A new family of continuous distributions: properties and applications

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Resumo

This article introduces generalized beta-generated (GBG) distributions. Sub-models include all classical beta-generated, Kumaraswamy-generated and exponentiated distributions. They are maximum entropy distributions under three intuitive conditions, which show that the classical beta generator skewness parameters only control tail entropy and an additional shape parameter is needed to add entropy to the center of the parent distribution. This parameter controls skewness without necessarily differentiating tail weights. The GBG class also has tractable properties: we present various expansions for moments, generating function and quantiles. The model parameters are estimated by maximum likelihood and the usefulness of the new class is illustrated by means of some real data sets.

Key Words:

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1 Introduction

There has been an increased interest in developing generalized families of distributions by introducing additional shape parameters to a baseline cumulative distribution. This mechanism has proved to be useful to make the generated distributions more flexible especially for studying tail properties than existing distributions and for improving their goodness-of-fit statistics to the data under study.

Let $G(x)$ be the cumulative distribution function (CDF) of a baseline distribution and $g(x) = dG(x)/dx$ be the associated probability density function (PDF) depending on a parameter vectot η. We present a generalized family with two additional shape parameters by transforming the CDF $G(x)$ according to two sequential important generators. These families are important for modeling data in several engineering areas. Many special distributions in these families are discussed by Tahir and Nadarajah (2015).

Marshall and Olkin (1997) pioneered a general method to expand a distribution G by adding an extra shape parameter. The CDF of their family (for $\theta > 0$) is

$$F_{\text{MO-G}}(x) = \frac{G(x)}{\theta + (1-\theta)G(x)} = \frac{G(x)}{1-(1-\theta)[1-G(x)]}, \quad x \in \mathbb{R}. \tag{1}$$

The density function corresponding to (1) is

$$f_{\text{MO-G}}(x) = \frac{\theta g(x)}{[\theta + (1-\theta)G(x)]^2}, \quad x \in \mathbb{R}. \tag{2}$$

For $\theta = 1$, $f_{\text{MO-G}}(x)$ is equal to $g(x)$ and, for different values of θ, $f_{\text{MO-G}}(x)$ can be more flexible than $g(x)$. The extra parameter θ is called "tilt parameter", since the HRF of the MO-G family is shifted below ($\theta > 1$) or above ($0 < \theta < 1$) of the baseline HRF. Equation (2) provides a useful mechanism to generate new distributions from existing ones. The advantage of this approach for constructing new distributions lies in its flexibility to model both monotonic and non-monotonic HRFs even when the baseline HRF may be monotonic. Tahir and Nadarajah (2015, Table 2) presented thirty distributions belonging to the MO-G family. Further, this family is easily generated from the baseline QF by $Q_{\text{MO-G}}(u) = Q_G\left(\theta u\left[\theta u + 1 - u\right]\right)$ for $u \in (0,1)$.

Marshall and Olkin considered the exponential and Weibull distributions for the baseline G and derived some structural properties of the generated distributions. The special case that G is an exponential distribution refers to a two-parameter competitive model to the Weibull and gamma distributions. A simple interpretation of (1) can be given as follows. Let $T_1, \ldots, T_N$ be a sequence of independent and identically distributed (i.i.d.) random variables with survival function (SF) $\overline{G}(x) = 1 - G(x)$, and let N be a positive integer random variable independent of the T_i's defined by the probability generating function (PGF) of a geometric distribution with parameter θ, say $\tau(z;\theta) = \theta z\left[1 - (1-\theta)z\right]^{-1}$. Then, the inverse of $\tau(z;\theta)$ becomes $\tau^{-1}(z;\theta) = \tau(z;\theta^{-1})$. We can verify that equation (1) comes from $1 - F_{\text{MO-G}}(x) = \tau(\overline{G}(x);\theta)$ for $0 < \theta < 1$ and $1 - F_{\text{MO-G}}(x) = \tau(\overline{G}(x);\theta^{-1})$ for $\theta > 1$. For both cases, $1 - F_{\text{MO-G}}(x)$ represents the SF of $\min\{T_1, \ldots, T_N\}$, where N has PGF $\tau(z;\cdot)$ with probability parameters θ or θ^{-1}.

Zografos and Balakrishnan (2009) and Ristić and Balakrishnan (2011) defined the gamma-G (Γ-G) family with an extra shape parameter $a > 0$ by the CDF (for $x \in \mathbb{R}$)

$$F_{\Gamma\text{-G}}(x) = \gamma_1\left(a, -\log\left[1 - G(x)\right]\right) = \frac{1}{\Gamma(a)}\,\gamma\left(a, -\log\left[1 - G(x)\right]\right), \tag{3}$$

where $\gamma(a,z) = \int_0^z t^{a-1}\,\mathrm{e}^{-t}\mathrm{d}t$ is the incomplete gamma function and $\gamma_1(a,z) = \gamma(a,z)/\Gamma(a)$ is the incomplete gamma function ratio. The PDF of the Γ-G family takes the form

$$f_{\Gamma\text{-G}}(x) = \frac{1}{\Gamma(a)}\left\{-\log[1 - G(x)]\right\}^{a-1}\,g(x), \quad x \in \mathbb{R}. \tag{4}$$

Each new Γ-G distribution can be determined from a given baseline distribution. For $a = 1$, the G distribution is a basic exemplar of the Γ-G family.

Zografos and Balakrishnan (2009) and Ristić and Balakrishnan (2011) presented a physical motivation for the Γ-G family: if $X_{L(1)}, \ldots, X_{L(n)}$ are lower record values from a sequence of independent random variables with common PDF $g(\cdot)$, then the PDF of the nth lower record value has the form (4). If Z is a gamma random variable with unit scale parameter and shape parameter $a > 0$, then $X = Q_G(1 - \mathrm{e}^{Z})$ has density (4). So, the Γ-G distribution is easily generated from the gamma distribution and the QF of G.

The rest of the paper is organized as follows. Section 2 describes the distribution and density of the new class of distributions called the *Marshall and Olkin-Gamma-G* (MO-Γ-G) family by combining the above generators. We also present some special models. The maximum likelihood estimation of the model parameters of the new family is described in Section 3. In Section 4, we provide some empirical applications to illustrate the potentiality of the proposed family. A variety of theoretical properties are considered in Section 5. Finally, conclusions are noted in Section 6.

2 The New Family

Let $X \sim$MO-Γ-G denote a random variable having the MO-Γ-G family with two extra shape parameters $\theta > 0$ and $a > 0$ and the baseline vector η. By combining Equations (1) and (3), the CDF of X has the form

$$F_X(x) = \frac{\gamma_1\left(a, -\log\left[1 - G(x)\right]\right)}{\theta + (1-\theta)\gamma_1\left(a, -\log\left[1 - G(x)\right]\right)}, \quad x \in \mathbb{R}. \tag{5}$$

By differentiating (6), we can write the PDF of X as

$$f_X(x) = \frac{\theta\left\{-\log[1 - G(x)]\right\}^{a-1} g(x)}{\Gamma(a)\ \left\{\theta + (1-\theta)\gamma_1\left(a, -\log\left[1 - G(x)\right]\right)\right\}^2}, \quad x \in \mathbb{R}. \tag{6}$$

Distribution	Baseline CDF	Generated PDF
Normal	$G(x) = \Phi(x)$	$f_X(x) = \frac{\theta\{-\log[1-\Phi(x)]\}^{a-1}\,\phi(x)}{\Gamma(a)\,\{\theta+(1-\theta)\gamma_1(a,-\log[1-\Phi(x)])\}^2}$
Logistic	$G(x) = \frac{1}{1+\mathrm{e}^{-\mathrm{x}}}$	$f_X(x) = \frac{\theta\,\mathrm{e}^{-\mathrm{x}}\left\{-\log[1-(1+\mathrm{e}^{-\mathrm{x}})^{-1}]\right\}^{\mathrm{a}-1}}{\Gamma(a)\,(1+\mathrm{e}^{-\mathrm{x}})^2\,\{\theta+(1-\theta)\gamma_1(\mathrm{a},-\log[1-(1+\mathrm{e}^{-\mathrm{x}})^{-1}])\}^2}$
Gumbel	$G(x) = 1 - \exp(-\mathrm{e}^{\mathrm{x}})$	$f_X(x) = \frac{\theta\,\exp(a\,x-\mathrm{e}^{\mathrm{x}})}{\Gamma(a)\,\{\theta+(1-\theta)\gamma_1(a,\mathrm{e}^{\mathrm{x}})\}^2}$
Log-Normal	$G(x) = \Phi(\log x)$	$f_X(x) = \frac{\theta\,\phi(\log x)\,\{-\log[1-\Phi(\log x)]\}^{a-1}}{\Gamma(a)\,x\,\{\theta+(1-\theta)\gamma_1(a,-\log[1-\Phi(\log x)])\}^2}$
Exponential	$G(x) = 1 - \exp(-\lambda x),\ \lambda > 0$	$f_X(x) = \frac{\theta\,\lambda^a\,x^{(a-1)}}{\Gamma(a)\,\{\theta+(1-\theta)\gamma_1(a,\lambda x)\}^2}$
Weibull	$G(x) = 1 - \exp(-(\lambda x)^{\gamma}),\ \lambda, \gamma > 0$	$f_X(x) = \frac{\theta\,\gamma\lambda^{a\,\gamma}x^{a\,\gamma-1}\exp\{-(\lambda\,\gamma)^{\gamma}\}}{\Gamma(a)\{\theta+(1-\theta)\gamma_1[a,(\lambda\,x)^{\gamma}]\}^2}$
Gamma	$G(x) = \gamma_1(\alpha, \beta x),\ \alpha,\ \beta > 0$	$f_X(x) = \frac{\theta\,\beta^{\alpha}\,x^{\alpha-1}\,\mathrm{e}^{-\beta\mathrm{x}}\,\{-\log[1-\gamma_1(\alpha,\beta\mathrm{x})]\}^{\mathrm{a}-1}}{\Gamma(a)\,\{\theta+(1-\theta)\gamma_1(a,-\log[1-\gamma_1(\alpha,\beta x)])\}^2}$
Pareto	$G(x) = 1 - \frac{1}{(1+x)^{\nu}},\ \nu > 0$	$f_X(x) = \frac{\theta\,\mathrm{e}^{-\mathrm{x}}\,[\nu\log(1+\mathrm{x})]^{\mathrm{a}-1}\,\mathrm{g(x)}}{\Gamma(a)\,(1+\mathrm{e}^{-\mathrm{x}})^2\,\{\theta+(1-\theta)\gamma_1(\mathrm{a},\nu\log[1+\mathrm{x}])\}^2}$

Tabela 1: Special Distributions in the MO-Γ-G family.
($\Phi(x)$ and $\phi(x)$ denote the standard normal distribution and density functions and $\gamma_1(a, z)$ is the incomplete gamma function ratio).

Figure 1 shows the performance of the density and hazard functions of the Marshall-Olkin Gamma Weibull (MO-Γ-W, for short). Note that this extension of the Weibull distribution provides more flexibility for density and hazard shapes.

(a) (b)

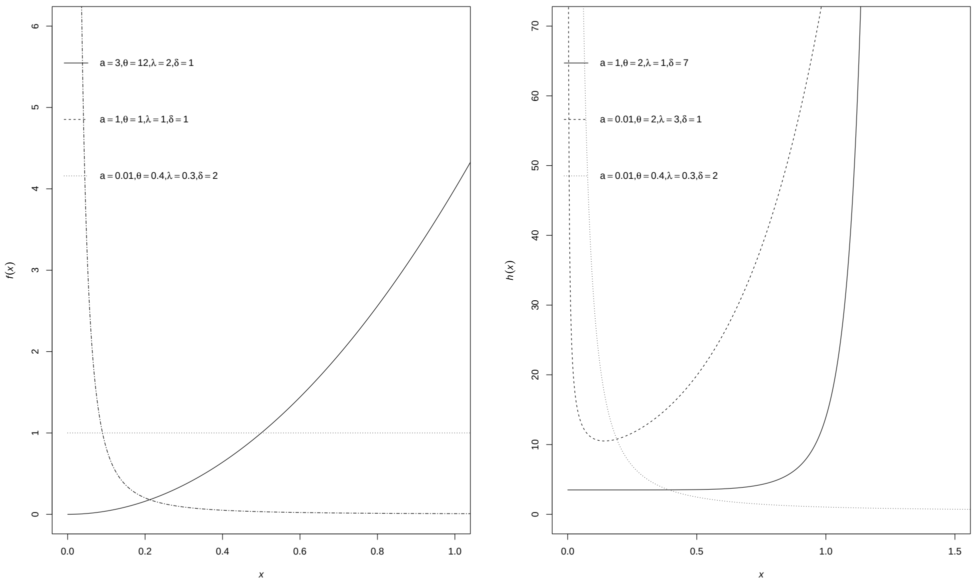

Figura 1: (a) Curves for the MO-Γ-W density; (b) Curves for the MO-Γ-W hazard;

2.1 Quantile Function

We can easily invert the cdf (6) to express the MOGa-G quantile function (qf), say $x = Q_X(u) = G^{-1}(u)$, in terms of the baseline qf $Q_G(u)$. Let u be a uniform $U(0,1)$ random variate. By using (6) and letting $z = \gamma_1\left(a, -\log[1-G(x)]\right)$, we obtain $z = z(u) = \theta u/[1-(1-\theta)u]$. Then, the qf of X has the form $x = Q_G\left(1 - \mathrm{e}^{-\mathrm{v}}\right)$, where $v = v(z)$ is a random variate generated from a gamma distribution (with shape parameter α and unit scale parameter) corresponding to the quantile z. This scheme is useful because of the existence of fast generators for uniform and gamma random variables.

3 Estimation

The MO-Γ-G family mentioned previously can be fitted to real data sets using the *AdequacyModel* package for the R statistical computing environment (`https://www.r-project.org/`). An important advantage of this package is that it is not necessary to define the log-likelihood function and that it computes the maximum likelihood estimates (MLE), their standard errors and the formal statistics presented in the nest section. We only need to provide the PDF and CDF of the distribution to be fitted to a data set. This *AdequacyModel* package uses the PSO (particle swarm optimization) method obtained by traditional global search approaches such as the quasi-Newton BFGS, Nelder-Mead and simulated-annealing methods to maximized the log-likelihood function. This method does not require initial values. More details are available at `https://rdrr.io/cran/AdequacyModel/`.

Here, we consider estimation of unknown parameters of the MO-Γ-G distribution by the method of maximum likelihood. If x is one observation from **??** and $\boldsymbol{\eta}$ a q-vector parameter vector specifying $G(.)$, thus the log-likelihood function, say $\log L = \log L(a, \theta, \boldsymbol{\eta})$ is

$$\begin{aligned}\ell(\boldsymbol{\Theta}) =& n\log(\theta) + n\log(\gamma) + n\,a\,\gamma\log(\lambda) + (a\gamma - 1)\sum_{i=1}^{n}\log(x_i) - \lambda^{\gamma}\sum_{i=1}^{n}\log(x_i) - n\log[\Gamma(a)] \\ & - 2\sum_{i=1}^{n}\log\{\theta + (1-\theta)\gamma_1[a, (\lambda\, x_i)^{\gamma}]\} \qquad (7)\end{aligned}$$

4 Applications

In this section, we provide two applications in order to show the performance of our proposed family with another ones. To do this, we consider the Weibull distribution as baseline and for competitive distributions, we consider: beta-Weibull (β-W), proposed by Famoye *et al.* (2005); Kumaraswamy Weibull (KW-W), see Cordeiro and Nadarajah (2010); Marshall-Olkin Weibull (MO-W), studied by Ahmed *et al.* (2017); Marshall-Olkin Extented Weibull (MOE-W), see the family proposed by Cordeiro *et al.* (2019) Exponentiated Weibull (exp-W), introduced by Mudholkar and Srivastava (1993); gamma Weibull (gamma-W), see Cordeiro *et al.* (2016) and exponentiated generalized Weibull (EG-W), see Oguntunde *et al.* (2015), with $a = 1$.

The log-likelihood for the Marshall-Olkin Gamma Weibull (MO-Γ-W) and considering one observation is given by

$$\begin{aligned}\ell(\boldsymbol{\Theta}) =& \log(\theta) + \log(\gamma) + (a\,\gamma)\log(\lambda) + (a\,\gamma - 1)\log(x) - (\gamma\, x)^{\gamma} - \log[\Gamma(a)] \\ & - 2\log\{\theta + (1-\theta)\gamma_1[a, (\lambda\, x)^{\gamma}]\} \qquad (8)\end{aligned}$$

where $\boldsymbol{\Theta} = (a, \theta, \lambda, \gamma)^T$. Besides that, the components of the score is given by

$$U_a(\boldsymbol{\Theta}) = \gamma\log(\lambda) + \gamma\log(x) - \psi^{(0)}(a) - \frac{2\left\{(1-\theta)A - (1-\theta)\psi^{(0)}(a)\gamma_1\left[a, (x\lambda)^{\gamma}\right]\right\}}{\theta\,\Gamma(a) + (1-\theta)\gamma_1\left[a, (x\lambda)^{\gamma}\right]},$$

$$U_\theta(\boldsymbol{\Theta}) = \frac{1}{\theta} - \frac{2\left\{\Gamma(a) - \gamma_1\left[a, (\lambda\, x)^{\gamma}\right]\right\}}{\theta\,\Gamma(a) + (1-\theta)\gamma_1\left[a, (\lambda\, x)^{\gamma}\right]},$$

$$U_\lambda(\boldsymbol{\Theta}) = \frac{\gamma}{\lambda}\left[a - (\lambda\, x)^{\gamma}\right] + \frac{2\gamma\,\lambda^{-1}(\lambda\, x)^{a\,\gamma}(1-\theta)\exp\{-(\lambda\, x)^{\gamma}\}}{\theta\,\Gamma(a) + (1-\theta)\gamma_1\left[a, (x\lambda)^{\gamma}\right]}$$

and

$$U_\gamma\boldsymbol{\Theta} = \frac{1}{\gamma} + a\log(\lambda) + a\log(x) - (\lambda x)^{\gamma}\log(\lambda x) + \frac{2(1-\theta)(\lambda x)^{\gamma\, a}\log(\lambda x)\exp\{-(\lambda x)^{\gamma}\}}{\theta\,\Gamma(a) + (1-\theta)\gamma_1\left[a, (x\lambda)^{\gamma}\right]},$$

where

$$A = G^{3,0}_{2,3}\left[(x\lambda)^{\gamma}\middle|\begin{matrix}1, 1\\ 0, 0, a\end{matrix}\right] + \log\left[(\lambda x)^{\gamma}\right]\gamma_1\left[a, (x\lambda)^{\gamma}\right].$$

where $\psi^{(n)}(x)$ is the n-th derivative of the digamma function and

$$A = \psi^{(0)}(a) - \log\left[(\lambda x)^{\gamma}\right] - G^{3,0}_{2,3}\left((\lambda x)^{\gamma}\middle|\, 0, \overset{1,1}{0}, a\right),$$

where $G_{p,q}^{m,n}\left(z\middle|\begin{matrix}a_1,\dots,a_p\\ b_1,\dots,b_q\end{matrix}\right)$ is the Meijer G function.

To provide the results, in this section we use the R software. In order to adjust the data sets consider here, we use the *AdequacyModel* package mentioned in Section 3. The method used here was the SANN method, that is a variant of simulated annealing (Belisle, 1992). To compare our proposed model with the other ones mentioned above, we use the Anderson Darling (A*) and Cramer Von Mises (W*) statistics, given by the *goodness.fit* function.

For the first data set, we consider a modification of the "FoodExpenditure" from the *betareg* package in R software, that refers on proportion of income spent on food for a random sample of 38 households in a large US city, according to the package information. Here, we consider the household expenditures for food. The modification is

$$data = FoodExpenditure_{food}/\#(FoodExpenditure_{food}),$$

where $FoodExpenditure_{food}$ is the random variable about the household expenditures for food and #(.) indicates the number of observations on the variable in question. The maximum likelihood estimators (MLE's) and their related standard errors (in parenthesis) are given in Table 2. Besides that, the (W*) and (A*) statistics are described in this table too. The results indicate that the proposed model has better performance than the other ones.

Tabela 2: Application 1

Model	a	θ	λ	γ	W^*	A^*
MO-Γ-W$(a,\theta,\lambda,\gamma)$	0.926134 (0.02626926)	1.379664 (0.22381086)	33.323073 (0.28530971)	25.398809 (0.08259864)	0.0339	0.2376
$\beta-$W$(a,\theta,\lambda,\gamma)$	9.92882 (0.02908181)	0.1700880 (0.02049663)	9.759469 (<0.0001)	1.530541 (<0.0001)	0.043567	0.2594618
KW-W$(a,\theta,\lambda,\gamma)$	0.04987575 (0.008090352)	99.99989793 (16.225905058)	1.07602954 (0.003156126)	23.40287099 (0.014646102)	1.330915	6.742609
MOE-W$(a,\theta,\lambda,\gamma)$	0.1366666 (0.1599182)	2.020436 (<0.0001)	62.72201 (<0.0001)	4.2956659 (0.7365535)	0.03541554	0.2579222
EGW(a,b,λ,γ)	5.6189861421 (0.0028147823)	6.1833138579 (0.0009807091)	1.2870816760 (0.1159810838)	1.3798562942 (0.1480747352)	0.03715255	0.2518879
MO-W(a,λ,γ)	0.15920715 (0.07170351)	- (-)	1.58609458 (0.13353716)	4.26713785 (0.16650667)	0.03456333	0.257339
exp-W(a,λ,γ)	6.1102948 (0.4222173)	- (-)	4.4680469 (0.3175876)	1.3858740 (0.1677722)	0.03726684	0.2523749
gamma-W(a,λ,γ)	5.751546 (0.0015006)	- (-)	10.0000 (0.0001163185)	1.208773 (0.00822782)	0.0879	0.6599

As a second application, we consider a data set that can be found in http://biostat.mc.vanderbilt.edu/wiki/ Main/DataSets. This data was collected in a pilot study about hypertension in the Dominican Republic in 1997. The observations are the systolic blood pressure of persons who came to medical clinics in several villages, for a variety of complaints. The MLEs of the model parameters and standard errors and the values of the statistics are listed in Table 3 for the previous models. Overall, by comparing the measures of these formal goodness-of-fit statistics, we conclude that the proposed distribution outperforms all distributions considered in Table 3.

Tabela 3: Application 2

Model	a	θ	λ	γ	W^*	A^*
MO-Γ-W$(a, \theta, \lambda, \gamma)$	9.629304 (0.006217876)	3.640779 (0.182495785)	6.260826 (0.024446563)	12.823429 (0.007965221)	0.5093	2.8076
$\beta-$W$(a, \theta, \lambda, \gamma)$	31.08471 (0.01279570)	47.14636 (<0.0001)	0.01698383 (0.00014535)	2.054037 (<0.0001)	0.7540428	4.279418
KW-W$(a, \theta, \lambda, \gamma)$	7363.281 (0.04194304)	0.03925762 (0.0004194304)	1.467640 (<0.0001)	0.6146940 (<0.0001)	0.5351	2.9617
MOE-W$(a, \theta, \lambda, \gamma)$	101.13471834 (46.782882038)	0.42386507 (0.100832102)	0.03095935 (0.003644092)	1.70240332 (0.168142275)	1.14418	6.644617
EGW(a, b, λ, γ)	0.2351691 (0.002540837)	140.0000 (3.278565)	0.4576370 (<0.0001)	0.7425738 (<0.0001)	0.8925	5.3338
MO-W(a, λ, γ)	173.2139 (0.00016394)	- (-)	0.02125455 (0.000212794)	1.601970 (0.0001398109)	1.476088	8.58705
exp-W(a, λ, γ)	69.02916 (0.08389090)	- (-)	0.02405120 (0.0002249256)	1.345536 (<0.0001)	0.8899261	5.094141
gamma-W(a, λ, γ)	9.11229459 (0.96330688)	- (-)	0.02617729 (0.00291875)	1.74641178 (0.08030313)	0.6227098	3.488268

5 Mathematical properties

In this section, we present some main mathematical properties for the MO-Γ-G family based on a general linear representation for its density given in the next section.

5.1 Linear Representation

A linear representation for the PDF the new family defined previously can be derived using the concept of exponentiated distributions. For an arbitrary baseline CDF $G(x)$, the exponentiated-G (exp-G) distribution with parameter $a > 0$, has CDF and PDF in the forms

$$\Pi_a(x) = G(x)^a \qquad \text{and} \qquad \pi_a(x) = a\, g(x)\, G(x)^{a-1},$$

respectively. The properties of the exponentiated distributions have been studied by many authors in recent years, see Mudholkar and Srivastava (1995) for exponentiated Weibull, Gupta *et al.* (1998) for exponentiated Pareto, Gupta and Kundu (2001) for exponentiated exponential and Nadarajah and Gupta (2007) for exponentiated gamma distribution. Tahir and Nadarajah (2015) cited almost thirty exponentiated distributions in their Table 1.

A linear representation for the PDF of the MO-Γ-G family in terms of exp-G densities is important to determine its mathematical properties from those of the exp-G distributions. They can follow from the papers described below.

First, the MO-G density (2) admits the linear combination (Barreto-Souza *et al.*, 2013)

$$f_{\text{MO-G}}(x) \;=\; \sum_{i=0}^{\infty} w_i^{\text{MO-G}}\, \pi_{i+1}(x),$$

where the coefficients are (for $i = 0, 1, \ldots$)

$$w_i^{\text{MO-G}} = w_i^{\text{MO-G}}(\theta) = \begin{cases} \dfrac{(-1)^i\,\theta}{(i+1)} \sum\limits_{j=i}^{\infty} \dbinom{j}{i} (j+1)\bar{\theta}^j, & \theta \in (0,1), \\ \theta^{-1}(1-\theta^{-1})^i, & \theta > 1, \end{cases}$$

and $\bar{\theta} = 1 - \theta$.

Second, the linear combination for the Γ-G density (4) was derived by Castellares and Lemonte (2015) as

$$f_{\Gamma\text{-G}}(x) \;=\; \sum_{i=0}^{\infty} w_i^{\Gamma\text{-G}}\; \pi_{a+i}(x),$$

where

$$w_i^{\Gamma\text{-G}} = w_i^{\Gamma\text{-G}}(a) = \frac{\varphi_i(a)}{(a+i)},$$

$$\varphi_0(a) = \frac{1}{\Gamma(a)}, \quad \varphi_i(a) = \frac{(a-1)}{\Gamma(a)}\,\psi_{i-1}(i+a-2), \quad i \geq 1,$$

and $\psi_{i-1}(\cdot)$ are the Stirling polynomials defined by

$$\psi_{n-1}(x) = \frac{(-1)^{n-1}}{(n+1)!}\Bigg[H_n^{n-1} - \frac{x+2}{n+2}H_n^{n-2} + \frac{(x+2)(x+3)}{(n+2)(n+3)}H_n^{n-3} - \cdots \\ + (-1)^{n-1}\frac{(x+2)(x+3)\cdots(x+n)}{(n+2)(n+3)\cdots(2n)}H_n^0\Bigg],$$

where H_n^m are positive integers given recursively by $H_{n+1}^m = (2n+1-m)H_n^m + (n-m+1)H_n^{m-1}$, with $H_0^0 = 1$, $H_{n+1}^0 = 1 \times 3 \times 5 \times \cdots \times (2n+1)$, $H_{n+1}^n = 1$.

6 Conclusions

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