Location-adjusted Wald statistics

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Outline

Outline

Wald statistic for scalar parameters

Data

$(y_i, x_i^\top) \qquad (i = 1, \ldots, n)$

$x_i = (x_{i1}, \ldots, x_{ik})^\top \in \Re^k$ is a vector of explanatory variables for y_i

Model

Independent random variables $Y_1, \ldots, Y_n$ with pdf/pmf $p_Y(y_i|x_i;\theta)$

Parameter $\theta \in \Theta \subset \Re^p$ with

$\theta = (\psi, \lambda^\top)^\top$, where $\psi \in \Re$ is of interest

Task

Draw inference about ψ

Wald statistic

Log-likelihood[1]

$$l(\theta) = \sum_{i=1}^{n} \log p_Y(y_i | x_i; \theta)$$

Wald statistic for testing $\psi = \psi_0$

$$t = \frac{\hat\psi - \psi_0}{\kappa(\hat\theta)} \overset{\text{appr}}{\sim} N(0, 1)$$

Maximum likelihood estimator (MLE)

$$\hat\theta = (\hat\psi, \hat\lambda^\top)^\top = \arg\max_{\theta \in \Theta} l(\theta)$$

Standard error

$\kappa(\theta)$ is the square root of the (ψ, ψ) element of the variance-covariance $\{i(\theta)\}^{-1}$ of the (asymptotic) null distribution of $\hat\theta$

$i(\theta)$ is typically taken to be the expected information $E\{\nabla l(\theta) \nabla l(\theta)^\top\}$ or some "robust" variant

[1] subject to usual regularity conditions; see, Pace and Salvan (1997, §4.3)

Wald statistic

Asymptotically equivalent alternatives

Signed root of the likelihood ratio statistic

$$r = \mathrm{sign}(\hat{\psi} - \psi_0)\{l(\hat{\psi}, \hat{\lambda}) - l(\psi_0, \hat{\lambda}_{\psi_0})\}^{1/2} \overset{\text{appr}}{\sim} N(0,1)$$

Signed root of the score statistic

$$s = \mathrm{sign}(\hat{\psi} - \psi_0)\frac{\partial l(\psi_0, \hat{\lambda}_{\psi_0})}{\partial \psi}\kappa(\psi_0, \hat{\lambda}_{\psi_0}) \overset{\text{appr}}{\sim} N(0,1)$$

where $\hat{\lambda}_{\psi_0} = \arg\max_\lambda l(\psi_0, \lambda)$ is the constrained MLE for λ

Pros of t

Computational convenience

```
z test of coefficients:

             Estimate Std. Error z value Pr(>|z|)
(Intercept) -0.37540    0.68957 -0.5444  0.58617
lullyes      1.43237    0.73414  1.9511  0.05105 .
day2        -0.11394    1.04442 -0.1091  0.91313
day3        -0.58487    1.13343 -0.5160  0.60584
day4        -1.71670    1.31233 -1.3081  0.19083
day5         1.82912    1.30168  1.4052  0.15996
day6         0.24783    0.94155  0.2632  0.79238
day7         0.94994    0.99256  0.9571  0.33854
day8         0.46505    0.96850  0.4802  0.63111
day9         0.88646    1.11872  0.7924  0.42813
day10        1.66815    1.05172  1.5861  0.11271
```

Cons of t

Inferential performance depends on the properties of $\hat{\theta}$ (bias, efficiency, etc)

Lack of reparameterization invariance

Reading accuracy IQ and dyslexia

Data

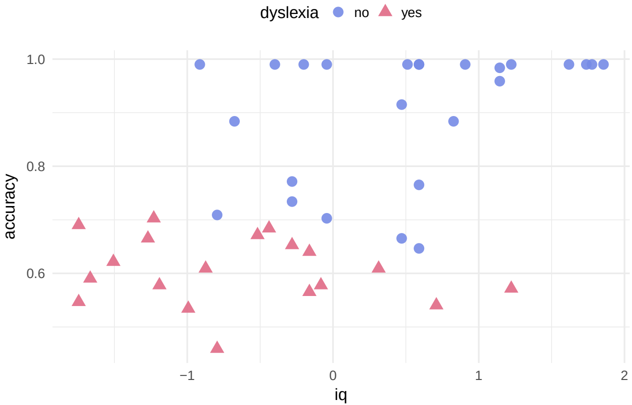
Reading accuracy for 44 nondyslexic and dyslexic Australian children[2]

Ages between 8 years+5 months and 12 years+3 months

Variables

`accuracy`	the score on a reading accuracy test
`iq`	standardized score on a nonverbal intelligent quotient test
`dyslexia`	whether the child is dyslexic or not

[2] data from Smithson and Verkuilen (2006)



Aim

Investigate the relative contribution of nonverbal IQ to the distribution the reading scores, controlling for the presence of diagnosed dyslexia

Reading accuracy IQ and dyslexia

Model

Score of the i-th child is from a Beta distribution with mean μ_i and variance $\mu_i(1-\mu_i)/(1+\phi_i)$ with

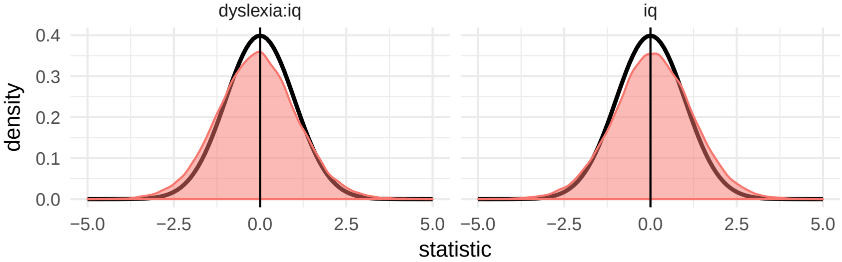
$$\log \frac{\mu_i}{1-\mu_i} = \beta_1 + \sum_{j=2}^{4} \beta_j x_{ij} \quad \text{and} \quad \log \phi_i = \gamma_1 + \sum_{j=2}^{3} \gamma_j x_{ij}$$

- x_{i2} takes value -1 if the ith child is dyslexic and 1 if not
- x_{i3} is the nonverbal IQ score, and
- $x_{i4} = x_{i2}x_{i3}$ is the interaction between `dyslexia` and `iq`

```
Call:
betareg(formula = accuracy ~ dyslexia * iq | dyslexia + iq, data = ReadingSkills,
    type = "ML")

Standardized weighted residuals 2:
    Min      1Q  Median      3Q     Max
-2.3900 -0.6416  0.1572  0.8524  1.6446

Coefficients (mean model with logit link):
            Estimate Std. Error z value Pr(>|z|)
(Intercept)   1.1232     0.1428   7.864 3.73e-15 ***
dyslexia     -0.7416     0.1428  -5.195 2.04e-07 ***
iq            0.4864     0.1331   3.653 0.000259 ***
dyslexia:iq  -0.5813     0.1327  -4.381 1.18e-05 ***

Phi coefficients (precision model with log link):
            Estimate Std. Error z value Pr(>|z|)
(Intercept)   3.3044     0.2227  14.835  < 2e-16 ***
dyslexia      1.7466     0.2623   6.658 2.77e-11 ***
iq            1.2291     0.2672   4.600 4.23e-06 ***
---
Signif. codes:  0 '***' 0.001 '**' 0.01 '*' 0.05 '.' 0.1 ' ' 1

Type of estimator: ML (maximum likelihood)
Log-likelihood:  65.9 on 7 Df
Pseudo R-squared: 0.5756
Number of iterations: 25 (BFGS) + 1 (Fisher scoring)
```

[3] see Grün, Kosmidis, and Zeileis (2012) for a range of modelling strategies and learning methods based on beta regression using the betareg R package

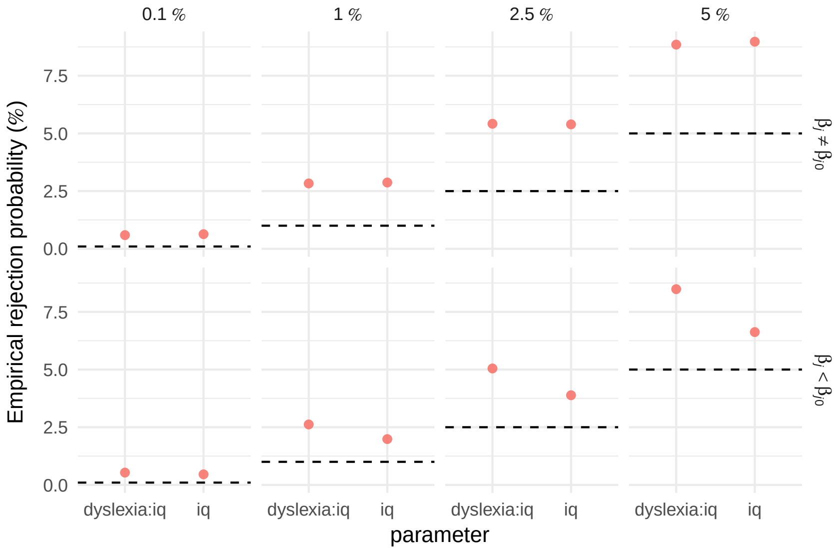
Null distribution of Wald statistic for $\beta_j = \beta_{0j}$



parameter	mean	sd
dyslexia:iq	-0.09	1.15
iq	0.08	1.15

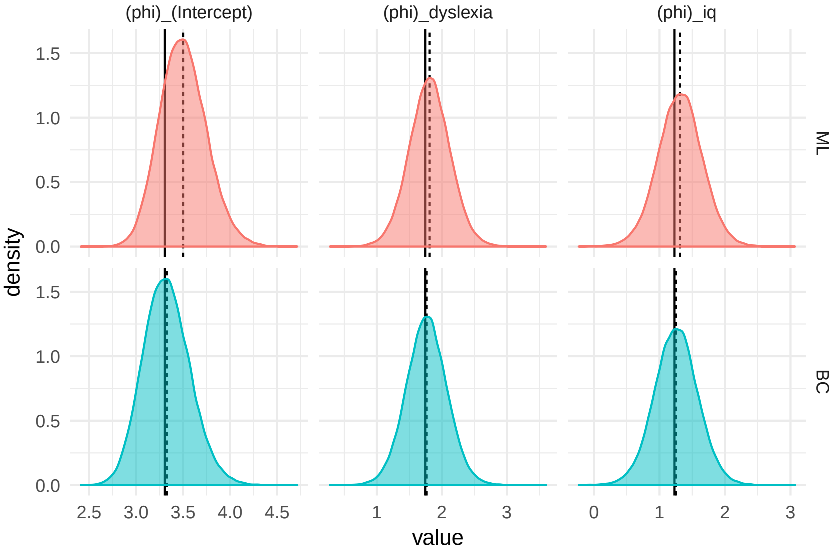
[4] figures based on 50 000 simulated samples under the maximum likelihood fit

Empirical null rejection probabilities



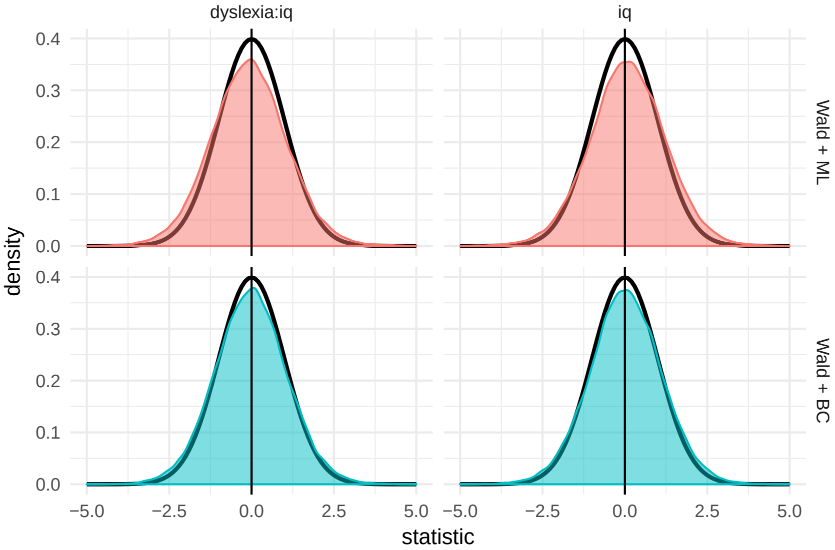
Empirical rejection probabilities are almost double the nominal level

MLE and bias corrected estimator[5]



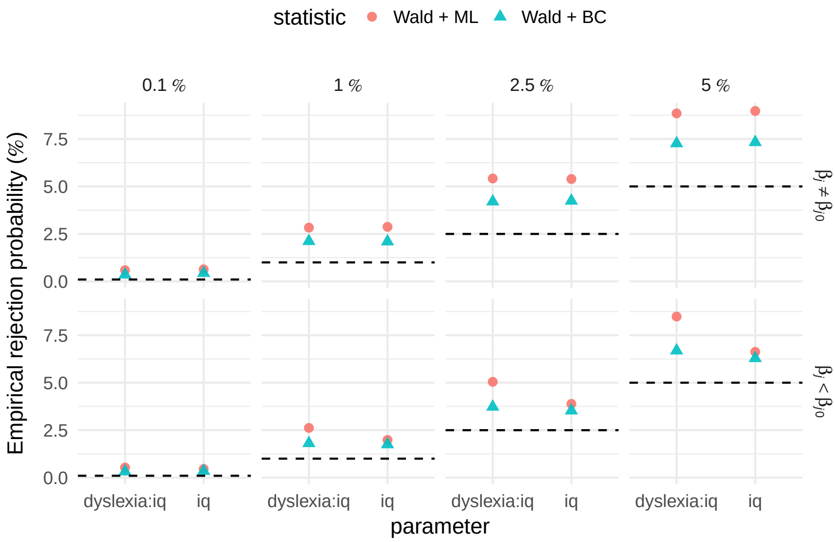
[5]with `type = BC` in the `betareg` call
see, Grün, Kosmidis, and Zeileis (2012) for details on bias correction

Null distribution of Wald statistic using BC estimators[6]



[6] Proposed in Kosmidis and Firth (2010)

Empirical null rejection probabilities



[7] figures based on 50 000 simulated samples under the maximum likelihood fit

Recap on Wald statistics with improved estimators

Use of improved estimators when forming Wald statistics can improve $N(0, 1)$ approximation and hence inferential performance[8]

But

Merely an observation, and in a few models

Rather indirect way to improving Wald inference

Better estimators in $t \not\Longrightarrow$ null distribution of t closer to $N(0, 1)$

[8] see, e.g., Kosmidis and Firth (2010)

Outline

Wald statistic as an estimator

Wald Transform

$$T(\theta; \psi_0) = \frac{\psi - \psi_0}{\kappa(\theta)}$$

$$\Downarrow$$

The Wald statistic
$t = T(\hat{\theta}; \psi_0)$
is the MLE of $T(\theta; \psi_0)$

Core idea

Bias reduction techniques to bring **asymptotic mean** of t "closer" to 0

Bias of t

Under regularity conditions[9] it can be shown that

$$E\{T(\hat{\theta}; \psi_0) - T(\theta; \psi_0)\} = B(\theta; \psi_0) + O\left(n^{-3/2}\right)$$

where

First-order bias of t

$$B(\theta; \psi_0) = b(\theta)^\top \nabla T(\theta; \psi_0) + \frac{1}{2}\mathrm{trace}\left[\{i(\theta)\}^{-1} \nabla \nabla^\top T(\theta; \psi_0)\right]$$

First-order bias of $\hat{\theta}$

$b(\theta)$ such that $E(\hat{\theta} - \theta) = b(\theta) + o\left(n^{-1}\right)$

[9] to guarantee that $T(\theta, \psi_0)$ is > 3 times differentiable wrt θ and $\hat{\theta}$ is consistent

Location-adjusted Wald statistic

Key result

The location-adjusted Wald statistic

$$t^* = T(\hat{\theta}; \psi_0) - B(\hat{\theta}; \psi_0)$$

has null expectation of order $O(n^{-3/2})$

Quantities in the bias of t

$i(\theta)$ and $b(\theta)$ are readily available for a wide range of models, including generalized linear and nonlinear models[10]

Gradient and Hessian of the Wald transform

$$\nabla T(\theta; \psi_0) = \left\{1_p - T(\theta; \psi_0)\nabla\kappa(\theta)\right\} / \kappa(\theta)$$

$$\nabla\nabla^\top T(\theta; \psi_0) = -\left[\nabla\kappa(\theta)\left\{\nabla T(\theta; \psi_0)\right\}^\top + \nabla T(\theta; \psi_0)\left\{\nabla\kappa(\theta)\right\}^\top + T(\theta; \psi_0)\nabla\nabla^\top\kappa(\theta)\right] / \kappa(\theta)$$

$\nabla\kappa(\theta)$ and $\nabla\nabla^\top\kappa(\theta)$ can be computed either analytically, or using automatic or numerical differentiation

[10] see, for example, Cook et al. (1986); Cordeiro and McCullagh (1991); Cordeiro and Vasconcellos (1997); Cordeiro and Toyama Udo (2008); Kosmidis and Firth (2009); Simas et al. (2010); Grün et al. (2012) etc

Example: Exponential with mean $e^{-\theta}$

Cornish-Fisher expansions (Hall, 1992, § 2.5) of the α-level quantiles q_α and q^*_α of the distribution of t and t^* in terms of the corresponding standard normal quantiles z_α are

$$q_\alpha = z_\alpha + n^{-1/2}\frac{z_\alpha^2 + 2}{6} - n^{-1}\frac{11z_\alpha^3 - 65z_\alpha}{144} + O\left(n^{-3/2}\right),$$
$$q^*_\alpha = z_\alpha + n^{-1/2}\frac{z_\alpha^2 - 1}{6} - n^{-1}\frac{11z_\alpha^3 - 65z_\alpha}{144} + O\left(n^{-3/2}\right),$$

provided that $\epsilon < \alpha < 1 - \epsilon$ for any $0 < \epsilon < 1/2$

$\rightarrow$ Quantiles of t^* are closer to those of $N(0,1)$ than t

Computational complexity and implementation

No extra matrix inversions (beyond $\{i(\theta)\}^{-1}$) or optimisation when computing t^*; only extra matrix multiplications

In its analytical form, t^* has the computational complexity $O(p^4)$, whence t has $O(p^3)$

Time complexity can be reduced drastically by exploiting sparsity in $i(\theta)$ in specific models and vectorising operations

Evaluation of t^* for each of the model parameters can be done post-fit and **in parallel**

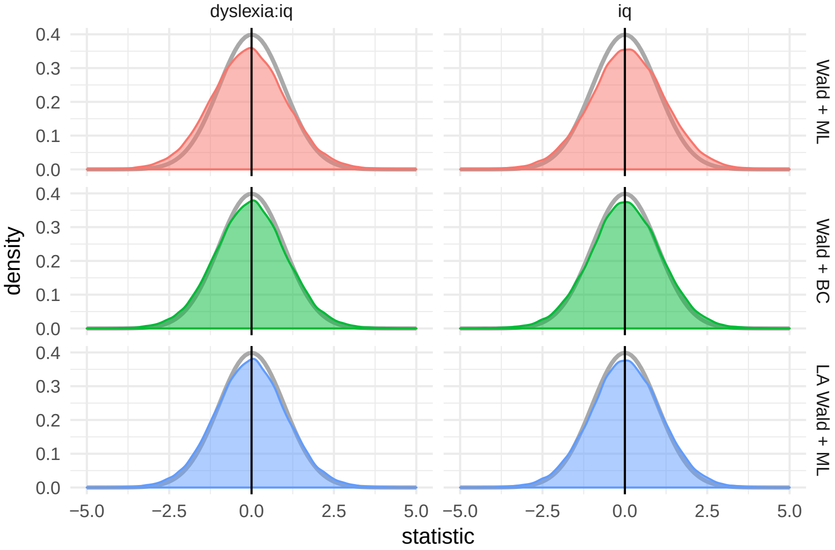
Implementation with numerical derivatives of $\kappa(\theta)$

As implemented in the **waldi** R package
`https://github.com/ikosmidis/waldi`[11]

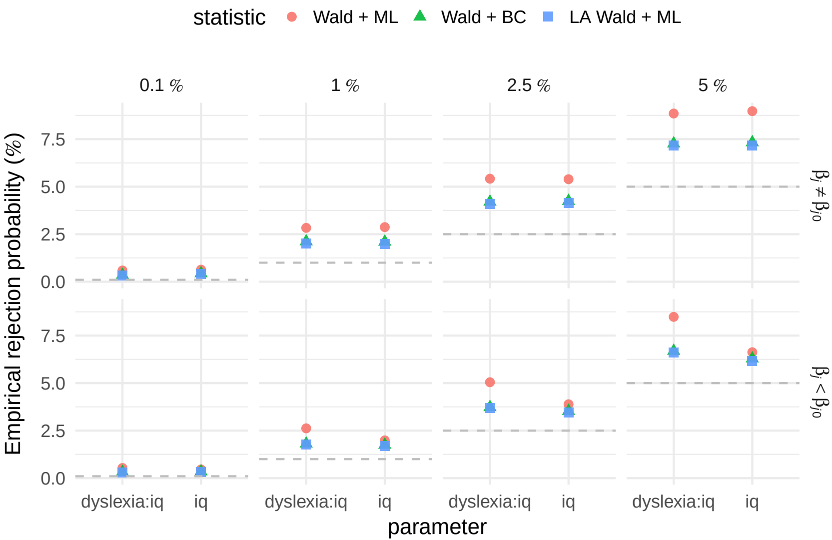
```
R> bias <- enrichwith::get_bias_function(object)
R> info <- enrichwith::get_information_function(object)
R>
R> t <- coef(summary(object))[, "z value"]
R> theta_hat <- coef(object)
R> b <- bias(theta_hat)
R> inverse_i_hat <- solve(info(theta_hat))
R>
R> kappa <- function(theta, j) {
+     inverse_i <- solve(info(theta))
+     sqrt(inverse_i[j, j])
+ }
R>
R> adjusted_t <- function(j) {
+     u <- numDeriv::grad(kappa, theta_hat, j = j)
+     V <- numDeriv::hessian(kappa, theta_hat, j = j)
+     a <- -t[j] * u
+     a[j] <- 1 + a[j]
+     t[j] - sum(a * b)/ses[j] +
+            (sum(inverse_i_hat * (tcrossprod(a, u)))/ses[j] +
+             0.5 * t[j] * sum(inverse_i_hat * V))/ses[j]
+ }
```

[11] Using R packages `enrichwith` (Kosmidis, 2017) and `numDeriv` (Gilbert and Varadhan, 2016)

Beta regression: Reading accuracy and dyslexia



Empirical null rejection probabilities

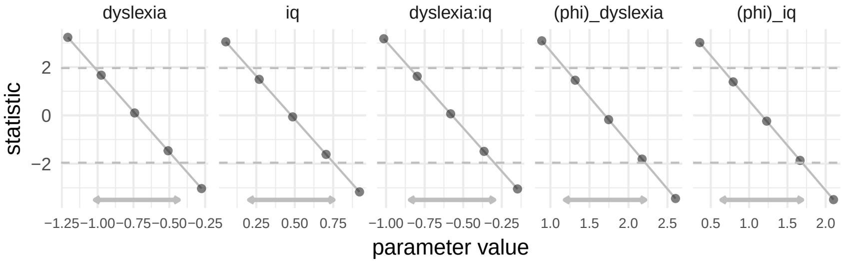


Confidence intervals based on t^*

$100(1-\alpha)\%$ confidence intervals based on t^* can be obtained by finding all ψ such that

$$|T(\hat{\theta};\psi) - B(\hat{\theta};\psi)| \le z_{1-\alpha/2}$$

where $z_{1-\alpha/2}$ is the $1-\alpha/2$ quantile of $N(0,1)$



Outline

Wald statistics with bias-corrected estimators

$$\tilde{t} = T(\tilde{\theta}; \psi_0)$$

with $\tilde{\theta}$ being a bias-corrected estimator with

$$E(\tilde{\theta} - \theta) = o(n^{-1})$$

Bias of $\tilde{t}$
$E\{T(\tilde{\theta}; \psi_0) - T(\theta; \psi_0)\} = \tilde{B}(\theta; \psi_0) + o(n^{-1/2})$ with

$$\tilde{B}(\theta; \psi_0) = \cancel{b(\theta)^\top \nabla T(\theta; \psi_0)} + \frac{1}{2}\mathrm{trace}\left[\{i(\theta)\}^{-1} \nabla\nabla^\top T(\theta; \psi_0)\right]$$

Use of bias-corrected estimators **eliminates a term**, but bias of Wald statistic is still $O(n^{-1/2})$

Models with categorical responses

Location-adjustment of $\tilde{t}$ is still fruitful

Categorical response models, where bias-correction leads to estimates that are **always finite** even in case where the MLE is infinite.

Outline

Lulling babies

Data

18 **matched pairs** of binomial observations on the effect of lulling on the crying of babies

Matching is per day and each day pair consists of the number of babies not crying out of a fixed number of control babies, and the outcome of lulling on a single child

Experiment involves 143 babies

Variables

`crying`	crying status of the baby (1 not crying; 0 crying)
`day`	the day of the experiment
`lull`	has the baby been lulled?

Aim: Test the effect of lulling on the crying of children

Logistic regression: lulling babies

Model

Y_{ij} is a Bernoulli random variable for the crying status of baby j in day i with probability μ_{ij} of not crying

$$\log \frac{\mu_{ij}}{1-\mu_{ij}} = \beta_i + \gamma z_{ij}$$

- z_{ij} is 1 if the jth child on day i was lulled, and 0 otherwise

Task

Test $\gamma = 0$ accounting for heterogeneity between days

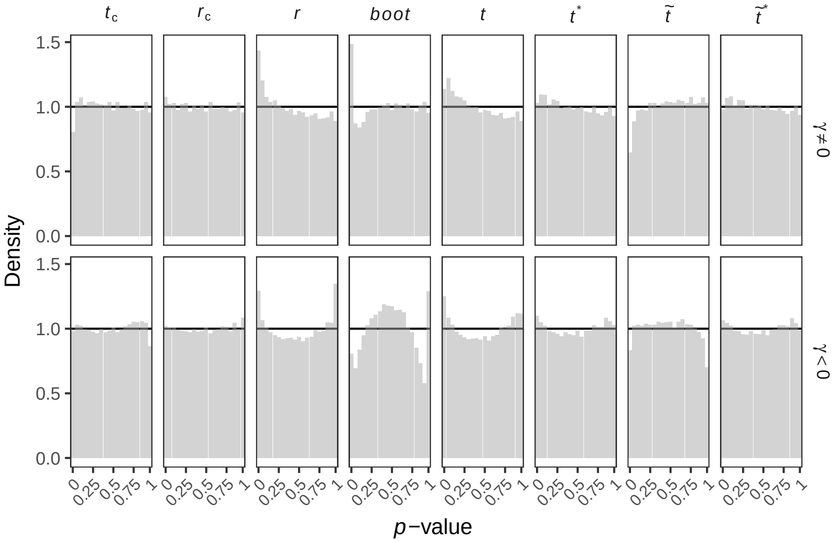
Testing for $\gamma = 0$

	t_c	r_c	r	t	t^*	$\tilde{t}$	$\tilde{t}^*$
statistic	1.8307	2.0214	2.1596	1.9511	1.9257	1.7362	1.9064
p-value	0.0671	0.0432	0.0308	0.0510	0.0541	0.0825	0.0566

t_c is the Wald statistic based on the maximum **conditional likelihood estimator**

r and r_c are the signed roots of the likelihood and conditional likelihood ratio statistics

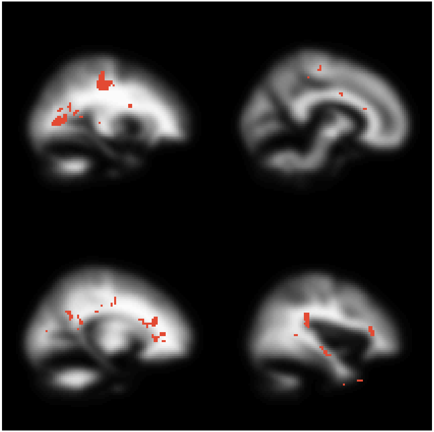
Empirical p-value distribution



[12] based on 50 000 samples from the mdoel with $\beta_1, \ldots, \beta_{18}$ set to their maximum likelihood estimates and $\gamma = 0$

Outline

Mass univariate regression for brain lesions



resolution: $91 \times 109 \times 91$ (902 629 voxels)

Sample

lesion maps for 50 patients[13]

Patient characteristics

multiple sclerosis type (MS)[14]

age

gender

disease duration (DD)

two disease severity measures (PASAT and EDSS)

Aim: Construct significance maps, highlighting voxels according to the evidence against the null hypothesis of no covariate effect

[14] from the supplementary material of Ge et al. (2014)

[15] 0 for relapsing-remitting and 1 for secondary progressive multiple sclerosis

Voxel-wise probit regressions

Lesion occurence in voxel j for patient i

$$Y_{ij} \sim \text{Bernoulli}(\pi_{ij})$$

Lesion probability

$$\Phi^{-1}(\pi_{ij}) = \beta_{j0} + \beta_{j1}\text{MS}_i + \beta_{j2}\text{age}_i + \beta_{j3}\text{gender}_i + \beta_{j4}\text{DD}_i + \beta_{j5}\text{PASAT}_i + \beta_{j6}\text{EDSS}_i$$

Results

Occurrence of infinite estimates

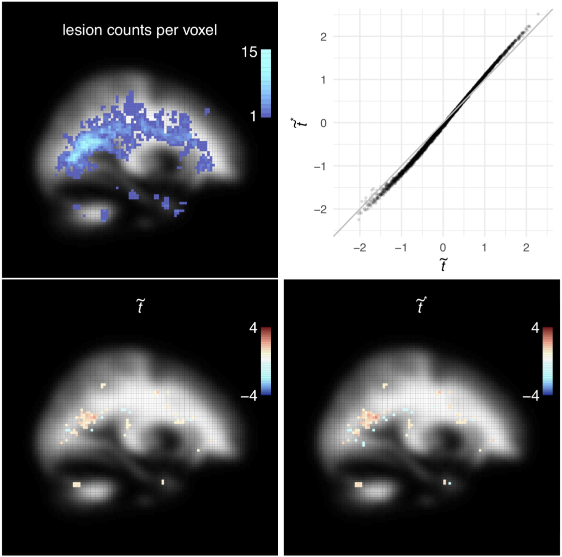
Covariate	Occurrence
MS	75.5%
age	63.7%
gender	78.3%
DD	63.7%
PASAT	63.6%
EDSS	63.2%

Failures in evaluation of r

Covariate	Occurrence
MS	19.2%
age	20.5%
sex	22.4%
DD	18.1%
PASAT	16.8%
EDSS	10.3%

[15] summaries based on voxels with lesion occurrence for at least one lesion across patients

Significance map for disease duration



18.9% of voxels with $|\tilde{t}| > 1$

24.8% of voxels with $|\tilde{t}^*| > 1$

Outline

Recap

Location-adjustment can deliver **substantial improvements** to Wald inference

Extra computational overhead is mainly due to matrix multiplications

Location adjustment with "robust"[16] variance-covariance matrices

Location adjustment with alternative estimators of estimator bias, including bootstrap and jackknife; particularly useful, e.g., for generalized linear mixed effects models

Extensions to other pivotal quantities, including Wald statistics for composite hypotheses, score statistics, or even directly p-values

[16] see, for example, MacKinnon and White (1985)

References I

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Location-adjusted Wald statistic

$$t^* = T(\hat{\theta}; \psi_0) - B(\hat{\theta}; \psi_0)$$

Preprint

Di Caterina C and Kosmidis I (2017). Location-adjusted Wald statistic for scalar parameters. *ArXiv e-prints*. `arXiv:1710.11217`

Software

waldi R package[17] (soon in CRAN!) for computing t^* for well-used models, including GLMs (`glm`, `brglm2`) and beta regression (`betareg`)

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[17] https://github.com/ikosmidis/waldi