Model and Data Uncertainty in Model Averaging in Dose Response Assessment

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Abstract

Food-borne infection is caused by intake of foods or beverages contaminated with microbial pathogens. Dose-response modeling is used to estimate exposure levels of pathogens associated with specific risks of infection or illness. When a single dose-response model is used and confidence limits on infectious doses are calculated, only data uncertainty is captured. We propose a method to estimate the lower confidence limit on an infectious dose by including model uncertainty and separating it from data uncertainty. The infectious dose is estimated by a weighted average of effective dose estimates from a set of dose-response models via a Kullback information criterion. The confidence interval for the infectious dose is constructed by the delta method, where data uncertainty is addressed by a bootstrap method. The actual coverage probabilities of the lower confidence limit are addressed under dose-response shapes that can be commonly found in real data sets. Our model-averaging method achieves coverage close to nominal in almost all cases, thus providing a useful and efficient tool for accurate calculation of lower confidence limits on infectious doses.

Key Words: bias-skewness correction; confidence limit; data uncertainty; food safety; Kullback information criterion

1. Introduction

Food safety is a critical issue in public health. The Centers for Disease Control and Prevention (CDC, 2011) estimated that there are 48 million illnesses, 128 thousand hospitalizations, and 3000 deaths due to foodborne pathogens every year in the United States. Among well-known pathogens, *Norovirus* and *Salmonella* are included in the top five pathogens causing domestically acquired foodborne illness and resulting in hospitalization and death. Another well-known pathogen, *E. coli* O157, was included in the top five pathogens causing domestically acquired foodborne illnesses resulting in hospitalization.

In order to control diseases caused by microbial contaminants in food, it is essential to assess their dose-response relationships as accurately as possible. However, definitive

dose-response data on humans at low levels of contamination likely to occur in practice are scarce to nonexistent. Hence, when sufficient animal or human data at high doses are available to allow dose-response modeling, allowable contamination levels of specific microorganisms in food can be derived using infectious dose (*ID*) levels derived from these models as "points of departure" for low-dose extrapolation. This approach to setting exposure levels is equivalent to the benchmark dose (BMD) approach used in chemical risk assessment (EPA, 2000). Reliable methods for deriving such ID_p levels ($0.01 \leq p \leq 0.10$) are essential, where ID_p is defined as a dose that causes a response (infection or illness) at a predetermined risk level, p.

Dose-response models with one, two, and three parameters have been proposed for dose-response modeling in microbial risk assessment (MRA) (Kodell et al, 2002; Marks et al, 1998; Moon et al., 2004). The simplest model is the one-parameter exponential model, $P(d;\ \alpha) = 1 - exp(-\alpha d), \alpha > 0,$ which can be derived from basic biological assumptions considering low numbers of pathogens as discrete particles (Haas, 1983). Even though the exponential model has low-dose linearity, it is often not flexible enough to provide an adequate fit to dose-response data on pathogenic microorganisms. The Beta-Poisson (BP) model (Haas, 1983; Furumoto and Mickey, 1967; Haas et al., 1999), which includes slightly more complex biology, has been used in MRA(WHO, 2001a, 2001b, 2002). However, the adequacy of the BP model as a potential "default" model for MRA has been questioned (Marks et al., 1998). It can be shown that the exponential and the BP models are dose-response pattern-specific so that they may not be suitable models under a certain dose-response pattern (e.g. a hypothetical sublinear pattern in Figure 1). Other two-parameter models include the Log-Normal (LN), the Log-Logistic (LL), and the Extreme-Value (EV) models (Pinsky, 2000). The best-known three-parameter model is the Weibull-Gamma (WG) model (Farber et al., 1996). However, three-parameter models require data at four or more dose levels, which may not be readily available for many microbial agents. It was noted that the BP and LL models are special cases of the WG model (Kodell et al, 2002).

Several dose-response models often provide reasonably good fits to the data in the experimental dose range but can yield very different infectious dose (*ID*) estimates in the low-dose range, even with infection rates as high as $0.01 \leq p \leq 0.10$. Hence, it is undesirable to choose only one model and estimate an *ID* based on the chosen model. In order to account for model uncertainty, model averaging (MA) methods have been proposed (FDA/FSIS, 2003; Moon et al., 2004, 2005; Bailer et al., 2005; Faes et al., 2007; Wheeler and Bailer, 2007; Namata et al., 2008).

Kang et al. (2000) used four two-parameter models to demonstrate how model uncertainty can be addressed in MRA using the Akaike information criterion (AIC) (Akaike, 1974) to average the individual-model *ID*s. Kodell et al. (2002) presented a general framework for generating dose-response models in the interest of deriving potential competitors for the three-parameter Weibull-Gamma (WG) model (Farber et al., 1996). Moon et al. (2004) suggested that two-parameter dose-response models for MRA reflected at least as much model uncertainty as three-parameter models. Moon et al. (2005) used maximum likelihood estimates (MLEs) from a binomial log-likelihood function to estimate parameters of dose-response models and integrated model uncertainty into estimating *ID*s using weights obtained from the Kullback information criterion (Cavanaugh, 1999; Kim and Cavanaugh, 2005) as a measure of model variability.

Faes et al. (2007) showed the necessity of model averaging by comparing the *ID* based on a set of fractional polynomials to the *ID* from the selected best model. Namata et al. (2008) also investigated model averaging in MRA using fractional polynomials, suggesting that the common two-parameter models were not sufficiently diverse to give appropriate confidence coverage in model averaging.

Bayesian model averaging provides a coherent approach for accounting for model uncertainty (Hoeting et al., 1999). Bailer et al. (2005) illustrated Bayesian model averaging with a simple Bayesian information criterion (BIC) approximation (Schwarz, 1978; Kass and Wasserman, 1995) under the assumption of the unit information prior on the parameter space for addressing uncertainty in the selection of models when generating risk estimates. Their emphasis was on the Bayesian analysis of model uncertainty to obtain a model-averaged summary.

Recently, Wheeler and Bailer (2007) investigated an alternative MA approach to estimate *ID*s (benchmark doses, or BMDs, in their terminology) based on a weighted "average model" and illustrated their method with dose-response lung cancer data on rats. Their *ID* estimates are obtained by solving the weighted average model, and the weights are determined by AIC. The lower confidence bound on the *ID* was obtained by the bootstrap percentile method. They investigated the coverage of their *ID* lower confidence estimates with linear and sublinear dose-response patterns.

We propose a new method to estimate the lower confidence limit (LCL) for an *ID* under the consideration of both model uncertainty and data uncertainty. Following Moon et al. (2004), four two-parameter dose-response models are considered a sufficiently diverse set of models. Parameters of the models are estimated by the method of maximum likelihood. The *ID* is estimated by a KIC-weighted average of effective dose (ED) estimates from the dose-response models. We introduce a new variance formula and calculate the variance of the *ID* estimate with separate components for model uncertainty and data uncertainty via a bootstrap method. The LCL for the *ID* is constructed assuming the *ID* is log-normally distributed. The delta method is used to approximate the variance of the log-*ID* estimate.

The use of the natural log transformation of *ID* estimates to normalize the *ID* distribution is not unusual. In a similar line of this study, Faes et al. (2007) used a log-normal assumption, and their equation (8) is similar to our idea, but with a different algebraic expression. The log transformation of *ID* estimates makes the distribution more symmetric. Even after the log-transformation, skewness still exists. By the same token, Figure 5 in Wheeler and Bailer (2007) showed a skewed distribution at low doses. However, our proposed method with nonparametric quantile estimation further adjusts the skewness via the BCa bootstrap method.

2. Methods

2.1 Infectious Dose Estimation

Let I denote the number of independent dose groups. Let n_i denote the number of independent subjects in the i-th group, $i = 1, 2, \ldots, I$. Let X_i denote the number of subjects infected or with symptoms in the i-th group. Assume that X_i has a binomial distribution with n_i and $P(d_i; \boldsymbol{\theta})$, where $P(d_i; \boldsymbol{\theta})$ is a dose-response model and $\boldsymbol{\theta}$ is a

Table I: Dose-response models for microbial risk assessment

Name	Model	Domain of Parameters
Beta Poisson (BP)	$P(d;\alpha,\beta) = 1 - \left(1 + \frac{d}{\beta}\right)^{-\alpha}$	$\alpha > 0, \beta > 0$
Log-normal (LN)	$P(d;\alpha,\beta) = \Phi\left(\frac{\ln(d) - \alpha}{\beta}\right)$	$-\infty < \alpha < \infty, \beta > 0$
Log-logistic (LL)	$P(d;\alpha,\beta) = \left[1 + \exp\left(-\frac{\ln d - \alpha}{\beta}\right)\right]^{-1}$	$-\infty < \alpha < \infty, \beta > 0$
Extreme-value (EV)	$P(d;\alpha,\beta) = 1 - \exp[-\exp(\alpha + \beta \ln d)]$	$-\infty < \alpha < \infty, \beta > 0$

parameter vector. The estimation of $\boldsymbol{\theta}$ is accomplished by maximizing the binomial log-likelihood function (Kodell et al., 2002), $l(\boldsymbol{\theta}) \propto \sum_{i=1}^{I}[X_i \ln\{P(d_i, \boldsymbol{\theta})\} + (n_i - X_i)\ln\{1 - P(d_i, \boldsymbol{\theta})\}]$.

In this paper, four two-parameter dose-response models are used as shown in Table I. The two-parameter models include the Beta Poisson (BP), log-normal (LN), log-logistic (LL), and extreme-value (EV) models. The maximization is performed by employing the SAS procedure NLMIXED (SAS code is available from the authors upon request). It is found that the dose levels may be scaled by a constant scale factor in order to achieve greater stability in the maximum likelihood estimates. The model fitting procedure is invariant to the transformation. Therefore, the estimates of the effective doses are the same regardless of scaling.

When the parameters of a dose-response model are estimated, the effective dose at a specific risk level p (ED_p) for a given model, which itself can be regarded as a fixed "parameter," can be estimated by substituting the "other" parameter estimates and solving for dose d. Every two-parameter model in this study has a closed-form solution for d. If a model has no closed-form solution, the bisection method, a simple and robust root-finding algorithm, can be employed to obtain the estimate of ED_p. Estimates of effective dose at risk levels of 1% (ED_{01}) or 10% (ED_{10}) represent how many microorganisms can produce a 1% or 10% increase in infection or illness, relative to the control response (usually assumed to be zero). These effective dose levels (ED_{01} and ED_{10}) correspond to the lower and upper limits of the risk range (1% and 10%) generally recommended for restricting the calculation of BMDs in chemical risk assessment (correspondingly, *ID*s in microbial risk assessment). For quantal data, an excess risk above background risk of 10% is known as the default benchmark risk (BMR) (Nordberg et al., 2007), which here we term the IDR.

In order to estimate an *ID*, effective doses obtained from the two-parameter models (BP, LN, LL and EV) are averaged using Kullback weights. We let m be the dimension of the parameter vector. The *KIC* is defined as $KIC = -2l(\widehat{\boldsymbol{\theta}}) + 3m$. We define the Kullback weight for the j-th model, denoted by $w_j^{KIC} = \exp(-\Delta_j^{KIC}/2)/\sum_{j=1}^{K}\exp(-\Delta_j^{KIC}/2)$, where $\Delta_j^{KIC} = KIC_j - \min(KIC_1, KIC_2, \ldots, KIC_K)$, the sum of the weights $\sum_{j=1}^{K} w_j^{KIC} = 1$, and K represents the total number of candidate dose-response models. The model with the

minimum KIC_j has the most contribution to the *ID* estimate. The point estimate of the *ID* is obtained as

$$\widehat{ID}_p = \sum_{j=1}^{K} w_j^{KIC} \widehat{ED}_{p,j}, \quad (1)$$

where $\widehat{ED}_{p,j}$ represents the effective dose estimate from the j-th model at p (1% or 10%) risk level.

2.2 Lower Confidence Limit on Infectious Dose

In this paper a method to construct a lower confidence limit (LCL) on an ID_p at excess infection rate IDR = p is proposed. We assume that the distribution of *ID* estimates is not symmetric but approximately log-normal (see also Faes et al., 2007). The proposed method incorporates model-averaging with KIC-weights (Moon et al., 2005) and our nonparametric quantile estimator z_α^* via the bias-corrected and accelerated (BCa) bootstrap method (Efron, 1987; Efron and Tibshirani, 1993) in order to adjust for both bias associated with empirical estimates of infectious doses (*ID*) and skewness in the sampling distributions of log-*ID*. The main contribution of the paper is the inclusion of model uncertainty and separation of the model uncertainty from data uncertainty to estimate the lower confidence limit on an *ID*.

It is reasonable to assume that each $\widehat{ED}_k$ has expectation ED_k and that the ED_k s themselves have expectation $ID = \sum_{k=1}^{K} w_k^{KIC} ED_k$ as similar to assumptions of Faes et al. (2007). In order to obtain an LCL on an *ID* at excess risk BMR, we propose the variance of *ID* as follows:

$$Var(ID) = \sum_{k=1}^{K} Var(w_k^{KIC} ED_k) + 2\sum_{j=1}^{K}\sum_{k=j+1}^{K} Cov\left(w_j^{KIC} ED_j, w_k^{KIC} ED_k\right), \quad (2)$$

where the variance term can be decomposed to

$$Var\left(w_k^{KIC} ED_k\right) = Var\left\{E\left(w_k^{KIC} ED_k \middle| M_k\right)\right\} + E\left\{Var\left(w_k^{KIC} ED_k \middle| M_k\right)\right\}, \quad (3)$$

and the covariance term can be obtained as

$$Cov\left(w_j^{KIC} ED_j, w_k^{KIC} ED_k\right) = E\left\{Cov\left[\left(w_j^{KIC} ED_j \middle| M_j\right), \left(w_k^{KIC} ED_k | M_k\right)\right]\right\}, \quad (4)$$

where M_k indicates the model k. The proposed formula is based on the law of total variance shown in basic statistics textbooks (Devore, 1991; Burnham and Anderson, 2002). It is also similar to one used in Faes et al. (2007). However, a main difference from the one in Faes et al. (2007) is that the weight vector w_k is treated as a random quantity inside the variance and covariance operators rather than a fixed quantity because even the weights vary from dataset to dataset. We note that there is no covariance among the model means other than the underlying variance itself because *ID* estimates for each model do not co-vary in any defined or measurable way. Hence, we consider only data uncertainty in the covariance term. Therefore, equation (2) with equations (3) and (4) can be rewritten as

$$Var(ID) = \sum_{k=1}^{K} Var\left\{E\left(w_k^{KIC} ED_k \middle| M_k\right)\right\} + \sum_{j=1}^{K}\sum_{k=1}^{K} E\left\{Cov\left[\left(w_j^{KIC} ED_j \middle| M_j\right), \left(w_k^{KIC} ED_k | M_k\right)\right]\right\}. \quad (5)$$

In this framework the first and second terms in equation (5) represent model uncertainty and data uncertainty, respectively. We estimate $Var(ID)$ via

bootstrapping by generating B bootstrap samples and by estimating w_k^{KIC} and ED_k for each dose-response model in each bootstrap sample.

To construct an LCL on an *ID*, we apply the delta method. We assume that $\ln(ID)$ is approximately normally distributed with mean $\ln(\widehat{ID})$ and variance $Var[\ln(\widehat{ID})]$. The $Var[\ln(\widehat{ID})]$ can be approximated by $\widehat{Var(ID)}/\widehat{ID}^2$ using the delta method. Then, an LCL on $\ln(ID)$ can be estimated as

$$\widehat{LCL} = \ln(\widehat{ID}) + z_\alpha^* \frac{\sqrt{\widehat{Var(ID)}}}{\widehat{ID}}, \tag{6}$$

where $\widehat{Var(ID)}$ is obtained by equation (5). A critical value z_α^* corresponding to the $100(1-\alpha)\%$ confidence level is estimated by

$$z_\alpha^* = \frac{\ln(\widehat{ID})_{(L)} - m}{\sqrt{v}}, \tag{7}$$

where $\ln(\widehat{ID})_{(L)}$ is the LCL estimate for $\ln(ID)$ from the BCa bootstrap method, and m and v are the mean and variance of log-*ID* estimates from the B bootstrap samples, respectively. An LCL on the *ID* is obtained by applying the anti-log transformation $\exp(\widehat{LCL})$.

The estimation of $\ln(\widehat{ID})_{(L)}$ is summarized as follows: First, from each bootstrap sample, B bootstrap estimates of *ID*'s are obtained. Next, the bias correction factor z_0 is obtained as $\hat{z}_0 = \Phi^{-1}(\#[\ln(\widehat{ID}_b) < \ln(\widehat{ID})]/B)$, where $\ln(\widehat{ID}_b)$ can be estimated from the B bootstrap samples ($b = 1, 2, \ldots, B$), and $\ln(\widehat{ID})$ can be estimated from the original data. Next, the acceleration factor a can be obtained as $\hat{a} = \sum_{b=1}^{B}[\overline{\ln(ID)^*} - \ln(\widehat{ID}_b)]^3 / \left[6\left(\sum_{b=1}^{B}[\overline{\ln(ID)^*} - \ln(\widehat{ID}_b)]^2\right)^{3/2}\right]$, where $\overline{\ln(ID)^*}$ is the mean of the bootstrap estimates $\ln(\widehat{ID}_b)$. Finally, we estimate $\ln(\widehat{ID})_{(L)}$ by calculating $L = \lfloor B \times \alpha_1 \rfloor$, where $\alpha_1 = \Phi[\hat{z}_0 + (\hat{z}_0 + z_\alpha)/(1 - \hat{a}(\hat{z}_0 + z_\alpha))]$.

This process is computationally intensive. Instead of using one thousand or more bootstrap samples, a smaller bootstrap sample size was determined by a simulation study. The mean differences between z_α^* obtained from a bootstrap sample size of 1000 and z_α^* obtained from bootstrap sample sizes less than 1000 were compared using both a *t*-statistic and a Wilcoxon rank sum statistic as a nonparametric alternative. The bootstrap sample size $B = 700$ was selected based on *p*-value > 0.10.

Table II: *Echovirus* 12 virus data (Teunis et al., 1996) from human volunteers

[1]Dose (d_i)	[2]Total (n_i)	[3]Infection (X_i)	[4]Probability (p_i)
330	50	15	0.3000
1000	20	9	0.4500
3300	26	19	0.7308
10000	12	12	1.0000

[1]Dose: ingested numbers of pfu (plaque forming units)
[2]Total: number of subjects at a certain dose
[3]Infection: number of subjects with infection (excretion of echovirus or seroconversion, or both)
[4]Probability: sample proportion (X_i/n_i)

3. ILLUSTRATIONS

3.1 Infectious Dose Estimation

The proposed method is illustrated with *Echovirus* 12 data (Teunis et al., 1996) from human volunteers (Table II). This section is not intended as a formal risk assessment for *Echovirus* 12, but is simply an illustration of the proposed method on a real dataset. In a formal risk assessment, it is likely that different *ID*s would be estimated for subpopulations of differential sensitivity (e.g., FAO/WHO, 2004). Consideration of that and other sources of uncertainty is beyond the scope of this paper.

One hundred and eight healthy human volunteers participated in a microbial dose-response experiment with *Echovirus* 12. The subjects were divided into four dose-level groups: 330, 1000, 3300, and 10000 pfu (plaque forming units). Then the subjects were exposed to the microbial pathogen, *Echovirus 12*, and their binary response (infected or not infected) was recorded. The estimates of ID_{01} and ID_{10} from the four two-parameter dose-response models and their Kullback weights were obtained by the methods illustrated in Section 2.1 (See Moon *et al*., 2005). Figure 1 shows a plot of the *Echovirus* 12 data overlaid with the four fitted two-parameter dose-response models. A superlinear pattern is apparent in the plotted data and in the fitted models. Since the lowest dose-level is 330 pfu with estimated infection probability 0.30 (See Table II), an extrapolation is inevitable even to estimate *ID* at risk levels $p = 0.01$ and $p = 0.10$.

As shown in Table III, the estimate of the *ID* at risk level $p = 0.01$ varies from 1.75 pfu (EV) to 17.11 pfu (LN). At risk level $p = 0.10$, it varies from 60.86 pfu (EV) to 104.26

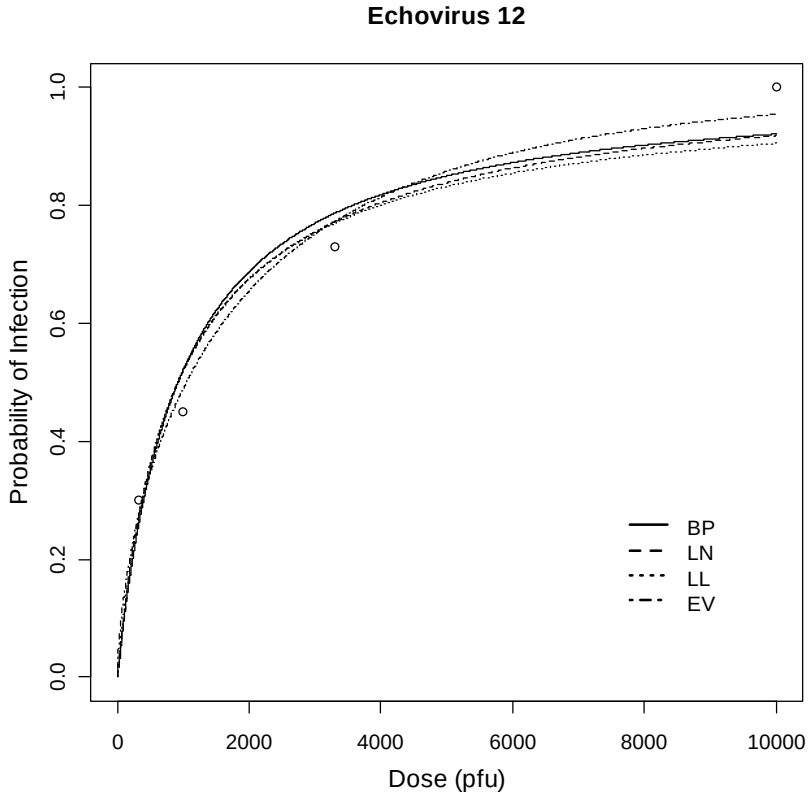

Figure 1. Four two-parameter dose-response models fitted to *Echovirus* 12 data

Table III. Estimates of ID_{01} and ID_{10} for *Echovirus 12* with Kullback Information Criterion and Kullback Weight

Model	m	$\widehat{ID}_{01}$	$\widehat{ID}_{10}$	ML Value	KIC_j	Δ_j^{KIC}	w_j^{KIC}
BP	2	9.5030×10^0	1.0426×10^2	–61.0565	128.1130	1.5706	0.1841
LN	2	1.7107×10^1	1.0239×10^2	–60.8679	127.7358	1.1934	0.2223
LL	2	7.1644×10^0	9.0319×10^1	–61.0262	128.0524	1.5100	0.1898
EV	2	1.7499×10^0	6.0855×10^1	–60.2712	126.5424	0.0000	0.4038

pfu (BP). Note that even though the models appear inseparable at low doses in Figure 1, in reality there is considerable model-to-model variation. In the estimation of ID_{01} and ID_{10}, the EV model is known to be conservative (Moon et al., 2004), and the result is also reflected in *Echovirus* 12 data. Using equation (1) in Section 2.1, the estimates of *ID* at risk levels $p = 0.01$ and $p = 0.10$ are 7.62 pfu and 83.67 pfu, respectively. Based on the maximum likelihood value, the EV model results in the best fit to the data; thus the estimate from the EV model has the most contribution in the estimation of *ID*. Faes *et al*. (2007), Wheeler and Bailer (2007), and Namata *et al*. (2008) also accommodated the different *ID* estimates from competitive models via model-averaging based on AIC weights as proposed by Buckland *et al*. (1997). The estimation of *ID* should not be sensitive to the choice of KIC or AIC in *Echovirus* 12 data because sample size is relatively large. In fact, the AIC-based estimates of *ID* are identical to KIC-based estimates because every model has the same number of parameters ($m = 2$) in this example.

We note that the BP model evaluated is an approximation. However, the parameter estimates ($\boldsymbol{\theta}$) are consistent with use of the approximation for the *Echovirus* data set. For the BP model shown in Table I it is known that if $(\alpha/\beta)\cdot d$ is very small (e.g. much less than one), $P(d \mid \alpha, \beta) \approx (\alpha/\beta)\cdot d$. Thus the relationship between the dose and the probability of infection (or illness) is approximately linear (with the slope of α/β) if the condition is met. In the *Echovirus* data, the estimated parameters for $(\alpha, \beta)$are (1.0563, 994.0256), and the estimated effective dose corresponding to $p = 0.01$ (ED_{01}) is 9.5032 from the BP model. From this result, at low dose, $\hat{\alpha}/\hat{\beta} \approx .001$ is an "approximate" slope for the approximate linear relationship between dose and probability of infection. We can see that $\hat{\alpha}/\hat{\beta}\widehat{ED}_{01} = (1.0563/994.0256) * 9.5032 = .0101 \approx .01$, which is consistent with the result for the *Echovirus* dataset.

3.2 Estimation of the Lower Confidence Limit on ID_p

Under the assumptions of unbiasedness and perfect symmetry (normality) of log-*ID* estimates (i.e., using $z_{0.05} = -1.645$), 95% lower limits of *ID* are estimated as 1.09 and 33.32 for $p = 0.01$ and $p = 0.10$, respectively, by equation (6) with $z_{0.05} = -1.645$. On the other hand, using the proposed variance formula with z_α^* as illustrated in Section 2.2, 95% lower limits for ID_{01} and ID_{10} are 0.71 and 28.04 pfu, respectively, by equations (6) and (7). Corresponding values of $z_{0.05}^*$ for $p = 0.01$ and $p = 0.10$ in equation (7) are estimated as –2.02 and –1.95, respectively. After correcting bias and skewness, the estimated 95% LCLs are lower than before the corrections.

It is not surprising to observe the substantially different estimates of LCLs between using z_α and using z_α^*. As shown in Figures 2 and 3, the sampling distribution of log-*ID* from bootstrap samples does not exhibit perfect symmetry. The bootstrap sample skewness is –0.49 and –0.67 for log-ID_{01} and log-ID_{10}, respectively. Furthermore, the estimate of *ID*

is known to be biased, and the sampling distributions of the number of infections are highly skewed especially in the low-dose group as discussed by Wheeler and Bailer (2007). As illustrated in the simulation study, corrections in bias and skewness are desired to improve the accuracy of coverage rate regardless of dose-response shapes, sample sizes, and risk levels. Obtaining an LCL on a *ID* takes about 30 minutes, mostly for the estimation of model parameters using the Newton-Raphson algorithm in SAS NLMIXED repeated in $B = 700$ bootstrap samples. However, the computation of z_α^* does not require additional time for the model-fitting procedure. The proposed method is a useful and efficient tool for the estimation of lower confidence limits with higher accuracy.

4. Discussion

The main contribution of the paper is the inclusion of model uncertainty, and the separation of model uncertainty from data uncertainty to estimate the lower confidence limit on an *ID*. The proposed method incorporates model-averaging with KIC-weights (Moon et al., 2005) and our nonparametric quantile estimator z_α^* via the BCa bootstrap method in order to adjust for both bias of the *ID* estimate and skewness in the sampling distributions of log-*ID*. We note that the estimation of a lower confidence limit on an *ID* should not be sensitive to the choice of KIC-weights or AIC-weights when the dose-response models have the same number of parameters. In our study, the number of parameters *m* is 2 for every model in model-averaging.

Although the typical BCa bootstrap method (without the proposed variance formula) accounts for both bias and skewness in data uncertainty to improve the bootstrap percentile method, it does not properly account for model uncertainty. As a result, the coverage rates do not meet the desired confidence level 0.95 (Moon et al., 2013). This result highlights the importance of properly accounting for model uncertainty in a model-averaging method. Our nonparametric quantile estimator z_α^* using the BCa bootstrap

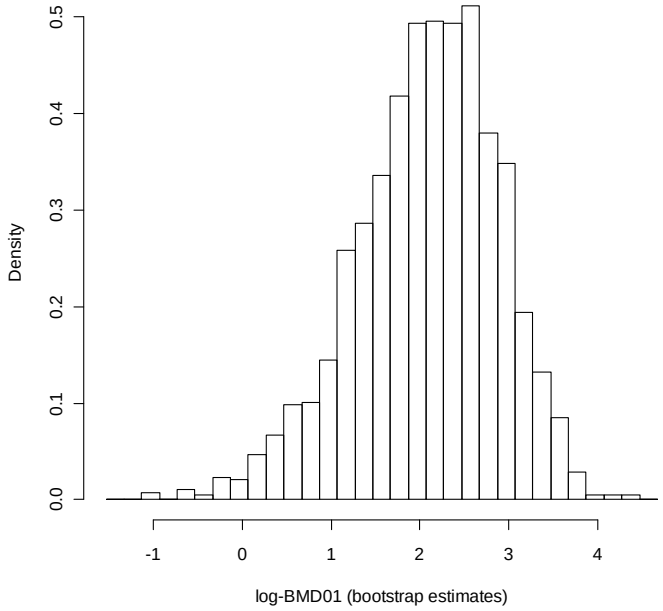

Figure 2. Sampling distribution of log(ID_{01}) from bootstrap samples

method consistently outperforms z_α from the normality assumption regardless of dose-response patterns, risk levels, and sample sizes. In other words, ignoring bias and skewness may lead to inaccurate coverage of the LCL on the *ID*.

In the Monte Carlo simulation study, the three configurations, superlinear, linear, and sublinear, represent various microbial risk patterns showing the relationship between the dose to an agent and the severity of associated adverse response relating to the food-borne contamination process. The coverage rates from the proposed method in various simulation settings are near the nominal level except for sublinear with $n_i = 10$ as shown in Moon et al. (2013). It may be due to a small sample size and an extreme sublinear pattern, which may lead to abnormally high estimates of ID_p. None of the four compared methods meets the nominal coverage for the sublinear pattern with the small sample size. For $n_i = 10$ and the given sublinear pattern, it is more likely to have zero binary responses in the low dose-levels, and the estimates of ID_p become abnormal.

The coverage of ID_{01} in the superlinear pattern with $n_i = 30$ appears to be lower than the nominal level (Moon et al., 2013). It may be due to inadequate representation of model uncertainty because every model is fitted close to each other and an indistinguishable amount of model variation exists near the 1% level (Moon et al., 2013). Elimination of wrong models may be a possible remedy, but the determination of wrong models is another challenge. Furthermore, if a wrong model exists, it has been already treated by the Kullback information criterion by imposing a small or negligible weight in the estimation as a penalty for the poor fit.

Another issue is the sample size for each dose group. In the superlinear pattern, the coverage for ID_{01} is consistently lower among the methods when the sample size per dose group is higher. In other words, increasing the sample size produces an inaccurate result, which is a contradiction to statistical common sense. Overall, the difference between the sample sizes $n_i = 10$ and $n_i = 30$ seems to be sensitive to patterns and/or risk levels. The exact relationship is still not revealed, and a further investigation is deferred to a future study.

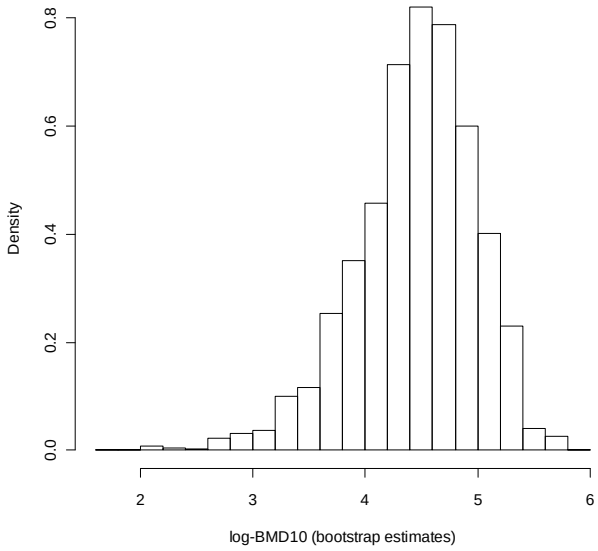

Figure 3. Sampling distribution of log(ID_{10}) from bootstrap samples

Moon et al. (2004) claimed that the two-parameter models (beta-Poisson, log-normal, log-logistic, and extreme value) reflected at least as much model uncertainty on average as the three-parameter models (Weibull gamma, exponential gamma, Weibull exponential, and shifted Weibull). Namata et al. (2008) studied model-averaging in MRA with $K = 40$ dose-response models including the same four two-parameter models in this study plus the family of fractional polynomial models with the combination of 3 negative powers and 4 positive powers. They claimed that the set of candidate models should be rich enough. We note that the richness may not be solely determined by the number of dose-response models considered in model-averaging. A set of fewer but diverse models may reflect as much model uncertainty in a more efficient manner. In our simulation study, we confirm that only the four two-parameter models are sufficient to account for model uncertainty in all three representative dose-response patterns.

An ideal combination of dose-response models is unknown, or it may not exist. Our proposed method can be applied with any set of dose-response models. We note that the number of models in model-averaging and the number of model parameters may significantly influence the computational process. We also note that an alternative approach may be model selection based on classification of a model as "mechanistic" or empirical. Mechanistic models are plausible because of their interpretability. However, the underlying assumptions are sometimes strong, and there may be circumstances in which the assumptions are not valid. A good mix of mechanistic and empirical models is also a key point in model-averaging (selection of model space). If data arise from the assumed mechanism, and the mechanistic model fits the data well, then the model will be highly weighted. If assumptions are not met, our estimates will be weighted more by empirical models.

In future studies, we may investigate if a subset of BP, LN, LL, and EV models performs well for all three representative patterns. We may be able to discover a pattern-specific subset of the two-parameter models or of any larger model spaces. For a simple illustration, if a real data set exhibits a sublinear pattern, the BP model appears to be eliminated because it is unable to fit the pattern. In the BP model, the second-derivative with respect to d (dose) is negative for all $\alpha > 0$, $\beta > 0$, and $d > 0$; hence it is unable to be concave upward for any data points (See Figure 1). We note that Teunis et al. (1996) suggest that the BP model is the best-fit model for the *Echovirus* data set, although we assert that the *Echovirus* data exhibit a superlinear dose-response pattern. If a general guideline can be developed for a preferable pattern-specific subset of dose-response models, it may contribute substantially to MRA and it can be widely used in practice.

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