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Bayesian Restricted Likelihood Methods: Conditioning on Insufficient Statistics in Bayesian Regression*

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Abstract. Bayesian methods have proven themselves to be successful across a wide range of scientific problems and have many well-documented advantages over competing methods. However, these methods run into difficulties for two major and prevalent classes of problems: handling data sets with outliers and dealing with model misspecification. We outline the drawbacks of previous solutions to both of these problems and propose a new method as an alternative. When working with the new method, the data is summarized through a set of insufficient statistics, targeting inferential quantities of interest, and the prior distribution is updated with the summary statistics rather than the complete data. By careful choice of conditioning statistics, we retain the main benefits of Bayesian methods while reducing the sensitivity of the analysis to features of the data not captured by the conditioning statistics. For reducing sensitivity to outliers, classical robust estimators (e.g., M-estimators) are natural choices for conditioning statistics. A major contribution of this work is the development of a data augmented Markov chain Monte Carlo (MCMC) algorithm for the linear model and a large class of summary statistics. We demonstrate the method on simulated and real data sets containing outliers and subject to model misspecification. Success is manifested in better predictive performance for data points of interest as compared to competing methods.

Keywords: Markov chain Monte Carlo, M-estimation, Robust regression.

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1 Introduction

Bayesian methods have provided successful solutions to a wide range of scientific problems, with their value having been demonstrated both empirically and theoretically. Bayesian inference relies on a model consisting of three elements: the prior distribution, the loss function, and the likelihood or sampling density. While formal optimality of Bayesian methods is unquestioned if one accepts the validity of all three of these elements, a healthy skepticism encourages us to question each of them. Concern about the prior distribution has been addressed through the development of techniques for subjective elicitation (Garthwaite et al., 2005; O'Hagan et al., 2006) and objective Bayesian methods (Berger, 2006). Concern about the loss function is reflected in, for example, the extensive literature on Bayesian hypothesis tests (Kass and Raftery, 1995).

The focus of this work is the development of techniques to handle imperfections in the likelihood $f(\boldsymbol{y}|\boldsymbol{\theta}) = L(\boldsymbol{\theta}|\boldsymbol{y})$. Concern for imperfections in the likelihood are reflected in work considering minimally informative likelihoods (Yuan and Clarke, 1999), sensitivities of inferences to perturbations in the model (Zhu et al., 2011), the specification of a class of models and the use of Bayesian model averaging over the class (Clyde and George, 2004), and considerations of such averaging when the specified class may not contain the so-called true data generating model (Bernardo and Smith, 2000; Clyde and Iversen, 2013; Clarke et al., 2013). In practice, the imperfections in a proposed likelihood often show themselves through the presence of outliers – cases not reflecting the phenomenon under study. There are three main solutions to Bayesian outlier-handling. The first is to replace the basic sampling density with a mixture model which includes one component for the "good" data and a second component for the "bad" data. With this approach, the good component of the sampling density is used for prediction of future good data. The second approach replaces the basic sampling density with a thick-tailed density in an attempt to discount outliers, yielding techniques that often provide solid estimates of the center of the distribution but do not easily translate to predictive densities for further good data. The third approach fits a flexible (typically nonparametric) model to the data, producing a Bayesian version of a density estimate for both good and bad data. In recent development, inference is made through the use of robust inference functions (Lee and MacEachern, 2014).

These traditional strategies all have their drawbacks. The outlier-generating processes may be transitory in nature, constantly shifting as the source of bad data changes. This prevents us from appealing to large-sample arguments to claim that, with enough data, we can nail down a model for both good and bad data combined. Instead of attempting to model both good and bad data, we propose a novel strategy for handling outliers. In a nutshell, we begin with a complete model as if all of the data are good. Rather than driving the move from prior to posterior by the full likelihood, we use only the likelihood driven by a few summary statistics which typically target inferential quantities of interest. We call this likelihood a restricted likelihood because conditioning is done on a restricted set of data; the set which satisfies the observed summary statistics. This is a formal update of the prior distribution based on the sampling density of the summary statistics.

The remainder of the paper is as follows: Section 2 introduces the Bayesian restricted likelihood and provides context with previous work, Section 3 demonstrates some advantages of the methods on simple examples, and Section 4 details a MCMC algorithm to apply the method to Bayesian linear models. This is a major contribution to the work providing an approach to apply the method on realistic examples. Many of the the technical proofs are in the Appendix 8 with `R` code available from the authors. Section 6 illustrates the method on a messy insurance industry data set containing many outliers with a novel twist on model evaluation. A discussion (Section 7) provides some final commentary on the new method.

2 Restricted Likelihood

2.1 Examples

To describe the use of the restricted likelihood, we begin with a pair of simple examples for the one-sample problem. For both, the model takes the data $\boldsymbol{y} = (y_1, \ldots, y_n)$ to be a random sample of size n from a continuous distribution indexed by a parameter vector $\boldsymbol{\theta}$, with pdf $f(y|\boldsymbol{\theta})$. The standard, or full, likelihood is $L(\boldsymbol{\theta}|\boldsymbol{y}) = \prod_{i=1}^{n} f(y_i|\boldsymbol{\theta})$.

The first example considers the case where a known subset of the data are known to be bad in

the sense of not informing us about $\boldsymbol{\theta}$. This case mimics the setting where outliers are identified and discarded before doing a formal analysis. Without loss of generality, we label the good cases 1 through $n-k$ and the bad cases $n-k+1$ through n. The relevant likelihood to be used to move from prior distribution to posterior distribution is clearly $L(\boldsymbol{\theta}|y_1,\ldots,y_{n-k}) = \prod_{i=1}^{n-k} f(y_i|\boldsymbol{\theta})$. For an equivalent analysis, we rewrite the full likelihood as the product of two pieces:

$$L(\boldsymbol{\theta}|\boldsymbol{y}) = \left(\prod_{i=1}^{n-k} f(y_i|\boldsymbol{\theta})\right)\left(\prod_{i=n-k+1}^{n} f(y_i|\boldsymbol{\theta})\right). \tag{1}$$

where the second factor may not actually depend on $\boldsymbol{\theta}$. We wish to keep the first factor and drop the second for better inference on $\boldsymbol{\theta}$.

The second example involves deliberate censoring of small and large observations. This is sometimes done as a precursor to the analysis of reaction time experiments (e.g., Ratcliff, 1993) where very small and large reaction times are physiologically implausible; explained by either anticipation or lack of attention of the subject. With lower and upper censoring times at t_1 and t_2, the post-censoring sampling distribution is of mixed form, with masses $F(t_1|\boldsymbol{\theta})$ at t_1 and $1-F(t_2|\boldsymbol{\theta})$ at t_2, and density $f(y|\boldsymbol{\theta})$ for $y \in (t_1,t_2)$. We adjust the original data y_i, producing $c(y_i)$ by defining $c(y_i) = t_1$ if $y_i \leq t_1$, $c(y_i) = t_2$ if $y_i \geq t_2$, and $c(y_i) = y_i$ otherwise. The adjusted update is performed with $L(\boldsymbol{\theta}|c(\boldsymbol{y}))$. Letting $g(t_1|\boldsymbol{\theta}) = F(t_1|\boldsymbol{\theta})$, $g(t_2|\boldsymbol{\theta}) = 1-F(t_2|\boldsymbol{\theta})$, and $g(y|\boldsymbol{\theta}) = f(y|\boldsymbol{\theta})$ for $y \in (t_1,t_2)$, we may rewrite the full likelihood as the product of two pieces

$$L(\boldsymbol{\theta}|\boldsymbol{y}) = \left(\prod_{i=1}^{n} g(c(y_i)|\boldsymbol{\theta})\right)\left(\prod_{i=1}^{n} f(y_i|\boldsymbol{\theta}, c(y_i)).\right), \tag{2}$$

Only the first part is retained the analysis. Several more examples are detailed in Lewis (2014).

2.2 Generalization

To generalize the approach in (1) and (2), we write the full likelihood in two pieces with a conditioning statistic $T(\boldsymbol{y})$, as indicated below:

$$L(\boldsymbol{\theta}|\boldsymbol{y}) \quad = \quad f(T(\boldsymbol{y})|\boldsymbol{\theta})\; f(\boldsymbol{y}|\boldsymbol{\theta}, T(\boldsymbol{y})). \tag{3}$$

Here, $f(T(\boldsymbol{y})|\boldsymbol{\theta})$ is the conditional pdf of $T(\boldsymbol{y})$ given $\boldsymbol{\theta}$ and $f(\boldsymbol{y}|\boldsymbol{\theta}, T(\boldsymbol{y}))$ is the conditional pdf of $\boldsymbol{y}$ given $\boldsymbol{\theta}$ and $T(\boldsymbol{y})$. In the dropped case example, the conditioning statistic is $T(\boldsymbol{y}) = (y_1, \ldots, y_{n-k})$. In the censoring example, the conditioning statistic is $T(\boldsymbol{y}) = (c(y_1), \ldots, c(y_n))$. We refer to $f(T(\boldsymbol{y})|\boldsymbol{\theta})$ as the restricted likelihood and $L(\boldsymbol{\theta}|\boldsymbol{y}) = f(\boldsymbol{y}|\boldsymbol{\theta})$ as the full likelihood.

Bayesian methods can make use of a restricted likelihood since $T(\boldsymbol{y})$ is a well-defined random variable with a probability distribution indexed by $\boldsymbol{\theta}$. This leads to the restricted likelihood posterior

$$\pi(\boldsymbol{\theta}|T(\boldsymbol{y})) \quad = \quad \frac{\pi(\boldsymbol{\theta}) f(T(\boldsymbol{y})|\boldsymbol{\theta})}{m(T(\boldsymbol{y}))}, \tag{4}$$

where $m(T(\boldsymbol{y}))$ is the marginal distribution of $T(\boldsymbol{y})$ under the prior distribution. Predictive statements for further (good) data rely on the model. For another observation, say y_{n+1}, we would have the predictive density

$$f(y_{n+1}|T(\boldsymbol{y})) = \int f(y_{n+1}|\boldsymbol{\theta}) \pi(\boldsymbol{\theta}|T(\boldsymbol{y})) \, d\boldsymbol{\theta}. \tag{5}$$

2.3 Literature review

Our motivation for the use of summary statistics in Bayesian inference is concern about outliers or, more generally, model misspecification. Specifically, the likelihood is not specified correctly and concentrating on using well chosen parts of the data can help improve the analysis (e.g., Wong and Clarke, 2004). Direct use of restricted likelihood for this reason appears in many areas of the literature. For example, the use of rank likelihoods is discussed by Savage (1969), Pettitt (1983, 1982), and more recently by Hoff et al. (2013). Lewis et al. (2012) make use of order statistics and robust estimators as choices for $T(\boldsymbol{y})$ in the location-scale setting. Asymptotic properties of restricted posteriors are studied by Doksum and Lo (1990), Clarke and Ghosh (1995), Yuan and Clarke (2004), and Hwang et al. (2005). The tenor of these asymptotic results is that, for a variety of conditioning statistics with non-trivial regularity conditions on prior, model, and likelihood, the posterior distribution resembles the asymptotic sampling distribution of the conditioning statistic.

Restricted likelihoods have also been used as practical approximations to a full likelihood. For example, Pratt (1965) appeals to heuristic arguments regarding approximate sufficiency to justify the

use of the restricted likelihood of the sample mean and standard deviation. Approximate sufficiency is also appealed to in the use of Approximate Bayesian Computation (ABC), which is related to our method. ABC is a collection of posterior approximation methods which has recently experienced success in applications to epidemiology, genetics, and quality control (see, for example, Tavaré et al., 1997; Pritchard et al., 1999; Marjoram et al., 2003; Fearnhead and Prangle, 2012). Interest typically lies in the full data posterior and ABC is used for computational convenience as an approximation. Consequently, effort is made to choose an approximately sufficient $T(\boldsymbol{y})$ and update to the ABC posterior by using the likelihood $L(\boldsymbol{\theta}|\mathcal{B}(\boldsymbol{y}))$, where $\mathcal{B}(\boldsymbol{y}) = \{\boldsymbol{y}^*|\rho(T(\boldsymbol{y}), T(\boldsymbol{y}^*)) < \epsilon\}$, ρ is a metric, and ϵ is a tolerance level. This is the likelihood conditioned on the collection of data sets that result in a $T(\cdot)$ within ϵ of the observed $T(\boldsymbol{y})$. With an approximately sufficient $T(\cdot)$ and a small enough ϵ, heuristically $L(\boldsymbol{\theta}|\mathcal{B}(\boldsymbol{y})) \approx L(\boldsymbol{\theta}|T(\boldsymbol{y})) \approx L(\boldsymbol{\theta}|\boldsymbol{y})$. Consequently, the ABC posterior approximates the full data posterior and efforts have been made to formalize what is meant by approximate sufficiency (e.g., Joyce and Marjoram, 2008). ABC is related to our method in that the conditioning is on something other than the data $\boldsymbol{y}$. However, we specifically seek to condition on an insufficient statistic to guard against misspecification in parts of the likelihood. Additionally, we develop methods where the conditioning is exact (i.e. $\epsilon = 0$).

This work extends the development of Bayesian restricted likelihood by arguing that deliberate choice of $T(\boldsymbol{y})$ on which to guide inference is sound practice. We also expand the class of conditioning statistics in which a formal Bayesian update can be achieved. Our methods do not rely on asymptotic properties, nor do they rely on approximate conditioning.

3 Illustrative Examples

Before discussing computational details, the method is applied to two simple examples on well known data sets to demonstrate its effectiveness in situations where outliers are a major concern. The full model in each case fits into the Bayesian linear regression framework discussed in Section 4.

The first example is an analysis of Simon Necomb's 66 measurements of the passage time of light (Stigler, 1977); two of which are significant outliers in the lower tail. The full model is a standard

location-scale Bayesian model also used in Lee and MacEachern (2014):

$$\beta \sim N(23.6, 2.04^2),\ \sigma^2 \sim IG(5, 10),\ y_i \overset{iid}{\sim} N(\beta, \sigma^2), i = 1, 2, \ldots, n = 66, \tag{6}$$

where y_i denotes the i^{th} (recoded) measurement of the passage time of light. β is interpreted as the passage time of light with the deviations $y_i - \beta_i$ representing measurement error. Four versions of the restricted likelihood are fit with conditioning statistics: 1) Huber's M-estimator for location with Huber's 'proposal 2' for scale 2) Tukey's M-estimator for location with Huber's 'proposal 2' for scale 3) LMS (least median squares) for location with associated estimator of scale and 4) LTS (least trimmed squares) for location with associated estimator of scale. The tuning parameters for the M-estimators are chosen to achieve 95% efficiency under normality (Huber and Ronchetti, 2009) and, for comparability, roughly 5% of the residuals are trimmed for LTS. Two additional approaches to outlier handling are fit: 1) the normal distribution is replaced with a t-distribution and, 2) the normal distribution is replaced with a mixture of two normals. The t-model assumes $y_i \overset{iid}{\sim} t_\nu(\beta, \sigma^2)$ with $\nu = 5$. The prior on σ^2 is $IG(5, \frac{\nu-2}{\nu}10)$ ensures that the prior on the variance is the same as the other models. The mixture takes the form: $y_i \overset{iid}{\sim} pN(\beta, \sigma^2) + (1-p)N(\beta, 10\sigma^2)$ with the prior $p \sim \text{beta}(20, 1)$ on the probability of belonging to the 'good' component.

The posterior of β under each model appears in Figure 1. The posteriors group into two batches. The normal model and restricted likelihood with LMS do not discount the outliers and have posteriors centered at low values of β. These posteriors are also quite diffuse. In contrast, the t model and the other restricted likelihood methods discount the outliers and have posterior centered at higher values. There is modest variation among these centers. Posteriors in this second group have less dispersion than those in the first group.

The pattern for predictive distributions differs (see bottom plot in Figure **??**). The normal and t models have widely dispersed predictive distributions. The other predictive distributions show much greater concentration. The restricted likelihood fits based on M-estimators (Tukey's and Huber's) are centered appropriately and are concentrated. The LMS predictive is concentrated, but it is poorly centered.

Overall, we find that the restricted likelihood methods based on M-estimators provide the most

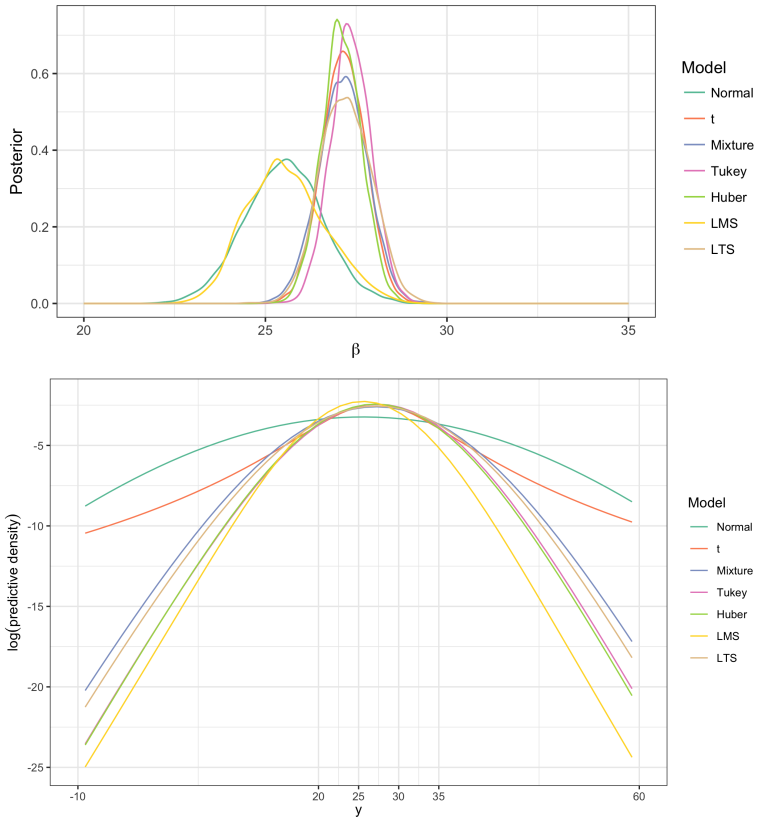

Figure 1: Results from the analysis of the speed of light data. Top: Posterior distributions of β under each model. Bottom: Log posterior predictive distributions under each model. The differences in the tails are emphasized in the bottom plot. The horizontal axis is strategically labeled to help compare the centers of the distributions in each of the plots.

attractive analysis for these data. They provide sharp and appropriate inference for parameters (β) and for prediction.

As a second example, a data set measuring the number of telephone calls in Belgium from 1950-1973 is analyzed. The outliers in this case are due to a change in measurement units on which calls were recorded for part of the data set. Specifically, for years 1964-1969 and parts of 1963 and 1970, the length of calls in minutes were recorded rather than the total number (Rousseeuw and Leroy,

1987). The full model is a standard normal Bayesian linear regression:

$$\boldsymbol{\beta} \sim N_2(\boldsymbol{\mu}_0, \boldsymbol{\Sigma}_0),\ \sigma^2 \sim IG(a, b), \boldsymbol{y} \sim N(X\boldsymbol{\beta}, \sigma^2 I), \tag{7}$$

where $\boldsymbol{\beta} = (\beta_0, \beta_1)^\top$, $\boldsymbol{y}$ is the vector of the logarithm of the number of calls, and X is the $n \times 2$ design matrix with a vector of 1's in the first column in the year covariate in the second. Prior parameters are fixed via a fit to the first 3 data points. In particular, the prior covariance for β is set to $\Sigma_0 = g\sigma_0^2(X_p^\top X_p)^{-1}$, with X_p the 3×2 design matrix for the first 3 data points, $g = n = 21$, and $\sigma_0 = 0.03$ and $\mu_0 = (1.87, 0.03)^\top$; the MLEs fit to the first three data points. This has the spirit of a unit information prior (Kass and Wasserman, 1995) but uses a design matrix for prior data not used in the fit. Finally $a = 2$ and $b = 1$.

Four models are compared: 1) the normal theory base model 2) A two component normal mixture model, 3) a t-model, and 4) a restricted likelihood model conditioning on Tukey's M-estimator for the slope and intercept with Huber's 'proposal 2' for scale. Each model is fit to the remaining 21 data points. The normal theory model is also fit a second time after removing observations 14-21 (years 1963 - 1970). This includes the obvious large outliers as well as the two smaller outliers at the beginning and end of this sequence of points caused by the change in measurement units. The mixture model assumes different mean regression functions and variances for each component, but keeps the same, relatively non-informative priors. The probability of belonging to the first component is given a beta$(5, 1)$ prior. The heavy-tailed model fixes the degrees of freedom to 5 with the same adjustment to the prior on σ^2 as above.

The data and 95% credible bands of the posterior predictive distribution under each model are displayed in Figure 2. The normal model fit to all cases results in a very wide posterior predictive distribution due to an inflated estimate of the variance. Since the t-model assumes the data are heavy-tailed, the posterior predictive distribution is also wide. Additionally, both of these are biased upward by the outliers in the upper tail of the assumed data distribution. On the other hand, the predictive distribution under the restricted likelihood approach is much more precise and is close to that of the normal theory fit to the non-outlying cases. It is also close to the two component mixture results where the predictive distribution is formulated using only the good component. The mixture

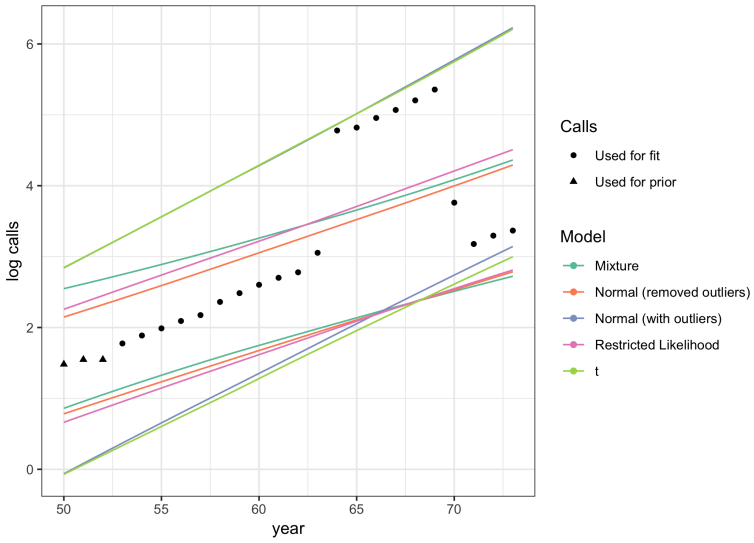

Figure 2: Pointwise posterior predictive intervals of log(calls) under the normal theory model fit to the non-outliers, the restricted likelihood model with Tukey's M-estimator for the slope and intercept with Huber's 'proposal 2' for scale, and a heavy-tailed t-distribution model. The first three data points were used to specify the prior with each model using the remaining 21 for fitting. The normal theory model was also fit after removing observations 14-20 (years 1963 - 1970).

model involves explicitly modeling the outlier generating mechanism. In more complex situations where the outlier generating mechanism is transient (i.e. ever changing and more complex than just a unit error in the recording), modeling the outliers becomes more difficult. As in classical robust estimation, the restricted likelihood approach avoids explicitly modeling the outliers.

4 Restricted Likelihood for the Linear Model

The simple examples in the previous section highlight that productive use of the restricted likelihood relies on a good choice of $T(\boldsymbol{y})$. This work focuses on robustness in linear models where natural choices include many used above: M-estimators in the tradition of Huber (1964), least median squares (LMS), and least trimmed squares (LTS). For these choices the restricted likelihood is not available in closed form, making computation of the restricted posterior a challenge. For low-dimensional statistics $T(\boldsymbol{y})$

and parameters $\boldsymbol{\theta}$, direct computational strategies described in Lewis (2014) can be used to estimate the restricted posterior conditioned on essentially any statistic. These strategies rely on density estimation $f(T(\boldsymbol{y})|\theta)$ using samples of $T(\boldsymbol{y})$ for many values of $\boldsymbol{\theta}$; a strategy which breaks down in higher dimensions. This section outlines a data augmented MCMC algorithm that can be applied to the Bayesian linear model when $T(\boldsymbol{y})$ consists of estimates of the regression coefficients and scale parameter.

4.1 The Bayesian linear model

We focus on the use of restricted likelihood for the Bayesian linear model with a standard formulation:

$$\begin{aligned} \boldsymbol{\theta} &= (\boldsymbol{\beta}, \sigma^2) \sim \pi(\boldsymbol{\theta}) \\ y_i &= x_i^\top \boldsymbol{\beta} + \epsilon_i, \text{ for } i = 1, \ldots, n \end{aligned} \tag{8}$$

where x_i and $\boldsymbol{\beta} \in \mathbb{R}^p$, $\sigma^2 \in \mathbb{R}^+$, and the ϵ_i are independent draws from a distribution with center 0 and scale σ. X denotes the design matrix whose rows are $x_i^\top$. For the restricted likelihood model, conditioning statistics are assumed to be of the form $T(\boldsymbol{y}) = (\boldsymbol{b}(X, \boldsymbol{y}), s(X, \boldsymbol{y}))$ where $\boldsymbol{b}(X, \boldsymbol{y}) = (b_1(X, \boldsymbol{y}), \ldots, b_p(X, \boldsymbol{y}))^\top \in \mathbb{R}^p$ is an estimator for the regression coefficients and $s(X, \boldsymbol{y}) \in \{0\} \cup \mathbb{R}^+$ is an estimator of the scale. Throughout, observed data and summary statistic is denoted by $\boldsymbol{y}_{obs}$ and $T(\boldsymbol{y}_{obs}) = (\boldsymbol{b}(X, \boldsymbol{y}_{obs}), s(X, \boldsymbol{y}_{obs}))$, respectively. Several conditions are imposed on the model and statistic to ensure validity of the MCMC algorithm:

C1. The $n \times p$ design matrix, X, whose i^{th} row is $x_i^\top$, is of full column rank.

C2. The ϵ_i are a random sample from some distribution which has a density with respect to Lebesgue measure on the real line and for which the support is the real line.

C3. $\boldsymbol{b}(X, \boldsymbol{y})$ is almost surely continuous and differentiable with respect to $\boldsymbol{y}$.

C4. $s(X, \boldsymbol{y})$ is almost surely positive, continuous, and differentiable with respect to $\boldsymbol{y}$.

C5. $\boldsymbol{b}(X, \boldsymbol{y} + X\boldsymbol{v}) = \boldsymbol{b}(X, \boldsymbol{y}) + \boldsymbol{v}$ for all $\boldsymbol{v} \in \mathbb{R}^p$.

C6. $\boldsymbol{b}(X, a\boldsymbol{y}) = a\boldsymbol{b}(X, \boldsymbol{y})$ for all constants a.

C7. $s(X, \boldsymbol{y} + X\boldsymbol{v}) = s(X, \boldsymbol{y})$ for all $\boldsymbol{v} \in \mathbb{R}^p$.

C8. $s(X, a\boldsymbol{y}) = |a|s(X, \boldsymbol{y})$ for all constants a.

Properties C5 and C6 of $\boldsymbol{b}$ are called *regression* and *scale equivariance*, respectively. Properties C7 and C8 are called *regression invariance* and *scale equivariance*. Many estimators satisfy the above properties, including simultaneous M-estimators (Huber and Ronchetti, 2009; Maronna et al., 2006) for which the R package `brlm` (`github.com/jrlewi/brlm`) is available to implement the MCMC described here. Further software development is required to extend the MCMC implementation beyond these M-estimators. The package also implements the direct computational methods described in Lewis (2014). These methods are effective in lower dimensional problems and were used in several of the examples in Section 3.

4.2 Computational strategy

The general style of algorithm we present is a data augmented MCMC targeting $f(\boldsymbol{\theta}, \boldsymbol{y}|T(\boldsymbol{y}) = T(\boldsymbol{y}_{obs}))$, the joint distribution of $\boldsymbol{\theta}$ and the full data given the summary statistic $T(\boldsymbol{y}_{obs})$. The Gibbs sampler (Gelfand and Smith, 1990) iteratively samples from the full conditionals 1) $\pi(\boldsymbol{\theta}|\boldsymbol{y}, T(\boldsymbol{y}) = T(\boldsymbol{y}_{obs}))$ and 2) $f(\boldsymbol{y}|\boldsymbol{\theta}, T(\boldsymbol{y}) = T(\boldsymbol{y}_{obs}))$. When $\boldsymbol{y}$ has the summary statistic $T(\boldsymbol{y}) = T(\boldsymbol{y}_{obs})$, the first full conditional is the same as the full data posterior $\pi(\boldsymbol{\theta}|\boldsymbol{y})$. In this case, the condition $T(\boldsymbol{y}) = T(\boldsymbol{y}_{obs})$ is redundant. This allows us to make use of conventional MCMC steps for this generation. For typical regression models, algorithms abound. Details of the recommended algorithms depend on details of the prior distribution and sampling density and we assume this can be done (see e.g., Liu, 1994; Liang et al., 2008).

For a typical model and conditioning statistic, the second full conditional $f(\boldsymbol{y}|\boldsymbol{\theta}, T(\boldsymbol{y}) = T(\boldsymbol{y}_{obs}))$ is not available in closed form. We turn to Metropolis-Hastings (Hastings, 1970), using the strategy of proposing full data $\boldsymbol{y} \in \mathcal{A} := \{\boldsymbol{y} \in \mathbb{R}^n | T(\boldsymbol{y}) = T(\boldsymbol{y}_{obs})\}$ from a well defined distribution with support $\mathcal{A}$ and either accepting or rejecting the proposal. Let $\boldsymbol{y}_p, \boldsymbol{y}_c \in \mathcal{A}$ represent the proposed and

current full data, respectively. Denote the proposal distribution for $\boldsymbol{y}_p$ by $p(\boldsymbol{y}_p|\boldsymbol{\theta}, T(\boldsymbol{y}_p) = T(\boldsymbol{y}_{obs})) = p(\boldsymbol{y}_p|\boldsymbol{\theta}, \boldsymbol{y}_p \in \mathcal{A}) = p(\boldsymbol{y}_p|\boldsymbol{\theta})$. The last equality follows from the fact that our $p(\cdot|\boldsymbol{\theta})$ assigns probability one to the event $\{\boldsymbol{y}_p \in \mathcal{A}\}$. These equalities still hold if the dummy argument $\boldsymbol{y}_p$ is replaced with $\boldsymbol{y}_c$. The conditional density is

$$f(\boldsymbol{y}|\boldsymbol{\theta}, \boldsymbol{y} \in \mathcal{A}) = \frac{f(\boldsymbol{y}|\boldsymbol{\theta}) I(\boldsymbol{y} \in \mathcal{A}|\boldsymbol{y}, \boldsymbol{\theta})}{\int_{\mathcal{A}} f(\boldsymbol{y}|\boldsymbol{\theta}) d\boldsymbol{y}} = \frac{f(\boldsymbol{y}|\boldsymbol{\theta})}{\int_{\mathcal{A}} f(\boldsymbol{y}|\boldsymbol{\theta}) d\boldsymbol{y}}$$

for $\boldsymbol{y} \in \mathcal{A}$. This includes both $\boldsymbol{y}_p$ and $\boldsymbol{y}_c$. The Metropolis-Hastings acceptance probability is the minimum of 1 and R where,

$$\begin{aligned} R &= \frac{f(\boldsymbol{y}_p|\boldsymbol{\theta}, \boldsymbol{y}_p \in \mathcal{A})}{f(\boldsymbol{y}_c|\boldsymbol{\theta}, \boldsymbol{y}_c \in \mathcal{A})} \frac{p(\boldsymbol{y}_c|\boldsymbol{\theta}, \boldsymbol{y}_c \in \mathcal{A})}{p(\boldsymbol{y}_p|\boldsymbol{\theta}, \boldsymbol{y}_p \in \mathcal{A})} && (9) \\ &= \frac{f(\boldsymbol{y}_p|\boldsymbol{\theta})}{\int_{\mathcal{A}} f(\boldsymbol{y}|\boldsymbol{\theta}) d\boldsymbol{y}} \frac{\int_{\mathcal{A}} f(\boldsymbol{y}|\boldsymbol{\theta}) d\boldsymbol{y}}{f(\boldsymbol{y}_c|\boldsymbol{\theta})} \frac{p(\boldsymbol{y}_c|\boldsymbol{\theta})}{p(\boldsymbol{y}_p|\boldsymbol{\theta})} && (10) \\ &= \frac{f(\boldsymbol{y}_p|\boldsymbol{\theta})}{f(\boldsymbol{y}_c|\boldsymbol{\theta})} \frac{p(\boldsymbol{y}_c|\boldsymbol{\theta})}{p(\boldsymbol{y}_p|\boldsymbol{\theta})}. && (11) \end{aligned}$$

For the models we consider, evaluation of $f(\boldsymbol{y}|\boldsymbol{\theta})$ is straightforward. Therefore, the difficulty in implementing this Metropolis-Hastings step manifests itself in the ability to both simulate from and evaluate $p(\boldsymbol{y}_p|\boldsymbol{\theta})$; the well defined distribution with support $\mathcal{A}$. We now discuss such an implementation method for the linear model in (8).

Construction of the proposal

Our computational strategy relies on proposing $\boldsymbol{y}$ such that $T(\boldsymbol{y}) = T(\boldsymbol{y}_{obs})$ where $T(\cdot) = (\boldsymbol{b}(X, \cdot), s(X, \cdot))$ satisfies the conditions C3-C8. It is not a simple matter to do this directly, but with the specified conditions, it is possible to scale and shift any $\boldsymbol{z}^*$ which generates a positive scale estimate to such a $\boldsymbol{y}$ via the following Theorem, whose proof is in the appendix.

Theorem 4.1. *Assume that conditions C4-C8 hold. Then, any vector $\boldsymbol{z}^* \in \mathbb{R}^n$ with conditioning statistic $T(\boldsymbol{z}^*)$ for which $s(X, \boldsymbol{z}^*) > 0$ can be transformed into $\boldsymbol{y}$ with conditioning statistic $T(\boldsymbol{y}) = T(\boldsymbol{y}_{obs})$ through the transformation*

$$\boldsymbol{y} = h(\boldsymbol{z}^*) := \frac{s(X, \boldsymbol{y}_{obs})}{s(X, \boldsymbol{z}^*)} \boldsymbol{z}^* + X \left(\boldsymbol{b}(X, \boldsymbol{y}_{obs}) - \boldsymbol{b}(X, \frac{s(X, \boldsymbol{y}_{obs})}{s(X, \boldsymbol{z}^*)} \boldsymbol{z}^*) \right).$$

Using the theorem, the general idea is to first start with an initial vector $\boldsymbol{z}^*$ drawn from a known distribution, say $p(\boldsymbol{z}^*)$, and transform via $h(\cdot)$ to $\boldsymbol{y} \in \mathcal{A}$. The proposal density $p(\boldsymbol{y}|\boldsymbol{\theta})$ is then a change-of-variables adjustment on $p(\boldsymbol{z}^*)$ derived from $h(\cdot)$. In general however, the mapping $h(\cdot)$ is many-to-one: for any $\boldsymbol{v} \in \mathbb{R}^n$ and any $c \in \mathbb{R}^+$, $c\boldsymbol{z}^* + X\boldsymbol{v}$ map to the same $\boldsymbol{y}$. This makes the change-of-variables adjustment difficult. We handle this by first noticing that the set $\mathcal{A}$ is an $n-p-1$ dimensional space: there are p constraints imposed by the regression coefficients and one further constraint imposed by the scale. Hence, we restrict the initial $\boldsymbol{z}^*$ to an easily understood $n-p-1$ dimensional space. Specifically, this space is the unit sphere in the orthogonal complement of the column space of the design matrix: $\mathbb{S} := \{\boldsymbol{z}^* \in \mathcal{C}^\perp(X) \mid ||\boldsymbol{z}^*|| = 1\}$, where $\mathcal{C}(X)$ and $\mathcal{C}^\perp(X)$ are the column space of X and its orthogonal complement, respectively. The mapping $h : \mathbb{S} \to \mathcal{A}$ is one-to-one and onto. The one-to-one property makes the change of variables more feasible. The onto property is important so that the support of the proposal distribution (i.e. the range of $h(\cdot)$) contains the support of the target $f(\boldsymbol{y}|\theta, y \in \mathcal{A})$; a necessary condition for convergence of the Metroplis-Hastings algorithm (in this case the supports are both $\mathcal{A}$). A formal theorem and proof is provided in the appendix.

Given the one-to-one and onto mapping $h : \mathbb{S} \to \mathcal{A}$, the general proposal strategy is summarized as follows:

1. Sample $\boldsymbol{z}^*$ from a distribution with known density on $\mathbb{S}$.

2. Set $\boldsymbol{y} = h(\boldsymbol{z}^*)$ and calculate the Jacobian of this transformation in two steps.

   (a) Scale from $\mathbb{S}$ to the set $\Pi(\mathcal{A}) := \{\boldsymbol{z} \in \mathbb{R}^n | \ \exists\ \boldsymbol{y} \in \mathcal{A}\ s.t.\ \boldsymbol{z} = Q\boldsymbol{y}\}$ with $Q = I - XX^\top$. [1] $\Pi(\mathcal{A})$ is the projection of $\mathcal{A}$ onto $\mathcal{C}^\perp(X)$ and, by condition C7, every element of this set has $s(X, \boldsymbol{z}) = s(X, \boldsymbol{y}_{obs})$. Specifically, set $\boldsymbol{z} = \frac{s(X,\boldsymbol{y}_{obs})}{s(X,\boldsymbol{z}^*)}\boldsymbol{z}^*$. There are two pieces of this Jacobian: one for the scaling and one for the mapping of the sphere onto $\Pi(\mathcal{A})$. The latter piece is given in equation (12).

   (b) Shift from $\Pi(\mathcal{A})$ to $\mathcal{A}$: $\boldsymbol{y} = \boldsymbol{z} + X\left(\boldsymbol{b}(X, \boldsymbol{y}_{obs}) - \boldsymbol{b}(X, \boldsymbol{z})\right)$. This shift is along the column

[1] We have used condition C1 to assume without loss of generality that the columns of X form an orthonormal basis for $\mathcal{C}(X)$ (i.e., $X^\top X = I$).

space of X to the unique element in $\mathcal{A}$. The Jacobian of this transformation is given by equation (13).

The final proposal distribution including the complete Jacobian is given in equation (14) with details in the next section. Before giving these details we provide a visualization in Figure 3 of each of the sets described above using a notional example to aid in the understanding of the strategy we are taking. In the figure, $n = 3$, $p = 1$, and the conditioning statistic is $T(\boldsymbol{y}) = (\min(\boldsymbol{y}), \sum(y_i - \min(\boldsymbol{y}))^2)$. The set $\mathcal{A}$ is depicted for $T(\boldsymbol{y}_{obs}) = (0, 1)$ which we describe as a "warped triangle" in light blue, with each side corresponding to a particular coordinate of $\boldsymbol{y}$ being the minimum value of zero. The other two coordinates are restricted by the scale statistic to lie on the quarter circle of radius one in the positive orthant. In this example, the column vector $X = \mathbf{1}$ (shown as a reference) spans $\mathcal{C}(X)$ and $\mathbb{S}$ is a unit circle on the orthogonal plane (shown in red). $\Pi(\mathcal{A})$ is depicted as the bowed triangle in dark blue. We will come back to this artificial example in the next section in an attempt to visualize the Jacobian calculations.

Evaluation of the proposal density

We now explain each step in computing the Jacobian described above.

Scale from $\mathbb{S}$ to $\Pi(\mathcal{A})$

The first step is constrained to $\mathcal{C}^{\perp}(X)$ and scales the initial $\boldsymbol{z}^*$ to $\boldsymbol{z} = \frac{s(X,\boldsymbol{y}_{obs})}{s(X,\boldsymbol{z}^*)}\boldsymbol{z}^*$. For the Jacobian, we consider two substeps: first, the distribution on $\mathbb{S}$ is transformed to that along a sphere of radius $r = \|\boldsymbol{z}\| = s(X, \boldsymbol{y}_{obs})/s(X, \boldsymbol{z}^*)$. By comparison of the volumes of these spheres, this transformation contributes a factor of $r^{-(n-p-1)}$ to the Jacobian. For the second substep, the sphere of radius r is deformed onto $\Pi(\mathcal{A})$. This deformation contributes an attenuation to the Jacobian equal to the ratio of infinitesimal volumes in the tangent spaces of the sphere and $\Pi(\mathcal{A})$ at $\boldsymbol{z}$. Restricting to $\mathcal{C}^{\perp}(X)$, this ratio is the cosine of the angle between the normal vectors of the two sets at $\boldsymbol{z}$. The normal to the sphere is its radius vector $\boldsymbol{z}$. The normal to $\Pi(\mathcal{A})$ is given in the following lemma.

Lemma 4.2. *Assume that conditions C1-C2, C4, and C7 hold and $\boldsymbol{y} \in \mathcal{A}$. Let $\nabla s(X, \boldsymbol{y})$ denote the*

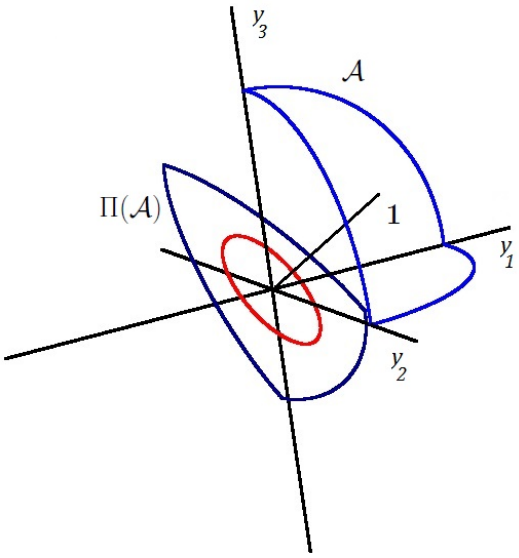

Figure 3: A depiction of $\mathcal{A}$, $\Pi(\mathcal{A})$, and the unit circle for the illustrative example where $b_1(\mathbf{1}, \boldsymbol{y}) = \min(\boldsymbol{y}) = 0$ and $s(\mathbf{1}, \boldsymbol{y}) = \sum(y_i - b_1(\mathbf{1}, \boldsymbol{y}))^2 = 1$. $\mathcal{A}$ is the combination of three quarter circles, one on each plane defined by $y_i = 0$. The projection of this manifold onto the deviation space is depicted by the bowed triangular shape in the plane defined by $\sum y_i = 0$. The circle in this plane represents the sample space for the intermediate sample $\boldsymbol{z}^*$. Also depicted is the vector $\mathbf{1}$, the design matrix for the location and scale setting.

gradient of the scale statistic with respect to the data vector evaluated at $\boldsymbol{y}$. *Then* $\nabla s(X, \boldsymbol{y}) \in \mathcal{C}^{\perp}(X)$ *and is normal to* $\Pi(\mathcal{A})$ *at* $\boldsymbol{z} = Q\boldsymbol{y}$ *in* $\mathcal{C}^{\perp}(X)$.

As a result of the lemma, the contribution to the Jacobian of this attenuation is

$$\cos(\gamma) = \frac{\nabla s(X, \boldsymbol{y})^{\top} \boldsymbol{z}}{\|\nabla s(X, \boldsymbol{y})\| \|\boldsymbol{z}\|}, \tag{12}$$

where γ is the angle between the two normal vectors. This step is visualized in Figure 4.2 for the notional location-scale example. The figure pictures only the $\mathcal{C}^{\perp}(X)$, which in this case is a plane. The unit sphere (here, the solid circle) is stretched to the dashed sphere contributing $r^{-(n-p-1)}$ to the Jacobian as seen in panel (a). In panel (b), the dashed circle is transformed onto $\Pi(\mathcal{A})$ contributing $\cos(\gamma)$ to the Jacobian. The normal vectors in panel (b) are orthogonal to the tangent vectors of $\Pi(\mathcal{A})$ and the circle.

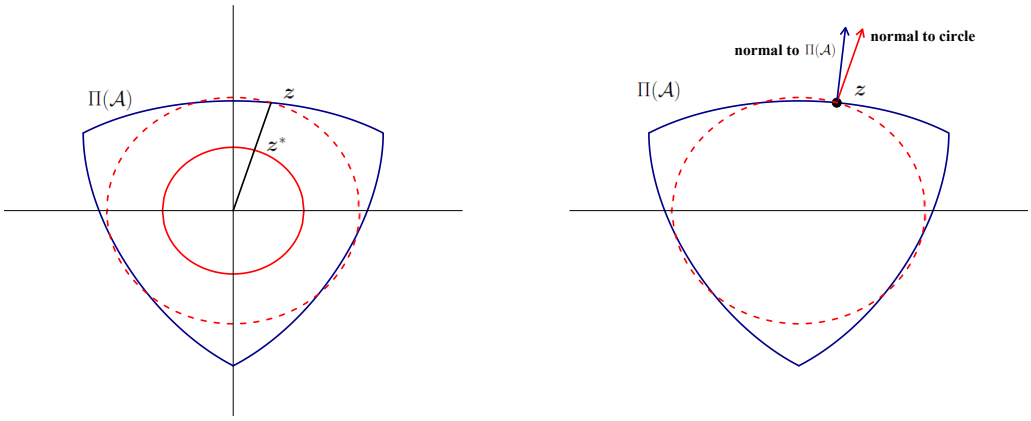

Figure 4: Visualization of the scaling from z^* to z. Left: the first substep scales z^* on the unit circle to the circle of radius $r = ||z||$; resulting in a change-of-variables transformation for the unit circle to that of radius r. The contribution to the Jacobian of this transformation is $r^{-(n-p-1)}$. Right: The second substep accounts for the the change-of-variables transformation from the circle of radius r to $\Pi(\mathcal{A})$. The normal vectors to these two sets are used to calculated the contribution to the Jacobian of this part of the transformation are shown in the figure.

Shift from $\Pi(\mathcal{A})$ to $\mathcal{A}$

The final piece of the Jacobian comes from the transformation from $\Pi(\mathcal{A})$ to $\mathcal{A}$. This step involves a shift of $\boldsymbol{z}$ to $\boldsymbol{y}$ along the column space of X. Since the shift depends on $\boldsymbol{z}$, the density on the set $\Pi(\mathcal{A})$ is deformed by the shift. The contribution of this deformation to the Jacobian is, again, the ratio of the infinitesimal volumes along $\Pi(\mathcal{A})$ at $\boldsymbol{z}$ to the corresponding volume along $\mathcal{A}$ at $\boldsymbol{y}$. The ratio is calculated by considering the volume of the projection of a unit hypercube in the tangent space of $\mathcal{A}$ at $\boldsymbol{y}$ onto $\mathcal{C}^{\perp}(X)$. Computational details are given in the following lemmas and subsequent theorem. Throughout, let $\mathcal{T}_y(\mathcal{A})$ and $\mathcal{T}_y^{\perp}(\mathcal{A})$ denote the tangent space to $\mathcal{A}$ at $\boldsymbol{y}$ and its orthogonal complement respectively. All gradients denote with ∇ are with respect to the data vector.

Lemma 4.3. *Assume that conditions C1-C5 and C7-C8 hold. Then the $p+1$ gradient vectors $\nabla s(X, \boldsymbol{y}), \nabla b_1(X, \boldsymbol{y}), \ldots, \nabla b_p(X, \boldsymbol{y})$ form a basis for $\mathcal{T}_y^{\perp}(\mathcal{A})$ with probability one.*

The lemma describes construction of a basis for $\mathcal{T}_y^{\perp}(\mathcal{A})$, leading to a basis for $\mathcal{T}_y(\mathcal{A})$. Both of these bases can be orthonormalized. Let $A = [a_1, \ldots, a_{n-p-1}]$ and $B = [b_1, \ldots, b_{p+1}]$ denote the matrices whose columns contain the orthonormal bases for $\mathcal{T}_y(\mathcal{A})$ and $\mathcal{T}_y^{\perp}(\mathcal{A})$, respectively. The columns in A define a unit hypercube in $\mathcal{T}_y(\mathcal{A})$ and their projections onto $\mathcal{C}^{\perp}(X)$ define a parallelepiped. We defer construction of A until later.

Lemma 4.4. *Assume that conditions C1-C5 and C7-C8 hold. Then the $n \times (n-p-1)$ dimensional matrix $P = QA$ is of full column rank.*

As a consequence of this lemma, the parallelepiped spanned by the columns of P is not degenerate (it is $n-p-1$ dimensional), and its volume is given by

$$\text{Vol}(P) := \sqrt{\det(P^\top P)} = \prod_{i=1}^{r} \sigma_i \tag{13}$$

where $r = \text{rank}(P) = n-p-1$ and $\sigma_1 \geq \sigma_2 \geq \cdots \geq \sigma_r > 0$ are the singular values of P (e.g., Miao and Ben-Israel (1992)). Combining Lemmas 4.3 and 4.4 above leaves us with the following result concerning the calculation of the desired Jacobian.

Theorem 4.5. *Assume that conditions C1-C5 and C7-C8 hold. Then the Jacobian of the transformation from the distribution along $\Pi(\mathcal{A})$ to that along $\mathcal{A}$ is equal to the volume given in (13).*

The proposal density

Putting all the pieces of the Jacobian together we have the following result. Any dependence on other variables, including current states in the Markov chain, is made implicit.

Theorem 4.6. *Assume that conditions C1-C8 hold. Let $\boldsymbol{z}^*$ be sampled on the unit sphere in $\mathcal{C}^{\perp}(X)$ with density $p(\boldsymbol{z}^*)$. Using the transformation of $\boldsymbol{z}^*$ to $\boldsymbol{y} \in \mathcal{A}$ described in Theorem 4.1, the density of $\boldsymbol{y}$ is*

$$p(\boldsymbol{y}) = p(\boldsymbol{z}^*) r^{-(n-p-1)} \cos(\gamma)\, Vol(P) \tag{14}$$

where $r = s(X, \boldsymbol{y}_{obs})/s(X, \boldsymbol{z}^*)$, and $\cos(\gamma)$ and $Vol(P)$ are as in equations (12) and (13), respectively.

A few details for computingthe needed quantities are worth further explanation. Computing Vol(P) involves finding an orthornormal matrix A whose columns span $\mathcal{T}_y(\mathcal{A})$. This matrix can be

found by supplementing B with a set of n linearly independent columns on the right, and apply Gram-Schmidt orthonormalization. This is $\mathcal{O}(n^3)$ and is infeasibly slow when n is large because it must be repeated at each iterate of the MCMC when a complete data set is drawn. However, using results related to *principal angles* found in Miao and Ben-Israel (1992) the volume (13) can be computed using only B. B is constructed by Gram-Schmidt orthogonalization of $\nabla s(X, \boldsymbol{y}), \nabla b_1(X, \boldsymbol{y}), \ldots, \nabla b_p(X, \boldsymbol{y})$, which is $\mathcal{O}(np^2)$; a considerable reduction in computational burden when $n \gg p$. The following corollary formally states how computation of A can be circumvented.

Corollary 4.7. *Let U be a matrix whose columns form an orthonormal basis for $\mathcal{C}(X)$ and set $Q = WW^\top$ where the columns of W form an orthonormal basis for $\mathcal{C}^\perp(X)$. Then the non-unit singular values of $U^\top B$ are the same as the non-unit singular values of $W^\top A$.*

The lemma implies the $\mathrm{Vol}(P)$ is the product of the singular values of $U^\top B$.

Second, the gradients of $\nabla s(X, \boldsymbol{y}), \nabla b_1(X, \boldsymbol{y}), \ldots, \nabla b_p(X, \boldsymbol{y})$ are easily computed. For example, below we consider M-estimators defined by the estimating equations:

$$\begin{aligned} \sum_{i=1}^{n} \psi\left(\frac{y_i - x_i^\top \boldsymbol{b}(\boldsymbol{y}, X)}{s(\boldsymbol{y}, X)}\right) &= 0 \\ \sum_{i=1}^{n} \chi\left(\frac{y_i - x_i^\top \boldsymbol{b}(\boldsymbol{y}, X)}{s(\boldsymbol{y}, X)}\right) &= 0, \end{aligned} \tag{15}$$

where ψ and χ are almost surely differentiable. Differentiating this system of equations with respect to each y_i can be used to find the gradients. In theory, finite differences could also be used.

5 Simulated Data

We study the performance of restricted likelihood methods in a hierarchical setting where the data are contaminated with outliers. Specifically, simulated data come from the following model:

$$\begin{aligned} &\theta_i \sim N(\mu, \tau^2), \ i = 1, 2, \ldots, 90 \\ &y_{ij} \sim (1 - p_i)N(\theta_i, \sigma^2) + p_i N(\theta_i, m_i\sigma^2), \ j = 1, 2, ..., n_i \end{aligned} \tag{16}$$

with $\mu = 0, \tau^2 = 1, \sigma^2 = 4$. The values of p_i, m_i, and n_i depend on the group and are formed using 5 replicates of the full factorial design over factors p_i, m_i, n_i with levels $p_i = .1, .2, .3$, $m_i = 9, 25$, and $n_i = 25, 50, 100$. This results in 90 groups that have varying levels of outlier contamination and sample size. We wish to build models that offer good prediction for the good portion of data within each group. The full model for fitting is a corresponding normal model without contamination:

$$\begin{aligned} &\mu \propto 1,\ \tau^2 \propto \tau^{-2}, \\ &\theta_i \sim N(\mu, \tau^2),\ \sigma_i^2 \sim IG(a_s, b_s),\ i = 1, 2, \ldots, 90, \\ &y_{ij} \sim N(\theta_i, \sigma_i^2),\ j = 1, 2, \ldots, n_i. \end{aligned} \tag{17}$$

For the restricted likelihood versions we condition on robust M-estimators of location and scale in each group: $T_i(y_{i1}, \ldots, y_{in_i}) = (\hat{\theta}_i, \hat{\sigma}_i^2), i = 1, 2, ..., 90$. These estimators are solutions to equation (15) (where $x_i \equiv 1$) with user specified ψ and χ functions designed to discount outliers. The two versions use Huber's and Tukey's ψ function, while both versions use Huber's χ function. The tuning parameters associated with these functions are chosen so that the estimators are 95% efficient under normally distributed data. These classical M-estimators are commonly used in robust regression settings (Huber and Ronchetti, 2009).

To complete the specification of model (17), a_s and b_s are fixed to a variety of values representing different levels of prior knowledge. For each we set $b_s = 4a_s c$ resulting in a prior mean for each σ_i^2 of $\frac{4ca_s}{a_s - 1}$, $a_s > 1$. The precision is $\frac{(a_s-1)^2(a_s-2)}{(4ca_s)^2}$; meaning larger a_s and smaller c result in a more informative prior. With $c = 1$ the shrinkage (for large a_s) is to the true value of $\sigma^2 = 4$. We consider $a_s = 1.25, 5, 10$ and $c = 0.5, 1, 2$ for a total of nine different priors whose densities are displayed in Figure 5. The vertical dashed line is at the known true value of $\sigma^2 = 4$.

$K = 30$ data sets are generated from (16). For each data set and each pair (a_s, c), the Bayesian models are fit using MCMC. The MCMC for the restricted likelihood version requires no computational details other than those described for the traditional Bayesian model in Section 4. This is because there are conditioning statistics for each group and the model's conditional independence between the groups allows the data augmentation described earlier to be performed independently within each group. That is, there is a separate Gibbs step for each group to generate the group level data matching the statistics for that group.

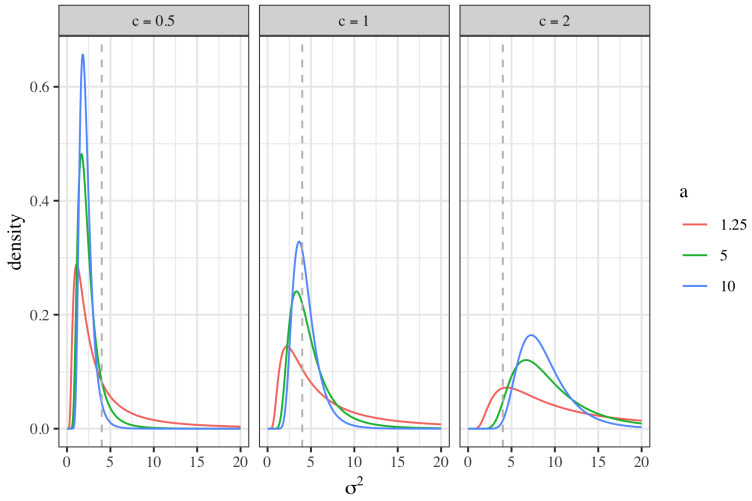

Figure 5: The nine different inverse gamma priors used for the group level variance parameters σ_j^2 in the simulation. The shape parameter values are $a_s = 1.25, 5, 10$. The scale parameter values are $b_s = 4a_s c$ with $c = 0.5, 1, 2$. The vertical dashed line in each panel is at the true value of $\sigma^2 = 4$ and is the variance of the good portion of the data in the simulation.

To assess predictive capability, the models are compared using Kullback-Leibler (KL) divergence from the distribution of good data to the posterior predictive distribution. Specifically, for the i^{th} group of the k^{th} simulated data set $\boldsymbol{y}_k$ compute:

$$KL_{ik}^{(M)} = \int \log \frac{f(\tilde{y}|\theta_i, \sigma^2)}{f_i(\tilde{y}|M, \boldsymbol{y}_k)} f(\tilde{y}|\theta_i, \sigma^2)\ dy \tag{18}$$

where M indexes the fitting model, $f(\tilde{y}|\theta_i, \sigma^2) = N(\tilde{y}|\theta_i, \sigma^2)$; the normal density function with mean θ_i and variance σ^2, evaluated at $\tilde{y}$. For the Bayesian models $f_i(\tilde{y}|M, \boldsymbol{y}_k) = \int f(\tilde{y}|\theta_i, \sigma_i^2)\pi(\theta_i, \sigma_i^2|M, \boldsymbol{y}_k)d\theta_i d\sigma_i^2$ where $\pi(\theta_i, \sigma_i^2|M, \boldsymbol{y}_k)$ is the posterior for the i^{th} group model parameters under model M for the k^{th} data set. M denotes either the full normal theory model (17) or one of the two restricted likelihood versions, along with specified a_s and c. For the classical robust fits, we set $f_i(\tilde{y}|M, \boldsymbol{y}_k) = N(\tilde{y}|\hat{\theta}_i, \hat{\sigma}_i^2)$ as a groupwise plug-in estimator for the predictive distribution. The classical fits are done separately for each group with no consideration of the hierarchical structure between the groups. The overall mean $\overline{KL}_{..}^{(M)} = \frac{1}{90K}\sum_{k=1}^{K}\sum_{i=1}^{90} KL_{ik}^{(M)}$ is used to compare the models, where smaller means correspond to better fits. Sampling variation is summarized with the standard error between the $K = 30$ replicates in the simulation: $SE(\overline{KL}_{\cdot k}^{(M)}) = \sqrt{\frac{1}{K(K-1)}\sum_{k=1}^{K}(\overline{KL}_{\cdot k}^{(M)} - \overline{KL}_{..}^{(M)})^2}$ where

$\overline{KL}_{\cdot k}^{(M)} = \frac{1}{90}\sum_{i=1}^{90} KL_{ik}^{(M)}$.

Figure 5 displays $\overline{KL}_{\cdot\cdot}^{(M)}$ with error-bars plus/minus one $SE(\overline{KL}_{\cdot k}^{(M)})$ for each $a_s = 1.25, 5, 10$ and $c = 0.5, 1, 2$. The values of a_s and c, do not effect the classical robust linear models. The average KL for the normal theory models range between 0.22 and 0.3 which is much worse than the robust methods and hence are left out of the figure. For $c = 0.5$ and $c = 1$, the results are in favor of the restricted likelihood methods with perhaps a slight advantage to the use of Tukey's location estimator over Huber's. This is likely due to the fact that Tukey's estimator trims extreme outliers completely in the estimation procedure (Huber and Ronchetti, 2009). For $c = 2$, the classical methods win, again with a slight advantage to Tukey's statistic. It is clear that the priors picture in Figure 5 matter. For $c = 2$, the priors subsequently add more mass significantly above $\sigma^2 = 4$ as a_s gets larger. Additionally the tuning parameters chosen for the location and scale estimators result in an upward bias in the estimate of σ^2. This bias depends on m and p. For example, for $m = 9$ and $p = .1$, Huber's version converges to roughly 4.8 asymptotically with n under the data generating model. This bias gets worse for more severe contamination levels in the simulation. For $c = 0.5$ and $c = 1$, the priors place more mass on smaller values as a_s grows. The larger values of a_s also result in better (smaller) average KL divergence. For $c = 1$ the prior concentration is roughly around the true $\sigma^2 = 4$. For $c = 0.5$ the prior concentration is roughly around half $\sigma^2 = 2$. Interestingly, this prior performs better in terms of KL divergence than having the seemingly more correct prior with $c = 1$. This prior does not capture good prior information on σ^2, but does compensates more for the upward bias in the estimates of scale. This highlights one potential effect of the choice of statistic and associated tuning parameters when applying the restricted likelihood.

It is also interesting to consider the effects of factors n, p, and m. We subset the results to a single prior ($a_s = 5$ and $c = 1$) to not confound the effects of the different priors. For each simulation k, the main effect averages of $KL_{ik}^{(M)}$ are found for each factor n, p, and m. Figure 7 displays the average of these main effects over the $K = 30$ simulations along with error bars plus/minus one standard error. For each group n, p, and m the restricted versions have better (lower) average KL divergences. The patterns also makes sense: average KL with get larger (worse) as the contamination gets more severe (larger m or larger p) and average KL gets smaller (better) as the sample size n grows.

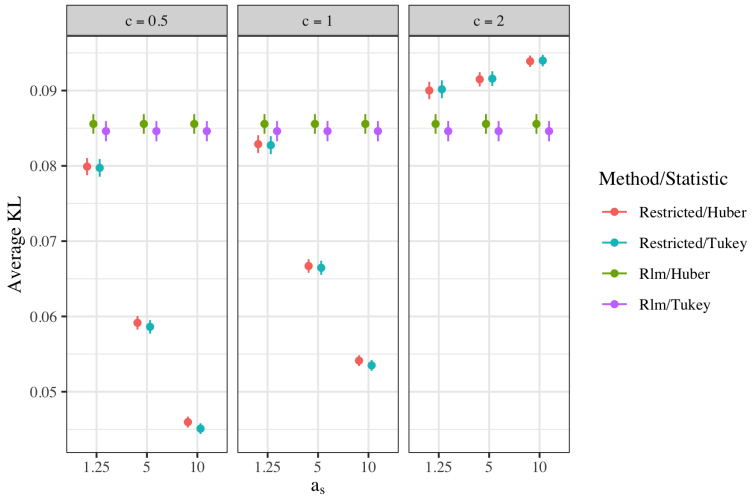

Figure 6: Average KL - divergence plus/minus one standard error for each value of a_s and c ($\overline{KL}_{..}^{(M)} \pm SE(\overline{KL}_{.k}^{(M)})$). Smaller values represent better fits. The panels correspond to $c = 0.5$ (left), $c = 1$ (middle), and $c = 2$ (right) with the values of a_s on the horizontal axis. The average KL for the normal theory model ranges between 0.31 and 0.37 and are left out of the figure.

This simulation shows the potential of the restricted likelihood while highlighting some interesting observations. Specifically, the choice of summary statistics, along with corresponding tuning parameters is important. For the tuning parameters, we applied the default choice of 95% efficiency at the normal. Under the simulation model here, this choice results in bias in the scale estimation which affects the performance of the method. These choices must be made when using both the classical and Bayesian methods. The Bayesian setting allows for the incorporation of informative prior information which, as shown in this example, can improve prediction if chosen wisely.

6 Real Data

We illustrate our methods with a pair of regression models for data from Nationwide Insurance Company, which concern prediction of the performance of insurance agencies. Nationwide sells many of its insurance policies through agencies which provide direct service to policy holders. The contractual agreements between Nationwide and these agencies vary. Our interest is the prediction of future per-

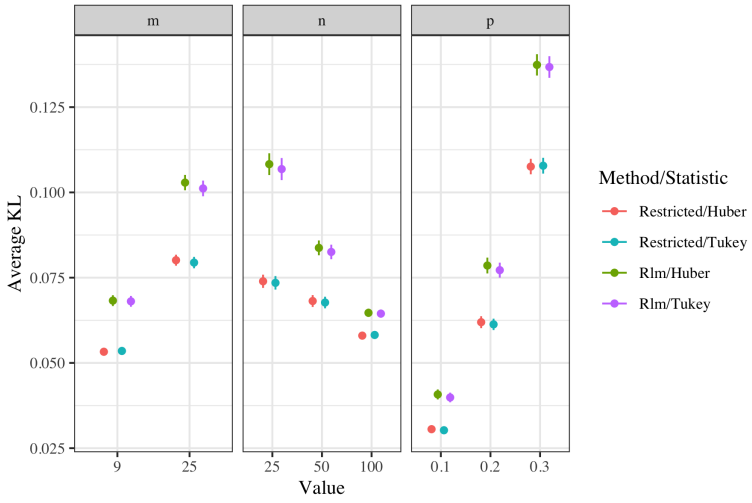

Figure 7: Average KL - divergence plus/minus one standard error grouped by the factors m (left), n (middle), and p (right). Results for a single prior ($a_s = 5$ and $c = 1$) are displayed to not confound the effects of the different priors.

formance of agencies where performance is measured by the total number of households an agency services ('household count'). The data are grouped by states with a varying number of agencies by state. Identifiers such as agency/agent names are removed. Likewise, state labels and agency types (identifying the varying contractual agreements) have been made generic to protect the proprietary nature of the data. As an exploratory view, a plot of the square root of household count in 2012, against that in 2010 is shown in Figure 8 for four states. The states have varying number of agencies and the different colors represent the varying types of contractual agreements as they stood in 2010 ('Type'). A significant number of agencies closed sometime before 2012, as represented by the 0 counts for 2012. Among the open agencies, linear correlations exists with strength depending on agency type and state. 'Type 1' agencies open in 2012 are of special interest. One could easily subset the analysis to only these agencies, removing the others. However, we leave them and use the data as a test bed for our techniques by fitting models that do not account for agency closures or contract type. Our expectation is that the restricted likelihood will facilitate prediction for the 'good' part of the data (i.e. open, 'type 1' agencies).

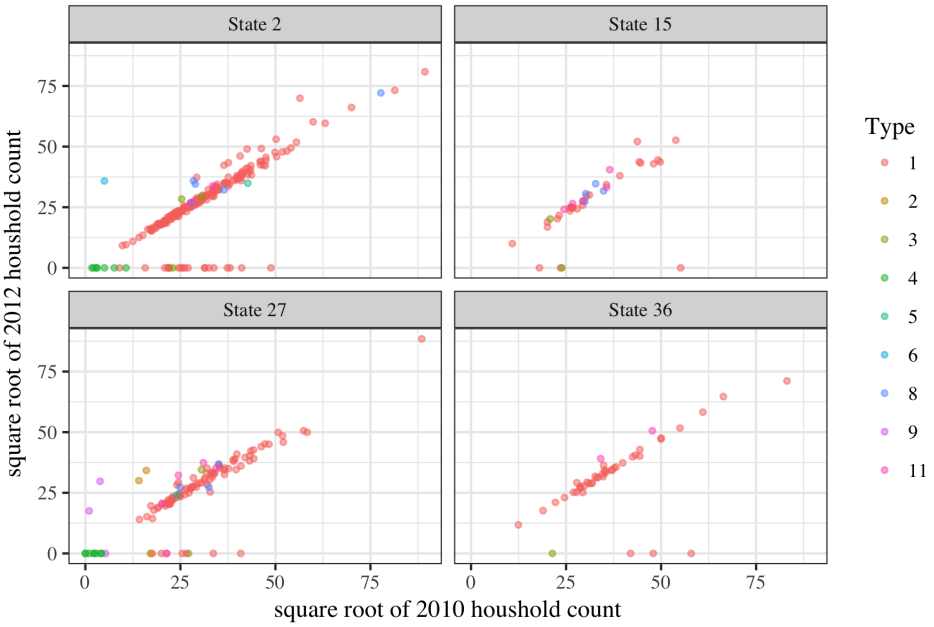

Figure 8: The square root of count in 2012 versus that in 2010 for four states. The colors represent the varying contractual agreements as they stood in 2010 ('Type'). Agencies that closed during the 2010-2012 period are represented by the zero counts for 2012.

6.1 State Level Regression model

The first analysis considered is based on individual regressions fit separately within states. The following normal theory regression model is used as the full model for a single state:

$$\beta \sim N(\mu_0, \sigma_0^2); \quad \sigma^2 \sim IG(a_0, b_0); \quad y_i = \beta x_i + \epsilon_i, \quad \epsilon_i \overset{iid}{\sim} N(0, \sigma^2), \ i = 1, \ldots, n, \tag{19}$$

where y_i and x_i are the square rooted household count in 2012 and 2010 for the i^{th} agency, respectively. The hyper-parameters a_0, b_0, μ_0 and σ_0^2 are all fixed and set from a robust regression fit to the corresponding state's data from the time period two years before. Specifically, Let $\hat{\beta}$ and $\hat{\sigma}^2$ be estimates from the robust linear regression of 2010 counts on 2008 counts. We fix $a_0 = 5$ and set $b_0 = \hat{\sigma}^2(a_0 - 1)$ so the prior mean is $\hat{\sigma}^2$. We set $\mu_0 = \hat{\beta}$ and $\sigma_0^2 = n_p se(\hat{\beta})^2$ where n_p is the number of agencies in the prior data set and $se(\hat{\beta})$ is the standard error of $\hat{\beta}$ derived from the robust regression. This prior is in the spirit of the Zellner's g-prior (Zellner, 1986; Liang et al., 2008); in general scaling

the prior variance $se(\hat{\beta})^2$ by a factor g. $g = n_p$ is analogous to the unit-information prior (Kass and Wasserman, 1995), with the difference that we are using a prior data set, not the current data set, to set the prior. The obvious reason this model is misspecified is due to omission of the contract type and agency closure information. Closing our eyes to these variables, many of the cases appear as outliers. Additionally, the model assumes equal variance within each state, an assumption whose worth is arguable (see Figure 8).

We compare four Bayesian models: the standard Bayesian normal theory model, two restricted likelihood models, both with simultaneous M-estimators, and a heavy-tailed model. For the restricted likelihood methods we use the same simultaneous M-estimators as in the simulation of Section 5 adapted to linear regression. The heavy-tailed model replaces the normal sampling density in (19) with a t-distribution with $\nu = 5$ degrees of freedom. The Bayesian models are all fit using MCMC, with the restricted versions using the algorithm presented in Section 4.2. We also fit the corresponding classical robust regressions and a least squares regression.

Method of model comparison

We wish to examine the performance of the models in a fashion that preserves the essential features of the problem. Since we are concerned with outliers and model misspecification, we understand that our models are imperfect and prefer to use an out-of-sample measure of fit. This leads us to cross-validation. We repeatedly split the data into training and holdout data sets; fitting the model to the training data and assessing performance on the holdout data.

The presence of numerous outliers in the data implies that both training and validation data will contain outliers. For this reason, the evaluation must be robust to a certain fraction of bad data. The two main strategies are to robustify the evaluation function (e.g., Ronchetti et al., 1997) or to retain the desired evaluation function and trim cases (Jung et al., 2014). Here, we pursue the trimming approach with log predictive density for the Bayesian models and log density from plug-in maximum likelihood for the classical fits used as the evaluation function.

The trimmed evaluation proceeds as follows in our context. The evaluation function for case i in

the holdout data is the log predictive density, say $\log(f(y_i))$, with the conditioning on the summary statistic suppressed. The trimming fraction is set at $0 \leq \alpha < 1$. To score a method, we first identify a base method. Denote the predictive density under this method by $f_b(y)$. Under the base method, $\log(f_b(y_i))$ is computed for each case in the holdout sample, say $i = 1, \ldots, M$. Order the holdout sample according to the ordering of $\log(f_b(y_i))$ and denote this ordering by $y^b_{(1)}, y^b_{(2)}, \ldots, y^b_{(M)}$. That is, for $i < j$ $\log(f_b(y^b_{(i)})) < \log(f_b(y^b_{(j)}))$. All of the methods are then scored on the holdout sample with the mean trimmed log marginal pseudo likelihood,

$$TLM_b(A) = (M - [\alpha M])^{-1} \sum_{i=[\alpha M]+1}^{M} \log(f_A(y^b_{(i)})),$$

where f_A corresponds to the predictive distribution under the method "A" being scored. In other words, the $[\alpha M]$ observations with the smallest values of $\log(f_b(y))$ are removed from the validation sample and all of the methods are scored using only the remaining $M - [\alpha M]$ observations. Larger values of $TLM_b(A)$ indicate better predictive performance. This process is advantageous to the base method since the smallest scores from this method are guaranteed to be trimmed. A method that performs poorly when it is the base method is discredited.

Comparison of predictive performance

'Type 1' agencies are of special interest to the company and so the evaluation of the TLM is done on only holdout samples of 'Type 1', wheres the training is done on agencies of all types. This is intended to demonstrate the robustness properties of the various methods. Models are fit to four states labelled State 2, 15, 27, and 36, with $n = 222, 40, 117$, and 46, representing a range of sample sizes. Fitting is done on $K = 50$ training/holdout samples with training sample sizes taken to be $.25n$ and $0.50n$. Holdout evaluation is done on the remaining ('Type 1') samples. For the data augmentation MCMC step under the restricted likelihood models, the acceptance rates range from 0.15 to 0.76 across the states, repetitions, and two versions of the model. The average $TLM_b(A)$ over the $K = 50$ training/holdout samples for the four states and seven methods are shown in Figure 9 where the base model is the Student-t model and $\alpha = 0.3$. Similar results are observed for other base models, with this one being most advantageous to the Student-t model. The error-bars

are plus/minus one standard deviation of the average $TLM_b(A)$ over the $K = 50$ training/holdout samples. It is clear the normal Bayesian model used as the full model (Normal) and the classical ordinary least squares fits (OLS) have poor performance due to the significant amount of outlier contamination in the data. In comparing our restricted methods to their corresponding classical methods, there is small, but consistent improvement across the states and training sample size. For state 2, the largest state with $n = 222$, the restricted and classical robust methods have similar performance especially for larger training sample size. This reflects the diminishing effect of the prior as the sample size grows. Notably, the Student-t model performs poorly in comparison for this state. The predictive distribution explicitly accounts for heavy-tailed values, resulting in poorer predictions of the 'good' data (i.e., the Type 1 agencies). Likewise, for State 27, another larger state, the Student-t model is outperformed by our restricted methods. For the other states, the Student-t performs similarly to our restricted methods for the smaller states (State 15 and 36) and smaller training sample size (25% of the sample). However, the performance is worse for the larger training sample size (50% of the sample). Intuitively, as more data is available for fitting, more outliers appear and the heavy-tailed model compensates for them by assuming they come from the tails of the model which is detrimental for prediction. The relative performance of the models are similar for various values of α as seen in Figure 10 which shows results for training sample size $0.5n$.

6.2 Hierarchical regression model

The previous analysis treated states independently. A natural extension is to reflect similar business environments between states using a hierarchical regression. The proposed model is:

$$\begin{aligned} &\beta \sim N_p(\mu_0, a\sigma_0^2); \quad \beta_j \overset{iid}{\sim} N_p(\beta, b\sigma_0^2); \quad \sigma_j^2 \sim IG(a_0, b_0); \\ &y_{ij} = x_{ij}\beta_j + \epsilon_{ij}, \quad \epsilon_{ij} \overset{iid}{\sim} N(0, \sigma_j^2), \ i = 1, \ldots, n_j, \ j = 1, \ldots, J \end{aligned} \tag{20}$$

where y_{ij} is the i^{th} observation of square rooted household count in 2012 in the j^{th} state, n_j is the total number of agencies in state j, and J is the number of states. x_{ij} is the square rooted household count in 2010 and β_j represents the individual regression coefficient vector for state j. The parameters μ_0, σ_0^2, a_0, and b_0 are fixed by fitting the regression $y_{ij} = x_{ij}\beta + \epsilon_{ij}$ using Huber's

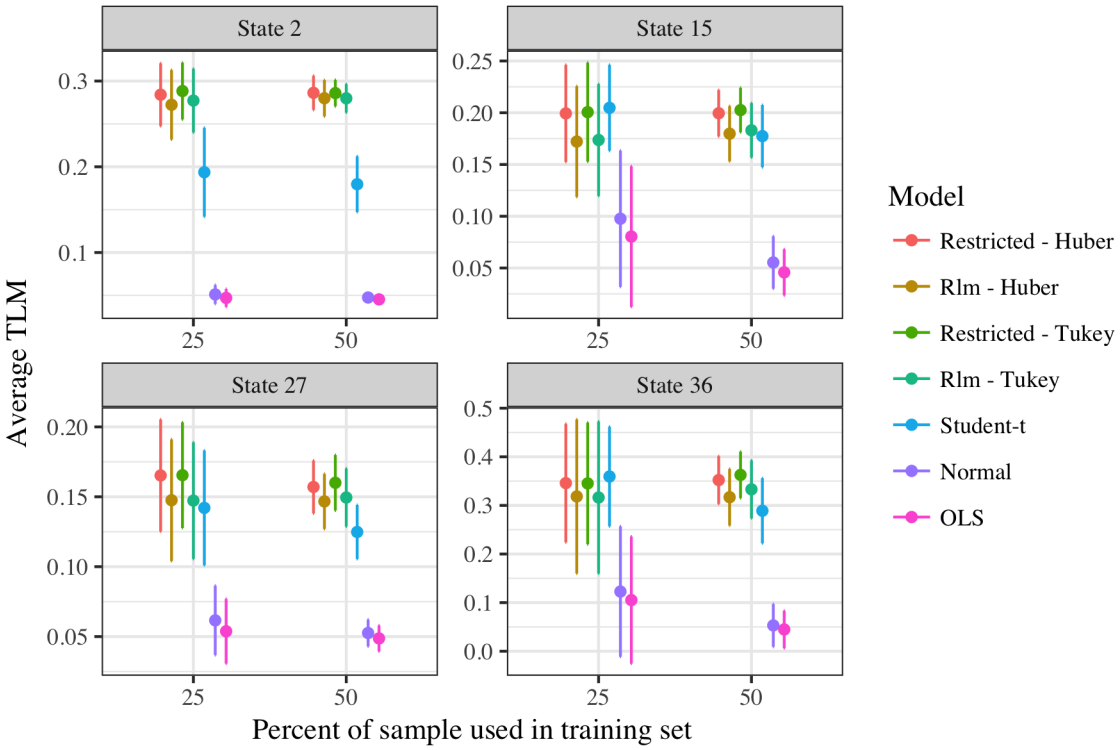

Figure 9: Average TLM plus/minus one standard deviation over $K = 50$ splits into training and holdout samples. The panels are for the different states 2, 15, 27, and 36, with $n = 222, 40, 117,$ and 46, respectively. The horizontal axis is the percent of n used in each training set. The color corresponds to the fitting model. Larger values of TLM are better.

M-estimators to the prior data set from two years before. Using the estimates from this model, we set $\mu_0 = \hat{\beta}$, $\sigma_0^2 = n_p se(\hat{\beta})^2$ ($n_p = 2996$ is the number of observations in the prior data set), $a_0 = 5$ and $b_0 = \hat{\sigma}^2(a_0 - 1)$. We constrain $a + b = 1$ in an attempt to partition the total variance between the individual β_j's and the overall β. We take $b \sim \text{beta}(v_1, v_2)$. Using the prior data set, we assess the variation between individual estimates of the β_j to set v_1 and v_2 to allow for a reasonable amount of shrinkage. To allow for dependence across the σ_j^2 we first take $(z_1, \ldots, z_J) \sim N_J(\mathbf{0}, \Sigma_\rho)$ with $\Sigma_\rho = (1 - \rho)\mathbf{I} + \rho\mathbf{1}\mathbf{1}^\top$. Then we set $\sigma_j^2 = H^{-1}(\Phi(z_j))$ where H is the cdf of an $IG(a_0, b_0)$ and Φ is the cdf of a standard normal. This results in the specified marginal distribution, while introducing correlation via ρ. We assume $\rho \sim \text{beta}(a_\rho, b_\rho)$ with mean $\mu_\rho = a_\rho/(a_\rho + b_\rho)$ and precision $\psi_\rho = a_\rho + b_\rho$.

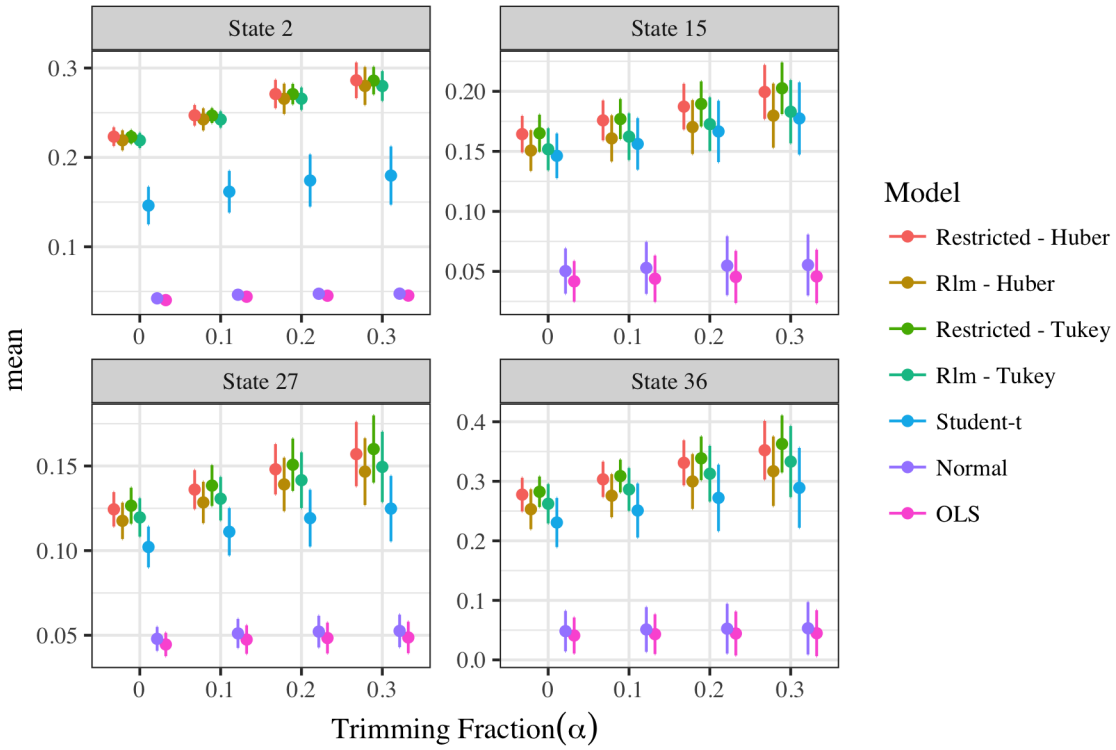

Figure 10: Average TLM plus/minus one standard deviation over $K = 50$ splits into training and holdout samples for several values of the trimming fraction α. The training sample size used is $0.5n$. Larger values of TLM are better.

The parameters μ_ρ and ψ_ρ are given beta and gamma distributions, with fixed hyperparameters. More details on setting prior parameters are given in the appendix.

Using the same techniques as in the previous section, we fit the normal theory hierarchical model above, a thick-tailed t version with $\nu = 5$ d.f., and two restricted likelihood versions (Huber's and Tukey's) of the model. For the restricted methods, we condition on robust regression estimates fit separately within each state. We also fit classical robust regression counterparts and a least squares regression separately within each state. Hierarchical models naturally require more data and so we include states having at least 25 agencies resulting in 22 states in total and $n = \sum_j n_j = 3180$ total agencies. For training data we take a stratified (by state) sample of

size $3180/2 = 1590$ where the strata sizes are $n_j/2$ (rounded to the nearest integer). The remaining data is used for a holdout evaluation using TLM computed separately within each state: $TLM_b(A)_j = (M_j - [\alpha M_j])^{-1} \sum_{i=[\alpha M_j]+1}^{M_j} \log(f_A(y^b_{(i)j}))$ where $y^b_{(1)j}, y^b_{(2)j}, ..., y^b_{(M_j)j}$ is the ordering of the M_j holdout observations within state j according to the log marginals under the base model b. For the non-Bayesian models, $f_A(y_{(i)j})^b$ is estimated using plug-in estimators for the parameters for state j. $TLM_b(A)_j$ is computed for each state for $K = 50$ splits of training and holdout sets. The Bayesian models are fit using MCMC, with the restricted versions applying the algorithm laid out in Section 4 and adapted to the hierarchical setting as described in Section 5. For the MH-step proposing augmented data, the acceptance rates for the two restricted likelihood models across all states and repetitions ranges from 0.24 to 0.74.

The average over states, $\overline{TLM}_b(A)_\cdot = \frac{1}{22}\sum_{j=1}^{22} TLM_b(A)_j$ for each of the K repetitions is summarized in Figure 11 for several trimming fractions using the Student-t as the base model. The points are the average of the $\overline{TLM}_b(A)_\cdot$ over the K repetitions with errorbars plus/minus one standard deviation over K. The normal theory fits, both Bayesian and classical, perform poorly and are left out of the figures. As the trimming fraction used for the TLM increases, so does TLM since more outliers are being trimmed. Similar patterns were seen in the individual state level regressions in Section 6.1. Despite being used as the base model to compute TLM, the Student-t doesn't perform well in comparison to the robust regressions. We attribute this to the assumption of heavier tails resulting in smaller log marginal values on average; emphasizing again that the t-model will do well to discount outlying observations but doesn't provide a natural mechanism for only predicting 'good' (i.e. non-outlying) data. On average, the classical robust regressions fit separately within each state slightly outperform the restricted likelihood hierarchical models. The hierarchical model does however reduce variance in predictions at the cost of an increase in bias induced by the prior distributions. This can be seen in Figure 12 which shows the standard deviation of $\overline{TLM}_b(A)_\cdot$ relative to the mean over the $K = 50$ repetitions. It is clear that the hierarchical model reduces sample-to-sample variation in comparison to the classical robust model counterparts.

It is also interesting to examine the results within each state. Figure 13 summarizes $TLM_b(A)_j$ for each state where the points and errorbars are the averages and plus/minus one standard deviation

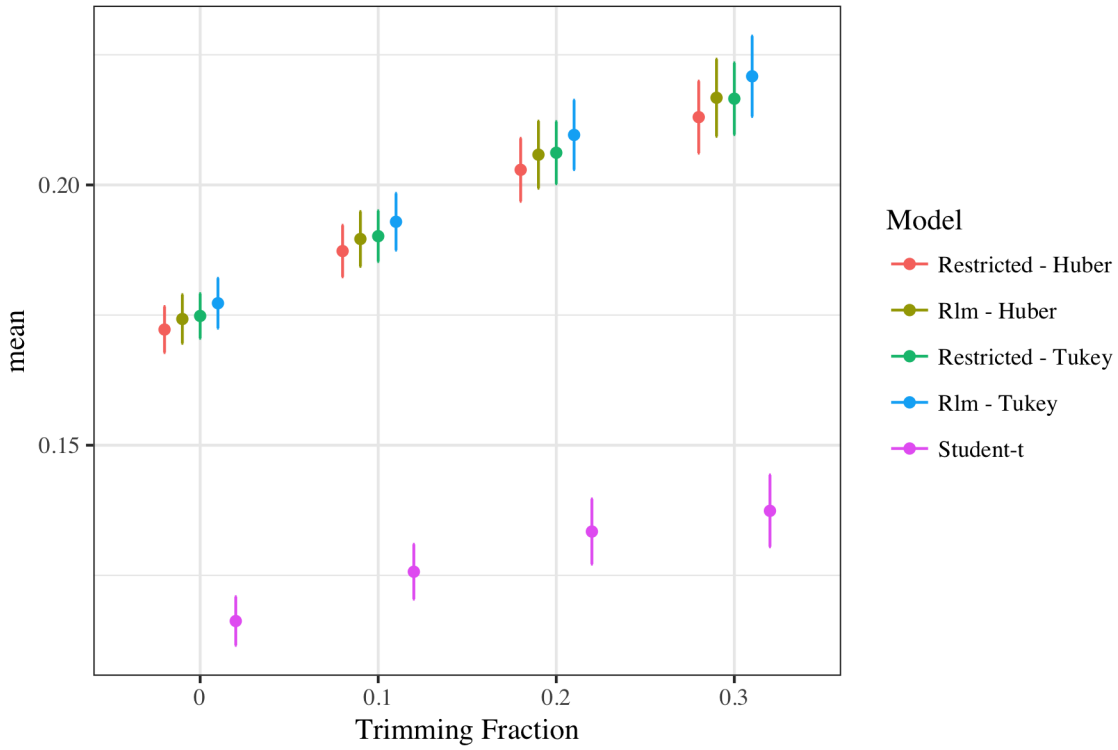

Figure 11: Hierarchical model results: $\overline{TLM}_b(A)$. plus/minus one standard deviation over $K = 50$ splits into training and holdout sets with the Student-t as the base model and several values of the trimming fraction α. Larger values of TLM are better.

of $TLM_b(A)_j$ over the $K = 50$ repetitions. The results are only given for the models using Tukey's M-estimators since, in general, these models performed the best on average. This is attributed to the corresponding ψ-function used in the M-estimation which trims large outliers from the data completely. Huber's version down-weights outliers, but never completely trims them. The states are ordered along the x-axis according to number of agencies within the state (shown in parentheses). In several of the smaller states, the restricted hierarchical model performs better on average with similar performance between the models in most of the larger states; a reflection of the decreased influence of the prior. It is clear that the prior is having the effect of pooling information across states; beneficial for reducing the sample-to-sample variance. Generally, the standard deviations are

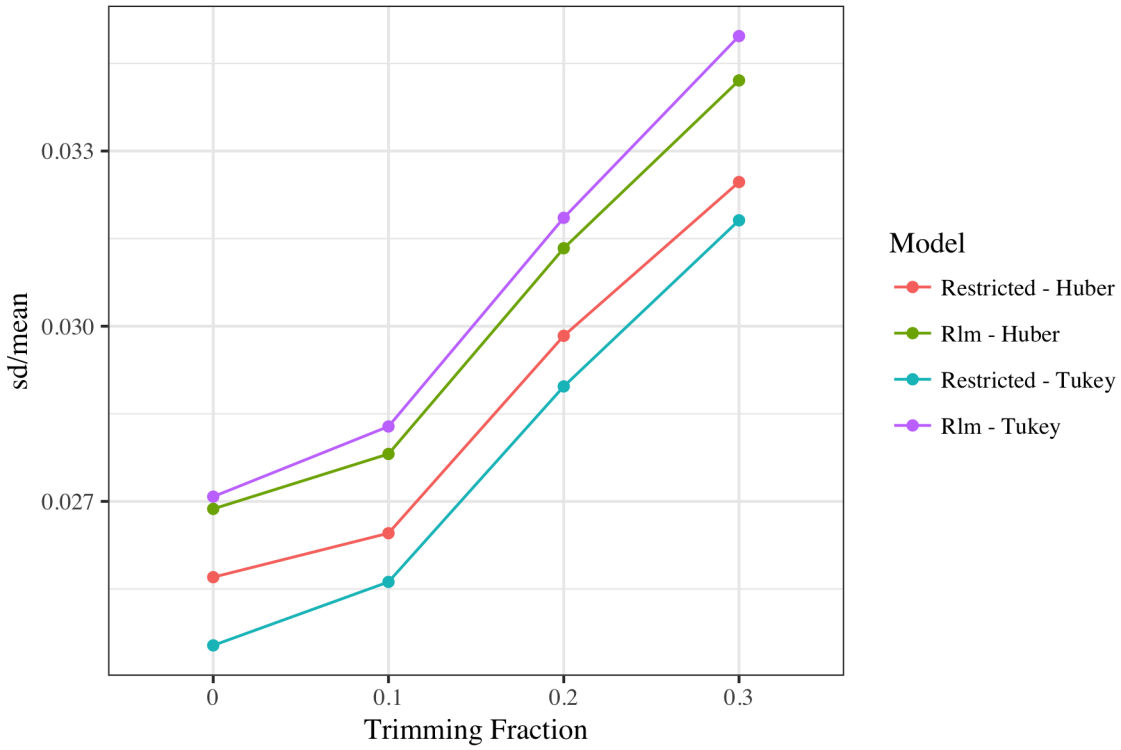

Figure 12: Hierarchical model results: standard deviation of $\overline{TLM}_b(A)$. relative to the mean over $K = 50$ splits into training and holdout sets for several values of the trimming fraction α. Results are given for the two restricted likelihood versions of the hierarchical model and their corresponding robust regression models fit separately within each state.

smaller for the hierarchical model for smaller states with virtually identical standard deviations for the larger states. Similar benefits are often seen for hierarchical models (e.g., Gelman, 2006).

7 Discussion

Many routine choices in an analysis react to the gap between reality and the statistical model, where a bit of set-up work improves inferential performance. Often, these choices can be recast in the framework of restricted likelihood presented here, lending them more formality and facilitating

development of theoretical results. A greater benefit of our framework is that it leads us to blend classical estimation with Bayesian methods. Here, we concentrate on outlier-prone settings where natural choices for conditioning statistic are classical robust estimators targeting the mean of the non-outlying data (e.g., M-estimators). The likelihood conditioned on these estimators is used to move from prior to posterior. The update follows Bayes' Theorem, conditioning on the observed estimators exactly. Computation is driven by MCMC methods, requiring only a supplement to existing algorithms by adding a Gibbs step to sample from the space of data sets satisfying the observed statistic. This step has additional computation costs. The estimator must be computed at every iteration to propose a new $\boldsymbol{y}$. This cost, as well as the need to find orthonormal bases derived from gradients of the estimators, adds some computational time compared to the full likelihood MCMC. The cost of finding the bases can be driven down by an understanding of the geometric space from which the samples are drawn as described in Section 4.2. Additionally, we have seen good mixing of the MCMC chains across a wide-variety of examples.

Outside of outlier-prone settings, we might condition on the results of a set of estimating equations, designed to enforce lexical preferences for those features of the analysis considered most important, yet still producing inferences for secondary aspects of the problem. For example, the computational strategies we devised here allow us to apply the method to inference on quantiles of a regression model by, for example, conditioning on estimates of these quantiles. This leads to questions regarding the choice of summary statistic to apply. For outliers, we have applied some obvious choices (M-estimators, LMS, LTS) which have shown value in classical robust estimation. Properties of such estimators are well known and can be used as a basis for selecting one. In general, great ingenuity has been used to create a wide variety of estimators designed to handle specific manifestations of a misspecified model; where outliers are just a specific case. The estimators are typically accompanied by asymptotic results on consistency and distribution. These results can be used as starting point for choosing appropriate statistics in specific settings. In general, we recommend a choice based on the analyst's understanding of the problem, model, reality, deficiencies in the model, inferences to be made, and the relative importance of various inferences.

The most general setting, of which outliers are a subset, is model misspecification. The traditional

view is that, if the model is inadequate, one should build a better model. In our empirical work, as data sets have become larger and more complex, we have bumped into settings where we cannot realistically build the perfect model. We ask the question "by attempting to improve our model through elaboration, will the overall performance of the model suffer?" If yes, we avoid the elaboration, retaining a model with some level of misspecification. Acknowledging that the model is misspecified implies acknowledging that the sampling density is incorrect, exactly as we do when outliers are present. In this sense, misspecified models and outliers are reflections of the same phenomenon, and we see the restricted likelihood as a method for dealing with this more general problem.

The framework allows us to retain many benefits of Bayesian methods: it requires a complete model for the data; it lets us combine various sources of information both through the use of a prior distribution and through creation of a hierarchical model; it guarantees admissibility of our decision rules among the class based on the summary statistic $T(\boldsymbol{y})$; and it naturally leads us to focus on predictive inference. The work does open a number of questions open to further work, including a need to revisit such issues as model selection, model averaging for predictive performance, and the role of diagnostics.

8 Appendix

8.1 Proofs

Proof of Theorem 4.1.

Proof.

$$s(X, \boldsymbol{y}) = s\left(X, \frac{s(X, \boldsymbol{y}_{obs})}{s(X, \boldsymbol{z}^*)}\boldsymbol{z}^* + X\left(\boldsymbol{b}(X, \boldsymbol{y}_{obs}) - \boldsymbol{b}(X, \frac{s(X, \boldsymbol{y}_{obs})}{s(X, \boldsymbol{z}^*)}\boldsymbol{z}^*)\right)\right) \quad (21)$$

$$= \frac{s(X, \boldsymbol{y}_{obs})}{s(X, \boldsymbol{z}^*)} s(X, \boldsymbol{z}^*) = s(X, \boldsymbol{y}_{obs}), \qquad \text{and} \quad (22)$$

$$\boldsymbol{b}(X, \boldsymbol{y}) = \boldsymbol{b}\left(X, \frac{s(X, \boldsymbol{y}_{obs})}{s(X, \boldsymbol{z}^*)}\boldsymbol{z}^* + X\left(\boldsymbol{b}(X, \boldsymbol{y}_{obs}) - \boldsymbol{b}(X, \frac{s(X, \boldsymbol{y}_{obs})}{s(X, \boldsymbol{z}^*)}\boldsymbol{z}^*)\right)\right) \quad (23)$$

$$= \boldsymbol{b}(X, \frac{s(X, \boldsymbol{y}_{obs})}{s(X, \boldsymbol{z}^*)}\boldsymbol{z}^*) + \boldsymbol{b}(X, \boldsymbol{y}_{obs}) - \boldsymbol{b}(X, \frac{s(X, \boldsymbol{y}_{obs})}{s(X, \boldsymbol{z}^*)}\boldsymbol{z}^*) \quad (24)$$

$$= \boldsymbol{b}(X, \boldsymbol{y}_{obs}) \quad (25)$$

□

Theorem 8.1. *The mapping $h : \mathbb{S} \to \mathcal{A}$ with h defined in Theorem 4.1 is one-to-one and onto.*

Proof. one-to-one: Let $z_1, z_2 \in \mathbb{S}$ with $h(z_1) = h(z_2)$. Rearrangement implies $z_1 = cz_2 + Xv$ for known $c \in \mathbb{R}$ and $v \in \mathbb{R}^p$ depending on $\boldsymbol{b}(X, \boldsymbol{y}_{obs})$, $s(X, \boldsymbol{y}_{obs})$, $\boldsymbol{b}(X, z_1)$, $s(X, z_1)$, $\boldsymbol{b}(X, z_2)$, $s(X, z_2)$. Given $z_2 \in \mathbb{S}$, $v \neq 0$ implies $z \notin \mathcal{C}^{\perp}(X)$ and $c \neq 1$ implies $||z_1|| \neq 1$. Thus $z_1 \in \mathbb{S}$ implies $c = 1$ and $v = 0$.

Onto: Let $\boldsymbol{y} \in \mathcal{A}$ and consider its projection onto $\mathcal{C}^{\perp}(X)$: $Q\boldsymbol{y}$ where $Q = I - XX^{\top}$. It is easy to show that $\boldsymbol{z}^* = Q\boldsymbol{y}/||Q\boldsymbol{y}|| \in \mathbb{S}$ and $h(\boldsymbol{z}^*) = \boldsymbol{y}$. □

Proof of Lemma 4.2.

Proof. We first show that $\nabla s(X, \boldsymbol{y}) \in \mathcal{C}^{\perp}(X)$. Recall that $H = I - Q$. By the regression invariance property C7, we have

$$s(X, \boldsymbol{y}) = s(X, Q\boldsymbol{y} + H\boldsymbol{y}) = s(X, Q\boldsymbol{y}). \quad (26)$$

Thus, by the chain rule $\nabla s(X, \boldsymbol{y}) = Q\nabla s(X, Q\boldsymbol{y}) = Q\nabla s(X, \boldsymbol{z})$. Hence $X^{\top}\nabla s(X, \boldsymbol{y}) = 0$ as desired. From equation (26), all vectors $\boldsymbol{z}' \in \Pi(\mathcal{A})$ satisfy $s(X, \boldsymbol{z}') = s(X, \boldsymbol{y}) = s(X, \boldsymbol{y}_{obs})$, and so all directional derivatives of s along each tangent $\boldsymbol{v}$ to $\Pi(\mathcal{A})$ in $\mathcal{C}^{\perp}(X)$ at $\boldsymbol{z}$ are equal to 0 (i.e., $\nabla s(X, \boldsymbol{z}) \cdot \boldsymbol{v} = 0$). Thus $\nabla s(X, \boldsymbol{z})$ is orthogonal to $\Pi(\mathcal{A})$ at $\boldsymbol{z}$. Since $\Pi(\mathcal{A})$ has dimension $n - p - 1$, $\nabla s(X, \boldsymbol{z})$ gives the unique (up to scaling and reversing direction) normal in the $n - p$ dimensional $\mathcal{C}^{\perp}(X)$. □

Proof of Lemma 4.3

Proof. Without loss of generality, assume the columns of X form an orthonormal basis for $\mathcal{C}(X)$ and likewise the columns of W form and orthonormal basis for $\mathcal{C}^{\perp}(X)$. With earlier notation, $H = XX^{\top}$ and $Q = WW^{\top}$. The set $\mathcal{A}$ is defined by the $p + 1$ equations $s(X, \boldsymbol{y}) = s(X, \boldsymbol{y}_{obs})$,

$b_1(X, \boldsymbol{y}) = b_1(X, \boldsymbol{y}_{obs}), \ldots, b_p(X, \boldsymbol{y}) = b_p(X, \boldsymbol{y}_{obs})$. Consequently, the gradients are orthogonal to $\mathcal{A}$. Let $\nabla \boldsymbol{b}(X, \boldsymbol{y})$ denote the $n \times p$ matrix with columns $\nabla b_1(X, \boldsymbol{y}), \ldots, \nabla b_p(X, \boldsymbol{y})$. We seek to show the $n \times (p+1)$ matrix $[\nabla \boldsymbol{b}(X, \boldsymbol{y}), \nabla s(X, \boldsymbol{y})]$ has rank $p+1$. Using property C5, we have that

$$\boldsymbol{b}(X, \boldsymbol{y}) = \boldsymbol{b}(X, Q\boldsymbol{y} + H\boldsymbol{y}) = \boldsymbol{b}(X, Q\boldsymbol{y}) + X^\top \boldsymbol{y}$$

Then $\nabla \boldsymbol{b}(X, \boldsymbol{y}) = Q \nabla \boldsymbol{b}(X, Q\boldsymbol{y}) + X$ and

$$[XX^\top, WW^\top]^\top [\nabla \boldsymbol{b}(X, \boldsymbol{y}), \nabla s(X, \boldsymbol{y})] = \begin{pmatrix} X & \mathbf{0} \\ WW^\top \nabla b(X, \boldsymbol{y}) & \nabla s(X, \boldsymbol{y}) \end{pmatrix} \tag{27}$$

The last column comes from Lemma 4.2. The matrix $[XX^\top, WW^\top]^\top$ is of full column rank (rank n), and so the rank of $[\nabla \boldsymbol{b}(X, \boldsymbol{y}), \nabla s(X, \boldsymbol{y})]$ is the same as the rank of the matrix on the right hand side of (27). This last matrix has rank $p+1$ since $\nabla s(X, \boldsymbol{y}) \neq \mathbf{0}$ by C8, and so does $[\nabla b(X, \boldsymbol{y}), \nabla s(X, \boldsymbol{y})]$. □

Proof of Lemma 4.4

Proof. P is the projection of the columns of A onto $\mathcal{C}^\perp(X)$. For this to result in a loss of rank, a subspace of $\mathcal{T}_y(\mathcal{A})$ must belong to $\mathcal{C}(X)$. Following property C5, for an arbitrary vector $X\boldsymbol{v} \in \mathcal{C}(X)$, $\boldsymbol{b}(X, \boldsymbol{y} + X\boldsymbol{v}) = \boldsymbol{b}(X, \boldsymbol{y}) + \boldsymbol{v}$. From the property, we can show that the directional derivative of $\boldsymbol{b}$ along $X\boldsymbol{v}$ with $\boldsymbol{v} \neq \mathbf{0}$ is $\boldsymbol{v}$, which is a nonzero vector. Hence $X\boldsymbol{v} \notin \mathcal{T}_y(\mathcal{A})$. □

Proof of Corollary 4.7

Proof. The corollary relies on a lemma and theorem from Miao and Ben-Israel (1992) which we restate slightly for brevity of presentation. The principal angles between subspaces pluck off a set of angles between subspaces, from smallest to largest. The number of such angles is the minimum of the dimensions of the two subspaces. Miao and Ben-Israel's first result (their Lemma 1) connects these principal angles to a set of singular values, and hence to volumes.

Lemma 8.2. *(Miao, Ben-Israel) Let the columns of $Q_L \in \mathbb{R}^{n \times l}$ and $Q_M \in \mathbb{R}^{n \times m}$ form orthonormal bases for linear subspaces L and M respectively, with $l \leq m$. Let $\sigma_1 \geq \cdots \geq \sigma_l \geq 0$ be the singular*

values of $Q_M^\top Q_L$. *Then* $\cos\theta_i = \sigma_i, i = 1, \ldots, l$ *where* $0 \leq \theta_1 \leq \theta_2 \leq \cdots \leq \theta_l \leq \frac{\pi}{2}$ *are the principal angles between* L *and* M.

Miao and Ben-Israel's second result (their Theorem 3) makes a match between the principal angles between a pair of subspaces and the principal angles between their orthogonal complements.

Theorem 8.3. *(Miao, Ben-Israel) The nonzero principal angles between subspace* L *and* M *are equal to the nonzero principal angles between* $L^\perp$ *and* $M^\perp$.

To establish the corollary, we appeal to Lemma 8.2 and Theorem 8.3. Translating Miao and Ben Israel's notation, we have $M = \mathcal{C}^\perp(X)$, $Q_M = W$, $L = \mathcal{T}_{\boldsymbol{y}}(\mathcal{A})$, and $Q_L = A$. By Theorem 8.3, the nonzero principal angles between $\mathcal{T}_{\boldsymbol{y}}(\mathcal{A})$ and $\mathcal{C}^\perp(X)$ are the same as the nonzero principal angles between $\mathcal{T}_{\boldsymbol{y}}^\perp(\mathcal{A})$ and $\mathcal{C}(X)$. By 8.2, the non-unit singular values of $W^\top A$ are the same as the non-unit singular values of $U^\top B$. □

8.2 Setting the hierarchical prior values

This section describes the how the prior parameters are set in Section 6.2. Using the previous data set from two years prior, we fit separate (robust) regressions to each state and a regression to the entirety of the data at once. Let the estimates for the fits to each state be $\hat{\beta}_1, \ldots, \hat{\beta}_J, \hat{\sigma}_1, \ldots, \hat{\sigma}_J$ and the estimates from the single regression be $\hat{\beta}$ and $\hat{\sigma}$. These are classical robust estimates using Tukey's regression and Huber's scale. For this sections, let n_j denote the number of observations in the j^{th} state (of the previous data set) and set $n_p = \sum n_j$.

First, consider v_1 and v_2 in the prior $b \sim \text{beta}(v_1, v_2)$. In the hierarchical model (20), $b = 0$ implies all the $\beta_j's$ are equal (no variation between states) and $b = 1$ implies the $\beta_j's$ vary about μ_0 according to $\Sigma_0 = n_p \cdot \text{se}(\hat{\beta})^2$ (see Section 6.1). We seek a prior measure for what we think b should be. Using the prior fit, a measure for uncertainty for β is $\Sigma_{\hat{\beta}} = \text{se}(\hat{\beta})^2$, the estimate of the variance from the single regression. For the $\beta_j's$, take $\delta_j = \hat{\beta}_j - \hat{\beta}$ and set the prior uncertainty to $\Sigma_\delta = n_p^{-1} \sum_j n_j \delta_j^2$. Consider $g = \Sigma_\delta / \Sigma_{\hat{\beta}}$ measuring of the amount of uncertainty between the $\beta_j's$ relative to that of β. Now in the prior, we heuristically set the uncertainty in the $\beta_j's$ ($b\Sigma_0$) to be

approximately equal to $g \cdot \Sigma_{\hat{\beta}}$. That is, $b\Sigma_0 \approx g \cdot \Sigma_{\hat{\beta}} = \frac{g}{n}\Sigma_0$, suggesting $b \approx \frac{g}{n}$. Thus, we set $E[b] = \frac{g}{n}$. The precision, $v_1 + v_2$, is set to 10, completing the specification for the prior on b.

Finally, recall $\rho \sim \text{beta}(a_\rho, b_\rho)$ with mean $\mu_\rho = a_\rho/(a_\rho + b_\rho)$ given a beta prior and precision $\psi_\rho = a_\rho + b_\rho$ given a gamma prior. There is little evidence of any strong correlation amongst estimates of σ_j^2 in the prior data set and we set the prior mean of μ_ρ equal to 0.2 and prior variance to .01. Noting $\text{var}(\rho|\mu_\rho, \psi_\rho) = \mu_\rho(1 - \mu_p)/(\psi_\rho + 1)$ we plug in $\mu_\rho = 0.2$ and $\text{var}(\rho|\mu_\rho, \psi_\rho) = 0.01$. Solving for ψ_ρ results in a value of 15. This is taken to be the mean of the gamma prior on ψ_ρ. Finally, we set the rate parameter for to 1 implying the variance of the gamma prior is equal to its the mean. With this specification, the prior on ρ has 80% of the central mass between roughly 0.03 and 0.42 and reflects our prior believe that there is likely only weak correlation amongst the σ_j^2's.

References

Berger, J. (2006). "The Case for Objective Bayesian Analysis." Bayesian Analysis, 1: 385–402. 2

Bernardo, J. M. and Smith, A. (2000). Bayesian Theory. John Wiley & Sons Ltd. 2

Clarke, B. and Ghosh, J. K. (1995). "Posterior Convergence Given the Mean." The Annals of Statistics, 23: 2116–2144. 5

Clarke, J. L., Clarke, B., Yu, C.-W., et al. (2013). "Prediction in M-complete Problems with Limited Sample Size." Bayesian Analysis, 8(3): 647–690. 2

Clyde, M. and George, E. I. (2004). "Model uncertainty." Statistical science, 81–94. 2

Clyde, M. A. and Iversen, E. S. (2013). "Bayesian model averaging in the M-open framework." Bayesian theory and applications. 2

Doksum, K. A. and Lo, A. Y. (1990). "Consistent and Robust Bayes Procedures for Location Based on Partial Information." The Annals of Statistics, 18: 443–453. 5

Fearnhead, P. and Prangle, D. (2012). "Constructing Summary Statistics for Approximate Bayesian Computation: Semi-Automatic Approximate Bayesian Computation." Journal of the Royal

Statistical Society: Series B, 74: 419–474. 6

Garthwaite, P. H., Kadane, J. B., and O'Hagan, A. (2005). "Statistical Methods for Eliciting Probability Distributions." Journal of the American Statistical Association, 100: 680–701. 2

Gelfand, A. E. and Smith, A. F. M. (1990). "Sampling-based approaches to calculating marginal densities." Journal of the American Statistical Association, 85: 398–409. 12

Gelman, A. (2006). "Multilevel (hierarchical) modeling: what it can and cannot do." Technometrics, 48(3): 432–435. 33

Hastings, W. K. (1970). "Monte Carlo sampling methods using Markov chains and their applications." Biometrika, 57: 97–109. 12

Hoff, P., Fosdick, B., Volfovsky, A., and Stovel, K. (2013). "Likelihoods for Fixed Rank Nomination Networks." Network Science, 1: 253–277. 5

Huber, P. and Ronchetti, E. (2009). Robust Statistics. Wiley Series in Probability and Statistics. Hoboken, New Jersey: John Wiley & Sons, Inc, 2nd edition. 7, 12, 20, 22

Huber, P. J. (1964). "Robust Estimation of a Location Parameter." The Annals of Mathematical Statistics, 35(1): 73–101. 10

Hwang, H., So, B., and Kim, Y. (2005). "On Limiting Posterior Distributions." Test, 14: 567–580. 5

Joyce, P. and Marjoram, P. (2008). "Approximately sufficient statistics and Bayesian computation." Statistical Applications in Genetics and Molecular Biology, 7(1). 6

Jung, Y., MacEachern, S., and Lee, Y. (2014). "Cross-validation via Outlier Trimming." In preparation. 26

Kass, R. E. and Raftery, A. E. (1995). "Bayes Factors." Journal of the American Statistical Association, 90: 773–795. 2

Kass, R. E. and Wasserman, L. (1995). "A reference Bayesian test for nested hypotheses and its relationship to the Schwarz criterion." Journal of the american statistical association, 90(431):

928–934. 9, 26

Lee, J. and MacEachern, S. N. (2014). "Inference functions in high dimensional Bayesian inference." Statistics and Its Interface, 7(4): 477–486. 2, 7

Lewis, J. (2014). "Bayesian Restricted Likelihood Methods." Ph.D. thesis, The Ohio State University. 4, 11, 12

Lewis, J., Lee, Y., and MacEachern, S. (2012). "Robust Inference via the Blended Paradigm." In JSM Proceedings, Section on Bayesian Statistical Science, 1773–1786. American Statistical Association. 5

Liang, F., Paulo, R., Molina, G., Clyde, M. A., and Berger, J. O. (2008). "Mixtures of g Priors for Bayesian Variable Selection." Journal of the American Statistical Association, 103: 410–423. 12, 25

Liu, J. S. (1994). "The Collapsed Gibbs Sampler in Bayesian Computations with Applications to a Gene Regulation Problem." Journal of the American Statistical Association, 89: 958–966. 12

Marjoram, P., Molitor, J., Plagnol, V., and Tavaré, S. (2003). "Markov Chain Monte Carlo without Likelihoods." Proceedings of the National Academy of Sciences of the United States of America, 100: 15324–15328. 6

Maronna, R., Martin, D., and Yohai, V. (2006). Robust Statistics: Theory and Methods. Wiley Series in Probability and Statistics. West Sussex, England: John Wiley & Sons, Ltd. 12

Miao, J. and Ben-Israel, A. (1992). "On Principal Angles Between Subspaces in $\mathbb{R}^n$." Linear Algebra and its Applications, 171: 81–98. 18, 19, 37

O'Hagan, A., Buck, C. E., Daneshkhah, A., Eiser, J. R., Garthwaite, P. H., Jenkinson, D. J., Oakley, J. E., and Rakow, T. (2006). Uncertain judgements: eliciting experts' probabilities. John Wiley & Sons. 2

Pettitt, A. N. (1982). "Inference for the Linear Model using a Likelihood Based on Ranks." Journal of the Royal Statistical Society. Series B, 44: 234–243. 5

— (1983). "Likelihood Based Inference Using Signed Ranks for Matched Pairs." Journal of the Royal Statistical Society. Series B, 45: 287–296. 5

Pratt, J. W. (1965). "Bayesian Interpretation of Standard Inference Statements." Journal of the Royal Statistical Society. Series B, 27: 169–203. 5

Pritchard, J. K., Seielstad, M. T., Perez-Lezaun, A., and Feldman, M. W. (1999). "Population Growth of Human Y Chromosomes: A Study of Y Chromosome Microsatellites." Molecular Biology and Evolution, 16: 1791–1798. 6

Ratcliff, R. (1993). "Methods for Dealing with Reaction Time Outliers." Psychological Bulletin, 114: 510. 4

Ronchetti, E., Field, C., and Blanchard, W. (1997). "Robust Linear Model Selection by Cross-Validation." Journal of the American Statistical Association, 92: 1017–1023. 26

Rousseeuw, P. J. and Leroy (1987). Robust regression and outlier detection. John wiley & sons. 8

Savage, I. R. (1969). "Nonparametric Statistics: A Personal Review." Sankhya: The Indian Journal of Statistics, Series A (1961-2002), 31: 107–144. 5

Stigler, S. M. (1977). "Do Robust Estimators Work with Real Data?" The Annals of Statistics, 5(6): 1055–1098. 6

Tavaré, S., Balding, D. J., Griffiths, R., and Donnelly, P. (1997). "Inferring Coalescence Times from DNA Sequence Data." Genetics, 145: 505–518. 6

Wong, H. and Clarke, B. (2004). "Improvement over Bayes prediction in small samples in the presence of model uncertainty." Canadian Journal of Statistics, 32(3): 269–283. 5

Yuan, A. and Clarke, B. (2004). "Asymptotic Normality of the Posterior Given a Statistic." The Canadian Journal of Statistics, 32: 119–137. 5

Yuan, A. and Clarke, B. S. (1999). "A minimally informative likelihood for decision analysis: illustration and robustness." Canadian Journal of Statistics, 27(3): 649–665. 2

Zellner, A. (1986). "On assessing prior distributions and Bayesian regression analysis with g-prior

distributions." In Bayesian Inference and Decision Techniques: Essays in Honor of Bruno de Finetti, 233. 25

Zhu, H., Ibrahim, J. G., and Tang, N. (2011). "Bayesian influence analysis: a geometric approach." Biometrika, 98(2): 307–323. 2

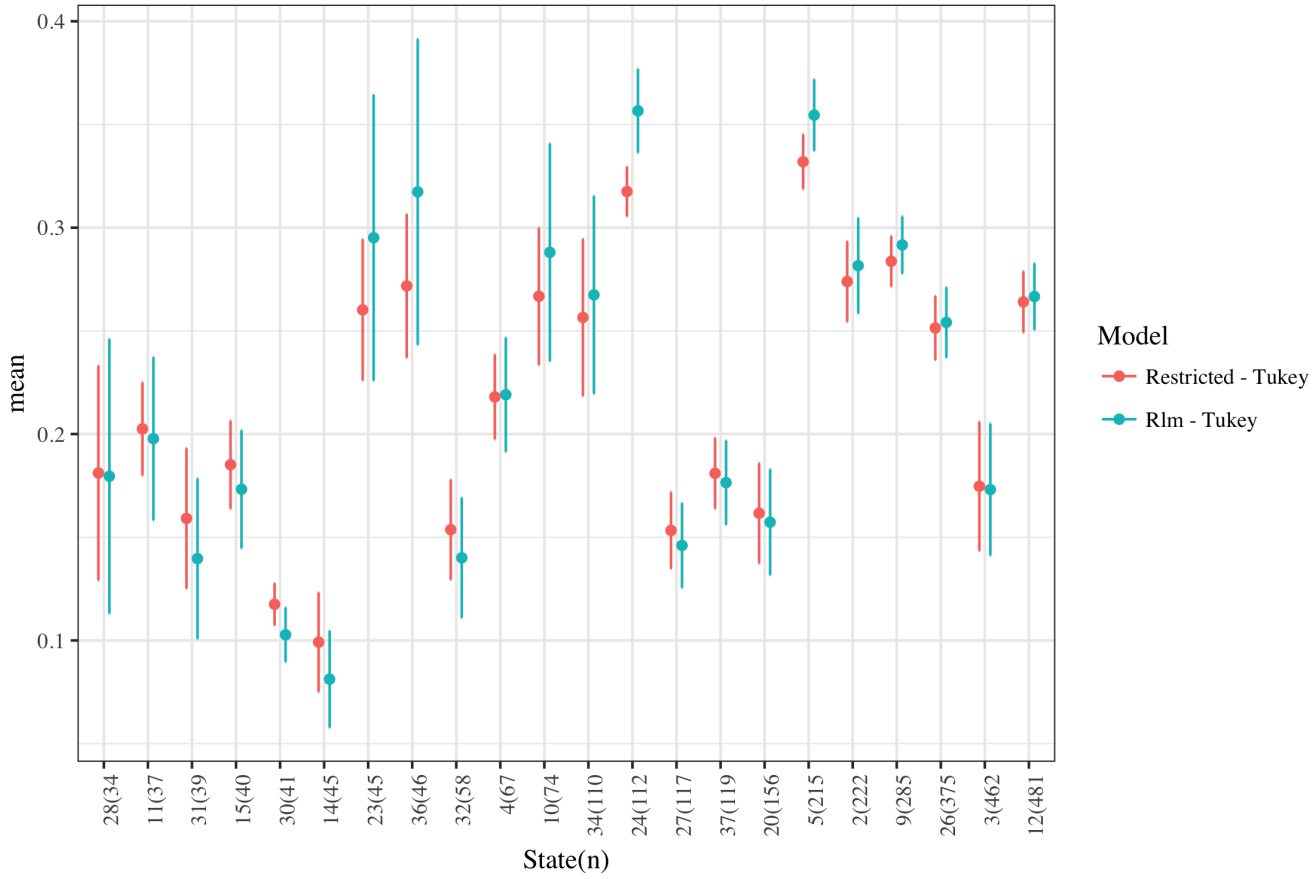

Figure 13: Hierarchical model results: $TLM_b(A)_j$ plus/minus one standard deviation over $K = 50$ repetitions for each state and $\alpha = 0.3$. The states are ordered along the x-axis according to number of agencies within the state (shown in parentheses). Results displayed are for the robust models using Tukey's M-estimators. Larger values of TLM are better.