The Marshall-Olkin-Gompertz-G family of distributions: properties and applications

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Abstract

We develop a new generalized family of the Gompertz-G distribution, namely, the Marshall-Olkin-Gompertz-G distribution. Statistical properties of the new proposed model are presented. Some special cases of the new family of distributions are presented. Maximum likelihood estimates of the model parameters are also determined. A simulation study was conducted to assess the performance of the maximum likelihood estimates. Applications to demonstrate the usefulness of the Marshall-Olkin-Gompertz-Weibull distribution to real data examples are provided.

Keywords: Gompertz-G distribution, Marshall-Olkin-G distribution, maximum likelihood estimation.

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1. Introduction

There is remarkable increase in the acceptability of generalized distributions over the past two decades. The motivation for the development of generalized distributions is their applicability to model heavy tailed data which is a prevalent phenomena in many lifetime analysis. Other areas of application of generalized distributions include finance, economics, hydrology and physics. Several attempts for generalizing classical distributions have been made. Earlier attempts include work by Gurvich et al. [22], Marshall and Olkin [26], Gupta and Kundu [21] and Eugene et al. [17].

Consequently, many generators for generalizing classical distributions have been proposed, these include gamma-G type I, type 2 and type 3 by Zografos and Balakrishnan [36], Ristić and Balakrishnan [30], Torabi and Hedesh [34], respectively, Weibull-G by Bourguignon et al. [6], Kumaraswamy-G (Kw-G) by Cordeiro et al. [14], type I half logistic-G family by Cordeiro et al. [12], odd log-logistic-G by Gleaton and Lynch [18], Gompertz-G by Alizadeh et al. [2], to mention a few.

Of interest in this paper are the generalized distributions by Alizadeh et al. [2], and Marshall and Olkin [26]. The Gompertz-G (Gom-G) distribution have cumulative distribution function (cdf) and probability density function (pdf) given by

\[ F_{\text{Gom-G}}(x; \theta, \gamma, \xi) = 1 - e^{\xi(1-[1-G(x;\xi)]^{-\gamma})} \] (1.1)

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and
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for $0 < u < 1$. Note that
\[
\log \left[ \frac{1 - u}{1 - u\delta} \right] = \frac{\theta}{\gamma} (1 - [1 - G(x; \xi)]^{-\gamma})
\]
so that
\[
G(x; \xi) = 1 - \left( 1 - \frac{\gamma}{\theta} \log \left[ \frac{1 - u}{1 - u\delta} \right] \right)^{-1/\gamma}.
\]
Therefore, the solution to the non-linear equation
\[
Q_X(u) = G^{-1} \left[ 1 - \left( 1 - \frac{\gamma}{\theta} \log \left[ \frac{1 - u}{1 - u\delta} \right] \right)^{-1/\gamma} \right]
\]
gives the quantile values of the MO-Gom-G family of distributions.

2.2. Stochastic orders

In this subsection, we present stochastic orders for the MO-Gom-G family of distributions. Suppose we have two random variables $X$ and $Z$ with distribution functions $F_X(r)$ and $F_Z(r)$, respectively, and $F_X(r) = 1 - F_X(z)$ the survival function, $X$ is stochastically smaller than $Z$ if $F_X(r) \leq F_Z(r)$ for all $r$. This is denoted by $X \leq_Z Z$. Hazard rate order and likelihood ratio order are stronger and are given by $X <_{hr} Z$ if $h_X(r) > h_Z(r)$ for all $r$, and $X <_{lr} Z$ if $h_X(r) / h_Z(r)$ is decreasing in $r$. From Shaked and Shanthikumar [31], we know that $X <_{lr} Z \Rightarrow X <_{hr} Z \Rightarrow X <_s Z$.

**Theorem 2.1.** Let $X_1 \sim MO - Gom - G(\theta, \gamma, \delta_1, \xi)$ and $X_2 \sim MO - Gom - G(\theta, \gamma, \delta_2, \xi)$. If $\delta_1 \leq \delta_2$, then
\[
f(x; \theta, \gamma, \delta_1, \xi) \quad \text{is decreasing in} \quad f(x; \theta, \gamma, \delta_2, \xi)
\]

**Proof.** Note that
\[
\frac{f(x; \theta, \gamma, \delta_1, \xi)}{f(x; \theta, \gamma, \delta_2, \xi)} = \frac{\delta_1}{\delta_2} \left( \frac{1 - \delta_1 [e^{\frac{\theta}{\gamma} (1 - [1 - G(x; \xi)]^{-\gamma})}]^3}{1 - \delta_2 [e^{\frac{\theta}{\gamma} (1 - [1 - G(x; \xi)]^{-\gamma})}]^3} \right),
\]
so that
\[
\frac{\partial}{\partial x} \left( \frac{f(x; \theta, \gamma, \delta_1, \xi)}{f(x; \theta, \gamma, \delta_2, \xi)} \right) = \frac{\delta_1}{\delta_2} (\delta_1 - \delta_2) \frac{1 - \delta_1 [e^{\frac{\theta}{\gamma} (1 - [1 - G(x; \xi)]^{-\gamma})}]^3}{1 - \delta_2 [e^{\frac{\theta}{\gamma} (1 - [1 - G(x; \xi)]^{-\gamma})}]^3}
\]
\[
\times \theta g(x; \xi) [1 - G(x; \xi)]^{-\gamma - 1} e^{\frac{\theta}{\gamma} (1 - [1 - G(x; \xi)]^{-\gamma})},
\]
which is $\leq 0$ if $\delta_1 \leq \delta_2$. Therefore, $X_1 <_{lr} X_2$, $X_1 <_{hr} X_2$ and $X_1 <_s X_2$. Hence, the random variables $X_1$ and $X_2$ are stochastically ordered.

2.3. Linear representation

We derive the linear representation of the MO-Gom-G family of distributions in this subsection. We make use of the general results by Barreto-Souza et al. [5]. Considering
\[
f_{MO-Gom-G}(x; \theta, \gamma, \delta, \xi) = \frac{\delta f_{Gom-G}(x; \theta, \gamma, \xi)}{(1 - \delta f_{Gom-G}(x; \theta, \gamma, \xi))^2},
\]
we can write equation (2.1) as
\[
f_{MO-Gom-G}(x; \theta, \gamma, \delta, \xi) = \frac{f_{Gom-G}(x; \theta, \gamma, \xi)}{\delta [1 - \delta f_{Gom-G}(x; \theta, \gamma, \xi)]^2}.
\]
where \( f_{\text{Gom-G}}(x; \theta, \gamma, \xi) \) and \( F_{\text{Gom-G}}(x; \theta, \gamma, \xi) \) are as given in equations (1.2) and (1.1), respectively. We also apply the series expansion
\[
(1 - z)^{-k} = \sum_{j=0}^{\infty} \frac{\Gamma(k+j)}{\Gamma(k)!} z^j,
\]
which is valid for \(|z| < 1\) and \(k > 0\). If \(\delta \in (0, 1)\), we obtain
\[
f_{\text{MO-Gom-G}}(x; \theta, \delta, \gamma, \xi) = f_{\text{Gom-G}}(x; \theta, \gamma, \xi) \sum_{j=0}^{\infty} \sum_{k=0}^{j} w_{j,k} F_{\text{Gom-G}}(x; \theta, \gamma, \xi)^{j-k},
\]
where \(w_{j,k} = w_{j,k}(\delta) = \delta(j+1)(1-\delta)^j(1-j)^{-k}(\frac{j}{k})\). For \(\delta > 1\), we have
\[
f_{\text{MO-Gom-G}}(x; \theta, \delta, \gamma, \xi) = f_{\text{Gom-G}}(x; \theta, \gamma, \xi) \sum_{j=0}^{\infty} v_{j} F_{\text{Gom-G}}(x; \theta, \gamma, \xi),
\]
where \(v_{j} = v_{j}(\delta) = \frac{(j+1)(1-1/\delta)}{\delta}\).

Thus, for \(\delta \in (0, 1)\), equation (2.1) becomes
\[
f_{\text{MO-Gom-G}}(x; \theta, \delta, \gamma, \xi) = \theta g(x; \bar{\xi})[1 - G(x; \bar{\xi})]^{-\gamma-1} e^{\frac{\delta}{\gamma}(1-[1-G(x;\bar{\xi})]^{-\gamma})} \\
\times \sum_{j=0}^{\infty} \sum_{k=0}^{j} w_{j,k}[1 - e^{\frac{\delta}{\gamma}(1-[1-G(x;\bar{\xi})]^{-\gamma})}]^{j-k}. 
\]

Using the following series expansions
\[
[1 - e^{\frac{\delta}{\gamma}(1-[1-G(x;\bar{\xi})]^{-\gamma})}]^{j-k} = \sum_{m=0}^{\infty} (-1)^m \binom{j-k}{m} e^{\frac{\delta}{\gamma}m(1-[1-G(x;\bar{\xi})]^{-\gamma})},
\]
\[
e^{\frac{\delta}{\gamma}(m+1)(1-[1-G(x;\bar{\xi})]^{-\gamma})} = \sum_{n=0}^{\infty} \frac{(\frac{\delta}{\gamma})^n (m+1)^n (1-[1-G(x;\bar{\xi})]^{-\gamma})^n}{n!},
\]
\[
(1 - [1 - G(x;\bar{\xi})]^{-\gamma})^n = \sum_{p=0}^{\infty} (-1)^p \binom{n}{p} [1 - G(x;\bar{\xi})]^{-\gamma p},
\]
\[
[1 - G(x;\bar{\xi})]^{-\gamma(p+1)-1} = \sum_{q=0}^{\infty} \left(-\gamma(p+1)-1\right)^q G^q(x;\bar{\xi}),
\]
we have for \(\delta \in (0, 1)\),
\[
f_{\text{MO-Gom-G}}(x; \theta, \gamma, \delta, \xi) = \sum_{j,m,n,p,q=0}^{\infty} \sum_{k=0}^{j} (-1)^{p+m} \frac{\gamma^{n+1}(m+1)^n}{\gamma^n n!} w_{j,k} \left(\frac{j-k}{m}\right) \binom{n}{p} \\
\times \left(-\gamma(p+1)-1\right)^q g(x; \bar{\xi}) G^q(x; \bar{\xi}) = \sum_{q=0}^{\infty} w_q^* g_q(x; \bar{\xi}).
\]

It follows that for \(\delta \in (0, 1)\), the MO-Gom-G family of distributions can be expressed as a linear combination of the exponentiated-G (Exp-G) distribution with power parameter \(q\) and linear component
\[
\sum_{j,m,n,p=0}^{\infty} \sum_{k=0}^{j} (-1)^{p+m} \frac{\gamma^{n+1}(m+1)^n}{(q+1)\gamma^n n!} w_{j,k} \left(\frac{j-k}{m}\right) \binom{n}{p} \left(-\gamma(p+1)-1\right)^q.
\]
and $g_q(x; \xi) = (q + 1)g(x; \xi)G^q(x; \xi)$.

Furthermore, for $\delta > 1$, equation (2.1) can be written as

$$f_{\text{MO-Gom-G}}(x; \theta, \gamma, \delta, \xi) = \theta g(x; \xi)[1 - G(x; \xi)^{-\gamma - 1}]e^\theta(1 - [1 - G(x; \xi)]^{-\gamma}) \sum_{j=0}^{\infty} v_j[1 - e^\theta(1 - [1 - G(x; \xi)]^{-\gamma})]^j.$$

Using the series expansion

$$[1 - e^\theta(1 - [1 - G(x; \xi)]^{-\gamma})]^j = \sum_{m=0}^{\infty} (-1)^m \frac{\theta^m(1 - [1 - G(x; \xi)]^{-\gamma})^m}{m!},$$

equations (2.3), (2.4), and (2.5), yields

$$f_{\text{MO-Gom-G}}(x; \theta, \gamma, \delta, \xi) = \sum_{j,m,n,p,q=0} (-1)^{m+p} \frac{\theta^{n+1}(m+1)^n}{(q+1)\gamma^n n!} v_j \binom{n}{p} \binom{m}{j} \binom{(n+q)}{p} (-\delta(p+1) - 1)^j,$$

(2.8)

Also, for $\delta > 1$, the MO-Gom-G family of distributions can be expressed as a linear combination of the Exp-G distribution with power parameter $q$ and linear component

$$v_q^* = \sum_{j,m,n,p=0}^{\infty} (-1)^{m+p} \frac{\theta^{n+1}(m+1)^n}{(q+1)\gamma^n n!} v_j \binom{n}{p} \binom{m}{j} \binom{(n+q)}{p} (-\delta(p+1) - 1)^j,$$

(2.9)

2.4. Moments and generating function

Let $X \sim \text{MO-Gom-G}(\theta, \gamma, \delta, \xi)$, then the $r$th moment can be obtained from equations (2.6) and (2.8), since the MO-Gom-G family of distributions can be expressed as a linear combination of Exp-G distribution.

- For $\delta \in (0, 1)$,
  $$E(X^r) = \sum_{q=0}^{\infty} w_q^* E(H_q^r),$$

where $w_q^*$ is as defined in equation (2.7) and $E(H_q^r)$ denotes the $r$th moment of $H_q$ which follows an Exp-G distribution with power parameter $q$.

- For $\delta > 1$
  $$E(X^r) = \sum_{q=0}^{\infty} v_q^* E(H_q^r),$$

where $v_q^*$ is as defined in equation (2.9) and $E(H_q^r)$ denotes the $r$th moment of $H_q$ which follows an Exp-G distribution with power parameter $q$. The incomplete moments can be obtained as follows:

- For $\delta \in (0, 1)$
  $$I_X(t) = \int_0^t x^r f_{\text{MO-Gom-G}}(x; \theta, \gamma, \delta, \xi) dx = \sum_{q=0}^{\infty} w_q^* I_q(t),$$

where $I_q(t) = \int_0^t x^r g_q(x; \xi) dx$ and $w_q^*$ is as defined in equation (2.7).

- For $\delta > 1$
  $$I_X(t) = \int_0^t x^r f_{\text{MO-Gom-G}}(x; \theta, \gamma, \delta, \xi) dx = \sum_{q=0}^{\infty} v_q^* I_q(t),$$
where \( I_q(t) = \int_0^1 x^q g_q(x; \xi) \, dx \) and \( v_q \) is as defined in equation (2.9). The moment generating function (mgf) of \( X \) is given by:

- For \( \delta \in (0, 1) \)
  \[
  M_X(t) = \sum_{q=0}^{\infty} w_q^* E(e^{tH_q}),
  \]
where \( E(e^{tH_q}) \) is the mgf of the Exp-G distribution with power parameter \( q \) and \( w_q^* \) is as defined in equation (2.7).

- For \( \delta > 1 \)
  \[
  M_X(t) = \sum_{q=0}^{\infty} v_q^* E(e^{tH_q}),
  \]
where \( E(e^{tH_q}) \) is the mgf of the Exp-G distribution with power parameter \( q \) and \( v_q^* \) is as defined in equation (2.9).

2.5. Entropy

Entropy is a measure of variation of uncertainty for a random variable \( X \) with pdf \( g(x) \). The two common measures of entropy are Shannon entropy [32] and Rényi entropy [29]. Shannon entropy is a special case of Rényi entropy. So in this paper we derive Rényi entropy of the MO-Gom-G family of distributions. By definition, Rényi entropy is given by

\[
I_R(\nu) = (1 - \nu)^{-1} \log \left[ \int_0^{\infty} g^\nu(x) \, dx \right],
\]
where \( \nu > 0 \) and \( \nu \neq 1 \). Using generalized binomial expansion (2.2), we get, for \( \delta \in (0, 1) \)

\[
f^\nu_{MO-Gom-G}(x; \theta, \delta, \gamma, \xi) = \frac{\delta^\nu f^\nu_{Gom-G}(x; \theta, \gamma, \xi)}{\Gamma(2\nu)} \sum_{j=0}^{\infty} (1 - \delta)^j \Gamma(2\nu + j) \frac{[1 - F_{Gom-G}(x; \theta, \gamma, \xi)]^j}{j!},
\]
and for \( \delta > 1 \)

\[
f^\nu_{MO-Gom-G}(x; \theta, \delta, \gamma, \xi) = \frac{\delta^\nu f^\nu_{Gom-G}(x; \theta, \gamma, \xi)}{\Gamma(2\nu)} \sum_{j=0}^{\infty} (\delta - 1)^j \Gamma(2\nu + j) \frac{F_{Gom-G}(x; \theta, \gamma, \xi)}{j!}.
\]
Thus, Rényi entropy for \( \delta \in (0, 1) \) and \( \delta > 1 \) are given by

\[
I_R(\nu) = (1 - \nu)^{-1} \log \left( \sum_{j=0}^{\infty} e_j \int_0^{\infty} f^\nu_{Gom-G}(x; \theta, \gamma, \xi)(1 - F_{Gom-G}(x; \theta, \gamma, \xi))^j \, dx \right)
\]
and

\[
I_R(\nu) = (1 - \nu)^{-1} \log \left( \sum_{j=0}^{\infty} h_j \int_0^{\infty} f^\nu_{Gom-G}(x; \theta, \gamma, \xi)F_{Gom-G}(x; \theta, \gamma, \xi) \, dx \right),
\]
where

\[
e_j = e_j(\delta) = \frac{\delta^\nu(1 - \delta)^j \Gamma(2\nu + j)}{\Gamma(2\nu)j!} \quad \text{and} \quad h_j = h_j(\delta) = \frac{(\delta - 1)^j \Gamma(2\nu + j)}{\delta^{\nu+1} \Gamma(2\nu)j!}.
\]
Now, for \( \delta \in (0, 1) \),

\[
f^\nu_{Gom-G}(x; \theta, \gamma, \xi)(1 - F_{Gom-G}(x; \theta, \gamma, \xi))^j = \theta^\nu g^\nu(x, \xi)[1 - G(x; \xi)]^{[-\gamma - 1] \nu} e_j^{(j+\nu)(1-[1-G(x; \xi)]=\nu)}.\]
Using the series expansion

\[ e^{\frac{\theta}{\gamma}}(1+\gamma)(1-[1-G(x;\xi)]^{-\gamma}) = \sum_{n=0}^{\infty} \frac{(\frac{\theta}{\gamma})^n(1 + \gamma)^n}{n!} (1 - [1 - G(x;\xi)]^{-\gamma})^n, \]

as well as equations (2.4) and (2.5), we get

\[ f^\nu_{G_{\theta\gamma\xi}}(x;\theta,\gamma,\xi)(1 - F_{G_{\theta\gamma\xi}}(x;\theta,\gamma,\xi))^j = \sum_{n,p,q=0}^{\infty} \frac{(-1)^p \theta^{\delta + \nu}(j + \nu)^n}{\gamma^n n!} \left(n\right) \left(-\delta(p + \nu) - \nu\right) q \ g(x;\xi) G^q(x;\xi). \]

Therefore, for \( \delta \in (0,1) \),

\[ I_R(\nu) = (1 - \nu)^{-1} \log \left\{ \sum_{j,n,p,q=0}^{\infty} \frac{(-1)^p \theta^{\delta + \nu}(j + \nu)^n}{\gamma^n n!} \left(n\right) \left(-\delta(p + \nu) - \nu\right) q \ g(x;\xi) G^q(x;\xi) \right\}, \]

for \( \nu > 0, \nu \neq 1 \), where \( I_{REG} = \frac{1}{1 - \nu} \log \left[ \int_0^\infty \left( \left[ \frac{\theta}{\gamma} + 1 \right] (G(x;\xi))^\delta g(x;\xi) \right)^\nu dx \right] \) is Rényi entropy of Exp-G distribution with power parameter \( \left( \frac{\theta}{\gamma} \right) \), and

\[ \eta_q = \sum_{j,n,p=0}^{\infty} \frac{(-1)^p \theta^{\delta + \nu}(j + \nu)^n}{\gamma^n n!} \left(n\right) \left(-\delta(p + \nu) - \nu\right) q \ \frac{1}{1 + \frac{\theta}{\gamma}}. \]

Also, for \( \delta > 1 \),

\[ f^\nu_{G_{\theta\gamma\xi}}(x;\theta,\gamma,\xi) F_{G_{\theta\gamma\xi}}(x;\theta,\gamma,\xi) = \theta^\nu g^\nu(x;\xi)[1 - G(x;\xi)]^{(-\gamma - 1)^\nu} \times e^{\theta(1-[1-G(x;\xi)]^{-\gamma})^\nu} [1 - e^{\theta(1-[1-G(x;\xi)]^{-\gamma})^\nu}]. \]

Using equations (2.3), the series expansion

\[ e^{\frac{\theta}{\gamma}}(m + \nu)(1 - [1 - G(x;\xi)]^{-\gamma})^\nu = \sum_{n=0}^{\infty} \left( \frac{\theta}{\gamma} \right)^n (1 - [1 - G(x;\xi)]^{-\gamma})^n, \]

as well as equations (2.4) and (2.5), we get

\[ f^\nu_{G_{\theta\gamma\xi}}(x;\theta,\gamma,\xi) F_{G_{\theta\gamma\xi}}(x;\theta,\gamma,\xi) = \sum_{m,n,p,q=0}^{\infty} \frac{(-1)^{m+p} \theta^{n+\gamma}(m + \nu)^n}{\gamma^n n!} \left(n\right) \left(m\right) \left(p\right) \times \left(-\gamma(p + \nu) - \nu\right) q \ g^\nu(x;\xi) G^q(x;\xi). \]

Hence, for \( \delta > 1 \),

\[ I_R(\nu) = (1 - \nu)^{-1} \log \left\{ \sum_{j,m,n,p,q=0}^{\infty} \frac{(-1)^{m+p} \theta^{n+\gamma}(m + \nu)^n}{\gamma^n n!} \left(n\right) \left(m\right) \left(p\right) \times \left(-\gamma(p + \nu) - \nu\right) \int_0^\infty g^\nu(x;\xi) G^q(x;\xi) dx \right\} = (1 - \nu)^{-1} \log \left[ \sum_{q=0}^{\infty} \eta_q e^{(1 - \nu) I_{REG}} \right], \]
for \( \nu > 0, \nu \neq 1 \), where \( I_{REG} = \frac{1}{1-\nu} \log \left[ \int_0^\infty \left( \frac{\nu}{\nu+1} \right) \gamma \left( G(x; \xi) \right) \gamma g(x; \xi) \right]^{\nu} dx \) is Rényi entropy of Exp-G distribution with power parameter \( \left( \frac{\nu}{\nu+1} \right) \), and

\[
\eta_q = \sum_{j,m,n,p=0}^\infty \frac{(-1)^m p^n \gamma (m + \nu)^n (j)}{\gamma n! j!} \left( \begin{array}{c} j \\ m \\ n \\ p \\ q \end{array} \right) \frac{1}{1 + \frac{q}{\nu}}.
\]

It follows that Rényi entropy of the MO-Gom-G family of distributions can be derived directly from Rényi entropy of the Exp-G distribution.

### 2.6. Distribution of order statistics

Suppose that \( X_1, X_2, \ldots, X_n \) are independent and identically distributed (i.i.d) random variables distributed according to equation (2.1), then the pdf of the \( i \)-th order statistic is given by

\[
f_{i:n}(x) = \delta_n f_{\text{Gom}-G}(x; \xi, \gamma) \sum_{j=0}^{n-i} \sum_{l=0}^{j} \sum_{k=0}^{l} \mathcal{U}_{j,l,k} f_{\text{Gom}-G}^{l-k+i-1}(x; \xi, \gamma),
\]

where

\[
\mathcal{U}_{j,l,k} = \mathcal{U}_{j,l,k}(\delta) = \frac{\delta n! (-1)^{l}(1-\delta)^{j}(-1)^{l-k}}{(i-1)! (n-i)!} \left( \begin{array}{c} j \\ l \\ k \\ i \\ j \end{array} \right).
\]

For \( \delta > 1 \),

\[
f_{i:n}(x) = \delta_n f_{\text{Gom}-G}(x; \xi, \gamma) \sum_{j=0}^{n-i} \sum_{l=0}^{j} \sum_{k=0}^{l} c_{j,l} f_{\text{Gom}-G}^{l-k+i-1}(x; \xi, \gamma),
\]

where

\[
c_{j,l} = c_{j,l}(\delta) = \frac{(-1)^{l} \delta^{-1} (1-\delta)^{j} n!}{\delta^{l+j+1} (i-1)! (n-i)!} \left( \begin{array}{c} l+j \\ j \end{array} \right).
\]

Therefore, for \( \delta \in (0, 1) \), we get

\[
f_{i:n} = \theta g(x; \xi, \gamma) \left[ 1 - G(x; \xi) \right]^{-\gamma - 1} e^{\theta \left( 1 - \left[ 1 - G(x; \xi) \right]^{-\gamma} \right)}
\times \sum_{j=0}^{n-i} \sum_{l=0}^{j} \sum_{k=0}^{l} \mathcal{U}_{j,l,k} \left( 1 - e^{\theta \left( 1 - \left[ 1 - G(x; \xi) \right]^{-\gamma} \right)} \right)^{j+l-k+i-1}.
\]

Using the generalized binomial series expansion

\[
(1 - [1 - G(x; \xi)]^{-\gamma})^{j+l-k+i-1} = \sum_{m=0}^{\infty} (-1)^m \left( \begin{array}{c} j+l-k+i-1 \\ m \end{array} \right) e^{\theta \left( 1 - [1 - G(x; \xi)]^{-\gamma} \right)}
\]

as well as equations (2.3), (2.4), and (2.5), we can write

\[
f_{i:n}(x) = \sum_{j,m,n,p,q=0}^{\infty} \sum_{l=0}^{n-i} \sum_{k=0}^{j} \frac{(-1)^{m+p} \gamma n + 1 (m + 1)^n}{\gamma n!} \mathcal{U}_{j,l,k} \left( \begin{array}{c} j+l-k+i-1 \\ m \end{array} \right) \left( \begin{array}{c} n \\ p \end{array} \right)
\times \left( \frac{-\gamma (p+1)}{q} \right) g(x; \xi) \left[ G(x; \xi) \right]^q = \sum_{q=0}^{\infty} U_q g_q (x; \xi),
\]
where \( g_q(x; \xi) = (q + 1)g(x; \xi)[G(x; \xi)]^q \) is an \( \text{Exp-G} \) distribution with power parameter \( q \) and

\[
U_q^* = \sum_{j,m,n,p=0}^{\infty} \sum_{k=0}^{n-i} \frac{(-1)^m+p\theta^{n+1}(m+1)^n}{(q+1)^{y^n n!}} U_{j,l,k}\left(\frac{j + 1 - k + i - 1}{m}\right)\left(\frac{n}{p}\right)\left(-\gamma(p + 1) - 1\right).
\]

Furthermore, for \( \delta > 1 \), we write

\[
f_{x:n}(x) = \theta g(x; \xi)[1 - G(x; \xi)]^{-\gamma - 1} e^{\frac{p}{\gamma}(1-[1-G(x;\xi)]^{-\gamma})} \sum_{j=0}^{\infty} \sum_{i=0}^{\infty} c_{j,i}(1 - e^{\frac{p}{\gamma}(1-[1-G(x;\xi)]^{-\gamma}))}) j + i - 1.
\]

By applying the generalized binomial series expansion

\[
(1 - [1 - G(x; \xi)]^{-\gamma}) j + i - 1 = \sum_{m=0}^{\infty} (-1)^m \left(\begin{array}{c} j + 1 + i - 1 \\ m \end{array}\right) e^{-\gamma \gamma_m(1-[1-G(x;\xi)]^{-\gamma}),}
\]

as well as equations (2.3), (2.4), and (2.5), we get

\[
f_{x:n}(x) = \sum_{j,m,n,p,q=0}^{\infty} \sum_{i=0}^{n-i} \frac{(-1)^m+p\theta^{n+1}(m+1)^n}{(q+1)^{y^n n!}} C_{j,i}\left(\frac{j + 1 + i - 1}{m}\right)\left(\frac{n}{p}\right)\left(-\gamma(p + 1) - 1\right)
\]

where \( g_q(x; \xi) = (q + 1)g(x; \xi)[G(x; \xi)]^q \) is an \( \text{Exp-G} \) distribution with power parameter \( q \) and

\[
C_q^* = \sum_{j,m,n,p=0}^{\infty} \sum_{i=0}^{n-i} \frac{(-1)^m+p\theta^{n+1}(m+1)^n}{(q+1)^{y^n n!}} C_{j,i}\left(\frac{j + 1 + i - 1}{m}\right)\left(\frac{n}{p}\right)\left(-\gamma(p + 1) - 1\right).
\]

3. Estimation

Let \( X_i \sim \text{MO – Gom – G} (\gamma, \delta, \theta, \xi) \) with the parameter vector \( \Psi = (\gamma, \delta, \theta, \xi)^T \), then the log-likelihood function \( \ell = \ell(\Psi) \) from a random sample of size \( n \) is given by

\[
\ell(\Psi) = n \log(2\delta) + \sum_{i=1}^{n} \log[g(x_i; \xi)] - (\gamma + 1) \sum_{i=1}^{n} \log[1 - G(x_i; \xi)]
\]

\[
+ \frac{\theta}{\gamma} \sum_{i=1}^{n} \left(1 - [1 - G(x_i; \xi)]^{-\gamma}\right) - 2 \sum_{i=1}^{n} \log[1 - \delta e^{\frac{p}{\gamma}(1-[1-G(x;\xi)]^{-\gamma})}].
\]

The elements of the score vector \( \mathbf{U}(\Psi) \) are given by

\[
\frac{\partial \ell}{\partial \delta} = \frac{n}{\delta} - 2 \sum_{i=1}^{n} \frac{e^{\frac{p}{\gamma}(1-[1-G(x_i;\xi)]^{-\gamma})}}{1 - \delta e^{\frac{p}{\gamma}(1-[1-G(x_i;\xi)]^{-\gamma})}},
\]

\[
\frac{\partial \ell}{\partial \theta} = \frac{n}{\gamma} + \frac{1}{\gamma} \sum_{i=1}^{n} \left(1 - [1 - G(x_i; \xi)]^{-\gamma}\right) + 2 \sum_{i=1}^{n} \frac{\delta(1 - [1 - G(x_i; \xi)]^{-\gamma}) e^{\frac{p}{\gamma}(1-[1-G(x;\xi)]^{-\gamma})}}{\gamma(1 - \delta e^{\frac{p}{\gamma}(1-[1-G(x_i;\xi)]^{-\gamma})}),
\]

\[
\frac{\partial \ell}{\partial \gamma} = \sum_{i=1}^{n} \frac{\theta(1 - G(x_i; \xi))^{-\gamma} \log[1 - G(x_i; \xi)] - (1 - 1 - G(x_i; \xi))^{-\gamma})}{\gamma^2}.
\]
where \( \Psi \) is a vector of \( p \) components, \( J_{\Psi} \) is a \( p \times p \) matrix, \( J_{\gamma\xi}(\Psi) \) is a \( p \times 1 \) matrix, and \( J_{\delta\xi}(\Psi) \) has \( p \times 1 \) components, respectively. Under the usual regularity conditions \( \Psi \) is asymptotically normal distributed, that is \( \Psi \sim N(0, I^{-1}(\Psi)) \) as \( n \to \infty \), where \( I(\Psi) \) is the expected information matrix. The asymptotic behavior remains valid if \( I(\Psi) \) is replaced by \( J(\Psi) \), the information matrix evaluated at \( \Psi \).

4. Some special cases

In this section, we present some special cases for the MO-Gom-G family of distributions by considering the Weibull, Kumaraswamy and Burr XII distributions as baseline distributions.


If we take the Weibull distribution as the baseline distribution with pdf and cdf given by \( g(x; \lambda) = \lambda x^{\lambda-1}e^{-x^\lambda} \) and \( G(x; \lambda) = 1 - e^{-x^\lambda} \), respectively, for \( \lambda > 0 \), we obtain the MO-Gom-W distribution with cdf, pdf and hrf given by

\[
F(x; \theta, \gamma, \delta, \lambda) = \frac{1 - e^{\frac{\theta}{\delta}(1-|e^{-x^\lambda}|^{-\gamma})}}{1 - \frac{\theta}{\delta}(1-|e^{-x^\lambda}|^{-\gamma})},
\]

\[
f(x; \theta, \gamma, \delta, \lambda) = \frac{\delta \lambda x^{\lambda-1}e^{-x^\lambda}[e^{-x^\lambda}]^{-\gamma} - e^{\frac{\theta}{\delta}(1-|e^{-x^\lambda}|^{-\gamma})}}{(1 - \frac{\theta}{\delta}(1-|e^{-x^\lambda}|^{-\gamma}))^2},
\]

\[
h(x; \theta, \gamma, \delta, \lambda) = \frac{\delta \lambda x^{\lambda-1}e^{-x^\lambda}[e^{-x^\lambda}]^{-\gamma} - e^{\frac{\theta}{\delta}(1-|e^{-x^\lambda}|^{-\gamma})}}{(1 - \frac{\theta}{\delta}(1-|e^{-x^\lambda}|^{-\gamma}))^2} \left[ 1 - \frac{1 - e^{\frac{\theta}{\delta}(1-|e^{-x^\lambda}|^{-\gamma})}}{1 - \frac{\theta}{\delta}(1-|e^{-x^\lambda}|^{-\gamma})} \right]^{-1},
\]

respectively, for \( \theta, \gamma, \delta \) and \( \lambda > 0 \).

Figure 1 shows plots of the pdfs and hrfs for the MO-Gom-W distribution. The pdf apply to heavy tailed data and with varying kurtosis. The hrf takes decreasing, increasing and bathtub shapes.

By taking the baseline distribution to be a Kumaraswamy distribution with pdf and cdf given by

\[ g(x; a, b) = abx^{a-1}(1-x^a)^{b-1} \]

and

\[ G(x; a, b) = 1 - (1-x^a)^b, \]

respectively, for \( a, b > 0 \), we obtain the MO-Gom-Kw distribution with cdf, pdf and hrf given by

\[
F_{\text{MO-Gom-G}}(x; \theta, \gamma, \delta, a, b) = 1 - \frac{1 - e^{\theta \gamma (1 - (1-x^a)^b - \gamma)}}{1 - \delta(e^{\gamma(1 - (1-x^a)^b - \gamma)})},
\]

\[
f_{\text{MO-Gom-G}}(x; \theta, \gamma, \delta, a, b) = \frac{\delta \theta abx^{a-1}(1-x^a)^{b-1}((1-x^a)^b - \gamma - 1)e^{\theta \gamma (1 - (1-x^a)^b - \gamma)}}{(1 - \delta(e^{\gamma (1 - (1-x^a)^b - \gamma)})^2},
\]

\[
h(x; \theta, \gamma, \delta, a, b) = \frac{\delta \theta abx^{a-1}(1-x^a)^{b-1}((1-x^a)^b - \gamma - 1)e^{\theta \gamma (1 - (1-x^a)^b - \gamma)}}{(1 - \delta(e^{\gamma (1 - (1-x^a)^b - \gamma)})^2}
\times \left[ 1 - \frac{1 - e^{\theta \gamma (1 - (1-x^a)^b - \gamma)}}{1 - \delta(e^{\gamma(1 - (1-x^a)^b - \gamma)})} \right]^{-1},
\]

respectively, for \( \theta, \gamma, \delta, a, b > 0 \).
or right-skewed data sets. The hrf exhibits increasing, bathtub and upside bathtub followed by bathtub shapes.

4.3. Marshall-Olkin-Gompertz-Burr XII (MO-Gom-BXII) distribution

If we consider the baseline distribution to be Burr XII distribution with pdf and cdf given by $g(x; c, k) = ckx^{c-1}(1 + x^{c})^{-k-1}$ and $G(x; c, k) = 1 - (1 + x^{c})^{-k}$, respectively, we obtain the MO-Gom-BXII distribution with cdf, pdf, and hrf given by

$$
F(x; \theta, \gamma, \delta, c, k) = \frac{1 - e^{\theta kx - [(1 + x^{c})^{-k} - \gamma]}}{1 - \delta e^{\theta kx - [(1 + x^{c})^{-k} - \gamma]}},
$$

$$
f(x; \theta, \gamma, \delta, c, k) = \frac{\delta \theta ckx^{c-1}(1 + x^{c})^{-k} - 1}{(1 - \delta e^{\theta kx - [(1 + x^{c})^{-k} - \gamma]})^2},
$$

$$
h(x; \theta, \gamma, \delta, c, k) = \frac{\delta \theta ckx^{c-1}(1 + x^{c})^{-k-1}[(1 + x^{c})^{-k} - \gamma] - 1}{(1 - \delta e^{\theta kx - [(1 + x^{c})^{-k} - \gamma]})^2}
\times \left[1 - \frac{1 - e^{\theta kx - [(1 + x^{c})^{-k} - \gamma]}}{1 - \delta e^{\theta kx - [(1 + x^{c})^{-k} - \gamma]}}\right]^{-1},
$$

respectively, for $\theta, \gamma, \delta, c, k > 0$. We obtain the MO-Gom-Log-Logistic (MO-Gom-LLoG) and MO-Gom-Lomax (MO-Gom-Lx) from the MO-Gom-BXII distribution by setting $k = 1$ and $c = 1$, respectively.

![Figure 3: Plots of the pdf and hrf for the MO-Gom-BXII distribution.](image)

Pdfs and hrfs plots for the MO-Gom-BXII distribution are shown in Figure 3. The pdf applies to heavy tailed data sets. The hrf exhibits decreasing, increasing, bathtub and upside bathtub shapes.

5. Simulation study

In this section, we present results of simulation study for the MO-Gom-W distribution. We fix the parameter $\theta = 1$ in the following sets of parameter values (I: $\gamma = 0.05, \delta = 1.0, \lambda = 0.5$), (II: $\gamma = 1.0, \delta = 0.5, \lambda = 0.01$), (III: $\gamma = 0.01, \delta = 1.1, \lambda = 0.5$), (IV: $\gamma = 0.01, \delta = 1.0, \lambda = 0.01$), (V: $\gamma = 1.0, \delta = 0.5, \lambda = 0.01$), and (VI: $\gamma = 0.01, \delta = 1.1, \lambda = 1.1$). We consider sample sizes $n= 25, 50, 100, 200, 400$ and 800, for $N=1000$ for each sample. We estimate the mean, average bias and root mean square error (RMSE). The bias and RMSE for the estimated parameter, say, $\hat{\lambda}$, say, are given by:

$$
\text{Bias}(\hat{\lambda}) = \frac{1}{N} \sum_{i=1}^{N} (\hat{\lambda}_i - \Delta), \quad \text{and} \quad \text{RMSE}(\hat{\lambda}) = \sqrt{\frac{1}{N} \sum_{i=1}^{N} (\hat{\lambda}_i - \Delta)^2},
$$

where $\Delta$ is the true parameter value.
respectively. From the results in Table 1, the mean values approximate the true parameter values, RMSE and bias decay towards zero for all the parameter values. We therefore, conclude that our model give consistent maximum likelihood estimates (MLEs).

Table 1: Monte Carlo simulation results for MO-Gom-W distribution: mean, RMSE, and average bias

<table>
<thead>
<tr>
<th>n</th>
<th>Mean RMS Bias</th>
<th>Mean RMS Bias</th>
<th>Mean RMS Bias</th>
</tr>
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<tr>
<td>25</td>
<td>0.089236</td>
<td>0.035689</td>
<td>0.039232</td>
</tr>
<tr>
<td>50</td>
<td>0.085115</td>
<td>0.036042</td>
<td>0.015115</td>
</tr>
<tr>
<td>100</td>
<td>0.084317</td>
<td>0.062569</td>
<td>0.006417</td>
</tr>
<tr>
<td>200</td>
<td>0.059659</td>
<td>0.015359</td>
<td>0.006569</td>
</tr>
<tr>
<td>400</td>
<td>0.050102</td>
<td>0.009864</td>
<td>0.001052</td>
</tr>
<tr>
<td>800</td>
<td>0.050612</td>
<td>0.007112</td>
<td>0.000612</td>
</tr>
</tbody>
</table>

6. Applications

In this section, we present two real data examples to demonstrate the flexibility of the MO-Gom-W distribution in data fitting. The maximum likelihood estimation technique is used to estimate the model parameters via the nlm package in R software [28]. The AdequacyModel package in R software [28] is used to assess model performance. We consider the following goodness-of-fit statistics: Crámer-von-Mises ($W^*$) and Anderson-Darling ($A^*$), as well as -2loglikelihood (-2 log L), Akaike Information Criterion (AIC), Consistent Akaike Information Criterion (AICC), Bayesian Information Criterion (BIC), Kolmogorov-Smirnov (K-S) statistic (and it’s p-value), and sum of squares (SS). The model with the smallest values of the goodness-of-fit statistics and a bigger p-value for the K-S statistic is regarded as the best model.

We present data analysis results in Tables 2 and 3, parameter estimates (standard errors in parentheses) and the goodness-of-fit statistics for the various models considered in this paper. We also provide fitted densities and probability plots (as described by Chambers et al. [7]) to demonstrate the flexibility of the MO-Gom-W distribution in data fitting compared to the selected competing models (see Figures 4 and 5, for details).

We compare the MO-Gom-W distribution to the following models: Gompertz distribution by Gompertz [19], generalized Gompertz (G-Gom) distribution by El-Gohary at ali. [16], Marshall-Olkin log-logistic (MO-LLog) distribution by Wenhao [20], Marshall-Olkin extended inverse Weibull (MO-IW) by
for $\alpha, \beta, \lambda > 0$,

$$f_{G-Gom}(x; \alpha, \beta, \lambda) = \alpha \beta e^{\lambda x} (1 - e^{-\lambda x})^{-\alpha - 1},$$

for $\alpha, \gamma, \lambda > 0$,

$$f_{MO-EW}(x; \alpha, \gamma, \lambda) = \frac{\alpha \gamma \lambda^\gamma (e^{-\gamma x} - e^{-\lambda x})}{(1 - e^{-\lambda x})^2},$$

for $\alpha, \theta, \lambda > 0$,

$$f_{MO-LLog}(x; \alpha, \beta, \gamma) = \frac{\alpha \beta \gamma x^{\beta - 1}}{(x^\beta + \gamma)^2},$$

for $\alpha, \beta, \gamma > 0$,

$$f_{EHLOW-TL-RL}(x; \alpha, \beta, \delta, \lambda, \gamma) = \frac{4 \alpha \beta \delta \lambda x^{\gamma - 1} (1 + x^\delta)^{\gamma - 1} (1 - (1 + x^\delta)^{\gamma - 1})^{\alpha \beta - 1}}{\lambda (1 - (1 + x^\delta)^{-\gamma})^{\alpha \beta - 1}} \times \exp(-t)(1 + \exp(-t))^{-2} [1 - \exp(-t)]^{\delta - 1},$$

where $t = \left[ \frac{1 - (1 + x^\delta)^{-\gamma}}{1 - (1 + x^\delta)^{-\gamma}} \right]^{\beta}$, for $\alpha, \beta, \delta, \lambda, \gamma > 0$,

$$f_{OGLH-LW}(x; \alpha, \beta, \lambda, \gamma) = \frac{2 \alpha \beta \lambda x^{\gamma - 1} (1 - e^{-\lambda x})^{\beta - 1} \exp \left\{ - \alpha \left[ \frac{1 - e^{-\lambda x}}{e^{-\lambda x}} \right]^\beta \right\}}{e^{-\gamma + 1} x^\gamma (1 + \exp \left\{ - \alpha \left[ \frac{1 - e^{-\lambda x}}{e^{-\lambda x}} \right]^\beta \right\}^2)}$$

for $\alpha, \beta, \lambda, \gamma > 0$,

$$f_{ELLOLL}(x; \alpha, \beta, \gamma, \theta, \lambda) = \frac{\alpha \theta^2 \gamma \lambda x^{\gamma - 1} (1 - e^{-\lambda x})^{\alpha \theta - 1} (1 - e^{-\lambda x})^{\alpha - 1}}{(1 + \beta) (1 - e^{-\lambda x})^{\alpha \theta - 1} (1 + \beta) (1 - e^{-\lambda x})^{\alpha - 1} + e^{-\alpha (\lambda x) \theta - 1}} \times \left( 1 - \beta \log \left[ \frac{1 - e^{-\lambda x}}{1 - e^{-\lambda x}} \right]^{\alpha \theta - 1} \right),$$

for $\alpha, \beta, \gamma, \theta, \lambda > 0$,

$$f_{KOL-LLog}(x; \alpha, \beta, \lambda, c) = ab \left[ \frac{\lambda^2}{(1 + \lambda)} \frac{c x^{e - 1}}{1 + x^e} \right] \exp(-\lambda z) \left[ 1 - \frac{\lambda + (1 + x^c)^{-1}}{1 + \lambda (1 + x^c)^{-1}} \exp(-\lambda z) \right]^{\alpha - 1} \times \left( 1 - \left[ 1 - \frac{\lambda + (1 + x^c)^{-1}}{1 + \lambda (1 + x^c)^{-1}} \exp(-\lambda z) \right] \right) a - b,$$

where $z = \frac{(1 - (1 + x^e)^{-1})}{(1 + x^e)^{-1}}$, $a, b, \lambda, c > 0$, and

$$f_{BOL-E}(x; \alpha, \beta, \lambda, \theta) = \frac{1}{B(a, b)} \left[ 1 - \frac{\lambda + e^{-\theta x}}{1 + \lambda} e^{-\theta x} \exp \left\{ -\lambda \left[ \frac{1 - e^{-\theta x}}{1 - e^{-\theta x}} \right] \right\} \right]^{\alpha - 1},$$
for \( a, b, \lambda, \theta > 0 \). We get the Gompertz distribution from G-Gom distribution by setting \( \alpha = 1 \). Also, we obtain the ELOW-TL-LLoG distribution from the ELOW-TL-BXII distribution by setting \( \gamma = 1 \). For the ELOWLOW distribution we considered the case when \( \alpha = 1 \).

### 6.1. Strengths of 1.5 cm glass fibres data

The first data set represents strengths of 1.5 cm glass fibres. The data set was also analyzed by Bourguignon et al. [6] and Smith and Naylor [33]. The data are 0.55, 0.93, 1.25, 1.36, 1.49, 1.52, 1.58, 1.61, 1.64, 1.68, 1.73, 1.81, 2.00, 0.74, 1.04, 1.27, 1.39, 1.49, 1.53, 1.59, 1.61, 1.66, 1.68, 1.76, 1.82, 2.01, 0.77, 1.11, 1.28, 1.42, 1.50, 1.54, 1.60, 1.62, 1.66, 1.69, 1.76, 1.84, 2.24, 0.81, 1.13, 1.29, 1.48, 1.50, 1.55, 1.61, 1.62, 1.66, 1.70, 1.77, 1.84, 0.84, 1.24, 1.40, 1.48, 1.51, 1.55, 1.61, 1.63, 1.67, 1.70, 1.78, 1.89.

The estimated variance-covariance matrix for the MO-Gom-W model based on glass fibres data set is

\[
\begin{bmatrix}
1.0655 \times 10^{-6} & 8.8229 \times 10^{-8} & 3.2278 \times 10^{-5} & -3.1485 \times 10^{-7} \\
8.8229 \times 10^{-8} & 7.3053 \times 10^{-9} & 2.6726 \times 10^{-6} & -2.6071 \times 10^{-8} \\
3.2278 \times 10^{-5} & 2.6726 \times 10^{-6} & 9.7776 \times 10^{-6} & -9.5380 \times 10^{-6} \\
-3.1485 \times 10^{-7} & -2.6071 \times 10^{-8} & -9.5380 \times 10^{-6} & 1.0402 \times 10^{-7}
\end{bmatrix}
\]

and the 95% confidence intervals for the model parameters are given by \( \gamma \in [8.3180 \pm 0.40020], \delta \in [21.4720 \pm 0.0002], \lambda \in [0.3134 \pm 0.0613] \) and \( \theta \in [0.0018 \pm 0.0006] \). We observe from the results presented in Table 2 that the MO-Gom-

#### Table 2: MLEs and goodness-of-fit statistics

<table>
<thead>
<tr>
<th>Model</th>
<th>( \gamma )</th>
<th>( \beta )</th>
<th>( \lambda )</th>
<th>( \theta )</th>
<th>(-2\log L)</th>
<th>AIC</th>
<th>AICC</th>
<th>BIC</th>
<th>W*</th>
<th>A*</th>
<th>K-S</th>
<th>p-value</th>
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<tr>
<td>MO-Gom-W</td>
<td>8.3180</td>
<td>21.4720</td>
<td>0.3134</td>
<td>0.0018</td>
<td>24.1</td>
<td>32.1</td>
<td>32.8</td>
<td>40.6</td>
<td>0.0868</td>
<td>0.5016</td>
<td>0.0970</td>
<td>0.5937</td>
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<td>OCGHLW-W</td>
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<td>(6.5710 \times 10^{-10})</td>
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<td>35.8</td>
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</table>

W distribution fits the glass fibres data set better than the other models since it has the lowest values for the goodness-of-fit statistics \( A^* \), \( W^* \) and K-S (and the largest p-value for the K-S statistic). Also, from fitted densities and probability plots shown in Figure 4, we observe that the MO-Gom-W model fit the data set better than the other models that were considered.
6.2. Turbocharger failure times data

The data represents failure times (10^3 h) of turbocharger of one type of engine as report by Xu et al [35]. The data is 1.6, 2.0, 2.6, 3.0, 3.5, 3.9, 4.5, 4.6, 4.8, 5.0, 5.1, 5.3, 5.4, 5.6, 5.8, 6.0, 6.1, 6.3, 6.5, 6.7, 7.0, 7.1, 7.3, 7.4, 7.6, 7.7, 7.8, 7.9, 8.0, 8.1, 8.3, 8.4, 8.5, 8.7, 8.8, 9.0.

The estimated variance-covariance matrix is

$$
\begin{bmatrix}
0.0018 & 0.0058 & -0.0384 & 1.5976 \times 10^{-4} \\
0.0058 & 1.0817 & -0.3088 & 5.6079 \times 10^{-3} \\
-0.0384 & -0.3088 & 0.8593 & -4.3962 \times 10^{-3} \\
0.0001 & 0.0056 & -0.0043 & 4.0375 \times 10^{-5}
\end{bmatrix}
$$

and the 95% confidence intervals for the model parameters are given by $\gamma \in [0.0181 \pm 0.0832]$, $\delta \in [0.7477 \pm 2.0386]$, $\lambda \in [2.3312 \pm 1.8169]$ and $\theta \in [0.0030 \pm 0.0125]$. 

Furthermore, results from the second example shown in Table 3 affirm that the MO-Gom-W model performs better than the other models considered since it has the lowest values for the goodness-of-fit statistics $A^*$, $W^*$ and K-S (and the largest p-value for the K-S statistic). Figure 5 shows the flexibility gained by using the MO-Gom-W model on turbocharger data set compared to the other models that were considered.
Table 3: MLEs and goodness-of-fit statistics.

<table>
<thead>
<tr>
<th>Model</th>
<th>(\gamma)</th>
<th>(\delta)</th>
<th>(\lambda)</th>
<th>(\theta)</th>
<th>(-2\log L)</th>
<th>AIC</th>
<th>AICC</th>
<th>BIC</th>
<th>W*</th>
<th>A*</th>
<th>K-S</th>
<th>p-value</th>
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<tr>
<td>MO-Gom-W</td>
<td>0.0181</td>
<td>0.7477</td>
<td>2.3312</td>
<td>0.0030</td>
<td>158.3</td>
<td>166.3</td>
<td>167.4</td>
<td>173.0</td>
<td>0.0230</td>
<td>0.1704</td>
<td>0.0821</td>
<td>0.9501</td>
</tr>
<tr>
<td>Gompertz</td>
<td>-</td>
<td>0.4908</td>
<td>-</td>
<td>(3.7553 \times 10^{-10})</td>
<td>302.4</td>
<td>306.4</td>
<td>306.7</td>
<td>309.8</td>
<td>0.1896</td>
<td>1.2707</td>
<td>0.7401</td>
<td>&lt; 2.2000 \times 10^{-16}</td>
</tr>
<tr>
<td>G-Gom</td>
<td>9.5146</td>
<td>0.4498</td>
<td>-</td>
<td>(2.9039 \times 10^{-4})</td>
<td>180.3</td>
<td>186.3</td>
<td>187.0</td>
<td>191.4</td>
<td>0.2757</td>
<td>1.7601</td>
<td>0.1542</td>
<td>0.2976</td>
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<tr>
<td>MO-LLoG</td>
<td>4.6678</td>
<td>4.8416</td>
<td>3.9474</td>
<td>-</td>
<td>177.4</td>
<td>183.4</td>
<td>184.1</td>
<td>188.5</td>
<td>0.2142</td>
<td>1.4072</td>
<td>0.1437</td>
<td>0.3807</td>
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<tr>
<td>MO-EW</td>
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<td>2.7877</td>
<td>1.1887</td>
<td>-</td>
<td>162.6</td>
<td>168.6</td>
<td>169.3</td>
<td>173.7</td>
<td>0.0496</td>
<td>0.3766</td>
<td>0.0918</td>
<td>0.8889</td>
</tr>
<tr>
<td>MO-IW</td>
<td>-</td>
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<td>0.9275</td>
<td>-</td>
<td>177.4</td>
<td>183.4</td>
<td>184.1</td>
<td>188.5</td>
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<td>1.4131</td>
<td>0.1438</td>
<td>0.3800</td>
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<tr>
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<td>0.2243</td>
<td>0.8263</td>
<td>159.1</td>
<td>167.1</td>
<td>168.2</td>
<td>173.8</td>
<td>0.0319</td>
<td>0.2646</td>
<td>0.1021</td>
<td>0.7985</td>
</tr>
<tr>
<td>EHLow-TL-LLoG</td>
<td>-</td>
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<td>0.2243</td>
<td>0.8263</td>
<td>159.1</td>
<td>167.1</td>
<td>168.2</td>
<td>173.8</td>
<td>0.0319</td>
<td>0.2646</td>
<td>0.1021</td>
<td>0.7985</td>
</tr>
<tr>
<td>OGLow-W</td>
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<td>0.3775</td>
<td>9.8384</td>
<td>0.7102</td>
<td>160.6</td>
<td>166.8</td>
<td>169.8</td>
<td>175.4</td>
<td>0.0353</td>
<td>0.2524</td>
<td>0.0856</td>
<td>0.9310</td>
</tr>
<tr>
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<td>21.9782</td>
<td>0.0630</td>
<td>9.9157</td>
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<td>172.3</td>
<td>173.5</td>
<td>179.1</td>
<td>0.0727</td>
<td>0.5442</td>
<td>0.1091</td>
<td>0.7278</td>
</tr>
<tr>
<td>ELOLLOW</td>
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<td>0.3775</td>
<td>9.8384</td>
<td>0.7102</td>
<td>160.6</td>
<td>166.8</td>
<td>169.8</td>
<td>175.4</td>
<td>0.0353</td>
<td>0.2524</td>
<td>0.0856</td>
<td>0.9310</td>
</tr>
<tr>
<td>KOL-LLoG</td>
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<td>0.0002</td>
<td>2.8682</td>
<td>165.1</td>
<td>171.3</td>
<td>172.9</td>
<td>175.8</td>
<td>0.0636</td>
<td>0.4815</td>
<td>0.1017</td>
<td>0.8027</td>
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<tr>
<td>BOL-E</td>
<td>0.2105</td>
<td>0.3251</td>
<td>0.0003</td>
<td>1.1911</td>
<td>159.0</td>
<td>167.0</td>
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<td>173.8</td>
<td>0.0343</td>
<td>0.2175</td>
<td>0.0832</td>
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</table>

7. Concluding remarks

We developed a new generalized distribution called the Marshall-Olkin-Gompertz-G (MO-Gom-G) family of distributions. We derived the statistical properties of the MO-Gom-G family of distributions. Maximum likelihood estimates of the model parameters were also derived. A simulation study was conducted to evaluate the consistency of the maximum likelihood estimates. A special case (MO-Gom-W) was applied to two real data examples to demonstrate the flexibility of the MO-Gom-G family of distributions.

Acknowledgment

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References

Appendix

The following url contains the simulation algorithm for the MO-Gom-W distribution https://drive.google.com/file/d/1e3ywmgm0Mx2YlrMUanFVJ7BjNjIubnLg/view?usp=sharing.