

On confidence intervals in regression that utilize uncertain prior information about a vector parameter

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Abstract

Consider a linear regression model with n -dimensional response vector, p -dimensional regression parameter $\boldsymbol{\beta}$ and independent normally distributed errors. Suppose that the parameter of interest is $\theta = \mathbf{a}^T \boldsymbol{\beta}$ where \mathbf{a} is a specified vector. Define the s -dimensional parameter vector $\boldsymbol{\tau} = \mathbf{C}^T \boldsymbol{\beta} - \mathbf{t}$ where \mathbf{C} and \mathbf{t} are specified. Also suppose that we have uncertain prior information that $\boldsymbol{\tau} = \mathbf{0}$. Part of our evaluation of a frequentist confidence interval for θ is the ratio (expected length of this confidence interval)/(expected length of standard $1 - \alpha$ confidence interval), which we call the scaled expected length of this interval. We say that a $1 - \alpha$ confidence interval for θ utilizes this uncertain prior information if (a) the scaled expected length of this interval is significantly less than 1 when $\boldsymbol{\tau} = \mathbf{0}$, (b) the maximum value of the scaled expected length is not too large and (c) this confidence interval reverts to the standard $1 - \alpha$ confidence interval when the data happen to strongly contradict the prior information. Let $\hat{\Theta} = \mathbf{a}^T \hat{\boldsymbol{\beta}}$ and $\hat{\boldsymbol{\tau}} = \mathbf{C}^T \hat{\boldsymbol{\beta}} - \mathbf{t}$, where $\hat{\boldsymbol{\beta}}$ is the least squares estimator of $\boldsymbol{\beta}$. We consider the particular case that that $E((\hat{\boldsymbol{\tau}} - \boldsymbol{\tau})(\hat{\Theta} - \theta)) = \mathbf{0}$, so that $\hat{\Theta}$ and $\hat{\boldsymbol{\tau}}$ are independent. We present a new $1 - \alpha$ confidence interval for θ that utilizes the uncertain prior information that $\boldsymbol{\tau} = \mathbf{0}$. The following problem is used to illustrate the application of this new confidence interval. Consider a 2^3 factorial experiment with 1 replicate. Suppose that the parameter of interest θ is a specified linear combination of the main effects. Assume that the three-factor interaction is zero. Also suppose that we have uncertain prior information that all of the two-factor interactions are zero. Our aim is to find a frequentist 0.95 confidence interval for θ that utilizes this uncertain prior information.

Keywords: Frequentist confidence interval; Prior information; Linear regression.

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1. Introduction

Suppose that the parameter of interest θ is a scalar and that we have uncertain prior information about the parameters of the model. Hodges and Lehmann (1952), Bickel (1984) and Kempthorne (1983, 1987, 1988) show how such uncertain prior information can be utilized in frequentist inference, mostly for point estimation of θ . A confidence interval for θ is said to be a $1 - \alpha$ confidence interval if it has infimum coverage probability $1 - \alpha$. We assess a $1 - \alpha$ confidence interval J by its scaled expected length, defined to be the ratio (expected length of J)/(expected length of the standard $1 - \alpha$ confidence interval for θ). The first requirement of a $1 - \alpha$ confidence interval that utilizes the uncertain prior information is that its scaled expected length is significantly less than 1 when the prior information is correct (Kabaila, 2009).

We classify confidence intervals that satisfy this first requirement into the following two groups. The first group consists of $1 - \alpha$ confidence intervals with scaled expected length that is less than or equal to 1 for all parameter values, so that these dominate the standard $1 - \alpha$ confidence interval. An example of such a confidence interval is the Stein-type confidence interval for the normal variance (see e.g. Maata and Casella, 1990 and Goutis and Casella, 1991). The second group consists of $1 - \alpha$ confidence intervals that satisfy this first requirement, when dominance of the usual $1 - \alpha$ confidence interval is not possible (the scaled expected length must exceed 1 for some parameter values). Some relevant admissibility results are provided by Kabaila, Giri and Leeb (2010) and Kabaila (2011). This second group includes the confidence intervals described by Pratt (1961), Brown et al (1995) and Puza and O'Neill (2006ab). This second group also includes $1 - \alpha$ confidence intervals that satisfy the additional requirements that (a) the maximum (over the parameter space) of the scaled expected length is not too much larger than 1 and (b) the confidence interval reverts to the standard $1 - \alpha$ confidence interval when the data happen to strongly contradict the prior information. Confidence intervals that utilize uncertain the prior information and satisfy these additional requirements have been proposed by Farchione and Kabaila (2008) and Kabaila and Giri (2009ab) (cf Kabaila and Giri, 2013).

Consider the linear regression model $\mathbf{Y} = \mathbf{X}\boldsymbol{\beta} + \boldsymbol{\varepsilon}$, where \mathbf{Y} is a random n -vector of responses, \mathbf{X} is a $n \times p$ matrix with linearly independent columns, $\boldsymbol{\beta}$ is an unknown parameter p -vector and $\boldsymbol{\varepsilon} \sim N(\mathbf{0}, \sigma^2 \mathbf{I}_n)$, where σ^2 is an unknown positive

parameter. Suppose that the parameter of interest is $\theta = \mathbf{a}^T \boldsymbol{\beta}$, where \mathbf{a} is a given p -vector ($\mathbf{a} \neq \mathbf{0}$). Let the s -dimensional parameter vector $\boldsymbol{\tau}$ be defined to be $\mathbf{C}^T \boldsymbol{\beta} - \mathbf{t}$ where \mathbf{C} is a specified $p \times s$ matrix ($s < p$) with linearly independent columns and \mathbf{t} is a specified s -vector. Suppose that \mathbf{a} does not belong to the linear subspace spanned by the columns of \mathbf{C} . Also suppose that previous experience with similar data sets and/or expert opinion and scientific background suggests that $\boldsymbol{\tau} = \mathbf{0}$. In other words, suppose that we have uncertain prior information that $\boldsymbol{\tau} = \mathbf{0}$. Our aim is to find a frequentist $1 - \alpha$ confidence interval for θ that utilizes this prior information. By “utilizes this prior information” we mean that (a) the scaled expected length of this interval is significantly less than 1 when $\boldsymbol{\tau} = \mathbf{0}$, (b) the maximum value of the scaled expected length is not too large and (c) this confidence interval reverts to the standard $1 - \alpha$ confidence interval when the data happen to strongly contradict the prior information.

Kabaila and Giri (2009a) have dealt with the case that $s = 1$, so we consider the case that $s \geq 2$. Let $\hat{\boldsymbol{\beta}}$ denote the least squares estimator of $\boldsymbol{\beta}$. Also let $\hat{\Sigma}^2 = (\mathbf{Y} - \mathbf{X}\hat{\boldsymbol{\beta}})^T(\mathbf{Y} - \mathbf{X}\hat{\boldsymbol{\beta}})/(n - p)$, $\hat{\Theta} = \mathbf{a}^T \hat{\boldsymbol{\beta}}$ and $\hat{\boldsymbol{\tau}} = \mathbf{C}^T \hat{\boldsymbol{\beta}} - \mathbf{t}$. We consider the particular case that $E((\hat{\boldsymbol{\tau}} - \boldsymbol{\tau})(\hat{\Theta} - \theta)) = \mathbf{0}$. An example of this particular case is the following. Consider a 2^3 factorial experiment with 1 replicate. For factorial experiments, it is a widely-held belief that the higher the order of interaction, the more likely it is to be negligible. Indeed, fractional factorial designs are based on this belief. Assume that the third-order interaction is zero. Also suppose that we have uncertain prior information that all of the second-order interactions are zero. In this case, $n - p = 1$ and $s = 3$. If the parameter of interest θ is a linear combination of the main effects then $E((\hat{\boldsymbol{\tau}} - \boldsymbol{\tau})(\hat{\Theta} - \theta)) = \mathbf{0}$.

In Section 2, we describe the new $1 - \alpha$ confidence interval for θ that utilizes the uncertain prior information that $\boldsymbol{\tau} = \mathbf{0}$. Define $\boldsymbol{\gamma} = (\text{Cov}(\hat{\boldsymbol{\tau}}))^{-1/2} \boldsymbol{\tau}$. The coverage probability and scaled expected length of this new confidence interval are even functions of $\|\boldsymbol{\gamma}\| = \sqrt{\boldsymbol{\gamma}^T \boldsymbol{\gamma}}$. In Section 3, we consider this 2^3 factorial experiment, when $1 - \alpha = 0.95$. Figure 2 presents graphs of the squared scaled expected length and the coverage probability of this new confidence interval (as functions of $\|\boldsymbol{\gamma}\|$). The infimum coverage probability is computed to be 0.95. To an excellent approximation, the coverage probability of the new confidence interval is equal to 0.95, throughout the parameter space. This figure demonstrates that this new confidence interval has excellent performance in terms of squared scaled expected length. When the prior

information is correct (i.e. $\|\boldsymbol{\gamma}\| = 0$), we gain since the square of the scaled expected length is 0.34707, which is much smaller than 1. The maximum value of the square of the scaled expected length is 1.0404, which is only slightly larger than 1. The new 0.95 confidence interval for θ coincides with the standard $1 - \alpha$ confidence interval when the data strongly contradicts the prior information. This is reflected in Figure 2 by the fact that the square of the scaled expected length approaches 1 as $\|\boldsymbol{\gamma}\| \rightarrow \infty$. In Section 4, we examine the effect on the performance of the new confidence interval of increasing s , for $n - p = 1$. The application of the new confidence interval is to the case that $n - p$ is small. As pointed out in Section 2, the ability of the new confidence interval to utilize the uncertain prior information comes from enhanced estimation of σ^2 . The smaller $n - p$ is, the larger will be this enhancement.

2. New confidence interval that utilizes the uncertain prior information

Our first step is to reduce the data to $(\hat{\Theta}, \hat{\boldsymbol{\tau}}, \hat{\Sigma}^2)$. Let $v_{11} = \mathbf{a}^T(\mathbf{X}^T \mathbf{X})^{-1} \mathbf{a}$ and $\mathbf{V}_{22} = \mathbf{C}^T(\mathbf{X}^T \mathbf{X})^{-1} \mathbf{C}$. Note that $\hat{\Theta} \sim N(\theta, \sigma^2 v_{11})$, $\hat{\boldsymbol{\tau}} \sim N(\boldsymbol{\tau}, \sigma^2 \mathbf{V}_{22})$ and $\hat{\Sigma}^2$ are independent random vectors. Let $m = n - p$. Define the quantile $t(m)$ by the requirement that $P(-t(m) \leq T \leq t(m)) = 1 - \alpha$ for $T \sim t_m$. The standard $1 - \alpha$ confidence interval for $\theta = \mathbf{a}^T \boldsymbol{\beta}$ is

$$I = [\hat{\Theta} - t(m)\sqrt{v_{11}} \hat{\Sigma}, \hat{\Theta} + t(m)\sqrt{v_{11}} \hat{\Sigma}].$$

The new confidence interval for θ that we will describe shortly is centered at $\hat{\Theta}$. The fact that $\hat{\Theta}$ and $\hat{\boldsymbol{\tau}}$ are independent suggests that the uncertain prior information that $\boldsymbol{\tau} = \mathbf{0}$ should not influence the point estimation of θ . However, this uncertain prior information can be used to enhance the estimation of σ^2 . In the absence of any prior information about $\boldsymbol{\tau}$, the standard estimator of σ^2 is $\hat{\Sigma}^2$ and $m\hat{\Sigma}^2/\sigma^2 \sim \chi_m^2$. However, if it known that $\boldsymbol{\tau} = \mathbf{0}$ then that standard estimator of σ^2 is

$$\tilde{\Sigma}^2 = \frac{m\hat{\Sigma}^2 + \hat{\boldsymbol{\tau}}^T \mathbf{V}_{22}^{-1} \hat{\boldsymbol{\tau}}}{m + s} = \hat{\Sigma}^2 \left(\frac{m + sF}{m + s} \right),$$

where $F = (\hat{\boldsymbol{\tau}}^T \mathbf{V}_{22}^{-1} \hat{\boldsymbol{\tau}}/s)/\hat{\Sigma}^2$ and $(m + s)\tilde{\Sigma}^2/\sigma^2 \sim \chi_{m+s}^2$. This suggests that the uncertain prior information that $\boldsymbol{\tau} = \mathbf{0}$ can be used to enhance the estimation of σ^2 by using the appropriate function of $\hat{\Sigma}^2$ and F .

This motivates us to consider a new confidence interval for θ of the form

$$J(d) = \left[\hat{\Theta} - \sqrt{v_{11}} \hat{\Sigma} d(\sqrt{F}), \hat{\Theta} + \sqrt{v_{11}} \hat{\Sigma} d(\sqrt{F}) \right],$$

where the function $d : [0, \infty) \rightarrow (0, \infty)$ is required to satisfy the following restrictions.

Restriction 1 d is a continuous function.

Restriction 2 $d(x) = t(m)$ for all $x \geq k$, where k is a specified positive number.

The first restriction implies that the endpoints of the confidence interval $J(d)$ are continuous functions of the data. Note that F is the usual F statistic for testing the null hypothesis $H_0 : \boldsymbol{\tau} = \mathbf{0}$ against the alternative hypothesis $H_1 : \boldsymbol{\tau} \neq \mathbf{0}$. Thus the second restriction implies that this confidence interval reverts to the usual $1 - \alpha$ confidence interval I when the data happen to strongly contradict the prior information that $\boldsymbol{\tau} = \mathbf{0}$.

Part of the evaluation of the confidence interval $J(d)$ consists of comparing it with the usual $1 - \alpha$ confidence interval I using the scaled expected length criterion (expected length of this confidence interval) / (expected length of I). Theorem 1, which is stated and proved in Appendix A, provides computationally-convenient expressions for the coverage probability and scaled expected length of $J(d)$. Define $\boldsymbol{\gamma} = (\text{Cov}(\hat{\boldsymbol{\tau}}))^{-1/2} \boldsymbol{\tau} = (1/\sigma) \mathbf{V}_{22}^{-1/2} \boldsymbol{\tau}$. According to this theorem, for given function d , both the coverage probability and the scaled expected length of $J(d)$ are functions of $\|\boldsymbol{\gamma}\|$. We denote this scaled expected length by $e(\|\boldsymbol{\gamma}\|; d)$. The numerical integration method used to evaluate this coverage probability is described in Appendix B.

Our aim is to find a function d satisfying Restrictions 1 and 2 and such that (a) the minimum of $P(\theta \in J(d))$ over $\|\boldsymbol{\gamma}\|$ is $1 - \alpha$ and (b) $e(0; d)$ is minimized subject to the restriction that $e(\|\boldsymbol{\gamma}\|; d) \leq \ell$ for all $\|\boldsymbol{\gamma}\|$, where $\ell \geq 1$ is chosen by the statistician prior to the analysis of the data. Theorem 2, which is stated and proved in Appendix C, provides a computationally-convenient expression for $e(0; d)$. We expect that, for small m , this constrained minimization will lead to a $1 - \alpha$ confidence interval for θ that has scaled expected length $e(\|\boldsymbol{\gamma}\|; d)$ that is substantially less 1 for $\boldsymbol{\gamma} = \mathbf{0}$.

We implement the coverage constraint $P(\theta \in J(d)) \geq 1 - \alpha$ for all $\|\boldsymbol{\gamma}\| \geq 0$ as follows. For any reasonable choice of the function d , $P(\theta \in J(d))$ converges to $1 - \alpha$ as $\|\boldsymbol{\gamma}\| \rightarrow \infty$. The constraints implemented in the computations are that $P(\theta \in J(d)) \geq 1 - \alpha$ for every $\|\boldsymbol{\gamma}\|$ in a judiciously-chosen finite set of values \mathcal{G} . That a given \mathcal{G} is adequate to the task is judged by checking numerically, at the

completion of the computations, that the coverage probability constraint is satisfied for all $\|\gamma\| \geq 0$ (cf. Farchione and Kabaila, 2012).

For computational feasibility, we specify the following parametric form for the function d . Suppose that x_1, \dots, x_q satisfy $0 = x_1 < x_2 < \dots < x_q = k$. We fully specify the function d by the vector $(d(x_1), \dots, d(x_{q-1}))$ as follows. The value of $d(x)$ for any $x \in [0, k]$ is specified by natural cubic spline interpolation for these given function values and $d(x_q) = t(m)$ (without any endpoint conditions on the first derivative of d). We call x_1, x_2, \dots, x_q the knots. Of course, the values of k , ℓ and knots x_i need to be judiciously-chosen and this will usually require some computational exploration.

3. Application to the analysis of data from a single-replicate 2^3 factorial experiment

Consider a 2^3 factorial experiment carried out without replication. Let Y denote the response and let x_1 , x_2 and x_3 denote the coded levels for each of the 3 factors, where the coded level takes either the value -1 or 1 . We assume the model

$$Y = \beta_0 + \beta_1 x_1 + \beta_2 x_2 + \beta_3 x_3 + \beta_{12} x_1 x_2 + \beta_{13} x_1 x_3 + \beta_{23} x_2 x_3 + \beta_{123} x_1 x_2 x_3 + \varepsilon$$

where $\beta_0, \beta_1, \beta_2, \beta_3, \beta_{12}, \beta_{13}, \beta_{23}, \beta_{123}$ are unknown parameters and $\varepsilon \sim N(0, \sigma^2)$, where σ^2 is an unknown positive parameter.

For factorial experiments it is commonly believed that higher order interactions are negligible (see e.g. Mead (1988, p.368) and Hinkelman & Kempthorne (1994, p.350)). Indeed, this type of belief is the basis for the design of fractional factorial experiments. Assume that $\beta_{123} = 0$. Also suppose that we have uncertain prior information that β_{12} , β_{13} and β_{23} are all zero. Thus $n - p = 1$. We consider the particular case that the parameter of interest θ is a linear combination of the main effects i.e. $\theta = a_1 \beta_1 + a_2 \beta_2 + a_3 \beta_3$. In this case, $E((\hat{\tau} - \tau)(\hat{\Theta} - \theta)) = \mathbf{0}$.

Of course, the properties of $J(d)$, resulting from the constrained minimization described in Section 2, depend on the values of k , ℓ , the knots x_i and $1 - \alpha$. We focus on the particular case that $k = 15$, $\ell = 1.02$, the knots are at $0, 1, 2, 3, 7, 12, 15$ and $1 - \alpha = 0.95$. When we compute the new confidence interval, we obtain the function d shown in Figure 1. All of the computations presented in the present paper were performed with programs written in MATLAB, using the optimization and statistics toolboxes. Consistent with the corollary stated in Appendix C, $d(x)$ takes values

larger than $t(m)$. Figure 2 presents graphs of the squared scaled expected length and the coverage probability (as functions of $\|\gamma\|$) of this new confidence interval. The squared scaled expected length and coverage probability computations were checked using Monte Carlo simulations.

The infimum coverage probability is computed to be 0.95. The upper panel of Figure 2 demonstrates that this new confidence interval has excellent performance in terms of squared scaled expected length. When the prior information is correct (i.e. $\|\gamma\| = 0$), we gain since the square of the scaled expected length is 0.34707, which is much smaller than 1. The maximum value of the square of the scaled expected length is 1.0404, which is only slightly larger than 1. The new 0.95 confidence interval for θ coincides with the standard $1 - \alpha$ confidence interval when the data strongly contradicts the prior information. This is reflected in the upper panel of Figure 2 by the fact that the square of the scaled expected length approaches 1 as $\|\gamma\| \rightarrow \infty$.

4. The effect on the performance of the new confidence interval of increasing the value of s

Suppose that the value of m is fixed and that we increase s . It seems plausible that the best possible performance of the new confidence interval for θ will increase as s increases i.e. as the amount of uncertain prior information increases. We have examined the truth of this plausible result as follows. We have considered $m = 1$ and $1 - \alpha = 0.95$, chosen $\ell = 1.02$ and the number of knots to be 7. For each $s = 1, 2, 3, 5$ and 7, we have chosen k and the knots so as to minimize the scaled expected length at $\|\gamma\| = 0$. We have obtained the following results:

s	Min sq sel	Max sq sel	Min CP	Max CP
1	0.80549	1.0414	0.95	0.95049
2	0.54698	1.0404	0.95	0.95030
3	0.34707	1.0404	0.95	0.95037
5	0.25151	1.0406	0.95	0.95034
7	0.19027	1.0404	0.95	0.95030

Table 1: Comparison of results for different values of s , for $m = 1$, $1 - \alpha = 0.95$ and $\ell = 1.02$. The column labels Min sq sel, Max sq sel, Min CP and Max CP denote the minimum squared scaled expected length, the maximum squared scaled expected length, the minimum coverage probability and the maximum coverage probability, respectively.

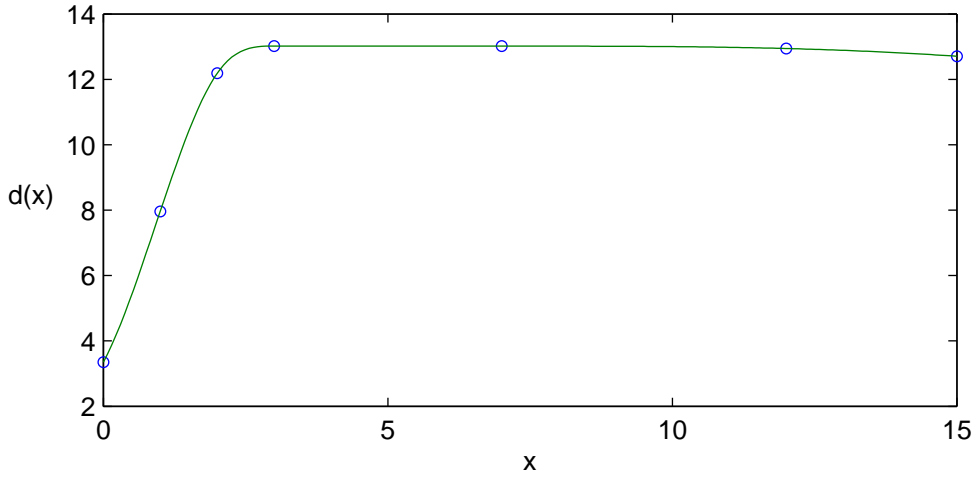


Figure 1: Plot of the function d for the new 0.95 confidence interval for θ when $s = 3$, $m = n - p = 1$ and $\ell = 1.02$. The knots are at 0, 1, 2, 3, 7, 12, 15.

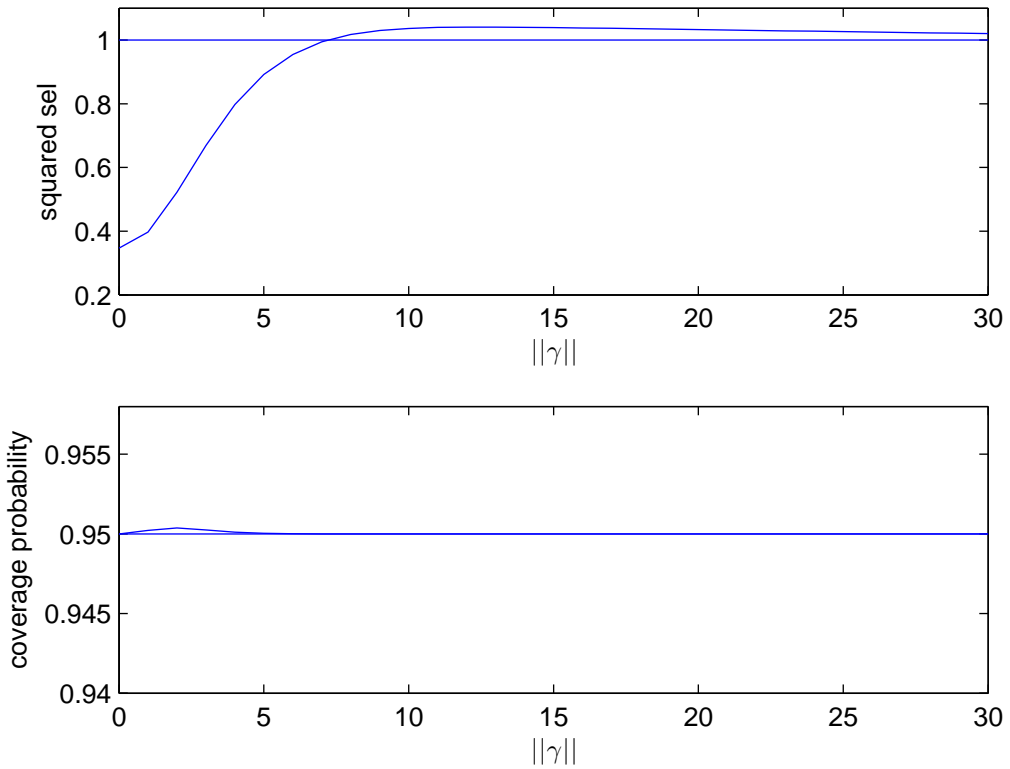


Figure 2: Plots of squared scaled expected length $e^2(\|\gamma\|; d)$ and coverage probability as function of $\|\gamma\|$ when $1 - \alpha = 0.95$, $s = 3$, $m = n - p = 1$ and $\ell = 1.02$.

If, for each s considered, we assume that the performance of the confidence interval in terms of the scaled expected length at $\|\boldsymbol{\gamma}\| = 0$ is about as good as it can be then this table tells us the following. As s increases, the amount of uncertain prior information increases and this leads to an improvement in the performance of this confidence interval.

5. Discussion

In this paper we have shown how to construct a frequentist confidence interval for the parameter of interest θ that utilizes the uncertain prior information that $\boldsymbol{\tau} = \mathbf{0}$. We have done this for the particular case that the covariances between the components of the least squares estimator of $\boldsymbol{\tau}$ and the least squares estimator of θ are all zero. Our practical experience with the computations of this new confidence interval, in a variety of circumstances, shows that the coverage probability needs to be computed with great accuracy for these computations to be successful. If we no longer restrict attention to the particular case that all of these covariances are zero then the construction of such a confidence interval necessitates the computation of coverage probabilities using more complicated methods of the type employed by Kabaila and Farchione (2012). However, the increased computation time of these methods would appear to make the constrained optimization not computationally practicable.

Appendix A. Theorem 1 and its proof

In this appendix, we state and prove Theorem 1, which provides new computationally-convenient expressions for the coverage probability and scaled expected length of the confidence interval $J(d)$. Define $G = (\hat{\Theta} - \theta)/(\sigma\sqrt{v_{11}})$ and $Q = (1/\sigma^2)\hat{\boldsymbol{\tau}}^T \mathbf{V}_{22}^{-1} \hat{\boldsymbol{\tau}}$. Also define $V = \sqrt{Q/s}$ and $W = \hat{\Sigma}/\sigma$. Now, (G, V) and W are independent random vectors. The assumption that $E((\hat{\boldsymbol{\tau}} - \boldsymbol{\tau})(\hat{\Theta} - \theta)) = \mathbf{0}$ implies that G and V are independent random variables. Thus, G , V and W are independent random variables. Note that $G \sim N(0, 1)$, Q has a noncentral χ^2 distribution with s degrees of freedom and noncentrality parameter $\|\boldsymbol{\gamma}\|^2 = \boldsymbol{\gamma}^T \boldsymbol{\gamma}$ and W has the same distribution as $\sqrt{\chi_m^2/m}$.

Theorem 1. *Let $f_V(\cdot; \|\boldsymbol{\gamma}\|)$ and f_W denote the probability density functions of V and W , respectively. Also let Φ denote the $N(0, 1)$ distribution function.*

(a) The coverage probability of $J(d)$ is equal to

$$1 - \alpha + 2 \int_0^k \int_0^\infty (\Phi(wd(x)) - \Phi(wt(m))) f_V(xw; \|\gamma\|) w f_W(w) dw dx. \quad (1)$$

For given function d , the coverage probability of $J(d)$ is a function of $\|\gamma\|$.

(b) The scaled expected length of $J(d)$ is equal to

$$1 + \frac{1}{t(m)E(W)} \int_0^\infty \int_0^k (d(x) - t(m)) f_V(xw; \|\gamma\|) dx w^2 f_W(w) dw. \quad (2)$$

For given function d , the scaled expected length of $J(d)$ is a function of $\|\gamma\|$.

Proof of part (a). It is straightforward to show that the coverage probability $P(\theta \in J(d))$ is equal to $P(-Wd(V/W) \leq G \leq Wd(V/W))$. By the law of total probability, this is equal to

$$P(-Wd(V/W) \leq G \leq Wd(V/W), V/W < k) + P(-t(m)W \leq G \leq t(m)W, V/W \geq k),$$

since $d(x) = t(m)$ for all $x \geq k$. Now

$$\begin{aligned} & P(-t(m)W \leq G \leq t(m)W, V/W \geq k) + P(-t(m)W \leq G \leq t(m)W, V/W < k) \\ &= P(-t(m)W \leq G \leq t(m)W) \\ &= P(-t(m) \leq G/W \leq t(m)) = 1 - \alpha \end{aligned}$$

Thus $P(\theta \in J(d))$ is equal to

$$\begin{aligned} & 1 - \alpha + P(-Wd(V/W) \leq G \leq Wd(V/W), V/W < k) \\ & \quad - P(-t(m)W \leq G \leq t(m)W, V/W < k) \\ &= 1 - \alpha + \int_0^\infty \int_0^{kw} (2\Phi(wd(v/w)) - 1) f_V(v) dv f_W(w) dw \\ & \quad - \int_0^\infty \int_0^{kw} (2\Phi(wt(m)) - 1) f_V(v) dv f_W(w) dw \\ &= 1 - \alpha + 2 \int_0^\infty \int_0^{kw} (\Phi(wd(v/w)) - \Phi(wt(m))) f_V(v) dv f_W(w) dw. \end{aligned}$$

Changing the variable of integration of the inner integral to $x = v/w$, we obtain

$$\begin{aligned} & 1 - \alpha + 2 \int_0^\infty \int_0^k (\Phi(wd(x)) - \Phi(wt(m))) f_V(xw; \|\gamma\|) dx w f_W(w) dw \\ &= 1 - \alpha + 2 \int_0^k \int_0^\infty (\Phi(wd(x)) - \Phi(wt(m))) f_V(xw; \|\gamma\|) w f_W(w) dw dx \end{aligned}$$

Proof of part (b). It is straightforward to show that the scaled expected length of $J(d)$ is equal to

$$\frac{E(W d(V/W))}{t(m) E(W)}. \quad (3)$$

We use the notation

$$\mathcal{I}(\mathcal{A}) = \begin{cases} 1 & \text{if } \mathcal{A} \text{ is true} \\ 0 & \text{if } \mathcal{A} \text{ is false} \end{cases}$$

where \mathcal{A} is an arbitrary statement. Since $\mathcal{I}(V/W < k) + \mathcal{I}(V/W \geq k) = 1$, $E(W d(V/W))$ is equal to

$$\begin{aligned} & E(W d(V/W) \mathcal{I}(V/W < k)) + E(W d(Q/W^2) \mathcal{I}(V/W \geq k)) \\ &= E(W d(V/W) \mathcal{I}(V/W < k)) + E(W t(m) \mathcal{I}(V/W \geq k)) \\ &= t(m) E(W) + E((d(V/W) - t(m)) W \mathcal{I}(V/W < k)). \end{aligned}$$

Thus the expression (3) for the scaled expected length is equal to

$$1 + \frac{1}{t(m) E(W)} E((d(V/W) - t(m)) W \mathcal{I}(V/W < k)) \quad (4)$$

$$= 1 + \frac{1}{t(m) E(W)} \int_0^\infty \int_0^{kw} \left(d\left(\frac{v}{w}\right) - t(m) \right) f_V(v; \|\gamma\|) dv w f_W(w) dw. \quad (5)$$

Changing the variable of integration of the inner integral to $x = v/w$, we obtain (2).

Appendix B. The method used to evaluate the coverage probability

In this appendix we describe the numerical integration method used to compute the coverage probability $P(\theta \in J(d))$, as given by (1). The function d , which is a cubic spline in the interval $[0, k]$, does not necessarily have a third derivative at each of the knots $x_2, \dots, x_q = k$. So we evaluate (1) by computing

$$1 - \alpha + 2 \sum_{i=1}^{q-1} \int_{x_i}^{x_{i+1}} \int_0^\infty (\Phi(wd(w)) - \Phi(wt(m))) f_V(xw; \|\gamma\|) w f_W(w) dw dx. \quad (6)$$

Each of the inner integrals is equal to

$$\begin{aligned} & \int_0^\infty (\Phi(wd(w)) - \Phi(wt(m))) f_V(xw; \|\gamma\|) w f_W(w) dw \\ &= E\left((\Phi(Wd(x)) - \Phi(Wt(m))) f_V(xW; \|\gamma\|) W\right), \end{aligned} \quad (7)$$

where W has pdf f_W . Let $Z = mW^2$, so that $Z \sim \chi_m^2$. Thus (7) is equal to

$$\begin{aligned} & E \left\{ \left(\Phi \left(\sqrt{\frac{Z}{m}} d(x) \right) - \Phi \left(\sqrt{\frac{Z}{m}} t(m) \right) \right) f_V \left(x \sqrt{\frac{Z}{m}}; \|\gamma\| \right) \sqrt{\frac{Z}{m}} \right\} \\ &= \frac{1}{\sqrt{m}} \int_0^\infty \left(\Phi \left(\sqrt{\frac{z}{m}} d(x) \right) - \Phi \left(\sqrt{\frac{z}{m}} t(m) \right) \right) f_V \left(x \sqrt{\frac{z}{m}}; \|\gamma\| \right) \sqrt{z} f_m(z) dz \end{aligned} \quad (8)$$

where f_m denotes the χ_m^2 pdf. It can be shown that

$$z^{1/2} f_m(z) = \frac{2^{1/2} \Gamma((m+1)/2)}{\Gamma(m/2)} f_{m+1}(z).$$

Thus (8) is equal to

$$\sqrt{\frac{2}{m}} \frac{\Gamma((m+1)/2)}{\Gamma(m/2)} \int_0^\infty \left(\Phi \left(\sqrt{\frac{z}{m}} d(x) \right) - \Phi \left(\sqrt{\frac{z}{m}} t(m) \right) \right) f_V \left(x \sqrt{\frac{z}{m}}; \|\gamma\| \right) f_{m+1}(z) dz. \quad (9)$$

Assuming that $d(x) > 0$ for all $x \in [0, k]$, we compute this as follows. Let F_m denote the χ_m^2 cdf. Now change the variable of integration to $u = F_{m+1}(z)$, so that (9) is equal to

$$\sqrt{\frac{2}{m}} \frac{\Gamma((m+1)/2)}{\Gamma(m/2)} \int_0^1 g(u; x, \|\gamma\|) du$$

where $g(u; x, \|\gamma\|)$ is defined to be

$$\left(\Phi \left(\sqrt{\frac{F_{m+1}^{-1}(u)}{m}} d(x) \right) - \Phi \left(\sqrt{\frac{F_{m+1}^{-1}(u)}{m}} t(m) \right) \right) f_V \left(x \sqrt{\frac{F_{m+1}^{-1}(u)}{m}}; \|\gamma\| \right) \quad (10)$$

for all $(x, u) \in [0, k] \times [0, 1)$ and the limit of (10) as u approaches 1 from below for $u = 1$ and all $x \in [0, k]$. Thus $g(1; x, \|\gamma\|) = 0$ for all $x \in [0, k]$.

Appendix C. Theorem 2 and its proof

The following theorem provides a computationally-convenient expression for the criterion $e(0; d)$.

Theorem 2. *The criterion $e(0; d)$ is equal to*

$$1 + \frac{2^{3/2} s^{s/2} \Gamma((s+m+1)/2)}{t(m) E(W) \Gamma(m/2) \Gamma(s/2)} \int_0^k (d(x) - t(m)) x^{s-1} \frac{m^{m/2}}{(sx^2 + m)^{(s+m+1)/2}} dx \quad (11)$$

where Γ denotes the gamma function.

Proof. The proof of this theorem uses Theorem 1 (b). It follows from (2) that

$$e(0; d) - 1 = \frac{1}{t(m) E(W)} \int_0^\infty \int_0^k (d(x) - t(m)) f_V(wx; 0) dx w^2 f_W(w) dw$$

Note that $f_V(v; 0) = 2sv f_s(sv^2)$, where f_s denotes χ_s^2 probability density function. Interchanging the order of integration, we obtain

$$e(0; d) - 1 = \frac{2s}{t(m) E(W)} \int_0^k (d(x) - t(m)) x \int_0^\infty f_s((sx^2)w^2) w^3 f_W(w) dw dx.$$

Lemma 1. For each $y > 0$,

$$\int_0^\infty f_s(yw^2) w^3 f_W(w) dw = \frac{2^{1/2} m^{m/2} y^{(s/2)-1} \Gamma((s+m+1)/2)}{(y+m)^{(s+m+1)/2} \Gamma(m/2) \Gamma(s/2)}. \quad (12)$$

Proof. Note that $f_W(w) = 2mw f_m(mw^2)$, where f_m denotes the χ_m^2 probability density function. Substituting the expressions for f_s and f_W , we obtain

$$\int_0^\infty f_s(yw^2) w^3 f_W(w) dw = \frac{m^{m/2} y^{(s/2)-1}}{2^{(s+m-2)/2} \Gamma(m/2) \Gamma(s/2)} \int_0^\infty e^{-(y+m)w^2/2} w^{s+m} dw.$$

By (A2.1.3) of Box and Tiao (1973), this is equal to the right-hand side of (12). \square

It follows from this lemma that

$$e(0; d) - 1 = \frac{2^{3/2} s^{s/2} \Gamma((s+m+1)/2)}{t(m)E(W) \Gamma(m/2) \Gamma(s/2)} \int_0^k (d(x) - t(m)) x^{s-1} \frac{m^{m/2}}{(sx^2 + m)^{(s+m+1)/2}} dx. \quad \square$$

Appendix D. Some simple results on confidence interval performance

In this appendix we consider the confidence interval

$$J(d) = \left[\hat{\Theta} - \sqrt{v_{11}} \hat{\Sigma} d(\sqrt{F}), \hat{\Theta} + \sqrt{v_{11}} \hat{\Sigma} d(\sqrt{F}) \right],$$

where $d : [0, \infty) \rightarrow (0, \infty)$. We make additional requirements of d , as needed. We state some simple results about the performance of this confidence interval. The proofs of these results are straightforward and are omitted, for the sake of brevity.

Theorems 3 and 4 concern the expected length of $J(d)$. Theorem 3 is used in the proof of Theorem 4.

Theorem 3. Suppose that $d_1(x) \geq d_2(x)$ for all $x \geq 0$. Then

$$E(\text{length of } J(d_1)) \geq E(\text{length of } J(d_2)) \text{ for all } \|\gamma\|.$$

Theorem 4. Suppose that $d_1(x) \geq d_2(x)$ for all $x \geq 0$ and that there exists $\epsilon > 0$ and an interval $[a, b]$ (where $0 \leq a < b$) such that $d_1(x) > d_2(x) + \epsilon$ for all $x \in [a, b]$. Then

$$E(\text{length of } J(d_1)) > E(\text{length of } J(d_2)) \text{ for all } \|\gamma\|.$$

Theorems 5 and 6 concern the coverage probability of $J(d)$. Theorem 5 is used in the proof of Theorem 6.

Theorem 5. *Suppose that $d_1(x) \geq d_2(x)$ for all $x \geq 0$. Then*

$$P(\theta \in J(d_1)) \geq P(\theta \in J(d_2)) \quad \text{for all } \|\gamma\|.$$

Theorem 6. *Suppose that $d_1(x) \geq d_2(x)$ for all $x \geq 0$ and that there exists $\epsilon > 0$ and an interval $[a, b]$ (where $0 \leq a < b$) such that $d_1(x) > d_2(x) + \epsilon$ for all $x \in [a, b]$. Then*

$$P(\theta \in J(d_1)) > P(\theta \in J(d_2)) \quad \text{for all } \|\gamma\|.$$

These theorems have the following three consequences.

Corollary. *Suppose that d is continuous. If $d(0) < t(m)$ and $d(x) \leq t(m)$ for all $x > 0$ then*

$$P(\theta \in J(d)) < 1 - \alpha \quad \text{for all } \|\gamma\|.$$

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