Efficient and Robust Fitting of Lognormal Distributions

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Abstract

In parametric modeling of loss distributions in actuarial science, a versatile choice with intermediate tail weight is the lognormal distribution. Surprisingly, however, the fitting of this model using estimators that are at once efficient and robust has not been seriously addressed in the extensive literature. Consequently, typical estimators of the lognormal mean and variance fail to be both efficient and robust. In particular, the highly efficient maximum likelihood estimators lack robustness because of their limited sensitivity to outliers in the sample. For the two-parameter lognormal estimation problem, the author considers the problem of efficient and robust joint estimation of the mean and variance of a normal model. He introduces generalized-median-type estimators that yield efficient and robust estimators of various parameters of interest in the lognormal model. The paper provides detailed treatment of the lognormal mean and discusses extension of the approach to the much more complicated problem of estimation for the three-parameter lognormal model.

1. Introduction and Preliminaries

In parametric modeling of loss distributions in actuarial science, the *lognormal* distribution is a versatile choice, with tail weight intermediate between that of the gamma and Pareto distributions. In its three-parameter form, $L(\mu, \sigma, \tau)$ is the distribution of

$$Y = \tau + e^X,$$

where τ represents a threshold value and X is a normal random variable with mean μ and standard deviation σ. See Daykin, Pentikäinen, and Pesonen (1994) for useful discussion and graphical illustrations and Klugman, Panjer, and Willmot (1998) for detailed treatment including methods of fitting. Applications arising in business and economics include modeling of firm sizes, incomes, stock prices, and lengths of service in labor turnover contexts, and the model serves many other kinds of applications as well. Complete books (Aitchison and Brown 1957; Crow and Shimizu 1988) as well as Chapter 14 of Johnson, Kotz, and Balakrishnan (1994) are dedicated to the theory and the diverse applications of the lognormal model.

In fitting a statistical model by estimation of parameters, two very important properties are desired of the estimators: *efficiency,* in the sense of small mean square error, and *robustness,* in the sense of low sensitivity to outliers in the data. An "outlier" is an observation sufficiently far afield from the bulk of the data that its representativeness of the underlying population is questionable. Surprisingly, however, the goal of finding estimators that are not only efficient but also robust has not been seriously addressed in the extensive literature on the lognormal model. This paper focuses on estimation of the *mean* of the lognormal distribution

$$\eta = \mathrm{E}\{Y\} = \tau + e^{\mu+\sigma^2/2}$$

and develops estimators of η meeting both of the above criteria.

Following Klugman, Panjer, and Willmot (1998) and others, attention is confined primarily to the *two-parameter* lognormal model corresponding to the case that the threshold parameter τ is known, as for example when τ represents a known deductible for claim amounts. (Section 3.3, however, briefly discusses extension of the

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results to the much more complicated problem of estimation for the three-parameter lognormal model.) Thus, setting $\tau = 0$ without loss of generality, consider the lognormal model $L(\mu, \sigma)$, defined by the cdf

$$F(y) = \Phi\left(\frac{\log y - \mu}{\sigma}\right), 0 < y < \infty,$$

where $-\infty < \mu < \infty$, $0 < \sigma < \infty$, and Φ denotes the standard normal cdf. A random variable Y thus has the distribution $L(\mu, \sigma)$ if $X = \log Y$ has the normal distribution $N(\mu, \sigma^2)$ with mean μ and standard deviation σ.

In principle, therefore, the fitting of a (two-parameter) lognormal model reduces simply to the fitting of a normal distribution; we will see, however, that our estimation goal central to the lognormal model corresponds in the associated normal model to a problem that has not received sufficient development. In particular, the problem of efficient and robust estimation of $\eta = e^{\mu+\sigma^2/2}$ clearly rests upon that of *simultaneously* efficient and robust *joint* estimation of μ and σ in the context of the corresponding model $N(\mu, \sigma^2)$. The latter problem has received only limited attention (see some general development in Hampel *et al.* 1986) that does not meet present needs. Rather, treatments of the model $N(\mu, \sigma^2)$ have developed excellent efficient and robust estimators of μ but have left σ to be estimated merely consistently as a nuisance parameter. This paper extends the methodology for the normal model in a way that serves such applications as efficient and robust estimation of the lognormal mean.

Here let us clarify that, although the normal model comes into play, our focus remains on the *lognormal* model in order to serve applications in which it is indeed the model of choice. Thus, for present purposes, the only relevant transformation is the logarithmic transformation. The Box-Cox power transformations, and various other transformations that arise in connection with the goal of exploring which kind of transformed normal model might fit a data set, are not relevant in the present context.

The efficiency criterion employed is based on the performance of the maximum likelihood estimator (MLE), whose asymptotic optimality in terms of variance provides a quantitative benchmark. Thus, for a competing estimator, the asymptotic relative efficiency (ARE) is defined as the limiting ratio of sample sizes at which that estimator and the MLE perform "equivalently." Precise formulation appears in Section 1.1.

For robustness, two interrelated measures are used. The breakdown point (BP) of an estimator is the greatest fraction of data values that may be corrupted without the estimator becoming uninformative about the target parameter. The gross error sensitivity (GES) approximately measures the maximum contribution to the estimation error that can be produced by a single outlying observation when the given estimator is used. From the discussion of the BP and GES measures in Section 1.2, it can be seen that, as the anticipated proportion of outliers increases, it suggests the use of an estimator with high BP. Therefore, it becomes increasingly important that the estimator have a low GES.

Since higher BP comes at a higher price in terms of reduced ARE, however, one should choose estimators with BP no higher than actually needed. In typical situations, the range 0.05 to 0.30 for BP provides very adequate protection. An effective general approach is to set a minimum acceptable BP and a maximum acceptable GES and then maximize ARE subject to these constraints.

In this spirit, we develop estimators for μ, σ, and η that offer very high ARE, along with adequately high BP and adequately low GES. Let us first examine the MLEs as candidates. For a data set $Y_1, \ldots, Y_n$ from the model $L(\mu, \sigma)$, transformation to the equivalent model $N(\mu, \sigma^2)$ yields the well-known MLEs of the location parameter μ and the scale parameter σ:

$$\hat{\mu}_{\rm ML} = n^{-1} \sum_{i=1}^{n} \log Y_i$$

and

$$\hat{\sigma}_{\rm ML} = \left(n^{-1} \sum_{i=1}^{n} (\log Y_i - \hat{\mu}_{\rm ML})^2\right)^{1/2}.$$

These yield the MLE of $\eta = e^{\mu+\sigma^2/2}$: $\hat{\eta}_{\rm ML} = e^{\hat{\mu}_{\rm ML}+\hat{\sigma}^2_{\rm ML}/2}$. While the estimators $\hat{\mu}_{\rm ML}$ and $\hat{\sigma}_{\rm ML}$ each possess the favorable properties of converging to their respective parameters and having minimal asymptotic variance, they fail to be robust, each having BP = 0 and GES = ∞ (the worst cases). Such sensitivity to outliers is undesirable; therefore, alternative estimators that give up some efficiency in return for a suitable degree of robustness are desired. (This nonrobustness of lognormal model-based estimators is also seen in a study by Myers and Pepin, 1990, in the context of estimation of population abundance using a lognormal distribution for the nonzero observations.)

For the parameter μ in $N(\mu, \sigma^2)$, a number of robust competitors to $\hat{\mu}_{ML}$ already exist that pay relatively small prices in terms of reduced efficiency. Trimmed means, the Hodges-Lehmann estimator, M-estimators, and others are discussed in Hampel et al. (1986). The present paper uses estimators of *generalized median* (GM) type (Serfling 1984), which offer excellent trade-offs between efficiency and robustness and have other attractive properties.

For the parameter σ in $N(\mu, \sigma^2)$, there are classical robust competitors to $\hat{\sigma}_{ML}$ based on the interquartile range and the median absolute deviation, but, while offering very high BP (0.50), these sacrifice too much efficiency. These and some trimmed standard-deviation-type competitors to $\hat{\sigma}_{ML}$ are improved upon, however, by estimators of Rousseeuw and Croux (1993) that also attain BP = 0.50 but give up a much smaller, though still substantial, amount of efficiency. The present paper, giving greater emphasis to ARE while relaxing somewhat the very stringent BP = 0.50 requirement, develops for σ new estimators of GM type that attain very high ARE, high enough BP, and low enough GES.

Formulation of the GM estimators for joint estimation of μ and σ in $N(\mu, \sigma^2)$ is carried out in Section 2, along with study of their BP, GES, and ARE performance measures. As the problem of efficient and robust fitting of normal models is very basic to statistical practice, these results are of general interest and have broad potential application. Particular application to the lognormal model is treated in Section 3, where GM competitors to $\hat{\eta}_{ML}$ are obtained that have very favorable ARE as well as attractive BP and GES. The Appendix provides miscellaneous details and proofs. Following some remarks below, the remainder of the present section is devoted to formulation of the ARE, BP, and GES measures.

REMARKS

(i) A more comprehensive study of robustness would examine the efficiencies of the estimators within a neighborhood of the target model. Of course, then one must define what is meant by "nearby," in the sense of a suitable metric (for which there are a number of standard choices). Such an extended treatment, however, entails technical development beyond the scope of the present paper.

(ii) The BP and GES correspond to particular features of the *maxbias curve,* a more sophisticated robustness measure introduced by Martin, Yohai, and Zamar (1989; see also Ferretti et al. 1999 for recent discussion and further references). While this curve does provide somewhat more information on the robustness of the estimators than BP and GES alone, its use in general requires technical development beyond the scope of the present paper. In any case, the GM estimators presented in this paper are competitive with any estimators obtained via alternative approaches and points of view.

(iii) The abbreviation "GM" used here (and elsewhere in the literature) for "generalized median estimators" is used alternatively in some other parts of the literature to denote "generalized M-estimators." We assume that this will present no difficulty for readers.

1.1 Efficiency Criterion: ARE

We start with the fact (see, e.g., Serfling 1980, Sect. 2.2, Prob. 2.P.8, and Sect. 3.3) that the bivariate MLE $(\hat{\mu}_{ML}, \hat{\sigma}_{ML})$ for (μ, σ) in $N(\mu, \sigma^2)$ is asymptotically bivariate normal with mean (μ, σ) and covariance matrix $n^{-1}\Sigma_0$, where

$$\Sigma_0 = \begin{bmatrix} \sigma^2 & 0 \\ 0 & \sigma^2/2 \end{bmatrix}.$$

That is,

$$n^{1/2}(\hat{\mu}_{ML} - \mu, \hat{\sigma}_{ML} - \sigma) \xrightarrow{d} N((0, 0), \Sigma_0),$$

as the sample size $n \to \infty$, where $\xrightarrow{d}$ denotes "converges in distribution" and $N((0, 0), \Sigma_0)$ denotes bivariate normal with mean (0, 0) and covariance matrix Σ_0.

As discussed in Serfling (1980, Sect. 4.1), for estimation of a d-variate parameter ξ by a

d-variate estimator $\hat{\xi}$, which is asymptotically d-variate normal with mean ξ and covariance matrix $n^{-1}\Sigma$, it follows that confidence ellipsoids for ξ based on the estimator $\hat{\xi}$ have volume proportional to $|\Sigma|^{1/d}$. Thus, the determinant $|\Sigma|$ plays in higher dimensions the role played by the variance in one dimension and is called the *generalized variance.* For two competing asymptotically d-variate normal estimators A and B, with the same mean vector ξ and respective covariance matrices Σ_A and Σ_B, it follows that the ratio of respective sample sizes n_A and n_B at which the estimators perform "equivalently" (that is, have confidence ellipsoids of equal volume) approaches a limit value,

$$\frac{n_A}{n_B} \to \left(\frac{|\Sigma_A|}{|\Sigma_B|}\right)^{1/d}, \tag{1}$$

as n_A and $n_B \to \infty$. This limit is then interpreted as the ARE of estimator "B" with respect to estimator "A." Of course, in the case of a one-dimensional estimator, we have $d = 1$ and the quantity in Equation (1) is just the ratio of asymptotic variance parameters.

Now consider an estimator $(\hat{\mu}, \hat{\sigma})$, which, like the MLE, is asymptotically bivariate normal with mean (μ, σ) but with some other covariance matrix Σ_1:

$$n^{1/2}(\hat{\mu} - \mu, \hat{\sigma} - \sigma) \xrightarrow{d} N((0, 0), \Sigma_1).$$

Applying Equation (1) with $d = 2$, we then obtain for the ARE of $(\hat{\mu}, \hat{\sigma})$, with respect to $(\hat{\mu}_{ML}, \hat{\sigma}_{ML})$:

$$\text{ARE}((\hat{\mu}, \hat{\sigma}), (\hat{\mu}_{ML}, \hat{\sigma}_{ML})) = \left(\frac{|\Sigma_0|}{|\Sigma_1|}\right)^{1/2}. \tag{2}$$

In the special case that Σ_1, like Σ_0, is of form

$$\Sigma_1 = \begin{bmatrix} v_{11} & 0 \\ 0 & v_{22} \end{bmatrix}, \tag{3}$$

Equation (2) becomes simply

$$\left(\frac{\sigma^2}{v_{11}} \times \frac{\sigma^2/2}{v_{22}}\right)^{1/2}; \tag{4}$$

that is,

$$\text{ARE}((\hat{\mu}, \hat{\sigma}), (\hat{\mu}_{ML}, \hat{\sigma}_{ML})) = (\text{ARE}(\hat{\mu}, \hat{\mu}_{ML}) \times \text{ARE}(\hat{\sigma}, \hat{\sigma}_{ML}))^{1/2}. \tag{5}$$

REMARKS

(i) All choices of estimators $(\hat{\mu}, \hat{\sigma})$ that we consider here for estimation of (μ, σ) in $N(\mu, \sigma^2)$ will have asymptotic covariance matrices of the form in Equation (3) and, hence, will satisfy Equation (5). This is because each estimator $\hat{\mu}$ considered will be both *odd,*

$$\hat{\mu}(-X_1, \ldots, -X_n) = -\hat{\mu}(X_1, \ldots, X_n),$$

and *translation equivariant,*

$$\hat{\mu}(X_1 + c, \ldots, X_n + c) = \hat{\mu}(X_1, \ldots, X_n) + c, \quad \text{for any } c,$$

while each estimator $\hat{\sigma}$ considered will be both *even,*

$$\hat{\sigma}(-X_1, \ldots, -X_n) = \hat{\sigma}(X_1, \ldots, X_n),$$

and *translation invariant,*

$$\hat{\sigma}(X_1 + c, \ldots, X_n + c) = \hat{\sigma}(X_1, \ldots, X_n), \quad \text{for any } c.$$

As seen in Randles and Wolfe (1979, Cor. 1.3.33), in the case of data from a symmetric distribution, any odd translation equivariant statistic and any even translation invariant statistic are *uncorrelated.* Thus, throughout, we will have $\text{Cov}\{\hat{\mu}, \hat{\sigma}\} = 0$.

(ii) Further, each estimator $\hat{\mu}$ considered will be *scale equivariant,*

$$\hat{\mu}(X_1/c, \ldots, X_n/c) = \hat{\mu}(X_1, \ldots, X_n)/c, \quad \text{for any } c > 0.$$

This property, together with translation equivariance, yields that v_{11} in Σ_1 must be of form $c_{11}\sigma^2$, where c_{11} is the value of v_{11} obtained in the case of standard normal data. In this case, the ARE does not depend on μ or σ:

$$\text{ARE}(\hat{\mu}, \hat{\mu}_{ML}) = \frac{1}{c_{11}}.$$

Likewise, each estimator $\hat{\sigma}$ considered also will be *scale equivariant,* which, together with translation invariance, yields that v_{22} in Σ_1 must be of form $c_{22}\sigma^2$, where c_{22} is the value of v_{22} obtained in the case of standard normal data. Consequently, again the ARE does not depend on μ or σ:

$$\mathrm{ARE}(\hat{\sigma}, \hat{\sigma}_{\mathrm{ML}}) = \frac{1}{2c_{22}}.$$

With these simplifications, Equations (4) and (5) reduce to

$$\mathrm{ARE}((\hat{\mu}, \hat{\sigma}), (\hat{\mu}_{\mathrm{ML}}, \hat{\sigma}_{\mathrm{ML}})) = \left(\frac{1}{c_{11}} \times \frac{1}{2c_{22}}\right)^{1/2}. \quad (6)$$

(iii) Such simplicity fails, however, to hold for estimators of the parameter $\eta = e^{\mu+\sigma^2/2}$. In particular, for $\hat{\eta} = e^{\hat{\mu}+\hat{\sigma}^2/2}$, $\log \hat{\eta}$ is translation equivariant but not scale equivariant. Consequently, as we will find in Section 3, the quantity $\mathrm{ARE}(\hat{\eta}, \hat{\eta}_{\mathrm{ML}})$ turns out to be a function of σ. In this case, the different choices of estimator $\hat{\eta}$ are compared with respect to the ARE criterion by comparing respective tables or plots of their ARE versus σ over a range of σ values.

1.2 Robustness Criteria: BP and GES

1.2.1 Breakdown Point

A popular and effective criterion for robustness of an estimator is its BP, loosely characterized as the largest proportion of sample observations that themselves may be corrupted without the estimator itself becoming corrupted beyond use. When the BP is well-defined as a quantity not depending on the particular sample values but only on the sample size n, then we typically take as our criterion its limit value as $n \to \infty$. The BP of an estimator measures the degree to which the estimator remains uninfluenced by the presence of outliers. We, thus, define BP as the largest proportion of sample observations that may be given arbitrary values without taking the estimator to a limit uninformative about the parameter being estimated.

In particular, for the *location* parameter μ in $N(\mu, \sigma^2)$, we define $\mathrm{BP}(\hat{\mu})$ to be the largest proportion of observations that may be given arbitrary values without taking $\hat{\mu}$ to $\pm\infty$. For the *scale* parameter σ, we define $\mathrm{BP}(\hat{\sigma})$ to be the largest proportion of observations that may be given arbitrary values without taking $\hat{\sigma}$ to either 0 or ∞.

It is readily seen that the estimators $\hat{\mu}_{\mathrm{ML}}$ and $\hat{\sigma}_{\mathrm{ML}}$ each have BP = 0 and, thus, are *nonrobust* in this sense. Clearly, estimators are desired that have *nonzero* breakdown points while possessing relatively high efficiency.

1.2.2 Gross Error Sensitivity

Associated with any estimator $\hat{\xi}$ of a parameter $\xi(G)$ associated with a distribution G is its *influence function* (IF), defined by

$$\mathrm{IF}(x) = \lim_{\lambda \downarrow 0} \frac{\xi((1-\lambda)G + \lambda\delta_x) - \xi(G)}{\lambda},$$

where δ_x denotes the distribution placing all mass at the point x. As the directional derivative of ξ at G in the direction of δ_x, $\mathrm{IF}(x)$ approximates the contribution to the total estimation error that is made by an observation located at x. It follows that, for the estimator $\hat{\xi}$ based on a sample $X_1, \ldots, X_n$ from G, a first order approximation to the estimation error is given in terms of the IF:

$$\hat{\xi} - \xi \doteq n^{-1} \sum_{i=1}^{n} IF(X_i).$$

See Hampel et al. (1986) for a full treatment of the IF and its general role in statistical inference, and see Marceau and Rioux (2001) for a nice exposition of the IF and its application in the context of estimators of excess of loss premiums and other quantities arising in risk theory.

The quantity $sup_x|\mathrm{IF}(x)|$ is called the GES. Thus, the maximum possible impact that any single outlier can produce on the estimation error for a sample size n is measured, approximately, by

$$n^{-1}\ \mathrm{GES}.$$

Clearly, estimators with relatively low GES are desired. Indeed, as a minimal requirement for robustness, the IF should be *bounded.* It is also clearly desirable that the IF be *smooth.*

In particular, for estimation of μ and σ in $N(\mu, \sigma^2)$, the estimator $\hat{\mu}_{\mathrm{ML}}$ has an IF $x - \mu$, and the estimator $\hat{\sigma}_{\mathrm{ML}}$ has an IF

$$\frac{(x-\mu)^2 - \sigma^2}{2\sigma},$$

yielding in each case GES = ∞ (*nonrobustness*).

2. GENERALIZED MEDIAN ESTIMATORS FOR μ AND σ IN $N(\mu, \sigma^2)$

2.1 Basic Formulation

Generalized median (GM) estimators fall within the class of "generalized L-estimators" which were introduced and investigated in Serfling (1984). In general, for estimation of a parameter ξ on the basis of a sample $Y_1, \ldots, Y_n$, GM estimators are defined as follows. For a given choice of integer J, a "kernel" $h(y_1, \ldots, y_J)$ is selected such that $h(Y_1, \ldots, Y_J)$ is *median unbiased* for ξ; that is, the median of the distribution of $h(Y_1, \ldots, Y_J)$ is ξ. The present paper confines attention to kernels that are invariant under permutations of their arguments. In this case, the corresponding GM estimator is given by taking the median of the evaluations $h(Y_{i_1}, \ldots, Y_{i_J})$ of the kernel h over all subsets of observations taken J at a time, that is, corresponding to all $\binom{n}{J}$ subsets $\{i_1, \ldots, i_J\}$ of distinct indices from $\{1, \ldots, n\}$. This yields for ξ the estimator

$$\hat{\xi}_{\mathrm{GM}} = \mathrm{Median}\{h(Y_{i_1}, \ldots, Y_{i_J})\}.$$

Different choices of J and kernel h lead to different GM estimators for ξ.

Letting H denote the cdf of a kernel evaluation $h(Y_1, \ldots, Y_J)$ and assuming differentiability of H at ξ, it follows readily from Serfling (1984) that the estimator $\hat{\xi}_{\mathrm{GM}}$ has an IF

$$\frac{J}{H'(\xi)}\left(\frac{1}{2} - w(y)\right), \tag{7}$$

where

$$w(y) = P\{h(y, Y_1, \ldots, Y_{J-1}) \le \xi\}. \tag{8}$$

Thus, the IF of a GM estimator is bounded. For many typical h, the function $w(\cdot)$ is smooth, thus giving the IF this additional favorable property as well.

Since $0 \le w(y) \le 1$ must hold, an upper bound for the GES is given by

$$\mathrm{GES} \le \frac{J}{2H'(\xi)}.$$

Typically, either $\inf_y w(y) = 0$ or $\sup_y w(y) = 1$ (or both), in which case the GES actually equals this upper bound.

Furthermore, the function $w(\cdot)$ is instrumental in obtaining the asymptotic distribution of the GM estimator. Serfling (1984) shows that $\hat{\xi}_{\mathrm{GM}}$ is asymptotically normal with mean ξ and variance

$$\frac{J^2\ \mathrm{Var}\{w(Y)\}}{[H'(\xi)]^2}\, n^{-1},$$

as $n \to \infty$.

2.2 GM Estimators for μ in $N(\mu, \sigma^2)$

We now specialize the GM approach to estimate μ in $N(\mu, \sigma^2)$, on the basis of a random sample $X_1, \ldots, X_n$. For any fixed integer $k \ge 1$ not depending on the sample size n, we introduce the kernel

$$h_1(x_1, \ldots, x_k) = k^{-1} \sum_{i=1}^{k} x_i.$$

Clearly, $h_1(X_1, \ldots, X_k)$ is median unbiased for μ. The particular choice of kernel h_1 is motivated by the fact that each evaluation $h_1(X_{i_1}, \ldots, X_{i_k})$ is the MLE of μ based on just the observations $X_{i_1}, \ldots, X_{i_k}$. Denote the corresponding GM estimator by $\hat{\mu}_{(k)}$. Then $\hat{\mu}_{(1)}$ is just the ordinary median of the data, and $\hat{\mu}_{(2)}$ is the well-known Hodges-Lehmann estimator. For general choice of k, the estimator $\hat{\mu}_{(k)}$ was introduced in Serfling (1984) as a particular example of a generalized L-statistic and has been further studied in Choudhury and Serfling (1988), Choudhury (1990), Chaudhuri (1992), Ambühl (2000), and Serfling (2000).

It is not difficult to obtain (see Appendix Section A.1) that $\hat{\mu}_{(k)}$ has asymptotic breakdown point

$$\mathrm{BP}(\hat{\mu}_{(k)}) = 1 - (1/2)^{1/k}$$

as $n \to \infty$. The random variable $h_1(X_1, \ldots, X_k)$ is found to have distribution $H_1 = N(\mu, \sigma^2/k)$, yielding

$$H_1'(\mu) = \sqrt{\frac{k}{2\pi}}\,\sigma^{-1}.$$

Also, it is readily derived that, for this kernel, the function in Equation (8) is given by

$$w(x) = \Phi\left(\frac{\mu - x}{\sqrt{k-1}\,\sigma}\right).$$

Thus, $\hat{\mu}_{(k)}$ has a smooth and bounded IF.

Noting that $w(x) \to 1$ or 0 as $x \to -\infty$ or $+\infty$, respectively, we obtain from Equation (7) the gross error sensitivity

$$\mathrm{GES}(\hat{\mu}_{(k)}) = \sqrt{\frac{\pi}{2}}\,\sqrt{k}\sigma = 1.2533\sqrt{k}\sigma.$$

To eliminate dependence on σ, we use a *standardized* version, GES* = GES/σ.

Also, from Chaudhuri (1992), we have $\mathrm{Var}\{w(X)\} = (2\pi)^{-1}\sin^{-1}(1/k)$, and it follows that $\hat{\mu}_{(k)}$ is asymptotically normal with mean μ and variance $c_{11k}\sigma^2 n^{-1}$, as $n \to \infty$ with k fixed, where

$$c_{11k} = k\,\sin^{-1}(1/k).$$

We, thus, arrive at

$$\mathrm{ARE}(\hat{\mu}_{(k)}, \hat{\mu}_{\mathrm{ML}}) = \frac{1}{c_{11k}}.$$

Evaluations of BP, GES*, ARE, and c_{11} for selected k are provided in Table 1.

REMARKS

(i) While $\hat{\mu}_{(1)}$ (the median) has very high BP and very low GES, its ARE of 0.64 is unacceptably low. For $k = 2, \ldots, 9$, however, the estimators $\hat{\mu}_{(k)}$ provide a spectrum of favorable choices, trading off BP and GES by degrees in return for improved ARE.

(ii) Somewhat competitive estimators are provided by trimmed means, which tend to favor GES more strongly at a greater sacrifice of BP or ARE. For example (see Hampel 1974), the 10% trimmed mean has BP = 0.10, GES = 1.60, and ARE = 0.943. For comparison, $\hat{\mu}_{(2)}$ (the Hodges-Lehmann estimator) has slightly worse GES but much better BP and slightly higher ARE of 0.955, and $\hat{\mu}_{(6)}$ has much worse GES but comparable BP and much better ARE.

(iii) As another type of competitor, one might also consider a particular M-estimator such as the Huber Proposal 2, H(1.3), for which the BP is 0.29 in agreement with $\hat{\mu}_{(2)}$, but whose ARE corresponds to that of the 10% trimmed mean at 0.943 (Huber 1981, p. 144).

Thus, GM estimators tend to offer somewhat more favorable trade-offs between ARE and BP than competing estimators, although it should be recognized that these are not the only criteria that we might choose for efficiency and robustness. For the purpose of estimating μ in the framework of the lognormal target problem considered in this paper, one could reasonably take a trimmed mean or an M-estimator instead of $\hat{\mu}_{(k)}$ without drastic change in the results.

2.3 GM Estimators for σ in $N(\mu, \sigma^2)$

For GM-type estimation of σ in $N(\mu, \sigma^2)$, it is convenient first to develop estimators for σ^2 and then to take square roots.

2.3.1 GM Estimators for σ^2

A kernel based on the maximum likelihood method is again utilized. For fixed integer $m \geq 2$ not depending on n, we use

$$\tilde{h}_2(x_1, \ldots, x_m) = m^{-1}\sum_{i=1}^{m}\left(x_i - m^{-1}\sum_{j=1}^{m} x_j\right)^2,$$

the "maximum likelihood kernel" for estimation of σ^2 on the basis of just m observations. It is readily seen that $m\tilde{h}_2(X_1, \ldots, X_m)/\sigma^2$ has cdf G_{m-1}, where G_ν denotes the chi-square distribu-

Table 1
BP($\hat{\mu}_{(k)}$), GES*($\hat{\mu}_{(k)}$), ARE($\hat{\mu}_{(k)}$, $\hat{\mu}_{\mathrm{ML}}$), and c_{11k} for $k = 1:9$

	k								
	1	2	3	4	5	6	7	8	9
BP	0.500	0.293	0.206	0.159	0.129	0.109	0.094	0.083	0.074
GES*	1.253	1.772	2.171	2.507	2.802	3.070	3.316	3.545	3.760
ARE	0.637	0.955	0.981	0.989	0.993	0.995	0.997	0.997	0.998
c_{11k}	1.571	1.047	1.020	1.011	1.007	1.005	1.003	1.003	1.002

tion with ν degrees of freedom. With M_ν denoting the median of G_ν, we define

$$h_2(x_1, \ldots, x_m) = \frac{m}{M_{m-1}} \tilde{h}_2(x_1, \ldots, x_m)$$

and, thus, have that $h_2(X_1, \ldots, X_m)$ is median unbiased for σ^2. An alternative expression for h_2 is

$$h_2(x_1, \ldots, x_m) = \frac{1}{mM_{m-1}} \sum_{1\le i<j\le m} (x_i - x_j)^2.$$

Denote the corresponding GM estimator by $\hat{\sigma}^2_{(m)}$. For $m = 2$, this reduces to

$$\frac{1}{2M_1} \text{Median}_{1\le i<j\le n}\{(X_i - X_j)^2\},$$

an estimator formulated by Shamos (1976) and Bickel and Lehmann (1979) and studied in detail by Rousseeuw and Croux (1993). For $m \ge 3$, however, $\hat{\sigma}^2_{(m)}$ has not been investigated previously in the literature.

The asymptotic breakdown point for fixed m as $n \to \infty$ is found in Section A.1 in the Appendix to be

$$\text{BP}(\hat{\sigma}^2_{(m)}) = 1 - (1/2)^{1/m}.$$

The random variable $h_2(X_1, \ldots, X_m)$ has cdf $H_2(z) = G_{m-1}(M_{m-1}z/\sigma^2)$, $z > 0$, yielding

$$H_2'(\sigma^2) = C_m\sigma^{-2},$$

where

$$C_m = \left(\frac{M_{m-1}}{2}\right)^{(m-1)/2} \frac{e^{-((M_{m-1})/2)}}{\Gamma\left(\frac{m-1}{2}\right)},$$

with $\Gamma(\cdot)$ denoting the gamma function. Values of M_m and C_m for selected m are provided in Section A.2 in the Appendix.

It is not difficult to see that the function in Equation (8) is given by $w(x) = w_0((x - \mu)/\sigma)$, where

$$w_0(z) = P\{h_2(z, Z_1, \ldots, Z_{m-1}) \le 1\}$$
$$= P\left\{\sum_{i=1}^{m-1} (z - Z_i)^2 + \sum_{1\le i<j\le m-1} \times (Z_i - Z_j)^2 \le m\, M_{m-1}\right\},$$

with $Z_1, \ldots, Z_{m-1}$ independent standard normal random variables. Thus, $\hat{\sigma}^2_{(m)}$ has a smooth and bounded IF.

Note that $w_0(z) \to 0$ as $z \to \pm\infty$. We then obtain from Equation (7) the gross error sensitivity

$$\text{GES}(\hat{\sigma}^2_{(m)}) = \frac{m}{2C_m}\sigma^2.$$

To eliminate the dependence on σ^2, we use GES* = GES/σ^2.

Also, $\text{Var}\{w(X)\} = \text{Var}\{w_0(Z)\}$ with Z standard normal. With this quantity denoted by ζ_m, $\hat{\sigma}^2_{(m)}$ is asymptotically normal with mean σ^2 and variance $\tilde{c}_{22m}\sigma^4 n^{-1}$, where

$$\tilde{c}_{22m} = \frac{m^2\zeta_m}{C_m^2}.$$

Values of ζ_m are provided in Section A.2 in the Appendix. Also, in Section A.3 of the Appendix, the maximum likelihood estimator $\hat{\sigma}^2_{\text{ML}}$ is asymptotically normal with mean σ^2 and variance $2\sigma^4 n^{-1}$. We, thus, arrive at

$$\text{ARE}(\hat{\sigma}^2_{(m)}, \hat{\sigma}^2_{\text{ML}}) = \frac{2}{\tilde{c}_{22m}}.$$

Evaluations of BP, GES*, ARE, and $\tilde{c}_{22}$ for selected m are provided in Table 2.

2.3.2 GM Estimators for σ

For the parameter σ, the GM and ML estimators are obtained by simply taking square roots of corresponding estimators for σ^2. Breakdown points clearly remain unchanged, but as shown in

Table 2

BP($\hat{\sigma}^2_{(m)}$), GES*($\hat{\sigma}^2_{(m)}$), ARE($\hat{\sigma}^2_{(m)}$, $\hat{\sigma}^2_{\text{ML}}$), and $\tilde{c}_{22m}$, for m = 2, 3, 5, 7, and 9

	m				
	2	3	5	7	9
BP	0.293	0.206	0.129	0.094	0.074
GES*	4.666	4.328	4.754	5.308	5.841
ARE	0.864	0.862	0.910	0.940	0.956
$\tilde{c}_{22m}$	2.314	2.320	2.198	2.128	2.091

Table 3
BP($\hat{\sigma}_{(m)}$), GES*($\hat{\sigma}_{(m)}$), ARE($\hat{\sigma}_{(m)}$, $\hat{\sigma}_{ML}$), and c_{22m}, for m = 2, 3, 5, 7, and 9

	m				
	2	3	5	7	9
BP	0.293	0.206	0.129	0.094	0.074
GES*	2.333	2.164	2.377	2.654	2.920
ARE	0.864	0.862	0.910	0.940	0.956
c_{22m}	0.579	0.580	0.549	0.532	0.523

Section A.4 of the Appendix, the GES values change as follows:

$$\mathrm{GES}(\hat{\sigma}_{(m)}) = \frac{1}{2\sigma}\,\mathrm{GES}(\hat{\sigma}^2_{(m)}) = \frac{1}{2}\,\mathrm{GES}^*(\hat{\sigma}^2_{(m)})\sigma.$$

Standardizing to $\mathrm{GES}^*(\hat{\sigma}_{(m)}) = \mathrm{GES}(\hat{\sigma}_{(m)})/\sigma$, we have

$$\mathrm{GES}^*(\hat{\sigma}_{(m)}) = \tfrac{1}{2}\mathrm{GES}^*(\hat{\sigma}^2_{(m)}).$$

Also, the asymptotic variance of $\hat{\sigma}_{(m)}$ is that of $\hat{\sigma}^2_{(m)}$ times the factor $1/4\sigma^2$ (see Section A.4 in the Appendix). Thus, $\hat{\sigma}_{ML}$ and $\hat{\sigma}_{(m)}$ are each asymptotically normal with mean σ and respective asymptotic variances $0.5\sigma^2 n^{-1}$ and $c_{22m}\sigma^2 n^{-1}$, where

$$c_{22m} = \frac{\tilde{c}_{22m}}{4}.$$

Consequently, the ARE remains the same:

$$\mathrm{ARE}(\hat{\sigma}_{(m)}, \hat{\sigma}_{ML}) = \mathrm{ARE}(\hat{\sigma}^2_{(m)}, \hat{\sigma}^2_{ML}).$$

The corresponding analogue of Table 2 is provided in Table 3.

REMARKS

(i) Well-known robust competitors to $\hat{\sigma}_{ML}$ given by suitably normalized versions of the interquartile range and the median absolute deviation have very favorable BPs of 0.25 and 0.50, respectively, and in common a very favorable GES of 1.27. Unfortunately, however, these estimators sacrifice too much efficiency, having in common ARE of only 0.37 with respect to $\hat{\sigma}_{ML}$. Considerably higher ARE is achieved by trimmed versions of $\hat{\sigma}_{ML}$, at the cost of somewhat lower BP. For example, the 10% trimmed standard deviation has BP = 0.10 and ARE = .78 (see Bickel and Lehmann 1976 and Janssen, Serfling, and Veraverbeke 1987). In turn, these estimators are substantially improved with respect to BP, while slightly improving ARE, by estimators introduced by Rousseeuw and Croux (1993), one of which is discussed in (ii) below. Alternatively, the GM estimators $\hat{\sigma}_{(m)}$ for m = 2, 3, 5, 7, and 9 provide a spectrum of favorable choices attaining much higher ARE by trading off BP and GES to some extent.

(ii) With respect to the estimator $\hat{\sigma}_{(2)}$, Rousseeuw and Croux (1993) consider modifications that achieve a high BP = 0.50 with only a moderately small sacrifice of ARE. Specifically, they replace the median of the $\binom{n}{2}$ pairwise interpoint differences $|X_i - X_j|$ by the j_n-th order statistic, where $j_n = \binom{n}{2}/4$, and they alter the constant factor to 2.222. In comparison with $\hat{\sigma}_{(2)}$, the resulting estimator has optimal BP and relatively low GES*, but at the costs of a reduction in ARE and an increase in small sample bias. Nevertheless, this estimator is discussed in later discussions and denoted by $\hat{\sigma}_{(1)}$. Augmenting the information in Table 3, Table 4 provides similar information for $\hat{\sigma}_{(1)}$.

(iii) If desired, the estimators $\hat{\sigma}_{(m)}$ for $m \geq 3$ could also be modified in the above vein to improve BP by replacing the median with another of the ordered kernel evaluations and changing the multiplicative constant. It is preferable however, to retain the median because of its simplicity and intuitive appeal, and because it yields higher ARE and lower small sample bias.

2.4 Joint Estimation of μ and σ

Let us now consider *joint* estimation of (μ, σ) and compare the estimator $(\hat{\mu}_{(k)}, \hat{\sigma}_{(m)})$ with the estimator $(\hat{\mu}_{ML}, \hat{\sigma}_{ML})$, for various choices of the pair (k, m). As discussed in Section 1.1, use Equation (6) for the ARE. In conjunction with the values in

Table 4
BP($\hat{\sigma}_{(1)}$), GES*($\hat{\sigma}_{(1)}$), ARE($\hat{\sigma}_{(1)}$, $\hat{\sigma}_{ML}$), and c_{221}

BP	GES*	ARE	c_{221}
0.500	2.069	0.823	0.610

Tables 1, 3, and 4, this yields the ARE values in Table 5 for selected pairs (k, m).

While the most favorable choice in Table 5 is $(k, m) = (9, 9)$, with ARE = 0.98, we must also take into account the corresponding BP value, 0.07, from Tables 1 and 3. If somewhat higher BP values are needed, at the cost of a small reduction in ARE, the choice $(k, m) = (5, 5)$ offers ARE = 0.95 with BP = 0.13. If considerably higher BP values are needed, the choice $(k, m) = (2, 2)$ offers BP = 0.29, but the ARE slips to 0.91. The choice $(k, m) = (1, 1)$ offers the best BP, 0.50, but the ARE drops sharply to 0.72.

Note also that GES* values worsen as ARE improves. With this in mind, the choice $(k, m) = (5, 5)$ offers very good overall balance with respect to the factors BP, GES, and ARE. There are also considerations of computational burden, however, discussed in Section 3.1.3.

3. APPLICATION TO LOGNORMAL MODELS

For the two-parameter lognormal model, estimation of the mean is treated in Section 3.1 and estimation of other target parameters in Section 3.2. Extension to the three-parameter lognormal model is treated briefly in Section 3.3.

3.1 Estimation of the Mean

Now let's address the problem of efficient and robust estimation of the mean

$$\eta = e^{\mu + \sigma^2/2}$$

of the lognormal distribution $L(\mu, \sigma)$, on the basis of a sample $Y_1, \ldots, Y_n$. The MLE is given by

$$\hat{\eta}_{ML} = e^{\hat{\mu}_{ML} + \hat{\sigma}^2_{ML}/2},$$

Table 5

ARE(($\hat{\mu}_{(k)}$, $\hat{\sigma}_{(m)}$), ($\hat{\mu}_{ML}$, $\hat{\sigma}_{ML}$)) for k = 1, 2, 3, 5, 7, and 9, and m = 1, 2, 5, 7, and 9

	m				
k	1	2	5	7	9
1	0.724	0.742	0.761	0.774	0.780
2	0.887	0.908	0.932	0.947	0.955
3	0.899	0.921	0.945	0.960	0.968
5	0.904	0.926	0.951	0.966	0.974
7	0.906	0.928	0.953	0.968	0.976
9	0.906	0.929	0.953	0.969	0.977

with $\hat{\mu}_{ML}$ and $\hat{\sigma}_{ML}$ as in Section 1, based on the transformed observations $X_i = \log Y_i$, $1 \le i \le n$, which have the $N(\mu, \sigma^2)$ distribution. Although efficient, this estimator of η inherits the nonrobustness of $\hat{\mu}_{ML}$ and $\hat{\sigma}_{ML}$ and their BP values of 0 and GES values of ∞. Therefore, utilizing the development in Section 2, we consider the competing estimators

$$\hat{\eta}_{(k,m)} = e^{\hat{\mu}_{(k)} + \hat{\sigma}^2_{(m)}/2}$$

and examine their BP and GES in Section 3.1.1, and ARE in Section 3.1.2. Summary discussion is provided in Section 3.1.3.

3.1.1 BP and GES

For the BP, we have in general

$$\text{BP}(\hat{\eta}_{(k,m)}) = \min\{\text{BP}(\hat{\mu}_{(k)}), \text{BP}(\hat{\sigma}_{(m)})\}.$$

In particular, the estimators $\hat{\eta}_{(1,1)}$, $\hat{\eta}_{(2,2)}$, $\hat{\eta}_{(5,5)}$, and $\hat{\eta}_{(9,9)}$ have BPs of 0.50, 0.29, 0.13, and 0.07, respectively.

For the GES, we have, as shown in Section A.4 of the Appendix,

$$\text{GES}(\hat{\eta}_{(k,m)}) = \eta(\text{GES}(\hat{\mu}_{(k)}) + \text{GES}(\hat{\sigma}_{(m)})\sigma).$$

In this case, standardizing does not completely eliminate the dependence on parameters. A partially standardized version, $\text{GES}^*(\hat{\eta}_{(k,m)}) = \text{GES}(\hat{\eta}_{(k,m)})/\eta\sigma(1 + \sigma)$, however, satisfies

$$\text{GES}^*(\hat{\eta}_{(k,m)}) = \frac{1}{1+\sigma}\text{GES}^*(\hat{\mu}_{(k)}) + \frac{\sigma}{1+\sigma}\text{GES}^*(\hat{\sigma}_{(m)}).$$

This quantity has limit $\text{GES}^*(\hat{\mu}_{(k)})$ as $\sigma \to 0$ and limit $\text{GES}^*(\hat{\sigma}_{(m)})$ as $\sigma \to \infty$. For the estimators $\hat{\eta}_{(1,1)}$, $\hat{\eta}_{(2,2)}$, $\hat{\eta}_{(5,5)}$, and $\hat{\eta}_{(9,9)}$, in particular, these pairs of limits are (1.253, 2.069), (1.772, 2.333), (2.802, 2.377), and (3.760, 2.920), respectively.

3.1.2 ARE

From standard results on transformations of asymptotically normal random variables, it follows (see Section A.3 in the Appendix) that if joint estimators $(\hat{\mu}, \hat{\sigma})$ of (μ, σ) satisfy

$$n^{1/2}(\hat{\mu} - \mu, \hat{\sigma} - \sigma) \xrightarrow{d} N((0, 0), [\sigma_{ij}]_{2\times2})$$

as $n \to \infty$, then the corresponding estimator $\hat{\eta} = e^{\hat{\mu}+\hat{\sigma}^2/2}$ satisfies

$$n^{1/2}(\hat{\eta} - \eta) \xrightarrow{d} N(0, \eta^2(\sigma_{11} + 2\sigma\sigma_{12} + \sigma^2\sigma_{22}))$$

as $n \to \infty$. For all estimators $(\hat{\mu}, \hat{\sigma})$ under consideration, we have

$$\sigma_{11} = c_{11}\sigma^2, \qquad \sigma_{12} = 0, \qquad \sigma_{22} = c_{22}\sigma^2$$

for numerical constants c_{11} and c_{22}, yielding

$$n^{1/2}(\hat{\eta} - \eta) \xrightarrow{d} N(0, \eta^2\sigma^2(c_{11} + c_{22}\sigma^2))$$

as $n \to \infty$. For the MLEs of μ and σ we have $c_{11} = 1$ and $c_{22} = 0.5$. Thus, for any estimator $\hat{\eta}$ satisfying the above conditions, the ARE is given by

$$\mathrm{ARE}(\hat{\eta}, \hat{\eta}_{\mathrm{ML}}) = \frac{1 + 0.5\sigma^2}{c_{11} + c_{22}\sigma^2}.$$

Note that this ARE converges to $\mathrm{ARE}(\hat{\mu}, \hat{\mu}_{\mathrm{ML}})$ as $\sigma \to 0$ and to $\mathrm{ARE}(\hat{\sigma}, \hat{\sigma}_{\mathrm{ML}})$ as $\sigma \to \infty$. In particular, for $\hat{\eta}_{(k,m)}$, we have

$$\mathrm{ARE}(\hat{\eta}_{(k,m)}, \hat{\eta}_{\mathrm{ML}}) = \frac{1 + 0.5\sigma^2}{c_{11k} + c_{22m}\sigma^2}.$$

Thus, utilizing Tables 1, 3 and 4, we obtain

$$\mathrm{ARE}(\hat{\eta}_{(1,1)}, \hat{\eta}_{\mathrm{ML}}) = \frac{1 + 0.5\sigma^2}{1.57 + 0.61\sigma^2},$$

which increases from 0.637 at $\sigma = 0$ to 0.820 as $\sigma \to \infty$;

$$\mathrm{ARE}(\hat{\eta}_{(2,2)}, \hat{\eta}_{\mathrm{ML}}) = \frac{1 + 0.5\sigma^2}{1.047 + 0.579\sigma^2},$$

which decreases from 0.955 to 0.864;

$$\mathrm{ARE}(\hat{\eta}_{(5,5)}, \hat{\eta}_{\mathrm{ML}}) = \frac{1 + 0.5\sigma^2}{1.007 + 0.549\sigma^2},$$

which decreases from 0.993 to 0.911; and

$$\mathrm{ARE}(\hat{\eta}_{(9,9)}, \hat{\eta}_{\mathrm{ML}}) = \frac{1 + 0.5\sigma^2}{1.002 + 0.523\sigma^2},$$

which decreases from 0.998 to 0.956. The best ARE is attained as $\sigma \to 0$ for $k = m = 1$, but as $\sigma \to \infty$ in the other cases. This is attributable to the considerable inefficiency of the ordinary median $\hat{\mu}_{(1)}$ compared with the estimators $\hat{\mu}_{(k)}$ for $k \geq 2$.

Table 6 exhibits for selected values of σ the ARE values for these four estimators.

3.1.3 Summary Discussion

A good overall estimator appears to be $\hat{\eta}_{(5,5)}$, which offers quite high ARE above 0.91 uniformly over σ, combined with favorable BP of 0.13 and acceptable standardized GES* within the range 2.4 to 2.8. The estimator $\hat{\eta}_{(9,9)}$, however, which offers ARE above 0.96 uniformly over σ, is more attractive if lower BP of 0.07 and higher GES* in the range 2.9 to 3.8 can be tolerated.

However, it should be noted that the estimators $\hat{\eta}_{(k,m)}$ become increasingly computationally intensive as k or m increase. If computational burden is a consideration, then the estimator $\hat{\eta}_{(2,2)}$ becomes attractive. It provides excellent BP of 0.29, GES* in the 1.8 to 2.3 range, and ARE above 0.86 uniformly over σ. Alternatively, the estimators $\hat{\eta}_{(k,m)}$ may be modified to eliminate the computational intensiveness, as discussed in Section A.5 of the Appendix.

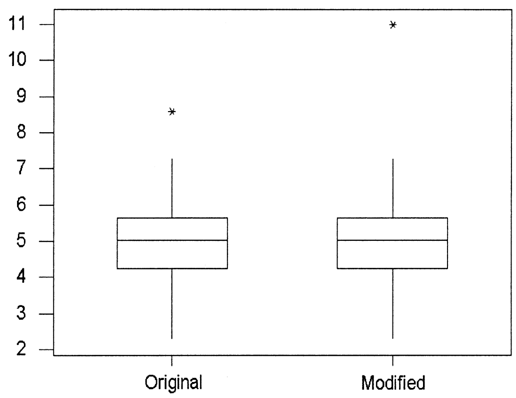
A quick illustration of the robustness of the GM estimators over the MLEs is provided by the following simple experiment. A random sample size of 100 from $N(5, 1)$ was taken, and the largest observation, already an "outlier," was increased in value to become a more extreme outlier, from 8.58 to 11.0. Standard boxplots of the original and modified samples are displayed in Figure 1, and

Table 6

ARE($\hat{\eta}_{(j,j)}$, $\hat{\eta}_{\mathrm{ML}}$) for j = 1, 2, 5, and 9, and for Selected σ

	σ						
	0	2.5	5.0	7.5	10.0	20.0	∞
(1, 1)	0.637	0.766	0.803	0.812	0.815	0.819	0.820
(2, 2)	0.955	0.884	0.870	0.866	0.865	0.864	0.864
(5, 5)	0.993	0.929	0.916	0.913	0.912	0.911	0.911
(9, 9)	0.998	0.966	0.959	0.957	0.957	0.956	0.956

Figure 1
Boxplots of Original and Modified Samples



the corresponding MLEs and $GM_{(2,2)}$ estimates of μ, σ, and η are listed in Table 7.

For $L(5, 1)$, we have $\eta = 244.7$. The presence of the outlier in the original sample results in the MLE slightly overestimating η, with a value of 249.6. The less efficient estimator $\hat{\eta}_{(2,2)}$ overestimates by a greater amount, with a value of 266.0. A rather moderate modification of the value of the single outlier in the original sample, however, influences a dramatic change in the value of $\hat{\eta}_{ML}$, from 249.6 to 282.9. However, the robust estimator $\hat{\eta}_{(2,2)}$ remains quite stable, with its value unchanged (although for some data sets its value would change somewhat, but not dramatically). We would expect similar results with the more efficient competitor $\hat{\eta}_{(5,5)}$, whose computation, however, requires for each of μ and σ taking the median of $\binom{100}{5} = 75{,}287{,}520$ kernel evaluations instead of just $\binom{100}{2} = 4{,}950$ as for $\hat{\eta}_{(2,2)}$. This computation may be carried out via an efficient algorithm or by the modified method described in Section A.5 of the Appendix.

It is of interest to investigate the small sample performance of the estimator $\hat{\eta}_{(2,2)}$, and a suitable study will be carried out elsewhere. One can somewhat anticipate the results from those of a simulation study by Brazauskas and Serfling (2001) for some other GM estimators in a different context. There it was found that the superiority of the GM estimators over various competitors remained valid even for small sample sizes $n = 10$ and 25, and that the specific ARE values are valid for sample size $n \geq 100$.

3.2 Other Target Parameters

Besides the mean, the *variance*

$$\theta = e^{2\mu+\sigma^2}(e^{\sigma^2} - 1)$$

is of interest, and insertion of efficient and robust estimates for μ and σ then yields estimates for θ that, likewise, are efficient and robust. In particular, the properties of the estimators $\hat{\theta}_{(k,m)}$ may be developed along the lines of Section 3.1.

Also, parameters such as the "limited expected value," $E\{X \wedge x\}$, and the "limited second moment," $E\{(X \wedge x)^2\}$, play important roles in actuarial practice. For extensive discussion and treatment, see Daykin, Pentikäinen, and Pesonen (1994) and Klugman, Panjer, and Willmot (1998). The latter authors remark (p. 73) on the greater flexibility offered by parametric modeling over empirical modeling. In particular, for the lognormal model, the following explicit formulas are readily derived:

$$E\{X \wedge x\} = e^{\mu+\sigma^2/2}\Phi\left(\frac{\log x - \mu - \sigma^2}{\sigma}\right) + x\left[1 - \Phi\left(\frac{\log x - \mu}{\sigma}\right)\right],$$

and

$$E\{(X \wedge x)^2\} = e^{2\mu+2\sigma^2}\Phi\left(\frac{\log x - \mu - 2\sigma^2}{\sigma}\right) + x^2\left[1 - \Phi\left(\frac{\log x - \mu}{\sigma}\right)\right].$$

Again, insertion of efficient and robust estimates for μ and σ yields efficient and robust estimates for the limited expected value and limited second moment.

Table 7
MLEs and $GM_{(2,2)}$ Estimates for Original and Modified Samples

Estimator	Original Sample	Modified Sample
$\hat{\mu}_{ML}$	4.98	5.00
$\hat{\sigma}_{ML}$	1.04	1.14
$\hat{\eta}_{ML}$	249.60	282.90
$\hat{\mu}_{(2,2)}$	5.00	5.00
$\hat{\sigma}_{(2,2)}$	1.08	1.08
$\hat{\eta}_{(2,2)}$	266.00	266.00

3.3 Extension to the Three-Parameter Lognormal Model

In the case of the three-parameter lognormal model $L(\mu, \sigma, \tau)$, even maximum likelihood estimation becomes highly problematic, and a variety of competing modified maximum likelihood approaches have been developed, along with other types of estimation methods (see Johnson, Kotz, and Balakrishnan 1994 for general discussion). The following estimators extend the foregoing methodology for the two-parameter case. First, let us denote the ordered Y_i's by

$$Y_{n1} \le Y_{n2} \le \cdots \le Y_{nn}$$

and estimate τ by

$$\hat{\tau}_n = Y_{n1} = \min\{Y_1, \ldots, Y_n\},$$

a natural estimator already considered in the literature. Next define transformed observations Z_i, $1 \le i \le n-1$ having ordered values given by

$$\begin{aligned} Z_{n-1,i} &= \log(Y_{n,i+1} - \hat{\tau}_n) \\ &= \log(Y_{n,i+1} - Y_{n1}), \quad 1 \le i \le n-1, \end{aligned}$$

and estimate μ and σ by the estimators $\hat{\mu}_{(k)}$ and $\hat{\sigma}_{(m)}$ based on the Z_i's as surrogates of the strictly normal variates $X_i = \log(Y_i - \tau)$, $1 \le i \le n$ used when τ is known. These estimators should retain the favorable combination of efficiency and robustness (in the case of $\hat{\tau}_n$, against *upper* outliers) established in the two-parameter case, but precise quantification of these properties is highly technical and deferred to a future investigation.

APPENDIX
PROOFS AND FURTHER DETAILS

A.1 BREAKDOWN POINTS

For efficient and robust estimation of the tail index of a Pareto distribution, generalized median estimators using appropriate kernels have been developed and studied in Brazauskas and Serfling (2000a,b). Arguments given there regarding breakdown points apply in similar fashion here and are given only briefly.

For the estimator $\hat{\mu}_{(k)}$, the relevant kernel $h(x_1, \ldots, x_k) = k^{-1}\sum_{i=1}^{k} x_i \to \infty$ if one or more arguments x_i are taken to $+\infty$. Thus, the GM estimator $\hat{\mu}_{(k)}$ can break down due to upper contamination unless the number M of upper contaminating observations is such that no more than half of the kernel evaluations contain a contaminating observation; that is,

$$\frac{\binom{n}{k} - \binom{n-M}{k}}{\binom{n}{k}} \le 0.5.$$

A similar argument applies in the case of lower contamination. Thus, we have

$$\mathrm{BP}(\hat{\mu}_{(k)}) = n^{-1} \max_{1\le M\le n} \left\{ M: \frac{\binom{n-M}{k}}{\binom{n}{k}} \ge \frac{1}{2} \right\} \to 1 - \left(\frac{1}{2}\right)^{1/k}, \quad n \to \infty.$$

For the estimator $\hat{\sigma}_{(m)}$, similar arguments apply to the relevant kernel, and the same BP is obtained.

A.2 SELECTED CONSTANTS

Table A1 lists values of M_{m-1}, C_m and ζ_m for selected m. The values of ζ_m have been obtained by numerical integration using MAPLE and some "tweaking," after first setting up the computations via certain technical reexpressions of the

Table A1
M_{m-1}, C_m and ζ_m for Selected m

	m				
	2	3	5	7	9
M_{m-1}	0.45494	1.38629	3.35669	5.34812	7.34412
C_m	0.21434	0.34657	0.52586	0.65941	0.77043
ζ_m	0.02658	0.03096	0.02432	0.01890	0.01532

problem. Specifically, it can be shown that, for $m = 2$,

$$w_0(z) = P\{(Z_0 - z)^2 \leq 2M_1\},$$

and for $m \geq 3$,

$$w_0(z) = P\left\{\left(Z_0 - \frac{z + U}{m - 1}\right)^2 \leq \frac{m}{(m - 1)^2} \times [(z + U)^2 + (m - 1)(M_{m-1} - z^2 - V)]\right\},$$

where $U = \sum_{i=1}^{m-2} Z_i$, $V = \sum_{i=1}^{m-2} Z_i^2$, and Z_0, $Z_1, \ldots, Z_{m-2}$ are independent $N(0, 1)$ random variables.

A.3 Asymptotic Normality of Transformed Random Variables

As noted in Section 1.1, $\hat{\sigma}_{\mathrm{ML}}$ is asymptotically normal, with mean σ and variance $(\sigma^2/2)n^{-1}$. Now apply the well-known "delta method," as follows. Given $\hat{\theta}$ asymptotically normal with mean θ and variance Δn^{-1}, if a function $g(\cdot)$ has nonzero derivative at θ, then $g(\hat{\theta})$ is asymptotically normal, with mean $g(\theta)$ and variance $[g'(\theta)]^2 \Delta n^{-1}$. In particular, applying this to $\hat{\sigma}_{\mathrm{ML}}$, with $g(x) = x^2$, we obtain that $\hat{\sigma}^2_{\mathrm{ML}}$ is asymptotically normal, with mean σ^2 and variance $2\sigma^4 n^{-1}$. Likewise, using $g(x) = \sqrt{x}$, we obtain that the asymptotic variance of $\hat{\sigma}_{(m)}$ is that of $\hat{\sigma}^2_{(m)}$ times the factor $1/(4\sigma^2)$. Similar considerations using transformations of asymptotically bivariate normal random variables yield the asymptotic results for $\hat{\eta}$ given in Section 3.1.2. See Serfling (1980, Sect. 3.3) for general treatment.

A.4 GES under Transformation

By standard theory on influence functions (Serfling 1980 and Hampel et al. 1986), it is readily seen that the influence functions of $\hat{\theta}$ and $g(\hat{\theta})$ are related by

$$\mathrm{IF}_{g(\hat{\theta})}(x) = g'(\theta)\mathrm{IF}_{\hat{\theta}}(x).$$

(This is also given as Proposition 4 in Marcel and Rioux 2001.) It follows that the corresponding GES values are related in the same fashion. This leads to the results on GES stated in Sections 2.3.2 and 3.1.1.

A.5 Computational Issues

For situations when the number $\binom{n}{k}$ of kernel evaluations needed for computation of $\hat{\mu}_{(k)}$, or the number $\binom{n}{m}$ of kernel evaluations needed for computation of $\hat{\sigma}^2_{(m)}$, is extremely large (in excess of 10^7), we reduce the computational burden by randomly choosing 10^7 kernel evaluations. Such an approach maintains a high degree of numerical accuracy (up to three decimal places) and renders the computational burden negligible.

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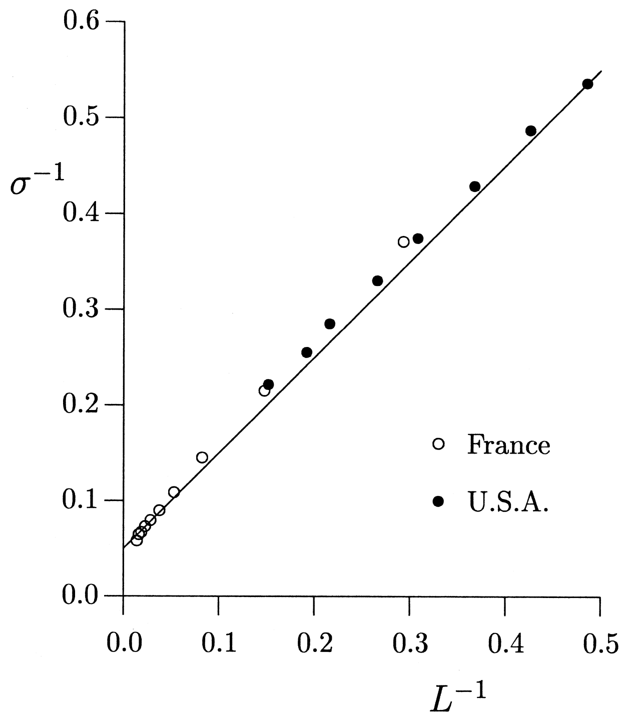
Discussions on this paper can be submitted until April 1, 2003. The author reserves the right to reply to any discussion. Please see the Submission Guidelines for Authors on the inside back cover for instructions on the submission of discussions.

Figure 1
Linear relation between σ^{-1} and L^{-1} in data from France and United States



Note: The line shown has an intercept of 0.05.

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"Efficient and Robust Fitting of Lognormal Distributions," Robert Serfling, October 2002

THIERRY DUCHESNE* AND JACQUES RIOUX†

We would first like to congratulate Professor Serfling for a very interesting paper. We also wish to thank him for arguing in favor of robust estimation in actuarial applications and for demonstrating that robustness does not necessarily come at a high price in efficiency; it has been our own contention for a while. In the abstract and introduction, Professor Serfling mentions that robust and efficient estimators for the lognormal model have not been seriously addressed in the literature. On the one hand, we tend to agree that there is certainly not a huge number of such studies available. On the other hand, we feel like he might have overlooked Duchesne, Rioux, and Luong (1997) (henceforth referred to as DRL).

In DRL, the authors show that the minimum Cramér-von Mises (MCVM) estimator, such as that presented in Hogg and Klugman (1984, Chap. 3), does have good efficiency and robustness properties. While DRL was of a more general scope and did not specifically concentrate on the lognormal model, it included it in its investigation, along with other important models used in actuarial loss modeling such as the gamma and Pareto distributions. Professor Serfling's paper prompted us to revisit our own investigation and compare the generalized median (GM) estimators for the lognormal to the MCVM estimator.

Preliminaries

Since we are considering only the lognormal model, the version of the MCVM estimator that we propose to use is the practically convenient version

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$$\hat{\theta}_{MCVM} = \arg\min_{\theta} \int_0^\infty \left[\Phi\left(\frac{\ln x - \mu}{\sigma}\right) - F_n(x)\right]^2 dF_n(x)$$

$$= \arg\min_{\theta} \sum_{i=1}^n \left[\Phi\left(\frac{\ln X_i - \mu}{\sigma}\right) - F_n(X_i)\right]^2, \quad (1)$$

where $\theta^T = (\mu, \sigma)$, Φ is the cdf of the standard normal distribution and F_n is the empirical cdf. The advantages of this approach in the lognormal case include that (1) it yields estimators with a bounded influence function (to be shown shortly), (2) the loss in efficiency is not dramatic (except for inference about σ under large true values of σ), (3) the estimates are found in a few iterations of the Newton-Raphson algorithm and, (4) the method can readily be adapted to grouped data (DRL, Sect. 5) or to right-censored data (by replacing F_n in Equation (1) by its product-limit correspondent), which is often of importance in actuarial applications.

Moreover, the estimator can be redefined by multiplying the integrand in Equation (1) by a nonnegative weight function $w(x; \theta)$ that allows the user to adjust the trade-off between efficiency and robustness, as shown by Boos (1981) in the case of inference about location parameters, and by DRL, in the case of the MCVM estimator for grouped data.

As Professor Serfling suggested, it is more convenient to work with the sample $Y_1, \ldots, Y_n$, where $Y_i = \ln X_i$. One easily sees that the MCVM estimator defined in Equation (1) also satisfies

$$\hat{\theta}_{MCVM} = \arg\min_{\theta} \int_{-\infty}^\infty \left[\Phi\left(\frac{y - \mu}{\sigma}\right) - F_n(y)\right]^2 dF_n(y), \quad (2)$$

where F_n now represents the empirical cdf based on $Y_1, \ldots, Y_n$. It is also easy to show that $\hat{\theta}_{MCVM}$ defined in Equation (2) is such that $\hat{\mu}_{MCVM}$ is translation equivariant and scale invariant, and that $\hat{\sigma}_{MCVM}$ is scale equivariant and translation invariant.

Estimation of μ, σ and $\eta = \exp\{\mu + \sigma/2\}$

Robustness

DRL derive an expression for the joint influence function of the MCVM estimators of general location and scale parameters. Let $\hat{\theta} = (\hat{\mu}, \hat{\sigma})^T$ denote the MCVM estimators of μ and σ. Let ϕ denote the density of the standard normal and Δ_x be the cdf of a random variable degenerated at x. Applying Equation (2.24) of DRL, we get the following influence function for $\hat{\theta}$:

$$IF(x) = \sigma \begin{pmatrix} \dfrac{\int_{-\infty}^{\infty} [\Phi(y) - \Delta_x(\sigma y + \mu)]\phi^2(y)\, dy}{\int_{-\infty}^{\infty} \phi^3(y)\, dy} \\ \dfrac{\int_{-\infty}^{\infty} y[\Phi(y) - \Delta_x(\sigma y + \mu)]\phi^2(y)\, dy}{\int_{-\infty}^{\infty} y^2\phi^3(y)\, dy} \end{pmatrix}$$

$$= \sigma \begin{pmatrix} \dfrac{\int_{-\infty}^{\infty} \Phi(y)\phi^2(y)\, dy - \int_{(x-\mu)/\sigma}^{\infty} \phi^2(y)\, dy}{\int_{-\infty}^{\infty} \phi^3(y);\, dy} \\ \dfrac{\int_{-\infty}^{\infty} y\Phi(y)\phi^2(y)\, dy - \int_{(x-\mu)/\sigma}^{\infty} y\phi^2(y)\, dy}{\int_{-\infty}^{\infty} y^2\phi^3(y)\, dy} \end{pmatrix}$$

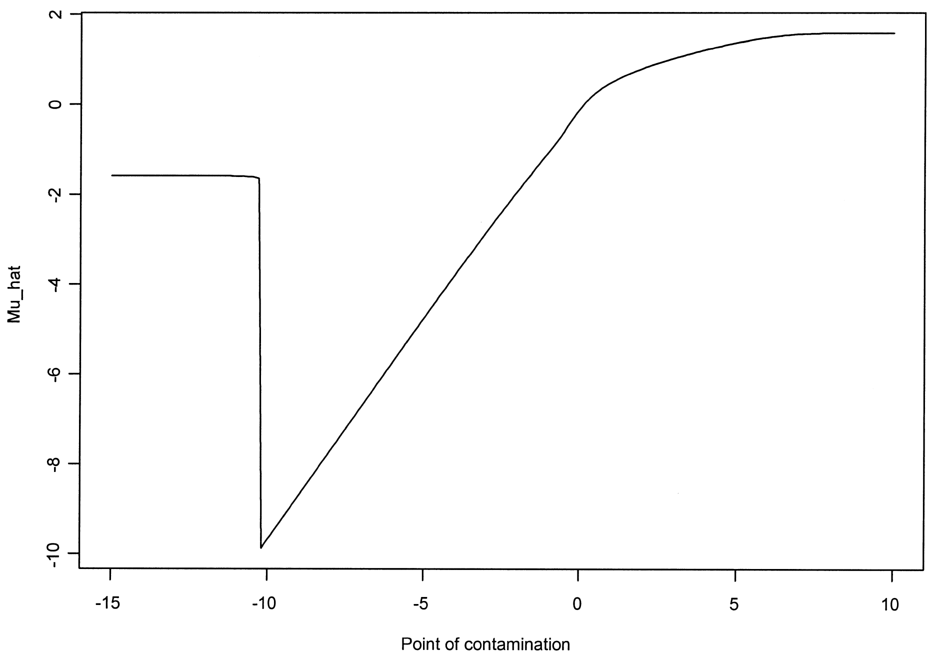
$$= \sigma \begin{pmatrix} [I_1 - I_2(x)]/I_3 \\ [I_4 - I_5(x)]/I_6 \end{pmatrix}.$$

The gross error sensitivity (GES) is given by $\sup_x |IF(x)|$. Let $\text{GES}^* = \sup_x |IF(x)|/\sigma$ be the standardized GES. Using software such as Maple or Mathematica, we get that $I_1 = 1/(4\sqrt{\pi})$, $I_3 = \sqrt{3}/(6\pi)$, $I_4 = \sqrt{3}/(12\pi)$ and $I_6 = \sqrt{3}/(18\pi)$. The function $I_2(x)$ strictly decreases from $1/(2\sqrt{\pi})$ to 0 as x goes from $-\infty$ to ∞, which leads to $\sup_x |(I_1 - I_2(x))/I_3| = \text{GES}^*(\hat{\mu}) = \sqrt{3\pi/2} \approx 1.53499$. As for the function $I_5(x)$, it increases from 0 to $1/(4\pi)$ as x goes from $-\infty$ to μ, then decreases from $1/(4\pi)$ back to 0 as x goes from μ to ∞. This yields $\sup_x |(I_4 - I_5(x))/I_6| = \text{GES}^*(\hat{\sigma}) = 3/2 = 1.5$.

In terms of inference about the lognormal mean $\eta = \exp\{\mu + \sigma/2\}$, we set $\hat{\eta} = \exp\{\hat{\mu} + \hat{\sigma}/2\}$ as our MCVM estimator of this mean. Using the formulas from Section 3.1.1 of Serfling's paper, we get that $\lim_{\sigma\to 0} \text{GES}^*(\hat{\eta}) = 1.53499$ and that $\lim_{\sigma\to\infty} \text{GES}^*(\hat{\eta}) = 1.5$. Hence, uniformly in σ, the MCVM estimator of the lognormal mean exhibits a strong robustness in terms of GES (with $\text{GES}^* \approx 1.5$), being only beat by the GM estimator based on $\hat{\mu}_{(1)}$ and $\hat{\sigma}_{(1)}$ for small values of σ.

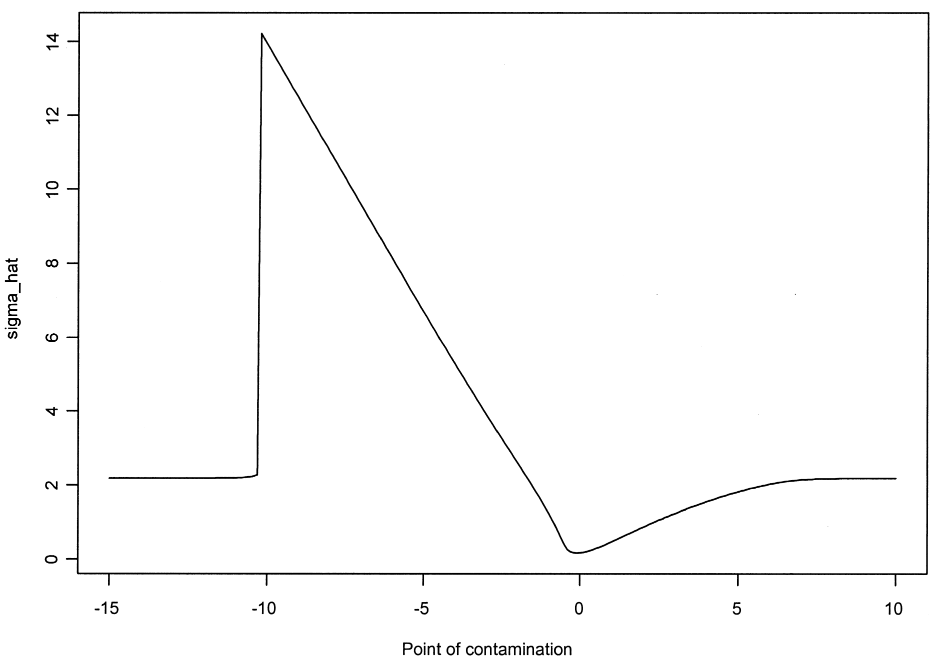
We did not derive analytically the breakdown point of the MCVM estimators for μ and σ. However, numerical evidence suggests that no breakdown occurs for $\hat{\mu}$ or $\hat{\sigma}$ with contamination levels as high as 0.5. Figures 1 and 2 display the behavior of $\hat{\mu}$ and of $\hat{\sigma}$, respectively, as a function of x_0, the contamination point, for a contamination level of 0.5. The graphs were obtained for an idealized sample of size 1,000 where the points are equi-

Figure 1
Value of the MCVM Estimator of μ.



Note: Based on a sample of size 1,000 from a contaminated normal(0, 1) distribution as a function of the point of contamination, with 50% contamination. When the point of contamination is too small or large, the MCVM statistic is minimized when the value of μ remains close to the uncontaminated data, and thus $\hat{\mu}$ is not dragged to $\pm\infty$.

Figure 2
Value of the MCVM Estimator of μ.



Note: Based on a sample of size 1,000 from a contaminated normal(0, 1) distribution as a function of the point of contamination, with 50% contamination. When the point of contamination is too small or large, the MCVM statistic is minimized when the value of σ remains close to the uncontaminated data, and thus $\hat{\sigma}$ is not dragged to ∞ or 0.

Table 1
Asymptotic Relative Efficiency of the MCVM Estimator of the Lognormal Mean for Different Values of the Scale Parameter σ

σ	0	2.5	5.0	7.5	10.0	20.0	∞
ARE($\hat{\eta}$)	0.914	0.699	0.665	0.657	0.654	0.651	0.650

distant (in the probability space) quantiles. We show the graphs for x_0 values in $[-15, +10]$. This provides ample coverage of the probability space and essentially can be thought of as $(-\infty, +\infty)$, as we have verified extensively. We have also verified that the behavior of the graphs is consistent over sample sizes varying from as small as 10 to as large as 100,000.

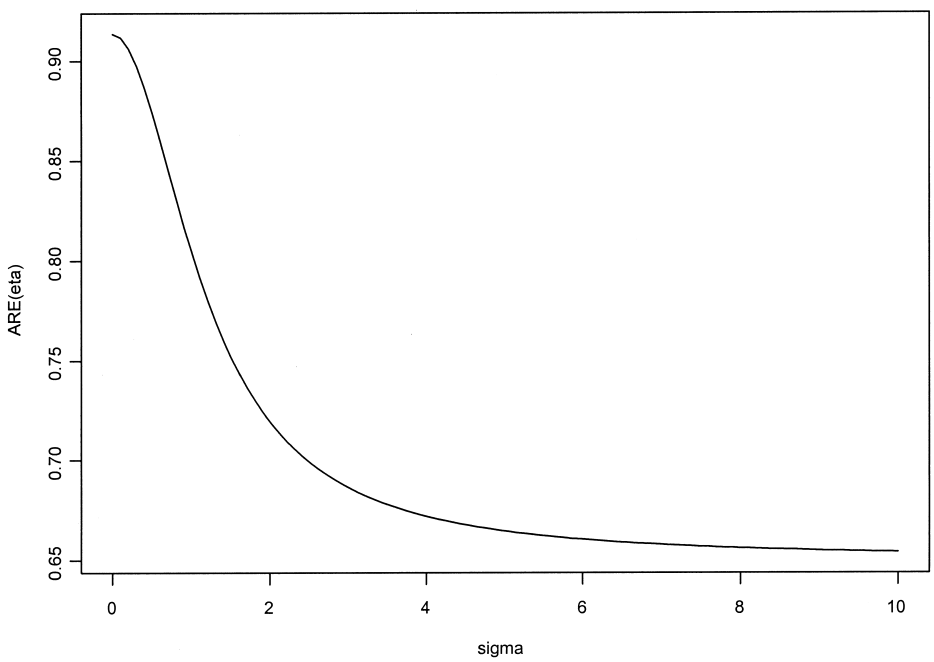
Efficiency

With the MCVM estimator emphasizing robustness so much, we would expect it to pay a high price in terms of asymptotic efficiency. Here we use the fact that $\text{AsVar}[\theta] = \text{Var}[IF(X)]/n$ to derive the asymptotic variance of $\hat{\theta}$. To evaluate $\text{Var}[IF(X)]$, we can again use Maple or Mathematica. We get that $E[I_2(X) - I_1] = E[I_5(X) - I_4] = E[(I_2(X) - I_1)(I_5(X) - I_4)] = 0$, as the general results of DLR and the equi- and in-variance properties would have led us to find. Numerical integration yields $E[(I_2(X) - I_1)^2] \simeq 0.009242108733$ and $E[(I_5(X) - I_4)^2] \simeq 0.7686917683$. We, thus, obtain the following variance for $IF(X)$:

$$\widehat{\text{Var}}[IF(X)] = \sigma^2\begin{pmatrix} 0.0092421/I_3^2 & 0 \\ 0 & 0.76869/I_6^2 \end{pmatrix}$$
$$= \sigma^2\begin{pmatrix} 1.0946 & 0 \\ 0 & 0.76869 \end{pmatrix}.$$

This corresponds to an asymptotic relative efficiency of 91.4% for $\hat{\mu}$, 65.0% for $\hat{\sigma}$ and $(1 + 0.5\sigma^2)/(1.0946 + 0.76869\sigma^2)$ for $\hat{\eta}$, which varies between 91.4% at $\sigma = 0$ and 65.0% as $\sigma \to \infty$ (see Table 1). The good news is that the efficiency of $\hat{\eta}$

Figure 3
Asymptotic Relative Efficiency of the MCVM Estimator of the Lognormal Mean η as a Function of the True Value of the Scale Parameter σ.



remains reasonably high under small values of σ, which are the values most often found in practice. The bad news is that, as we can see in Figure 3, the efficiency drops rapidly as the value of σ increases.

In summary, the MCVM estimator of the parameter μ behaves like a compromise between the GM estimators with $k = 1$ and $k = 2$, whereas the MCVM estimator of the parameter σ puts much more emphasis on robustness than any of the GM estimators considered by Serfling, thereby being much less efficient. In terms of inference about the lognormal mean η, this means that efficiency is traded off for more robustness as the true value of the scale parameter increases.

Conclusion

It appears that the GM estimators do provide a very good balance between robustness and efficiency. Minimum distance methods such as the MCVM estimator recalled here provide more robustness automatically and have decent efficiency. Presumably, efficiency of the MCVM estimator can be much improved with the right choice of weight function, while maintaining good robustness properties.

We would like to thank Professor Serfling again for helping raise the awareness of robustness properties in loss distribution modeling among the actuarial community with his fine paper.

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AUTHORS' REPLY*

The treatment of minimum Cramér-von Mises (MCVM) estimation for lognormal parameters in Duchesne, Rioux, and Luong (1997) had indeed, regrettably, escaped my attention, and I am very grateful to Professors Duchesne and Rioux for their comparison between the MCVM estimators in that paper with the generalized median (GM) estimators in my paper.

My goal (see p. 96) was to develop estimators offering very high asymptotic relative efficiency along with an adequately high breakdown point and adequately low gross error sensitivity. That is, *high efficiency* was favored, subject to the constraint of maintaining *adequate robustness*. In this spirit, it was shown that the GM estimators provide excellent options of this kind. However, I also mentioned that other criteria could be considered and that other robust estimation approaches such as trimmed methods and M-estimation could, in principle, be competitive. Certainly, MCMV, as a well-established approach, could well have been included in this list.

For the lognormal problem, in particular, Professors Duchesne and Rioux show that, if efficiency is emphasized less stringently and robustness emphasized more, then MCVM estimators offer competitive alternatives to the GM estimators. Furthermore, as they mention, the weaker efficiency of the MCVM approach might be improved by incorporating suitable weight functions. While the degree of possible improvement is uncertain, it would be very interesting to see the MCVM approach carried to its full potential in the lognormal problem.

It is important, as Professors Duchesne and Rioux point out, to "raise the awareness of robustness properties in loss distribution modeling." I agree that this should and can be accomplished without undue loss of efficiency. In this regard, my personal orientation has been that one should determine what level of robustness is *adequate* and then pursue the *highest possible efficiency* subject to that chosen robustness level. If we injudiciously overemphasize robustness, then we pay too high a price in terms of loss of efficiency, and unnecessarily.

From this viewpoint, the GM estimators provide a reasonably attractive and conceptually appealing solution, although certainly not exclusively so. Further investigations should yield superior estimators. I heartily thank Professors Duchesne and Rioux for providing a very important analysis that broadens our perspective and stimulates further interest.

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