of looks. Nadarajah [14] defined gamma-exponential distribution whose margins have the gamma and the exponential distributions. Nadarajah [15], by using two independent gamma variables, constructed a bivariate distribution which has gamma and beta distributions as its marginals. By using conditional approach (see Section 5.6 of Balakrishnan and Lai [1], Nagar, Nadarajah and Okorie [19]), Nagar, Zarrazola and Sánchez [18] constructed a bivariate distribution whose marginal laws are gamma and Macdonald. Piboongungon, Aalo, Iskander and Efthymoglou [20] derived the bivariate correlated generalised gamma fading distribution and have indicated its use in radar signal processing and communications.

The bivariate gamma distribution has also been defined as the joint distribution \( Z_1^2 \) and \( Z_2^2 \), where both \( Z_1 \) and \( Z_2 \) are standard normal variables with the correlation coefficient \( \rho \) (Vere-Jones [27], Maejima and Ueda [11]). Saboor and Ahemad [23] introduced a bivariate gamma-type density function of two variables involving a confluent hypergeometric function. Bondesson [2] reviewed some results for generalized gamma convolutions and derived new bivariate gamma distributions from shot-noise models.

For a review of known bivariate distributions, we refer the readers to Mardia [12], Kotz, Balakrishnan and Johnson [8], and Balakrishnan and Lai [1]. For an excellent review on univariate and bivariate gamma distributions the reader is referred to Saboor, Provost and Ahmad [22]. For matrix variate generalization of the gamma distribution one can consult Gupta and Nagar [6].

In this paper, we introduce a bivariate gamma distribution whose marginals are finite mixtures of gamma distributions and study its properties. This is the first bivariate distribution of its kind and is suitable for bivariate data with negative correlation. We organize our article as follows. In Section 2 we propose the bivariate gamma distribution and discuss some of its properties. Sections 3 and 4 deal with several results such as moments, correlation coefficients, measure of skewness, moment generating function, etc. Entropies such as Rényi and Shannon are derived in Section 5. Distributions of sum, quotient and product and many other distributional results are obtained in Section 6. Estimation of parameters and Fisher information matrix are discussed in Section 7. Finally, in the last section, simulation study is conducted to evaluate the performance of maximum likelihood method.

2. The bivariate gamma distribution

The random variables \( X_1 \) and \( X_2 \) are said to have a bivariate gamma distribution with parameters \( \alpha, \beta \) and \( k \), denoted by \( (X_1, X_2) \sim \text{BGa}(\alpha, \beta, k) \), if their joint density is given by

\[
f(x_1, x_2; \alpha, \beta, k) = C(\alpha, \beta, k)(x_1 x_2)^{\alpha-1}(x_1 + x_2)^k \exp \left[ -\frac{1}{\beta}(x_1 + x_2) \right],
\]

where \( x_1 > 0, x_2 > 0, \alpha > 0, \beta > 0, k \in \mathbb{N}_0 \) and \( C(\alpha, \beta, k) \) is the normalizing constant.

By integrating the joint density of \( X_1 \) and \( X_2 \) over its support set, the normalizing constant is derived as

\[
[C(\alpha, \beta, k)]^{-1} = \int_0^\infty \int_0^\infty (x_1 x_2)^{\alpha-1}(x_1 + x_2)^k \exp \left[ -\frac{1}{\beta}(x_1 + x_2) \right] dx_1 dx_2.
\]

Now, expanding \((x_1 + x_2)^k\) using binomial theorem and integrating \( x_1 \) and \( x_2 \), we obtain

\[
[C(\alpha, \beta, k)]^{-1} = \sum_{j=0}^k \binom{k}{j} \int_0^\infty \int_0^\infty x_1^{\alpha+j-1} x_2^{\alpha+k-j-1} \exp \left[ -\frac{1}{\beta}(x_1 + x_2) \right] dx_1 dx_2
\]

\[
= \beta^{2\alpha+k} \Gamma^2(\alpha) \sum_{j=0}^k \binom{k}{j} (\alpha)_j (\alpha)_{k-j}.
\]

Stat., Optim. Inf. Comput. Vol. 8, December 2020
Finally, using Lemma A.1, we get

\[ [C(\alpha, \beta, k)]^{-1} = \beta^{2\alpha + k} \Gamma^2(\alpha)(2\alpha)_k \]  

(3)

and

\[ C(\alpha, \beta, k) = \frac{\Gamma(\alpha\beta)}{\beta^{2\alpha + k} \Gamma^2(\alpha) \Gamma(2\alpha + k)}. \]  

(4)

An alternative way to compute (2) is to substitute \( s = x_1 + x_2 \) and \( r = x_1/(x_1 + x_2) \) and integrate \( s \) and \( r \) by using gamma and beta integrals. Since this approach works for all \( k > 0 \), we will use it to compute Shannon entropy.

Let us now briefly discuss the shape of (1). The first order derivatives of \( \ln f(x_1, x_2; \alpha, \beta, k) \) with respect to \( x_1 \) and \( x_2 \) are

\[ f_{x_1}(x_1, x_2) = \frac{\partial \ln f(x_1, x_2; \alpha, \beta, k)}{\partial x_1} = \alpha - 1 \frac{1}{x_1} + \frac{k}{x_1 + x_2} - \frac{1}{\beta} \]  

(5)

and

\[ f_{x_2}(x_1, x_2) = \frac{\partial \ln f(x_1, x_2; \alpha, \beta, k)}{\partial x_2} = \alpha - 1 \frac{1}{x_2} + \frac{k}{x_1 + x_2} - \frac{1}{\beta}, \]  

(6)

respectively. Setting (5) and (6) to zero, the only stationary point of (1) is obtained as

\[ a = x_{10} = x_{20} = \frac{\beta(2\alpha + k - 2)}{2}, \]

where \( 2\alpha + k - 2 > 0 \). Computing second order derivatives of \( \ln f(x_1, x_2; \alpha, \beta, k) \), from (5) and (6), we get

\[ f_{x_1x_1}(x_1, x_2) = \frac{\partial^2 \ln f(x_1, x_2; \alpha, \beta, k)}{\partial x_1^2} = - \frac{\alpha - 1}{x_1^2} - \frac{k}{(x_1 + x_2)^2}, \]  

(7)

\[ f_{x_1x_2}(x_1, x_2) = \frac{\partial^2 \ln f(x_1, x_2; \alpha, \beta, k)}{\partial x_1 \partial x_2} = - \frac{k}{(x_1 + x_2)^2}, \]  

(8)

and

\[ f_{x_2x_2}(x_1, x_2) = \frac{\partial^2 \ln f(x_1, x_2; \alpha, \beta, k)}{\partial x_2^2} = - \frac{\alpha - 1}{x_2^2} - \frac{k}{(x_1 + x_2)^2}. \]  

(9)

Further, from (7), (8) and (9), we get

\[ f_{x_1x_1}(a, a) = - \frac{4\alpha + k - 4}{(2\alpha + k - 2)^2 \beta^2}, \]

\[ f_{x_1x_2}(a, a) = - \frac{k}{(2\alpha + k - 2)^2 \beta^2}, \]

\[ f_{x_2x_2}(a, a) = - \frac{4\alpha + k - 4}{(2\alpha + k - 2)^2 \beta^2} \]

and finally

\[ f_{x_1x_1}(a, a)f_{x_2x_2}(a, a) - [f_{x_1x_2}(a, a)]^2 = \frac{8(\alpha - 1)}{(2\alpha + k - 2)^3 \beta^4}. \]

Now, observe that
• If $\alpha > 1$, then $f_{x_1x_2}(a,a)f_{x_2x_2}(a,a) - [f_{x_1x_2}(a,a)]^2 > 0$, $f_{x_1x_1}(a,a) < 0$ and $f_{x_2x_2}(a,a) < 0$ and therefore $(a,a)$ is a maximum point.
• If $0 < \alpha < 1$ and $2\alpha + k - 2 > 0$, then $f_{x_1x_1}(a,a)f_{x_2x_2}(a,a) - [f_{x_1x_2}(a,a)]^2 < 0$, and therefore $(a,a)$ is a saddle point.

Figure 1 illustrates the shape of the pdf (1) for selected values of $\alpha$ and $\beta$ and $k$.

It can easily be observed that $(X_1, X_2)$ and $(X_2, X_1)$ are identically distributed and hence $X_1$ and $X_2$ are exchangeable.

A distribution is said to be negatively likelihood ratio dependent if the density $f(x_1, x_2)$ satisfies

$$f(x_1, x_2)f(x_1^+, x_2^+) \leq f(x_1, x_2^+)f(x_1^+, x_2)$$

for all $x_1 > x_1^+$ and $x_2 > x_2^+$ (Lehmann [9], Tong [28]). One can check that the bivariate distribution defined by the density (1) is negatively likelihood ratio dependent.

By integrating $x_2$ in (1) the marginal density of $X_1$ is obtained as

$$f_{X_1}(x_1) = C(\alpha, \beta, k) \int_0^\infty (x_1x_2)^{\alpha-1}(x_1 + x_2)^k \exp \left[ -\frac{1}{\beta}(x_1 + x_2) \right] \, dx_2.$$  \hspace{1cm} (10)

Substituting $x_2/x_1 = z$ in (10), the marginal density of $X_1$ is rewritten as

$$f_{X_1}(x_1) = C(\alpha, \beta, k)x_1^{2\alpha - 1} \exp \left( -\frac{x_1}{\beta} \right) \int_0^\infty z^{\alpha-1}(1 + z)^k \exp \left( -\frac{x_1z}{\beta} \right) \, dz. \hspace{1cm} (11)$$

Now, Writing $(1 + z)^k$ using binomial theorem and integrating $z$ in (11), the marginal density of $X_1$ is derived as

$$f_{X_1}(x_1) = C(\alpha, \beta, k)x_1^{\alpha + k - 1} \beta^\alpha \exp \left( -\frac{x_1}{\beta} \right) \sum_{j=0}^k \binom{k}{j} (\alpha + j) \left( \frac{x_1}{\beta} \right)^{-j}$$

$$= C(\alpha, \beta, k)x_1^{\alpha - 1} \beta^{\alpha + k} \exp \left( -\frac{x_1}{\beta} \right) \sum_{j=0}^k \binom{k}{j} (\alpha + k - j) \left( \frac{x_1}{\beta} \right)^j. \hspace{1cm} (12)$$

Likewise, the marginal density of $X_2$ is obtained as

$$f_{X_2}(x_2) = C(\alpha, \beta, k)x_2^{\alpha + k - 1} \beta^\alpha \exp \left( -\frac{x_2}{\beta} \right) \sum_{j=0}^k \binom{k}{j} (\alpha + j) \left( \frac{x_2}{\beta} \right)^{-j}. \hspace{1cm} (13)$$

Thus, the marginal density of $X_1$ is a finite mixture of gamma densities. Figure 2 shows some plots of the marginal density of $X_1$ for $\beta = 2, k = 0, 1, \ldots, 20$ and some values of $\alpha$. Substituting $u = z/(1 + z)$ with $dz = (1 - u)^{-2}du$ in (11), one gets

$$f_{X_1}(x_1) = C(\alpha, \beta, k)x_1^{2\alpha + k - 1} \exp \left( -\frac{x_1}{\beta} \right)$$

$$\times \int_0^1 u^{\alpha-1}(1 - u)^{-(\alpha + k + 1)} \exp \left( -\frac{x_1u}{\beta(1 - u)} \right) \, du. \hspace{1cm} (14)$$

Now, writing

$$(1 - u)^{-(\alpha + k + 1)} \exp \left( -\frac{x_1u}{\beta(1 - u)} \right) = \sum_{j=0}^\infty u^j L_j^{(\alpha+k)} \left( \frac{x_1}{\beta} \right)$$
Figure 1. Pdf (1) with contour plots for some selected values of parameters.

in (14) and integrating $u$, the density $f_{X_1}(x_1)$, in series involving generalized Laguerre polynomials, is derived as

$$f_{X_1}(x_1) = C(\alpha, \beta, k)x_1^{2\alpha+k-1}\exp\left(-\frac{x_1^2}{\beta}\right)\sum_{j=0}^{\infty} \frac{1}{\alpha+j} L_j^{(\alpha+k)}\left(x_1^2\beta\right), \quad x_1 > 0.$$  (15)
where \( L_j^{(\alpha)}(\cdot) \) is the generalized Laguerre polynomial (see Appendix for the definition).

From the joint pdf (1) and the marginal density of \( X \) given in (12), the conditional pdf of \( X_2 \) given \( X_1 = x_1 \) is given by

\[
f(x_2 | x_1) = \frac{x_2^{\alpha-1}(x_1 + x_2)^k \exp(-x_2/\beta)}{\sum_{i=0}^{k} \binom{k}{i} \beta^\alpha \Gamma(\alpha + i) x_1^{i}}. \tag{16}\]

Also, the conditional pdf of \( X_1 \) given \( X_2 = x_2 \) is given by

\[
f(x_1 | x_2) = \frac{x_1^{\alpha-1}(x_1 + x_2)^k \exp(-x_1/\beta)}{\sum_{i=0}^{k} \binom{k}{i} \beta^\alpha \Gamma(\alpha + i) x_2^{i}}. \tag{17}\]

3. Moments

By definition

\[
E(X_1^m X_2^n) = C(\alpha, \beta, k) \int_0^\infty \int_0^\infty x_1^m x_2^n (x_1 x_2)^{\alpha-1} (x_1 + x_2)^k \exp \left[ -\frac{1}{\beta} (x_1 + x_2) \right] dx_1 dx_2.
\]

Substituting \( s = x_1 + x_2 \) and \( r = x_1 / (x_1 + x_2) \) with the Jacobian \( J(x_1, x_2 \rightarrow r, s) = s \) in the above integral, one gets

\[
E(X_1^m X_2^n) = C(\alpha, \beta, k) \int_0^1 r^{\alpha + m - 1} (1 - r)^{\alpha + n - 1} dr \int_0^\infty s^{2\alpha + m + n + k - 1} \exp \left( -\frac{s}{\beta} \right) ds \\
= C(\alpha, \beta, k) \frac{\Gamma(\alpha + m) \Gamma(\alpha + n)}{\Gamma(2\alpha + m + n)} \beta^{2\alpha + m + n + k} \Gamma(2\alpha + m + n + k),
\]

where the last line has been obtained by using beta and gamma integrals. Finally, simplifying the above expression, we get

\[
E(X_1^m X_2^n) = \beta^{m+n} \frac{\Gamma(\alpha + m) \Gamma(\alpha + n)(2\alpha + m + n)_k}{\Gamma^2(\alpha)(2\alpha)_k}.
\]

Further, substituting appropriately in the above expression, one gets

\[
E[(X_1 X_2)^h] = \beta^{2h} \frac{\Gamma^2(\alpha + h)(2\alpha + 2h)_k}{\Gamma^2(\alpha)(2\alpha)_k},
\]

\[
E(X_1 X_2) = \frac{\beta^2 \alpha(2\alpha + k)(2\alpha + k + 1)}{2(2\alpha + 1)},
\]

\[
E(X_1^2 X_2) = \frac{\beta^3 \alpha(2\alpha + k)(2\alpha + k + 1)(2\alpha + k + 2)}{4(2\alpha + 1)},
\]

\[
E(X_1^3 X_2) = \frac{\beta^4 \alpha(2\alpha + 2)(2\alpha + k)(2\alpha + k + 1)(2\alpha + k + 2)(2\alpha + k + 3)}{4(2\alpha + 1)(2\alpha + 3)},
\]

\[
E(X_1^4 X_2) = \frac{\beta^5 \alpha(2\alpha + 1)(2\alpha + k)(2\alpha + k + 1)(2\alpha + k + 2)(2\alpha + k + 3)}{4(2\alpha + 1)(2\alpha + 3)},
\]

\[
E(X_i) = \frac{\beta(2\alpha + k)}{2}, \quad i = 1, 2,
\]

Stat., Optim. Inf. Comput. Vol. 8, December 2020
A BIVARIATE GAMMA DISTRIBUTION

Figure 2. Plots of pdf (12) for some selected values of parameters.

\[
E(X_i^2) = \frac{\beta^2(\alpha + 1)(2\alpha + k)(2\alpha + k + 1)}{2(2\alpha + 1)}, \quad i = 1, 2,
\]

\[
E(X_i^3) = \frac{\beta^3(\alpha + 2)(2\alpha + k)(2\alpha + k + 1)(2\alpha + k + 2)}{4(2\alpha + 1)}, \quad i = 1, 2,
\]

Stat., Optim. Inf. Comput. Vol. 8, December 2020
and
\[ E(X_i^4) = \frac{\beta^4(\alpha + 2)(\alpha + 3)(\alpha + k)(2\alpha + k + 1)(2\alpha + k + 2)(2\alpha + k + 3)}{4(2\alpha + 1)(2\alpha + 3)}. \]

Further, variances, covariances, correlation and several higher central moments are derived as
\[ \mu_{11} = -k\beta^2(2\alpha + k) \]
\[ \mu_{20} = \mu_{02} = \frac{\beta^2(2\alpha + k)(4\alpha + k + 2)}{4(2\alpha + 1)}, \quad i = 1, 2, \]
\[ \mu_{30} = \mu_{03} = \frac{\beta^3(2\alpha + k)(8\alpha + 3k + 4)}{4(2\alpha + 1)}, \]
\[ \text{corr}(X_1, X_2) = -\frac{k}{4\alpha + k + 2}, \]
\[ \beta_{1i} = \sqrt{\frac{4(2\alpha + 1)(8\alpha + 3k + 4)^2}{(2\alpha + k)(4\alpha + k + 2)^3}}, \quad i = 1, 2, \]
\[ \mu_{21} = -\frac{\beta^3k(2\alpha + k)}{4(2\alpha + 1)}, \]
\[ \mu_{31} = -\frac{3\beta^4k(2\alpha + k)(2\alpha + k + 2)(4\alpha + k + 4)}{16(2\alpha + 1)(2\alpha + 3)}, \]
\[ \mu_{22} = \frac{(k + 2\alpha)[3k^3 + 2k^2(6 + 7\alpha) + 4k\alpha(11 + 8\alpha) + 8\alpha(2\alpha + 1)(2\alpha + 3)]\beta^4}{16(2\alpha + 1)(2\alpha + 3)}, \]

where
\[ \mu_{ij} = E[(X_1 - \mu)^i(X_2 - \mu)^j]. \]

4. Moment Generating Function

By definition, the joint mgf of \(X_1\) and \(X_2\) is given by
\[
M_{X_1, X_2}(t_1, t_2) = C(\alpha, \beta, k) \int_0^\infty \int_0^\infty (x_1x_2)^{\alpha-1}(x_1 + x_2)^k \
\exp \left[ t_1x_1 + t_2x_2 - \frac{1}{\beta}(x_1 + x_2) \right] \, dx_1 \, dx_2. \tag{18}
\]

Substituting \(x_1 = rs\) and \(x_2 = s(1 - r)\) in (18) with the Jacobian \(J(x_1, x_2 \rightarrow r, s) = s\) and integrating \(s\), we get
\[
M_{X_1, X_2}(t_1, t_2) = C(\alpha, \beta, k)\beta^{2\alpha + k} \Gamma(2\alpha + k) \
\times \int_0^1 \left[ r(1-r)^{\alpha-1}[r(1 - t_1\beta) + (1 - r)(1 - t_2\beta)]^{-(2\alpha + k)} \right] \, dr, \tag{19}
\]
where \(1 - t_1\beta > 0\) and \(1 - t_2\beta > 0\). Now, writing
\[
[r(1 - t_1\beta) + (1 - r)(1 - t_2\beta)]^{-(2\alpha + k)} = (1 - t_2\beta)^{-(2\alpha + k)} \left[ 1 - r \left(1 - \frac{1 - t_1\beta}{1 - t_2\beta} \right) \right]^{-(2\alpha + k)}, \quad \frac{1 - t_1\beta}{1 - t_2\beta} < 1,
\]
in (19) and integrating \( r \), we get

\[
M_{X_1,X_2}(t_1, t_2) = C(\alpha, \beta, k) \beta^{2\alpha+k}(1-t_2\beta)^{-2(2\alpha+k)} \Gamma(2\alpha + k) \\
\times \left[ \int_0^1 \left[ \frac{(1-r)r^{\alpha-1}}{r(1-r)(1 - \frac{1-t_1\beta}{1-t_2\beta})} \right]^{-(2\alpha+k)} dr \right].
\]

where the last line has been obtained by using the integral representation of the Gauss hypergeometric function given in (A.1). Finally, substituting for \( C(\alpha, \beta, k) \) and simplifying, we get

\[
M_{X_1,X_2}(t_1, t_2) = (1-t_2\beta)^{-2(2\alpha+k)} F\left(\alpha, 2\alpha + k; 2\alpha; 1 - \frac{1-t_1\beta}{1-t_2\beta}\right),
\]

which is the mgf of a gamma random variable with shape parameter \( 2\alpha + k \) and scale parameter \( \beta \).

5. Entropies

In this section, exact forms of Rényi and Shannon entropies are derived for the bivariate gamma distribution defined in this article.

Let \((\mathcal{X}, \mathcal{B}, \mathcal{P})\) be a probability space. Consider a pdf \( f \) associated with \( \mathcal{P} \), dominated by \( \sigma \)-finite measure \( \mu \) on \( \mathcal{X} \). Denote by \( H_{SH}(f) \) the well-known Shannon entropy introduced in Shannon [24]. It is defined by

\[
H_{SH}(f) = -\int_{\mathcal{X}} f(x) \log f(x) \, d\mu.
\]

One of the main extensions of the Shannon entropy was defined by Rényi [21]. This generalized entropy measure is given by

\[
H_R(\eta, f) = \frac{\log G(\eta)}{1-\eta} \quad \text{for } \eta > 0 \text{ and } \eta \neq 1,
\]

where

\[
G(\eta) = \int_{\mathcal{X}} f^n d\mu.
\]

The additional parameter \( \eta \) is used to describe complex behavior in probability models and the associated process under study. Rényi entropy is monotonically decreasing in \( \eta \), while Shannon entropy (21) is obtained from (22) for \( \eta \uparrow 1 \). For details see Nadarajah and Zografos [17], Zografos and Nadarajah [32] and Zografos [31].

**Theorem 5.1**

For the bivariate gamma distribution defined by the pdf (1), the Rényi and the Shannon entropies are given by

\[
H_R(\eta, f) = \frac{1}{1-\eta} \left[ \eta \ln C(\alpha, \beta, k) + \eta(2\alpha + k - 2) + 2 \ln \left( \frac{\beta}{\eta} \right) \right.
\]

\[
+ 2 \ln \Gamma[\eta(\alpha - 1) + 1] + \ln \Gamma[\eta(2\alpha + k - 2) + 2] - \ln \Gamma[\eta(2\alpha - 2) + 2] \right].
\]
and

\[ H_{SH}(f) = -\ln C(\alpha, \beta, k) - [(2\alpha + k - 2) \ln \beta (2\alpha + k) + 2(\alpha - 1)\psi(\alpha) + (2\alpha + k - 2)\psi(2\alpha + k) - (2\alpha - 2)\psi(2\alpha)], \]

respectively, where \( \psi(\alpha) = \Gamma'(\alpha)/\Gamma(\alpha) \) is the digamma function.

**Proof**

For \( \eta > 0 \) and \( \eta \neq 1 \), using the joint density of \( X_1 \) and \( X_2 \) given by (1), we have

\[
G(\eta) = \int_0^{\infty} \int_0^{\infty} f^\eta(x_1, x_2; \alpha, \beta, k) \, dx_2 \, dx_1
\]

\[
= [C(\alpha, \beta, k)]^\eta \int_0^{\infty} \int_0^{\infty} (x_1 x_2)^{\eta(\alpha - 1)} (x_1 + x_2)^{\eta k} \exp \left[ -\frac{\eta}{\beta} (x_1 + x_2) \right] \, dx_2 \, dx_1
\]

\[
= [C(\alpha, \beta, k)]^\eta \int_0^{\infty} \int_0^{1} [r(1 - r)]^{\eta(\alpha - 1)} s^{\eta(2\alpha - 2) + \eta k + 1} \exp \left( -\frac{\eta}{\beta} s \right) \, dr \, ds,
\]

where the last line has been obtained by substituting \( s = x_1 + x_2 \) and \( r = x_1/(x_1 + x_2) \). Finally, evaluating above integrals by using gamma and beta integrals and simplifying the resulting expression, we get

\[
G(\eta) = [C(\alpha, \beta, k)]^\eta \frac{\Gamma^2[\eta(\alpha - 1) + 1] \Gamma[\eta(2\alpha + k - 2) + 2]}{\Gamma[\eta(2\alpha - 2) + 2]} \left( \frac{\beta}{\eta} \right)^{\eta(2\alpha + k - 2) + 2}.
\]

Now, taking logarithm of \( G(\eta) \) and using (22) we get \( H_R(\eta, f) \). The Shannon entropy is obtained from \( H_R(\eta, f) \) by taking \( \eta \uparrow 1 \) and using L’Hôpital’s rule. \( \square \)

6. Sum, Quotient and Product

In this section we derive the distributions of \( X_1 + X_2, X_1/(X_1 + X_2), X_1 X_2, \) and \( X_1/X_2 \) when \( X_1 \) and \( X_2 \) follow a bivariate gamma distribution defined in (1).

**Theorem 6.1**

Let \( (X_1, X_2) \sim \text{BGa}(\alpha, \beta, k) \), and define \( R = X_1/(X_1 + X_2) \) and \( S = X_1 + X_2 \). Then, \( R \) and \( S \) are independent, the distribution of \( R \) is beta with both the parameters \( \alpha \) and the distribution of \( S \) is gamma with shape parameter \( 2\alpha + k \) and scale parameter \( \beta \).

**Proof**

Substituting \( x_1 = rs \) and \( x_2 = s(1 - r) \) with the Jacobian \( J(x_1, x_2 \rightarrow r, s) = s \), in the joint density of \( X_1 \) and \( X_2 \), we obtain the joint density of \( R \) and \( S \) as

\[
C(\alpha, \beta, k)[r(1 - r)]^{\alpha - 1} s^{2\alpha + k - 1} \exp \left( -\frac{s}{\beta} \right),
\]

where \( 0 < r < 1 \) and \( s > 0 \). Now, from (23), the desired result is obtained. \( \square \)

**Corollary 6.1.1**

Both \( X_1/X_2 \) and \( X_2/X_1 \) have inverted beta distribution with parameters \( \alpha \) and \( \alpha \).

**Theorem 6.2**

Let \( (X_1, X_2) \sim \text{BGa}(\alpha, \beta, k) \), and define \( P = X_1 X_2 \). Then, the density of \( P \) is given by

\[
C(\alpha, \beta, k) 2\beta^{\alpha + k/2 - 1} \sum_{j=0}^{k} \binom{k}{j} K_{k-2j} \left( 2 \sqrt{\frac{p}{\beta}} \right), \quad p > 0.
\]
Proof
Transforming $X_1 = X$ and $P = X_1X_2$ with the Jacobian $J(x_1, x_2 \rightarrow p) = 1/x$ in the joint density of $X_1$ and $X_2$ and integrating $x$, we obtain the density of $P$ as
\[
C(\alpha, \beta, k)p^{\alpha-1}\int_0^\infty \frac{1}{x} \left(x + \frac{p}{x}\right)^k \exp \left[-\frac{1}{\beta} \left(x + \frac{p}{x}\right)\right] \, dx
\]
\[
= C(\alpha, \beta, k)p^{\alpha-1}\sum_{j=0}^k \binom{k}{j} \frac{p^j}{j!} \int_0^\infty x^{k-2j-1} \exp \left[-\frac{1}{\beta} \left(x + \frac{p}{x}\right)\right] \, dx.
\]
(24)
Now, using the integral (Gradshteyn and Ryzhik [4, Eq. 3.471.9]),
\[
\int_0^\infty \exp \left(-az - \frac{b}{z}\right) z^{\nu-1} \, dz = 2 \left(\frac{b}{a}\right)^{\nu/2} K_{\nu}(2\sqrt{ab}), \quad a > 0, \quad b > 0,
\]
where $K_{\nu}$ is the modified Bessel function of the second kind, we obtain the desired result.

Next two theorems deal with bivariate distributions of $(X_1/Y, X_2/Y)$ and $(X_1/U, X_2/U)$, where $(X_1, X_2) \sim \text{BGa}(\alpha, \beta, k)$, $Y \sim \text{Ga}(\nu, \beta)$ and $U \sim \text{B}(a, b)$.

Theorem 6.3
Let $(X_1, X_2) \sim \text{BGa}(\alpha, \beta, k)$, and $Y \sim \text{Ga}(\nu, \beta)$ be independent. Then, the joint density of $Z_1 = X_1/Y$ and $Z_2 = X_2/Y$ is given by
\[
\frac{\Gamma(2\alpha)\Gamma(2\alpha+k+\nu)}{\Gamma^2(\alpha)\Gamma(2\alpha+k)\Gamma(\nu)} (z_1 + z_2)^{\alpha-1}(z_1 + z_2)^k y^{2\alpha+k+\nu-1} \exp \left[-\frac{(1 + z_1 + z_2)y}{\beta}\right], \quad z_1 > 0, \quad z_2 > 0.
\]

Proof
Transforming $X_1 = Z_1Y$ and $X_2 = Z_2Y$ with the Jacobian $J(x_1, x_2 \rightarrow z_1, z_2) = y^2$ in the joint density of $(X_1, X_2)$ and $Y$, the joint density of $(Z_1, Z_2)$ and $Y$ is obtained as
\[
\frac{C(\alpha, \beta, k)}{\Gamma(\nu)\beta^\nu} (z_1 z_2)^{\alpha-1}(z_1 + z_2)^k y^{2\alpha+k+\nu-1} \exp \left[-\frac{(1 + z_1 + z_2)y}{\beta}\right], \quad z_1 > 0, \quad z_2 > 0, \quad y > 0.
\]
Now, integrating $y$ by using gamma integral, we get the desired result.

For $k = 0$, the variables $X_1$, $X_2$ and $Y$ are independent gamma with scale parameter $\beta$ and therefore $(X_1/Y, X_2/Y)$ has a Dirichlet type 2 distribution.

Theorem 6.4
Let $(X_1, X_2) \sim \text{BGa}(\alpha, \beta, k)$, and $U \sim \text{B}(a, b)$ be independent. Then, the joint density of $Z_1 = X_1/U$ and $Z_2 = X_2/U$ is given by
\[
\frac{\beta^{-2\alpha+k}\Gamma(2\alpha)\Gamma(2\alpha+k+a)\Gamma(a+b)}{\Gamma^2(\alpha)\Gamma(2\alpha+k)\Gamma(2\alpha+k+a+b)\Gamma(a)} (z_1 z_2)^{\alpha-1}(z_1 + z_2)^k
\]
\[
\times \Phi \left(2\alpha + k + a; 2\alpha + k + a + b; -\frac{z_1 + z_2}{\beta}\right), \quad z_1 > 0, \quad z_2 > 0.
\]

Proof
Transforming $X_1 = Z_1U$ and $X_2 = Z_2U$ with the Jacobian $J(x_1, x_2 \rightarrow z_1, z_2) = u^2$ in the joint density of $(X_1, X_2)$ and $U$, the joint density of $(Z_1, Z_2)$ and $U$ is obtained as
\[
\frac{C(\alpha, \beta, k)}{B(a, b)} (z_1 z_2)^{\alpha-1}(z_1 + z_2)^k u^{2\alpha+k+a-1}(1 - u)^{b-1} \exp \left[-\frac{(z_1 + z_2)u}{\beta}\right],
\]
where $z_1 > 0$, $z_2 > 0$, and $0 < u < 1$. Now, integrating $u$ by using integral representation of confluent hypergeometric function given in (A.2), we get the desired result.
7. Estimation

Let \((X_{11}, X_{12}), \ldots, (X_{n1}, X_{n2})\) be a random sample from BGa\((\alpha, \beta, k)\). The log-likelihood function, denoted by \(l(\alpha, \beta)\), is given by

\[
l(\alpha, \beta) = n \left[ \ln \Gamma(2\alpha) - (2\alpha + k) \ln \beta - 2 \ln \Gamma(\alpha) - \ln \Gamma(2\alpha + k) \right] + (\alpha - 1) \sum_{i=1}^{n} (\ln x_{i1} + \ln x_{i2}) + k \sum_{i=1}^{n} \ln(x_{i1} + x_{i2}) - \frac{1}{\beta} \sum_{i=1}^{n} (x_{i1} + x_{i2}).
\]

Now, differentiating \(l(\alpha, \beta)\) w.r.t. \(\alpha\), we get

\[
\frac{\partial l(\alpha, \beta)}{\partial \alpha} = n \left[ 2\psi(2\alpha) - 2 \ln \beta - 2\psi(\alpha) - 2\psi(2\alpha + k) \right] + \sum_{i=1}^{n} (\ln x_{i1} + \ln x_{i2}).
\]

Using the duplication formula for digamma function, namely,

\[
2\psi(2z) = \ln 4 + \psi(z) + \psi \left( z + \frac{1}{2} \right)
\]

we obtain

\[
\frac{\partial l(\alpha, \beta)}{\partial \alpha} = n \left[ \ln 4 + \psi \left( \alpha + \frac{1}{2} \right) - \psi(\alpha) - 2 \ln \beta - 2\psi(2\alpha + k) \right] + \sum_{i=1}^{n} (\ln x_{i1} + \ln x_{i2}).
\]

Further,

\[
\frac{\partial l(\alpha, \beta)}{\partial \beta} = -\frac{n(2\alpha + k)}{\beta} + \frac{1}{\beta^2} \sum_{i=1}^{n} (x_{i1} + x_{i2}),
\]

\[
\frac{\partial l(\alpha, \beta)}{\partial \alpha \partial \beta} = -\frac{2n}{\beta},
\]

\[
\frac{\partial^2 l(\alpha, \beta)}{\partial \alpha^2} = n \psi_1 \left( \alpha + \frac{1}{2} \right) - n\psi_1(\alpha) - 4n\psi_1(2\alpha + k),
\]

\[
\frac{\partial^2 l(\alpha, \beta)}{\partial \beta^2} = \frac{n(2\alpha + k)}{\beta^2} - \frac{2}{\beta^3} \sum_{i=1}^{n} (x_{i1} + x_{i2}),
\]

\[
E \left[ \frac{\partial l(\alpha, \beta)}{\partial \alpha \partial \beta} \right] = -\frac{2n}{\beta},
\]

\[
E \left[ \frac{\partial^2 l(\alpha, \beta)}{\partial \alpha^2} \right] = n \psi_1 \left( \alpha + \frac{1}{2} \right) - n\psi_1(\alpha) - 4n\psi_1(2\alpha + k),
\]

\[
E \left[ \frac{\partial^2 l(\alpha, \beta)}{\partial \beta^2} \right] = -\frac{n(2\alpha + k)}{\beta^2}.
\]
For a given observation vector \((x_1, x_2)\), the Fisher information matrix for the bivariate distribution given by the density (1) is defined as

\[
\begin{pmatrix}
-\psi_1(\alpha + 1/2) + \psi_1(\alpha) + 4\psi_1(2\alpha + k) & 2/\beta \\
2/\beta & (2\alpha + k)/\beta^2
\end{pmatrix}.
\]

Further

\[
\frac{\partial l(\alpha, \beta)}{\partial \beta} = -\frac{n(2\alpha + k)}{\beta} + \frac{1}{\beta^2} \sum_{i=1}^{n} (x_{i1} + x_{i2}) = 0
\]
gives

\[(2\alpha + k)\beta = \bar{x}_1 + \bar{x}_2
\]
and

\[
\frac{\partial l(\alpha, \beta)}{\partial \alpha} = n [2\psi(2\alpha) - 2 \ln \beta - 2\psi(\alpha) - 2\psi(2\alpha + k)] + \sum_{i=1}^{n} (\ln x_{i1} + \ln x_{i2}) = 0
\]
gives

\[\psi(2\alpha + k) - \psi(2\alpha) + \ln \beta + \psi(\alpha) = \frac{1}{2} \ln(\bar{x}_1 \bar{x}_2),
\]
where \(\bar{x}_i = \prod_{j=1}^{n} x_{ij}^{1/n}, i = 1, 2\). Further, using

\[\psi(z + N) - \psi(z) = \sum_{j=0}^{N-1} \frac{1}{z + j}
\]
we have

\[\sum_{j=0}^{k-1} \frac{1}{2\alpha + j} + \ln \beta + \psi(\alpha) = \frac{1}{2} \ln(\bar{x}_1 \bar{x}_2).
\]

Thus, by solving numerically (25) and (26), the MLEs of \(\alpha\) and \(\beta\) can be obtained.

8. Simulation

In this section a simulation study is conducted to evaluate the performance of maximum likelihood method. Samples of size \(n = 30, 50, 200, 500\) from Equation (1) for selected values of parameters are generated by MCMC methods (Gibbs Metropolis, Markov Chain Monte Carlo Metropolis, Metropolis, Metropolis gaussian, random walk Metropolis and Metropolis-Hastings). For \(\alpha = 6, \beta = 2\) and \(k = 1, 4, 8\) that \(\rho = -\frac{1}{27}, \rho = -\frac{4}{7}\) and \(\rho = -\frac{8}{3}\) respectively, the random walk Metropolis algorithm method has better results. When \(\alpha = 0.75, \beta = 2, k = 1, 4, 8\) that \(\rho = -\frac{1}{6}, \rho = -\frac{2}{7}\) and \(\rho = -\frac{3}{8}\), the Gibbs sampling method provides better results and is used to simulate samples.

For each sample, MLEs for \(\alpha, \beta\) and \(k\) based on the numerical procedures are computed. This procedure is repeated five hundred times and \((\hat{\alpha}, \hat{\beta}, \hat{k})\), the bias (Ab) and the mean squared error (MSE) are obtained by using Monte Carlo method. The results are reported in Tables (1) and (2). Figures 3, 4 and 5 show the simulation data and contour plots for \(\alpha = 6, \beta = 2\) and \(k = 1, 4, 8\) with \(n = 200\). Figure 6 shows pairs style of random walk Metropolis method for \(\alpha = 6, \beta = 2\) and \(k = 1\) with \(n = 200\) and Figure 7 exhibits pairs style of Gibbs sampling method for \(\alpha = 0.75, \beta = 2\) and \(k = 4\) with \(n = 200\).
9. Multivariate generalization

The multivariate generalization of (1) can be defined as follows:

\[ C_n(\alpha, \beta, k)(x_1 x_2 \cdots x_n)^{\alpha-1}(x_1 + x_2 + \cdots + x_n)^k \exp \left[ -\frac{1}{\beta}(x_1 + x_2 + \cdots + x_n) \right], \tag{27} \]

where \( x_1 > 0, x_2 > 0, \ldots, x_n > 0 \) and \( C(\alpha, \beta, k) \) is the normalizing constant given by

\[ C_n(\alpha, \beta, k) = \frac{\Gamma(n\alpha)}{\beta^{n\alpha+k} \Gamma(\alpha) \Gamma(n\alpha+k)}. \tag{28} \]
Figure 4. Simulation data and contour plots for different methods for $\alpha = 6$, $\beta = 2$ and $k = 4$ with $n = 200$.

**Appendix**

The Gauss hypergeometric function, denoted by $F(a; b; c; z)$, and confluent hypergeometric function, denoted by $\Phi(b; c; z)$, for $\text{Re}(c) > \text{Re}(b) > 0$, are defined as (see Luke [10]),

\[
F(a; b; c; z) = \frac{1}{B(b, c - b)} \int_{0}^{1} t^{b-1}(1 - t)^{c-b-1} \frac{\exp(zt)}{(1-zt)^a} dt, \quad |\text{arg}(1-z)| < \pi, \tag{A.1}
\]

and

\[
\Phi(b; c; z) = \frac{1}{B(b, c - b)} \int_{0}^{1} t^{b-1}(1 - t)^{c-b-1} \exp(zt) dt. \tag{A.2}
\]
Figure 5. Simulation data and contour plots for different methods for $\alpha = 6$, $\beta = 2$ and $k = 8$ with $n = 200$.

Using the series expansion of $(1 - zt)^{-a}$ in (A.1) and $\exp(zt)$ in (A.2), the following series representations of the hypergeometric functions can be obtained:

$$F(a; b; c; z) = \sum_{k=0}^{\infty} \frac{(a)_k (b)_k}{(c)_k} \frac{z^k}{k!}, \quad |z| < 1, \quad (A.3)$$

and

$$\Phi(b; c; z) = \sum_{k=0}^{\infty} \frac{(b)_k}{(c)_k} \frac{z^k}{k!}, \quad (A.4)$$

where the Pochhammer symbol $(a)_n$ is defined by $(a)_n = a(a+1) \cdots (a+n-1) = (a)_{n-1}(a+n-1)$ for $n = 1, 2, \ldots$, and $(a)_0 = 1$. 

*Stat., Optim. Inf. Comput.* Vol. 8, December 2020
Also, under suitable conditions, we have (Luke [10, Eq. 3.6(10)]),

$$
\int_0^1 z^{\alpha-1}(1-z)^{\beta-1} F_p^q(a_1, \ldots, a_p; b_1, \ldots, b_q; zy) \, dz \\
= \frac{\Gamma(\alpha)\Gamma(\beta)}{\Gamma(\alpha + \beta)} F_{q+1}^{q+1}(a_1, \ldots, a_p, \alpha; b_1, \ldots, b_q, \alpha + \beta; y).
$$

(A.5)

**Lemma A.1**

For $a > 0, b > 0$ and $k \in \mathbb{N}$, we have

$$
\sum_{i=0}^{k} \binom{k}{i} (a)_i (b)_{k-i} = (a + b)_k.
$$
Figure 7. Pairs style of Gibbs sampling method for $\alpha = 0.75$, $\beta = 2$ and $k = 4$ with $n = 200$.

**Proof**

Writing $(1 - \theta)^{-(a+b)}$ as $(1 - \theta)^{-a}(1 - \theta)^{-b}$ and using power series expansion, for $0 < \theta < 1$, we get

\[
(1 - \theta)^{-a}(1 - \theta)^{-b} = \sum_{i=0}^{\infty} \sum_{j=0}^{\infty} \frac{(a)_i (b)_j}{i! j!} \theta^{i+j}
\]

\[
= \sum_{k=0}^{\infty} \sum_{i+j=k} \frac{(a)_i (b)_j}{i! j!}
\]

\[
= \sum_{k=0}^{\infty} \sum_{i=0}^{k} \frac{(a)_i (b)_{k-i}}{i! (k-i)!}
\]
and

\[(1 - \theta)^{-(\alpha+b)} = \sum_{k=0}^{\infty} \frac{(a+b)k}{k!} \theta^k.\]

Now, comparing the coefficients of \(\theta^k\), we get the desired result. \(\square\)

The generating function of the generalized Laguerre polynomial is

\[(1 - t)^{-(\alpha+1)} \exp \left(-\frac{zt}{1-t}\right) = \sum_{j=0}^{\infty} t^j L_j^{(\alpha)}(z).\]

Finally, we define the gamma, beta type 1 and beta type 2 distributions. These definitions can be found in Johnson, Kotz and Balakrishnan [7].

**Definition A.1**

A random variable \(X\) is said to have a gamma distribution with parameters \(\theta ( > 0), \kappa ( > 0)\), denoted by \(X \sim \Gamma(\kappa, \theta)\), if its pdf is given by

\[\{\theta^\kappa \Gamma(\kappa)\}^{-1} x^{\kappa-1} \exp \left(-\frac{x}{\theta}\right), \quad x > 0.\quad (A.6)\]

Note that for \(\theta = 1\), the above distribution reduces to a standard gamma distribution and in this case we write \(X \sim \Gamma(\kappa)\).

**Definition A.2**

A random variable \(X\) is said to have a beta type 1 distribution with parameters \((a, b)\), \(a > 0, b > 0\), denoted as \(X \sim B_1(a, b)\), if its pdf is given by

\[\{B(a, b)\}^{-1} x^{a-1}(1-x)^{b-1}, \quad 0 < x < 1,\quad (A.7)\]

where \(B(a, b)\) is the beta function.

**Definition A.3**

A random variable \(X\) is said to have a beta type 2 (inverted beta) distribution with parameters \((a, b)\), denoted as \(X \sim B_2(a, b)\), \(a > 0, b > 0\), if its pdf is given by

\[\{B(a, b)\}^{-1} x^{a-1}(1+x)^{-(a+b)}, \quad x > 0.\quad (A.8)\]

**Acknowledgement**

The authors are grateful to the editor and reviewer for their constructive and helpful comments and suggestions.

**REFERENCES**

1. N. Balakrishnan and Chin-Diew Lai, *Continuous bivariate distributions*, Second edition, Springer, Dordrecht, 2009.
2. Lennart Bondesson, *Univariate and bivariate generalized gamma convolutions*, Journal of Statistical Planning and Inference, vol. 139, no. 11, pp. 3759–3765, 2009.
3. F. Chatelain and J.-Y. Tourneret, *Bivariate gamma distributions for multisensor sar images*, 2007 IEEE International Conference on Acoustics, Speech and Signal Processing - ICASSP ’07, Honolulu, HI, pp. III-1237-III-1240, 2007.
4. I. S. Gradshteyn and I. M. Ryzhik, *Table of integrals, series, and products*, Translated from the Russian, Translation edited and with a preface by Daniel Zwillinger and Victor Moll, Eighth edition, Elsevier/Academic Press, Amsterdam, 2015.
5. Arjun K. Gupta and Saralees Nadarajah, *Sums, products and ratios for McKay’s bivariate gamma distribution*, Mathematical and Computer Modelling, vol. 43, no. 1–2, 185–193, 2006.
Table 1. MLE of simulation with $\alpha = 6$, $\beta = 2$ and $k = 1, 4, 8$

| $\alpha$ | $\beta$ | $k$ | $n$ | $\hat{\alpha}$ | $\hat{\beta}$ | $\hat{\alpha}$ | $\hat{\beta}$ | $MSE(\hat{\alpha})$ | $MSE(\hat{\beta})$ | corr |
|---|---|---|---|---|---|---|---|---|---|---|
| 6 | 2 | 1 | 30 | 8.040032 | 1.3521385 | 8.9860000 | 2.0490032 | -0.6478615 | 7.9860000 | 26.7930544 | 1.1103721 | 0.4534917 |
| 50 | 7.1348074 | 1.5613466 | 5.5640000 | 1.1348074 | -0.4386534 | 4.5640000 | 12.2150004 | 0.7738958 | 0.4534917 |
| 200 | 6.3116649 | 1.8737130 | 2.2340000 | 0.3116649 | -0.1262870 | 1.2340000 | 2.3459757 | 0.3360376 | 0.300168 |
| 500 | 6.12136864 | 2.01486647 | 1.1140000 | 0.12136864 | 0.01486647 | 0.1140000 | 1.00220167 | 0.18091314 | -0.2487341 |
| 6 | 2 | 4 | 30 | 8.547385 | 1.3830973 | 15.0960000 | 2.547385 | -0.6169027 | 11.0960000 | 40.5799509 | 1.088138 | 0.2069484 |
| 50 | 7.7117254 | 1.5676554 | 9.8386000 | 1.7117254 | -0.4323446 | 5.8386000 | 19.4659369 | 0.8266115 | -0.3564699 |
| 200 | 6.6526296 | 1.7460889 | 7.0020000 | 0.6526296 | -0.2539111 | 3.0020000 | 6.3696766 | 0.4541197 | 0.05095428 |
| 500 | 6.14778027 | 1.95136312 | 4.5280000 | 0.14778027 | -0.04863688 | 0.5280000 | 0.89259685 | 0.12620644 | 0.2078054 |
| 6 | 2 | 8 | 30 | 8.6204338 | 1.3615853 | 21.9680000 | 2.6204338 | -0.6384147 | 13.9680000 | 33.5312927 | 1.0200545 | -0.4879986 |
| 50 | 6.16577221 | 1.94406686 | 4.6820000 | 0.16577221 | -0.0593314 | 0.6820000 | 1.26158982 | 0.15624082 | -0.09814335 |
| 200 | 6.49986645 | 1.87237296 | 9.7100000 | 0.49986645 | -0.12762704 | 1.7100000 | 3.0631038 | 0.25328489 | 0.1303431 |
| 500 | 6.23188687 | 1.96264484 | 8.6040000 | 0.23188687 | -0.03735516 | 0.6040000 | 1.12598240 | 0.14648568 | -0.3217912 |
Table 2. MLE of simulation with $\alpha = 0.75$, $\beta = 2$ and $k = 1, 4, 8$

| $\alpha$ | $\beta$ | $k$ | $n$ | $\hat{\alpha}$ | $\hat{\beta}$ | $\hat{k}$ | $MSE(\hat{\alpha})$ | $MSE(\hat{\beta})$ | $MSE(\hat{k})$ | corr |
|----------|---------|-----|-----|-----------------|----------------|-------|--------------------|-----------------|-----------------|------|
| 0.75 2 1 | 30      | 0.8503846 | 3.1098761 | 3.14200000 | 0.1003846 | -0.6961239 | 2.14200000 | 0.1559368 | 0.9908157 | -0.8786252 |
| 50      | 0.79555454 | 1.54565705 | 1.95400000 | 0.04555454 | -0.45434295 | 0.95400000 | 0.05917405 | 0.72396201 | -0.09172372 |
| 200     | 0.76396678 | 1.91391820 | 1.21000000 | 0.01396678 | -0.08608180 | 0.21000000 | 0.01574066 | 0.39915125 | -0.3923851 |
| 500     | 0.754664660 | 1.960157607 | 1.09600000 | 0.004664660 | -0.039842393 | 0.09600000 | 0.005574255 | 0.162469098 | -0.3316668 |
| 0.75 2 4 | 30      | 1.08576145 | 0.99054874 | 14.6400000 | 0.33576145 | -1.1095126 | 10.6400000 | 1.29107613 | 1.46295908 | -0.6355327 |
| 50      | 0.93929384 | 1.25268403 | 9.74800000 | 0.18929384 | -0.74731597 | 5.74800000 | 0.36158134 | 1.03915630 | 0.2254778 |
| 200     | 0.78231536 | 1.76455788 | 5.21800000 | 0.03231536 | -0.23544212 | 1.21800000 | 0.04112525 | 0.45339322 | -0.1839379 |
| 500     | 0.757407912 | 1.909911596 | 4.50200000 | 0.007407912 | -0.090088404 | 0.50200000 | 0.014233611 | 0.227038802 | -0.4332585 |
| 0.75 2 8 | 30      | 0.757407912 | 1.909911596 | 4.50200000 | 0.007407912 | -0.090088404 | 0.50200000 | 0.014233611 | 0.227038802 | -0.4892968 |
| 50      | 0.90523165 | 1.55580000 | 12.7000000 | 0.15523165 | -0.44419893 | 4.70000000 | 0.25281686 | 0.76774585 | 0.3654835 |
| 200     | 0.78490199 | 1.86441825 | 9.21000000 | 0.03490199 | -0.13558175 | 1.21000000 | 0.02967985 | 0.29849273 | -0.5264573 |
| 500     | 0.76590872 | 1.93414852 | 8.55600000 | 0.01590872 | -0.06185148 | 0.55600000 | 0.01145940 | 0.14910234 | -0.4487656 |
6. A. K. Gupta and D. K. Nagar, *Matrix variate distributions*, Chapman & Hall/CRC, Boca Raton, 2000.
7. N. L. Johnson, S. Kotz and N. Balakrishnan, *Continuous univariate distributions*, vol. 2, Second Edition, John Wiley & Sons, New York, 1994.
8. S. Kotz, N. Balakrishnan and N.L. Johnson, *Continuous multivariate distributions*, vol. 1 , Second Edition, John Wiley & Sons, New York, 2000.
9. E. L. Lehmann, *Some concepts of dependence*, Annals of Mathematical Statistics, vol. 37, pp. 1137–1153, 1966.
10. Y. L. Luke, *The special functions and their approximations, Volume 1*, Academic Press, New York, 1969.
11. Makoto Maejima and Yohei Ueda, *A note on a bivariate gamma distribution*, Statistics & Probability Letters, vol. 80, no. 23–24, pp. 1991–1994, 2010.
12. K. V. Mardia, *Families of bivariate distributions*, Griffin’s Statistical Monographs and Courses, No. 27, Hafner Publishing Co., Darien, Conn., 1970.
13. Saralees Nadarajah, *Reliability for some bivariate gamma distributions*, Mathematical Problems in Engineering, vol. 2005, no. 2, pp. 151–163, 2005.
14. Saralees Nadarajah, *The bivariate gamma exponential distribution with application to drought data*, Journal of Applied Mathematics and Computing, vol. 24, no. 1-2, pp. 221–230, 2007.
15. Saralees Nadarajah, *A bivariate distribution with gamma and beta marginals with application to drought data*, Journal of Applied Statistics, vol. 36, no. 3-4, pp 277–30, 2009.
16. Saralees Nadarajah and Arjun K. Gupta, *Some bivariate gamma distributions*, Applied Mathematics Letters, vol. 19, no. 8, 767–774, 2006.
17. S. Nadarajah and K. Zografos, *Expressions for Rényi and Shannon entropies for bivariate distributions*, Information Sciences, vol. 170, no. 2-4, pp. 173–189, 2005.
18. Daya K. Nagar, Edwin Zarrazola and Luz Estela Sánchez, *A bivariate distribution whose marginal laws are gamma and Macdonald*, International Journal of Mathematical Analysis, vol. 10, no. 10, pp. 455–467, 2016.
19. Daya K. Nagar, S. Nadarajah and Idika E. Okorie, *A new bivariate distribution with one marginal defined on the unit interval*, Annals of Data Science, vol. 4, no. 3, pp. 405–420, 2017.
20. T. Piboongungon, V. Aalo, C. Iskander and G. Efthymoglou, *Bivariate generalised gamma distribution with arbitrary fading parameters*, Electronics Letters, vol. 41, pp. 709–710, 2005.
21. A. Rényi, *On measures of entropy and information*, in Proceedings of the Fourth Berkeley Symposium on Mathematical Statistics and Probability, Volume 1: Contributions to the Theory of Statistics, edited by J. Neyman, University of California Press, Berkeley, California, pp. 547–561, 1961.
22. Abdus Saboor, Serge B. Provost and Munir Ahmad, *Univariate and bivariate gamma-type distributions*, Lambert Academic Publishing, Saarbrucken, 2010.
23. Abdus Saboor and Munir Ahmad, *A bivariate gamma-type probability function using a confluent hypergeometric function of two variables*, Pakistan Journal of Statistics, vol. 28, no. 1, 81–91, 2012.
24. C. E. Shannon, *A mathematical theory of communication*, Bell System Technical Journal, vol. 27, pp. 379–423, 623–656, 1948.
25. O. E. Smith and S. I. Adelfangt, *Gust model based on the bivariate gamma probability distribution*, Journal of Spacecraft and Rockets, vol. 18, no. 6, pp. 545–549, 1981.
26. O. E. Smith, S. I. Adelfang, and J. D. Tubbs, *A bivariate gamma probability distribution with application to gust modeling*, Space Science Laboratory, George C. Marshall Space Flight Center, NASA TM-82483, July 1982.
27. D. Vere-Jones, *The infinite divisibility of a bivariate gamma distribution*, Sankhya Ser. A, vol. 29, 421–422, 1967.
28. Yung Liang Tong, *Probability inequalities in multivariate distributions*, Probabilities and Mathematical Statistics, Academic Press [Harcourt Brace Jovanovich, Publishers], New York-London-Toronto, Ont., 1980.
29. S.Yue, T. B. M. J. Ouarda, and B. Bobée, *A review of bivariate gamma distribution for hydrological application*, Journal of Hydrology, vol. 246, , no. 1-4, pp. 1–18, 2001.
30. L. Zhang and V. Singh, *Copulas and their applications in water resources engineering*, Cambridge: Cambridge University Press, 2019.
31. K. Zografos, *On maximum entropy characterization of Pearson’s type II and VII multivariate distributions*, Journal of Multivariate Analysis, vol. 71 , no. 1, pp. 67–75, 1999.
32. K. Zografos and S. Nadarajah, *Expressions for Rényi and Shannon entropies for multivariate distributions*, Statistics & Probability Letters, vol. 71 , no. 1, pp. 71–84, 2005.