Data Reduction for Efficient Probit Regression

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The probit model is a special case of the generalized linear model (GLM) described in [McCullagh and Nelder, 1989]. It is a statistical method for analyzing binary datasets, which we introduce in the following definition.

Definition 1 (Dataset)**.** *Let $\mathcal{D} = \{(x_i, y_i)\}_{i=1}^n$ be a set containing $n \in \mathbb{N}$ pairs of observations $x_i \in \mathbb{R}^d$, $y_i \in \{0, 1\}$. We call $\mathcal{D}$ a d-dimensional (binary) dataset.*

We can use this definition of a dataset (we will omit the term binary from now on since we will only be dealing with binary datasets in this work), to describe a whole range of possible scenarios that can be subjected to statistical analysis. For example, the x_i could represent some information of a patient, such as blood pressure or weight, and the y_i could indicate the presence or the absence of a heart disease.

In situations like this, we are often interested in modeling the relationship between the explanatory quantities x_i and the outcomes y_i. We need models, that can help us to answer questions about the data such as "Which factors increase/decrease the risk of suffering from a heart disease?", or "How likely is it, that a given patient will suffer from a heart disease?". The probit model is one of many approaches to model such a relationship in a probabilistic manner. It is described in detail in references like [McCullagh and Nelder, 1989], [Agresti, 2015] or [Fahrmeir et al., 2013].

We will outline the core assumptions of the probit model below, but instead of directly starting with its GLM formulation, we introduce it as a so-called latent variable model, which enables us to naturally arrive not only at its GLM specification, but also at a powerful sampling algorithm that enables us to efficiently apply the probit model in the realm of bayesian data analysis.

2.1 Introduction as a Latent Variable Model

When using a probit model to analyze a d-dimensional dataset $\mathcal{D} = \{(x_i, y_i)\}_{i=1}^n$, we implicitly make a set of assumptions about how the data was generated. Since it is reasonable to assume, that there is a degree of randomness involved in the data generating process, we model the y_i as realizations of independent random variables Y_i, which is the first assumption of the probit model.

The second assumption is that there is a hidden random quantity Y_i^* that is associated with each Y_i such that it directly determines its outcome:

$$Y_i = \begin{cases} 1, & \text{if } Y_i^* > 0 \\ 0, & \text{if } Y_i^* \leq 0 \end{cases} \tag{1}$$

The Y_i^* are also assumed to be independent from each other and, as already noted, unobservable, which is the reason why the Y_i^* are also called latent variables and why the probit model can also be thought of as a latent variable model.

The third and final assumption of the probit model defines the distribution of the Y_i^* and its part of the relationship between the non-random explanatory quantities x_i

and the outcomes y_i. In order to describe this relationship more concisely, we put all the observations x_i inside of a matrix $X \in \mathbb{R}^{n \times d}$ in such a way, that the i-th row of X corresponds to x_i. In the literature, this matrix X is often called the *model matrix* (see for example [Agresti, 2015]). We do the same with the Y_i^* and put them in a random vector Y^* as well, such that Y_i^* constitutes the i-th element of Y^*.

We are now ready for the third assumption of the probit model: The explanatory variables x_i influence Y_i^* in the form of a classical linear model:

$$Y^* = X\beta + \epsilon, \quad \epsilon \sim \mathcal{N}(0, \sigma^2 I), \tag{2}$$

where $\beta \in \mathbb{R}^d$ is the parameter vector of the linear model, ϵ is a normal distributed vector with independent components of mean zero and variance σ^2, and $I \in \mathbb{R}^{n \times n}$ is the $n \times n$ identity matrix. It follows directly that Y^* is also normal distributed: $Y^* \sim \mathcal{N}(X\beta, \sigma^2 I)$.

These three assumptions are already a complete specification of the probit model and are summarized in the following definition as a brief recapitulation:

Definition 2 (Probit Model)**.** *A d-dimensional binary dataset $\mathcal{D} = \{(x_i, y_i)\}_{i=1}^n$ with model matrix $X \in \mathbb{R}^{n \times d}$ was generated by a probit model with parameters $\beta \in \mathbb{R}^d$ and $\sigma \in \mathbb{R}_{>0}$, if the following three assumptions are true:*

1. *The observations $y_1, ..., y_n$ are realizations of independent binary random variables $Y_1, ..., Y_n$.*
2. *The outcomes of $Y_1, ..., Y_n$ are determined by hidden continuous random variables $Y_1^*, ..., Y_n^*$ by thresholding: If $Y_i^* > 0$, then $Y_i = 1$, and if $Y_i^* \leq 0$, then $Y_i = 0$.*
3. *The vector of hidden variables Y^* follows a multivariate normal distribution: $Y^* \sim \mathcal{N}(X\beta, \sigma^2 I)$, where $\beta \in \mathbb{R}^d$ and $\sigma \in \mathbb{R}_{>0}$ are the model parameters.*

Based on this definition, it is straight forward to determine the distribution of the response variables Y_i. We can calculate the probability $P(Y_i = 1)$ like this:

$$P(Y_i = 1) = P(Y_i^* > 0) = 1 - P(Y_i^* \leq 0) = 1 - P\left(\frac{Y_i^* - x_i^T\beta}{\sigma} \leq -\frac{x_i^T\beta}{\sigma}\right) = \Phi\left(\frac{x_i^T\beta}{\sigma}\right),$$

where $\Phi(\cdot)$ is the cumulative distribution function of the standard normal distribution:

$$\Phi(x) = \int_{-\infty}^{x} \frac{1}{\sqrt{2\pi}} e^{-\frac{1}{2}z^2} dz.$$

The result $P(Y_i = 1) = \Phi\left(\frac{x_i^T\beta}{\sigma}\right)$ leads us to an interesting observation: Both parameters β and σ are unknown model parameters and every value of σ can be compensated by a corresponding scaling of β. This means that, because we can't observe the hidden variables Y_i^*, it is impossible to determine which β and which σ generated the data without any prior knowledge. We can only draw conclusions with regard to the scaled parameter $\frac{1}{\sigma}\beta$. In this situation, we say that β and σ are *not identifiable*.

For this reason, in literature like [Fahrmeir et al., 2013] or [Agresti, 2015], it is often argued, that without the loss of generality, we can assume that $\sigma = 1$ and arrive at

$$P(Y_i = 1) = \Phi(x_i^T \beta).$$

Conversely, since Y_i is binary, it follows that

$$P(Y_i = 0) = 1 - P(Y_i = 1) = 1 - \Phi(x_i^T \beta) = \Phi(-x_i^T \beta),$$

and we arrive at the model equations:

$$Y_i \sim Bin(1, \pi_i), \quad \pi_i = \Phi(x_i^T \beta), \tag{3}$$

where $Bin(1, \pi_i)$ is a Bernoulli distribution with success probability $\pi_i = \Phi(x_i^T \beta)$.

2.2 A Special Case of the Generalized Linear Model

The final equations of the probit model that we arrived at in equation 3 are a special case of a more general model concept, the generalized linear model (GLM), that we briefly touch on below.

Generalized linear models consist of three components. The first one is the so called *random component*, a set of $n \in \mathbb{N}$ independent random variables $\{Y_i\}_{i=1}^n$. In GLMs, the distribution of these random variables is assumed to be a member of the *exponential family*, a broad family of probability distributions that encompasses the normal distribution, the binomial distribution and many others. It is characterized in more detail in [Agresti, 2015].

The second component of a GLM is the *linear predictor*. Just like in the probit model, we also assume that we are presented with some fixed observations $\{x_i \in \mathbb{R}^d\}_{i=1}^n$, that are assumed to have some explanatory power with regard to the Y_i. We thus call these observations the explanatory quantities. The linear predictor is used to relate the explanatory quantities to the distribution of the Y_i by linearly combining them as follows:

$$\eta_i = x_i^T \beta,$$

where $\eta_i \in \mathbb{R}$ denotes the linear predictor related to observation x_i and $\beta \in \mathbb{R}^d$ is the unknown parameter vector of the GLM that has to be estimated when fitting the model.

The third component of a GLM is the so called *link function*. This is a monotonic, differentiable and invertible function g that connects the linear predictor η_i to the distribution of the Y_i like this:

$$g(E[Y_i]) = \eta_i.$$

We are thus using the link function g to transform the expected value $E[Y_i]$ in such a way that it can be predicted by a linear model, hence the name *generalized linear models*.

Equivalently, we can also characterize this relationsthip by using the inverse function $h = g^{-1}$, also called the *response function*:

$$E[Y_i] = h(\eta_i).$$

We are now ready to establish the connection between the probit model and the generalized linear model. As we saw in equation 3, the assumptions of the probit model imply that the Y_i follow independent binomial distributions with a success probability of $\pi_i = \Phi(x_i^T \beta)$. The binomial distribution is a member of the exponential family, so we can also think of the Y_i as the random component of a GLM.

It also follows directly from the binomial distribution that $E[Y_i] = \pi_i$, thus we have from the probit model equations that $\pi_i = E[Y_i] = \Phi(x_i^T \beta)$, and equivalently $\Phi^{-1}(E[Y_i]) = x_i^T \beta$. Thus, we can think of Φ as the response function of a GLM and Φ^{-1}, the quantile function of the standard normal distribution, as the link function. The function Φ^{-1} is also known as the *probit function*, hence the name probit model.

2.3 Parameter Estimation

The parameters of generalized linear models and therefore the parameters of the probit model are usually estimated by using the *maximum likelihood method*. This method seeks to maximize the likelihood that some observed dataset $\mathcal{D} = \{(x_i, y_i)\}_{i=1}^n$ was generated under the assumptions of the model, given some parameter vector $\beta \in \mathbb{R}^d$.

To make notation a little easier, we also put the outcomes y_i in a vector $y \in \{0,1\}^n$ such that y_i is the i-th component of y. In the same way, we also put the random variables Y_i inside of a random vector Y.

In the probit model, the likelihood function is given as

$$\mathcal{L}(\beta) = P(Y = y|\beta) = \prod_{i=1}^{n} P(Y_i = y_i|\beta), \tag{4}$$

because the Y_i are independent. By using a little trick, we can write $P(Y_i = y_i|\beta)$ as a single expression by combining the equations $P(Y_i = 1) = \Phi(x_i^T \beta)$ and $P(Y_i = 0) = \Phi(-x_i^T \beta)$ from section 2.1 like this:

$$P(Y_i = y_i|\beta) = \Phi[(2y_i - 1)x_i^T \beta],$$

which works because $2y_i - 1 = 1$ for $y_i = 1$ and $2y_i - 1 = -1$ for $y_i = 0$. This enables us to arrive at the likelihood

$$\mathcal{L}(\beta) = \prod_{i=1}^{n} P(Y_i = y_i|\beta) = \prod_{i=1}^{n} \Phi[(2y_i - 1)x_i^T \beta] = \prod_{i=1}^{n} \Phi(-z_i^T \beta). \tag{5}$$

Here, we introduced the new vector $z_i = -(2y_i - 1)x_i$, which will simplify the notation later on.

The maximum likelihood estimate for β is then given by

$$\hat{\beta} \in \underset{\beta \in \mathbb{R}^d}{\operatorname{argmax}}\ \mathcal{L}(\beta), \tag{6}$$

and for $n \to \infty$ it holds that $E[\hat{\beta}] = \beta$ [Fahrmeir et al., 2013].

However, for finite sample sizes, the existence of $\hat{\beta}$ cannot be guaranteed and is dependent on the observed data. An overview of the conditions for the existence and uniqueness of $\hat{\beta}$ is given in [Demidenko, 2001]. In particular, there is one important condition shown in [Lesaffre and Kaufmann, 1992], that is related to the concept of linear separability, which we introduce in the following definition.

Definition 3 (Linear separability)**.** *Let $\mathcal{D} = \{(x_i, y_i)\}_{i=1}^n$ be a d-dimensional binary dataset. Let $S_0 = \{i \in [n] : \ y_i = 0\}$ and $S_1 = \{i \in [n] : \ y_i = 1\}$. If there exists a $\beta \in \mathbb{R}^d \setminus \{0\}$ such that*

$$\forall i \in S_0 : \ x_i^T \beta \leq 0 \quad \textit{and} \quad \forall i \in S_1 : \ x_i^T \beta \geq 0,$$

then we call $\mathcal{D}$ linearly separable.

Intuitively speaking, a dataset is linearly separable if there exists a hyperplane that perfectly separates the datapoints labeled with 1 from the datapoints labeled with 0. This property of a dataset is a both sufficient and necessary condition for the existence of the maximum likelihood estimate $\hat{\beta}$, as stated in the following theorem.

Theorem 1 ([Lesaffre and Kaufmann, 1992])**.** *Let $\mathcal{D} = \{(x_i, y_i)\}_{i=1}^n$ be a d-dimensional binary dataset. The maximum likelihood estimate $\hat{\beta}$ for the parameter β of the probit model exists if and only if $\mathcal{D}$ is not linearly separable.*

In [Haberman, 1974], it was further shown, that if the maximum likelihood estimate exists and the model matrix X has full column rank, i.e. $rank(X) = d$, then it is also unique. It now remains to explore, how the maximum likelihood optimization problem can be solved in such a case.

2.3.1 Finding the Maximum Likelihood Estimate

For the reason that the likelihood function $\mathcal{L}(\beta)$ is numerically inconvenient to maximize, the natural logarithm is often applied as a transformation to simplify the optimization problem:

$$\ell(\beta) = \ln \mathcal{L}(\beta) = \sum_{i=1}^{n} \ln \Phi(-z_i^T \beta). \tag{7}$$

Since we later wish to interpret ℓ as a loss function, we prefer to minimize the negative value of ℓ rather than maximizing:

$$f(\beta) = -\ell(\beta) = \sum_{i=1}^{n} \ln \left(\frac{1}{1 - \Phi(z_i^T \beta)} \right) = \sum_{i=1}^{n} g(z_i^T \beta). \tag{8}$$

Here, we define $g(x) = \ln \left(\frac{1}{1-\Phi(x)} \right)$ and call it the *probit loss*, i.e. the loss-function that determines how much each z_i contributes to the total loss $f(\beta)$ for a given value of β.

At this point, we could already elaborate on the minimization of $f(\beta)$, but there is one more generalization that we have to make, which will later be needed when applying

the theory of data reduction to the probit model: We have to introduce positive sample weights $w_1, ...w_n$, alternatively specified by the weight vector $w \in \mathbb{R}^n_{>0}$, that give a positive weight to each datapoint in the objective function:

$$f_Z^w(\beta) = \sum_{i=1}^{n} w_i g(z_i^T \beta). \tag{9}$$

Here, we also introduced the subscript Z, which refers to the matrix $Z \in \mathbb{R}^{n \times d}$, where the i-th row of Z is given by z_i, but if Z and w are clear from the context, we will usually omit it and simply refer to f_Z^w by f.

The optimization of f is usually done by applying the Newton-Raphson algorithm, an iterative procedure that starts at some initial guess $\beta^{(0)}$ and successively updates it like this:

$$\beta^{(t)} = \beta^{(t-1)} - \left(\frac{\partial^2 f(\beta^{(t-1)})}{\partial \beta \partial \beta^T} \right)^{-1} \cdot \frac{\partial f(\beta^{(t-1)})}{\partial \beta}, \tag{10}$$

where $\left(\frac{\partial^2 f(\beta^{(t-1)})}{\partial \beta \partial \beta^T} \right)^{-1}$ refers to the inverse of the hessian matrix of f, evaluated at $\beta^{(t-1)}$, and $\frac{\partial f(\beta^{(t-1)})}{\partial \beta}$ refers to the gradient of f, evaluated at $\beta^{(t-1)}$. The idea behind this procedure is, broadly speaking, to approximate f locally around $\beta^{(t)}$ as a second degree taylor-polynomial and then analytically find the minimum of this polynomial. The minimum of this local polynomial approximation of f is then iteratively used as the basis for the next step of the Newton-Raphson algorithm.

It remains to find the gradient as well as the hessian matrix of f. Because f is a sum of the function g evaluated at different points, it makes sense to first determine the derivative of g. This can be accomplished by using the chain rule as follows:

$$\begin{aligned}
\frac{d}{dx} g(x) &= \frac{d}{dx} \ln \left(\frac{1}{1 - \Phi(x)} \right) \\
&= (1 - \Phi(x)) \cdot \frac{d}{dx} \left(\frac{1}{1 - \Phi(x)} \right) \\
&= (1 - \Phi(x)) \cdot \frac{(-1)}{(1 - \Phi(x))^2} \cdot \frac{d}{dx} (1 - \Phi(x)) \\
&= \frac{(-1)}{1 - \Phi(x)} \cdot (-1) \cdot \phi(x) \\
&= \frac{\phi(x)}{1 - \Phi(x)},
\end{aligned} \tag{11}$$

where $\phi(x)$ is the density function of the standard normal distribution function:

$$\phi(x) = \frac{1}{\sqrt{2\pi}} e^{-\frac{1}{2}x^2}.$$

We can use this result to calculate the gradient of f:

$$\begin{aligned}\frac{\partial}{\partial\beta}f(\beta) &= \frac{\partial}{\partial\beta}\sum_{i=1}^{n} w_i g(z_i^T\beta) \\ &= \sum_{i=1}^{n} w_i z_i g'(z_i^T\beta) \\ &= \sum_{i=1}^{n} w_i z_i \frac{\phi(z_i^T\beta)}{1-\Phi(z_i^T\beta)}\end{aligned} \tag{12}$$

Next, we need to determine the hessian matrix of f. In order to do this, we again start by finding the second derivative of g, this time using the quotient rule:

$$\begin{aligned}\frac{d^2}{dx^2}g(x) &= \frac{d}{dx}\frac{\phi(x)}{1-\Phi(x)} \\ &= \frac{\phi'(x)(1-\Phi(x)) - \phi(x)\cdot(-1)\cdot\phi(x)}{(1-\Phi(x))^2} \\ &= \frac{(-1)\cdot x\cdot\phi(x)(1-\Phi(x)) - \phi(x)\cdot(-1)\cdot\phi(x)}{(1-\Phi(x))^2} \\ &= \frac{[\phi(x)]^2 - x\cdot\phi(x)\cdot(1-\Phi(x))}{(1-\Phi(x))^2} \\ &= \left(\frac{\phi(x)}{1-\Phi(x)}\right)^2 - x\cdot\frac{\phi(x)}{1-\Phi(x)} \\ &= \frac{\phi(x)}{1-\Phi(x)}\left(\frac{\phi(x)}{1-\Phi(x)} - x\right) \\ &= g'(x)\cdot(g'(x)-x)\end{aligned} \tag{13}$$

We can now use this result to find the hessian matrix of f:

$$\begin{aligned}\frac{\partial^2}{\partial\beta\partial\beta^T}f(\beta) &= \sum_{i=1}^{n}\frac{\partial^2}{\partial\beta\partial\beta^T}w_i g(z_i^T\beta) \\ &= \sum_{i=1}^{n} w_i z_i z_i^T g'(z_i^T\beta)(g'(z_i^T\beta) - z_i^T\beta) \\ &= \sum_{i=1}^{n} w_i z_i z_i^T \frac{\phi(z_i^T\beta)}{1-\Phi(z_i^T\beta)}\left(\frac{\phi(z_i^T\beta)}{1-\Phi(z_i^T\beta)} - z_i^T\beta\right).\end{aligned} \tag{14}$$

Because it can be shown, that $f(\beta)$ is a convex function [Wedderburn, 1976], and that the Newton Raphson algorithm converges to a global optimum when applied to a convex function [Nocedal and Wright, 2006], the optimization procedure converges to the maximum likelihood estimate $\hat{\beta}$ under the condition that the data is not linearly separable.

2.4 The Bayesian Perspective

The most fundamental difference between the bayesian approach to the probit model and the frequentist approach that was discussed above, is the assumption, that the model parameter β is not a fixed value, but a random variable with a probability distribution. The goal of bayesian data analysis is to draw conclusions about the distribution of the model parameter and to update these conclusions after observing more and more data.

A detailed overview of the principles of bayesian data analysis would certainly go beyond the scope of this work, but the interested reader will find a comprehensive reference in [Gelman et al., 2013]. In this section, we will merely touch on the most essential concepts, which are required in order to understand the bayesian view on the probit model.

2.4.1 Prior and Posterior Distributions

To characterize the prior uncertainty about the model parameter β, the first step of bayesian data analysis is to specify a so called *prior distribution*. In the probit model, one common choice that is also described in [Fahrmeir et al., 2013] is to assume that

$$\beta \sim \mathcal{N}(\mu_\beta, \Sigma_\beta), \tag{15}$$

i.e. β follows a normal distribution with mean $\mu_\beta \in \mathbb{R}^d$ and covariance matrix $\Sigma_\beta \in \mathbb{R}^{d\times d}$.

We can think of μ_β and Σ_β as a way to include prior knowledge into the model. If such knowledge is not present, we can choose μ_β and Σ_β in a more general fashion, perhaps we decide to set $\mu_\beta = 0$ and $\Sigma_\beta = \sigma_\beta \cdot I$ for a large value of σ_β, which would be an example of an *uninformative* prior because of the relatively unrestrictive assumptions. Alternatively, we could even go as far and also specify prior distributions on μ_β and Σ_β, which would lead us into the realm of hierarchical models (see [Gelman et al., 2013] for more details). But this would definitely go beyond the scope of this work, which is why we assume from now on that the values of μ_β and Σ_β are specified beforehand in a reasonable manner.

The next step in the process of bayesian data analysis is to determine how we should update our initial prior assumptions about β after we observed some new data represented by the dataset $\mathcal{D} = \{(x_i, y_i)\}_{i=1}^n$. Ultimately, the goal is to determine the *posterior distribution* of β given the new data, represented by the probability density function $p(\beta|Y = y)$, where y is the vector of observations and Y is the random vector that we assumed to have generated these observations in the probit model.

We can find the posterior distribution by making use of the bayes rule:

$$p(\beta|Y = y) = \frac{p(Y = y|\beta)p(\beta)}{p(Y = y)}. \tag{16}$$

This relationship tells us, that in order to arrive at the posterior distribution, there are three different parts that we have to combine.

The first part is the likelihood function $p(Y = y|\beta)$, which we already dealt with in section 2.3, and the second part is the prior density function $p(\beta)$, that we assumed to be normal.

The third and most challenging part to compute is the quantity $p(Y = y)$. We can see why it is so challenging by writing it out:

$$\begin{aligned} p(Y = y) &= \int p(Y = y|\beta)p(\beta)d\beta \\ &= \int \prod_{i=1}^{n} \Phi(-z_i^T\beta) \frac{1}{\sqrt{(2\pi)^d \det \Sigma_\beta}} \exp\left(-\frac{1}{2}(\beta - \mu_\beta)^T \Sigma_\beta^{-1}(\beta - \mu_\beta)\right) d\beta \end{aligned} \tag{17}$$

This infinite integral over all possible values of β is impossible to solve analytically, which means that it's impossible to exactly compute the posterior distribution for the probit model. But luckily, encountering an intractable integral like this is quite common in bayesian data analysis, so there are workarounds that still allow us to analyze the posterior distribution, even though we are unable to determine it exactly.

The first consideration is, that we could also analyze the posterior distribution if we had a large enough sample of it available instead. When the sample size is big enough, the Glivenko-Cantelli theorem tells us that the empirical posterior distribution converges to the true posterior distribution [Vaart, 1998]. This allows us to analyze the posterior distribution by analyzing a large enough sample of it, but the problem of how to obtain such a sample still remains.

In practice, instead of directly sampling from $p(\beta|Y = y)$, the posterior distribution can be approximated by so called Markov chain Monte Carlo (MCMC) methods. One such method that works particularly well for the probit model is the Gibbs sampler, which is described in the next section.

2.4.2 Gibbs Sampling in the Probit Model

Gibbs sampling is an iterative tool for drawing samples from probability distributions, that was first applied in the context of bayesian inference by [Gelfand and Smith, 1990] and has been adapted to the probit model by [Albert and Chib, 1993] using the idea of *data augmentation*, which was first introduced in [Tanner and Wong, 1987]. We describe this idea in the following section, as it yields an efficient algorithm for sampling from the posterior distribution of the probit model.

Remember, that the probit model has the following components: The vector of latent variables Y^* that follows a linear model $Y^* \mid \beta \sim \mathcal{N}(X\beta, 1)$, where we assume that $\sigma = 1$ for reasons of identifiability (see section 2.1) and the random vector Y that produces the observed outcomes y by thresholding: If $Y_i^* > 0$, then $Y_i = 1$ and $Y_i = 0$ otherwise. We also assumed a normal prior distribution: $\beta \sim \mathcal{N}(\mu_\beta, \Sigma_\beta)$.

Now, imagine that we knew the outcomes of the latent variable vector Y^*. The conditional distribution of β given the realization y^* of the latent variables can be shown to be normal [Albert and Chib, 1993]:

$$\beta \mid Y^* = y^* \ \sim \ \mathcal{N}(b, B), \tag{18}$$

where $b = (\Sigma_\beta^{-1} + X^TX)^{-1}(\Sigma_\beta^{-1}\mu_\beta + X^Ty^*)$ and $B = (\Sigma_\beta^{-1} + X^TX)^{-1}$. From this distribution, it is possible to sample efficiently.

The problem is, that in reality we can't observe the latent variables and therefore we don't know the realizations y^*. Here, an important finding by [Albert and Chib, 1993] comes into play: If we could observe β and see the realization $\tilde{\beta}$, then we could determine the conditional distribution of the latent variable vector Y^*:

$$Y_i^* \mid \beta = \tilde{\beta},\ Y_i = y_i \ \sim \begin{cases} \mathcal{N}(x_i^T\tilde{\beta}, 1) \text{ truncated at the left by 0}, & \text{if } y_i = 1 \\ \mathcal{N}(x_i^T\tilde{\beta}, 1) \text{ truncated at the right by 0}, & \text{if } y_i = 0 \end{cases} \tag{19}$$

This means, that given a realization $\tilde{\beta}$ and the observed values in y, the latent variables follow a truncated normal distribution, from which it is also possible to sample efficiently.

These two observations bring us directly to the Gibbs sampling algorithm for the probit model. The first step of this procedure is to determine a starting value $\tilde{\beta}^{(0)}$. [Albert and Chib, 1993] suggest that this could for example be the maximum likelihood estimate, that we already discussed in section 2.3.

The next step of the Gibbs sampling algorithm is to use this value $\tilde{\beta}^{(0)}$ to sample a realization $y^{*(1)}$ from the latent variable vector Y^*, by using the conditional distribution in equation 19. Given $y^{*(1)}$, it is then possible to sample a new value $\tilde{\beta}^{(1)}$ from the normal distribution in equation 18, which starts a new cycle. These two sampling steps, which can both be carried out efficiently, are repeated until the desired amount of samples is reached. See Algorithm 1 for the full algorithm.

To sum up, we augmented the observed data by incorporating the hidden variables Y^* to arrive at a two-stage procedure that draws alternating samples from the conditional distributions of β and Y^*, hence the name data augmentation.

Algorithm 1: Gibbs Sampler for the Probit Model

Input: Dataset $\mathcal{D} = \{(x_i, y_i)\}_{i=1}^n$ with model matrix X, prior mean $m \in \mathbb{R}^d$, prior covariance matrix $M \in \mathbb{R}^{d\times d}$, sample size $k \in \mathbb{N}$

Output: A sample $\beta_1, ..., \beta_k$ from the posterior distribution

Set $B = (M^{-1} + X^TX)^{-1}$
Initialize $\beta_0 = \hat{\beta}$, where $\hat{\beta}$ is the MLE for β computed on $\mathcal{D}$
for $j = 1, ..., k$ **do**
  for $i = 1, ..., n$ **do**
    if $y_i = 1$ **then**
      Sample $y_i^{*(j)}$ from $\mathcal{N}(x_i^T\beta_{j-1}, 1)$ truncated at the left by 0
    else if $y_i = 0$ **then**
      Sample $y_i^{*(j)}$ from $\mathcal{N}(x_i^T\beta_{j-1}, 1)$ truncated at the right by 0
  Set $y^{*(j)} = \left(y_1^{*(j)}, ..., y_n^{*(j)}\right)^T$
  Set $b^{(j)} = B\left(M^{-1}m + X^Ty^{*(j)}\right)$
  Sample β_j from $\mathcal{N}\left(b^{(j)}, B\right)$
return $\beta_1, ..., \beta_k$

3 Coresets and Sensitivity Sampling

In this work, we are using the method of coresets (see for example [Munteanu and Schwiegelshohn, 2018]) to approach the problem of data reduction for the probit model. The idea behind coresets is, that when given a dataset $\mathcal{D}$, we are interested in selecting only a small subset of observations $\mathcal{C} \subseteq \mathcal{D}$, such that the objective function evaluated on the (possible reweighted) subset $\mathcal{C}$ does not differ too much from the objective function evaluated on the original dataset $\mathcal{D}$.

This approach will allow us to estimate the model parameters efficiently on the ideally much smaller set $\mathcal{C}$, when a full optimization on $\mathcal{D}$ could already be infeasible for big datasets. We are thus following the paradigm of *sketch-and-solve*, i.e. first reducing the size of the original dataset and then solving the optimization problem on the reduced dataset.

In order to work out a formal definition of when we call a subset $\mathcal{C} \subseteq \mathcal{D}$ a coreset in the context of probit regression, we first have to slightly extend the concept of the model matrix, as we will need it for the coreset definition.

Definition 4 (Scaled model matrix)**.** *Let $\mathcal{D} = \{(x_i, y_i)\}_{i=1}^n$ be a d-dimensional dataset. Let $z_i = (2y_i - 1)x_i$ for all i in $[n]$. Then we call the matrix $Z \in \mathbb{R}^{n \times d}$, where the i-th row consists of the vector z_i for all $i \in [n]$, the scaled model matrix of $\mathcal{D}$.*

This definition of the scaled model matrix is nothing particularly new, it just formalizes the concept of factoring the labels into the model matrix, which we already encountered when dealing with the parameter estimation in section 2.3.

We are now ready for the coreset definition:

Definition 5 (Coreset)**.** *Let $\mathcal{D} = \{(x_i, y_i)\}_{i=1}^n$ be a d-dimensional dataset with scaled model matrix $Z \in \mathbb{R}^{n \times d}$ and a vector of positive sample weights $w \in \mathbb{R}^n_{>0}$. Let $\mathcal{C} \subseteq \mathcal{D}$ be a subset of $\mathcal{D}$ of size $|\mathcal{C}| = k$ with scaled model matrix $C \in \mathbb{R}^{k \times d}$ and a vector of positive sample weights $u \in \mathbb{R}^k_{>0}$. Let $\frac{1}{2} > \epsilon > 0$. We call $\mathcal{C}$ a $(1+\epsilon)$-coreset of $\mathcal{D}$ for probit regression, if*

$$(1-\epsilon) f_Z^w(\beta) \leq f_C^u(\beta) \leq (1+\epsilon) f_Z^w(\beta) \quad \forall \beta \in \mathbb{R}^d,$$

where $f_Z^w(\beta) = \sum_{i=1}^n w_i g(z_i^T \beta)$ is the weighted objective function of the probit model.

The size parameter $k = |\mathcal{C}|$ of a coreset usually depends on the desired approximation quality ϵ, as well as on specific problem characteristics, such as the number of observations n as well as the dimensionality d of the dataset. When constructing coresets, we are interested in keeping this parameter low in comparison to the total size of the dataset, i.e. we usually require that at least $k \in O(\log n)$, so that the data reduction is actually meaningful.

In the next section, we will investigate if there are any guarantees that can be given regarding the coreset size without imposing any further restrictions on the dataset. We will find out, that in the general case, it can't be guaranteed that a reasonably small coreset always exists. As a consequence, we will later confine our research to a specific class of datasets that we will call μ-complex, for which small upper bounds on the coreset size can be derived.

3.1 Lower Bounds for Coreset Size in the General Case

The first result that we will take a look at in the following theorem shows, that there are some datasets, for which no sufficiently small coresets of a size of at maximum $k \in O(\log n)$ can be found.

Theorem 2. *There exists a d-dimensional dataset $\mathcal{D}$ of size $|\mathcal{D}| = n$, such that any $(1+\epsilon)$-coreset $\mathcal{C}$ of $\mathcal{D}$ for probit regression has a size $k = |\mathcal{C}|$ of at least $k \in \Omega\left(\frac{n}{\log n}\right)$.*

Proof. We can construct such a dataset by showing how coresets can be used in a communication protocol for the so called INDEX communication game to encode a message. The same technique was also used in [Munteanu et al., 2018] to find lower bounds for coresets of logistic regression and is here slightly adapted for probit regression.

The INDEX game consists of two players, Alice and Bob. Alice is given a random binary string $m \in \{0,1\}^n$ of n bits and Bob is given an index $i \in [n]$. The goal is for Alice to send a message to Bob that allows Bob to obtain the value m_i of Alice's binary string m. It was shown in [Kremer et al., 1999], that the minimum length of a message sent by Alice that still allows Bob to obtain m_i with constant probability is in $\Omega(n)$ bits. We will now see, how a coreset for probit regression can be used to encode such a message.

The first step is for Alice to convert her binary string m into a two-dimensional dataset $\mathcal{D}$ as follows: For each entry m_j of her binary string where $m_j = 1$, she adds a point

$$x_j = \left(\cos\left(2\pi\frac{j}{n}\right), \sin\left(2\pi\frac{j}{n}\right)\right)^T$$

to her set $\mathcal{D}$ and labels it with $y_j = 1$, ending up with the dataset

$$\mathcal{D} = \{(x_j, 1)\}_{j \in \{i \in [n]:\ m_i = 1\}}.$$

As we can see, all of these points are on the unit circle and all of them are labeled with 1.

The next step for her is to construct a $(1+\epsilon)$-coreset $\mathcal{C}$ of $\mathcal{D}$ for probit regression with sample weights $u \in \mathbb{R}^k_{>0}$ and to transmit both the coreset and the weight vector to Bob, which requires $O(\log(n))$ space for each point and weight.[1] We will later see, how large the size $|\mathcal{C}| = k$ of this coreset must be, so that Bob can still obtain the value of m_i with constant probability.

As soon as Alice's coreset $\mathcal{C}$ arrives at Bob, Bob can use it to obtain the value of m_i. To do this, Bob first adds two new points

$$q_1 = \left(\cos\left(2\pi\frac{i-0.5}{n}\right), \sin\left(2\pi\frac{i-0.5}{n}\right)\right)^T$$

and

$$q_2 = \left(\cos\left(2\pi\frac{i+0.5}{n}\right), \sin\left(2\pi\frac{i+0.5}{n}\right)\right)^T$$

[1] TODO: Why is this a reasonable assumption.

to the set and labels both points with 0 (see figure 1), i.e. Bob now has the dataset

$$\mathcal{C}' = \mathcal{C} \cup \{(q_1, 0)\} \cup \{(q_2, 0)\}.$$

Next, he uses this new dataset $\mathcal{C}'$ with scaled model matrix C' to minimize the weighted objective function $f^u_{C'}$ of the probit model, by using the Newton-Raphson optimization algorithm.

Taking a look at figure 1, it becomes evident, that Bobs points q_1 and q_2 are linearly separable from the other points if and only if Alice didn't add a point x_i, i.e. if $m_i = 0$. He can use the results of the optimization procedure to make a distinction between the two cases (which then allows him to determine the value of m_i) like this:

In the case of $m_i = 1$, Bobs points are not linearly separable from Alices original points, which means that there must occur at least one misclassification at a cost of $g(0) = \log(2)$ for the original loss function. Because Bobs dataset $\mathcal{C}'$ allows him to obtain a $(1 \pm \epsilon)$-approximation of the original cost function, he can check if the Newton-Raphson algorithm converges to a cost of at least $(1 - \epsilon)\log(2) \geq \frac{1}{2}\log(2)$. In this case, he knows that Alice must have added the point x_i, which means that $m_i = 1$.

Conversely, if at any point during the optimization procedure the cost function drops below $\frac{1}{2}\log(2)$ and approaches zero, Bob knows that Alice didn't add the point x_i, because his dataset $\mathcal{C}'$ is linearly separable. This will allow him to conclude that $m_i = 0$.

Let us now see how big the size k of Alice's coreset must be for this protocol to work with constant probability. In [Kremer et al., 1999] it was shown, that the minimum length of a message that Alice must send in order for the protocol to work is in $\Omega(n)$ bits. Since each of the points that Alice created can be encoded in $\log(n)$ space, it follows from the lower bound that $\Omega(n) \subseteq \Omega(k\log(n))$, so k must be in $\Omega\left(\frac{n}{\log(n)}\right)$.

We can conclude, that if there existed a $(1 + \epsilon)$-coreset of $\mathcal{D}$ for probit regression with size $k \in o\left(\frac{n}{\log(n)}\right)$, it would contradict the minimum message length of the INDEX communication game, which proves the theorem. □

In the proof of theorem 2, we have already encountered such a "degenerate" dataset, for which no sublinear sized coreset can be found, consisting of only positive labels. The next step of this work is to introduce a new complexity measure for datasets in the context of probit regression, that allows us to specify a broad class of datasets for which sublinear sized coresets do exist and to show how such coresets can be constructed for this class. The idea behind this complexity measure goes back to the work in [Munteanu et al., 2018], where the authors introduced a similar measure to describe a class of datasets that allow the construction of sublinear coresets in the context of logistic regression. We adapt this measure to the context of probit regression in the following definition.

Definition 6. *(μ-complexity) Let $\mathcal{D}$ be a d-dimensional dataset of size $|\mathcal{D}| = n$ with scaled model matrix $Z \in \mathbb{R}^{n \times d}$, where $z_i \in \mathbb{R}^d$ constitutes the i-th row of Z and let $w \in \mathbb{R}^n_{>0}$ be a vector of positive weights. Let $I^+_\beta = \{i \in [n] : \ w_i z_i^T \beta > 0\}$ and let*

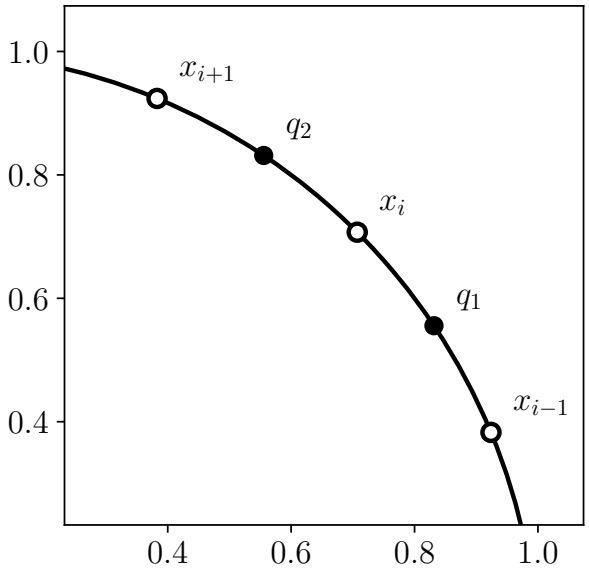

Figure 1: Bob places two points q_1 and q_2 in such a way on the unit circle, that they can be linearly seperated from the other points if and only if Alice didn't place a point at x_i.

$I_\beta^- = \{i \in [n] : \ w_i z_i^T \beta < 0\}$. *Let*

$$\mu_w(\mathcal{D}) = \sup_{\beta \in \mathbb{R}^d \setminus \{0\}} \frac{\sum_{i \in I_\beta^+} w_i (z_i^T \beta)^2}{\sum_{i \in I_\beta^-} w_i (z_i^T \beta)^2}.$$

We call the dataset $\mathcal{D}$ with weight vector w μ-complex, if there exists a $\mu \in \mathbb{R}$, such that $\mu_w(\mathcal{D}) \leq \mu$.

The following lemma will be helpful later on:

Lemma 1. *Let $\mathcal{D}$ be a d-dimensional and μ-complex dataset of size $|\mathcal{D}| = n$ with scaled model matrix $Z \in \mathbb{R}^{n \times d}$ and weight vector $w \in \mathbb{R}^n_{>0}$ like in definition 6. The following relationship holds for all $\beta \in \mathbb{R}^d$:*

$$\mu^{-1} \sum_{i \in I_\beta^-} w_i (z_i^T \beta)^2 \leq \sum_{i \in I_\beta^+} w_i (z_i^T \beta)^2 \leq \mu \sum_{i \in I_\beta^-} w_i (z_i^T \beta)^2.$$

Proof. If $\mathcal{D}$ with weights w is μ-complex, then

$$\begin{aligned} &\frac{\sum_{i \in I_\beta^+} w_i (z_i^T \beta)^2}{\sum_{i \in I_\beta^-} w_i (z_i^T \beta)^2} \leq \mu_w(\mathcal{D}) \leq \mu \\ \iff &\sum_{i \in I_\beta^+} w_i (z_i^T \beta)^2 \leq \mu \sum_{i \in I_\beta^-} w_i (z_i^T \beta)^2, \end{aligned}$$

which proves the second inequality.

Considering that the labeling of a dataset is arbitrary, i.e. we could always switch the 1 labels for the 0 labels and vice versa (if we flip the sign of β accordingly), the following relationship is true as well:

$$\begin{aligned}
&\frac{\sum_{i\in I_\beta^-} w_i(z_i^T\beta)^2}{\sum_{i\in I_\beta^+} w_i(z_i^T\beta)^2} \leq \mu_w(\mathcal{D}) \leq \mu \\
\iff &\sum_{i\in I_\beta^-} w_i(z_i^T\beta)^2 \leq \mu \sum_{i\in I_\beta^+} w_i(z_i^T\beta)^2 \\
\iff &\mu^{-1} \sum_{i\in I_\beta^-} w_i(z_i^T\beta)^2 \leq \sum_{i\in I_\beta^+} w_i(z_i^T\beta)^2,
\end{aligned}$$

which proves the first inequality. □

There is a close relationship between μ and the linear separability of a dataset as shown in the following theorem.

Theorem 3. *Let $\mathcal{D}$ be a d-dimensional dataset of size $|\mathcal{D}| = n$ like in definition 6 and let $w \in \mathbb{R}^n_{>0}$ be a vector of positive weights. Then, the dataset $\mathcal{D}$ with weight vector w is μ-complex if and only if $\mathcal{D}$ is not linearly separable.*

Proof. We first prove the "⇒" direction, i.e. we show that if $\mathcal{D}$ is μ-complex, then it is not linearly separable. We do this by proving the equivalent contraposition that if $\mathcal{D}$ is linearly separable, then it is not μ-complex.

Let $S_0 = \{i \in [n] : \ y_i = 0\}$ and $S_1 = \{i \in [n] : \ y_i = 1\}$ like in definition 3. If $\mathcal{D}$ is

linearly separable, then there exists a $\beta \in \mathbb{R}^d \setminus \{0\}$, such that

$$
\begin{aligned}
&\forall i \in S_0 :\ x_i^T \beta \leq 0 \quad \text{and} \quad \forall i \in S_1 :\ x_i^T \beta \geq 0 \\
\iff \\
&\forall i \in S_0 :\ (-1) x_i^T \beta \geq 0 \quad \text{and} \quad \forall i \in S_1 :\ x_i^T \beta \geq 0 \\
\iff \\
&\forall i \in S_0 :\ (2y_i - 1) x_i^T \beta \geq 0 \quad \text{and} \quad \forall i \in S_1 :\ (2y_i - 1) x_i^T \beta \geq 0 \\
\iff \\
&\forall i \in S_0 :\ z_i^T \beta \geq 0 \quad \text{and} \quad \forall i \in S_1 :\ z_i^T \beta \geq 0 \\
\iff \\
&\forall i \in [n] :\ z_i^T \beta \geq 0 \\
\iff \\
&I_\beta^- = \{i \in [n] :\ w_i z_i^T \beta < 0\} = \emptyset \\
\iff \\
&\sum_{i \in I_\beta^-} w_i (z_i^T \beta)^2 = 0 \\
\Rightarrow \\
&\mu_w(\mathcal{D}) \geq \frac{\sum_{i \in I_\beta^+} w_i (z_i^T \beta)^2}{\sum_{i \in I_\beta^-} w_i (z_i^T \beta)^2} = \infty,
\end{aligned}
$$

which means that $\mathcal{D}$ is not μ-complex.

It now remains to prove the "$\Leftarrow$" direction, i.e. to show that if $\mathcal{D}$ is not linearly separable, then it is μ-complex. Again, we do this by proving the equivalent contraposition that if $\mathcal{D}$ is not μ-complex, then it is linearly separable.

The first step in order to do so is to show that we can restrict the supremum in $\mu_w(\mathcal{D})$ to finite β with $\|\beta\| = 1$:

$$
\begin{aligned}
\mu_w(\mathcal{D}) &= \sup_{\beta \in \mathbb{R}^d \setminus \{0\}} \frac{\sum_{i \in I_\beta^+} w_i (z_i^T \beta)^2}{\sum_{i \in I_\beta^-} w_i (z_i^T \beta)^2} \\
&= \sup_{\beta \in \mathbb{R}^d \setminus \{0\}} \frac{\sum_{i \in I_\beta^+} \frac{1}{\|\beta\|^2} w_i (z_i^T \beta)^2}{\sum_{i \in I_\beta^-} \frac{1}{\|\beta\|^2} w_i (z_i^T \beta)^2} \\
&= \sup_{\beta \in \mathbb{R}^d \setminus \{0\}} \frac{\sum_{i \in I_\beta^+} w_i \left(z_i^T \frac{\beta}{\|\beta\|} \right)^2}{\sum_{i \in I_\beta^-} w_i \left(z_i^T \frac{\beta}{\|\beta\|} \right)^2} \\
&= \sup_{\tilde{\beta} \in \mathbb{R}^d,\ \|\tilde{\beta}\| = 1} \frac{\sum_{i \in I_\beta^+} w_i \left(z_i^T \tilde{\beta} \right)^2}{\sum_{i \in I_\beta^-} w_i \left(z_i^T \tilde{\beta} \right)^2},
\end{aligned}
$$

which lets us conclude that even in the supremum, both expressions $\sum_{i \in I_\beta^+} w_i (z_i^T \beta)^2$ and $\sum_{i \in I_\beta^-} w_i (z_i^T \beta)^2$ are finite. This means that if $\mathcal{D}$ is not μ-complex, then the denominator must be zero, i.e. it must hold that there exists a $\beta \in \mathbb{R}^d \setminus \{0\}$ such that

$$\sum_{i \in I_\beta^-} w_i (z_i^T \beta)^2 = 0.$$

From here, we can follow the same chain of equivalences that we showed when proving the "⇒"-direction of the theorem, which leads us directly to the fact, that $\mathcal{D}$ in this case must be linearly separable, which concludes the proof. □

As we already noted in section 2.3, linear separability is also closely related to the existence of the maximum likelihood estimate in the probit model. The next theorem uses the relationship between μ and linear separability to show the connection between μ and the existence of the maximum likelihood estimate.

Theorem 4. *Let $\mathcal{D}$ be a d-dimensional dataset. Then, the maximum likelihood estimate $\tilde{\beta}$ for the probit model exists if and only if $\mathcal{D}$ is μ-complex.*

Proof. This is a direct corollary from theorem 1 and theorem 3. □

In the following parts of this work, we will derive efficient upper bounds on the coreset size for μ-complex datasets. In order to do this, we first introduce a theoretic framework that we use for the coreset construction which is based on the concept of sensitivities.

3.2 The Sensitivity Framework

The sensitivity framework, which was first introduced by [Feldman and Langberg, 2011] (see also [Feldman et al., 2020] for a detailed overview), is a method for constructing provably small coresets by randomly sampling observations from a dataset according to a probability distribution, that emphasizes observations, which have a greater impact on the objective function.

Instead of representing a dataset $\mathcal{D} = \{(x_i, y_i)\}_{i=1}^n$ as a set of labeled datapoints, the sensitivity framework represents each point as a function that describes its contribution to the objective function. Recall, that in 2.3, we defined the weighted objective function of the probit model as $f_Z^w(\beta) = \sum_{i=1}^n w_i g(z_i^T \beta)$. We will now associate each datapoint (x_i, y_i) with the function $g_i(\beta) := g(z_i^T \beta) = g((2y_i - 1)x_i^T \beta)$, that describes its contribution to the total loss. That way, we can equivalently represent a dataset in the context of probit regression as a collection of loss functions $F = \{g_1, ..., g_n\}$.

The idea behind the sensitivity framework is to draw a random sample from this set of functions, where the sampling probability of each function is proportional to its worst-case contribution to the total loss for any $\beta \in \mathbb{R}^d$. This worst-case importance is also called sensitivity and was first introduced in [Langberg and Schulman, 2010]:

Definition 7 ([Langberg and Schulman, 2010]). *Let $F = \{g_1, ..., g_n\}$ be a set of functions, $g_i : \mathbb{R}^d \to \mathbb{R}_{\geq 0},\ i \in [n]$ and let $w \in \mathbb{R}^n_{>0}$ be a vector of positive weights. The sensitivity of g_i for $f_w(\beta) = \sum_{i=1}^n w_i g_i(\beta)$ is defined as*

$$\varsigma_i = \sup_{\beta \in \mathbb{R}^d,\ f_w(\beta) > 0} \frac{w_i g_i(\beta)}{f_w(\beta)}.$$

The total sensitivity, i.e. the sum of the sensitivities is $\mathfrak{S} = \sum_{i=1}^n \varsigma_i$.

The true sensitivity ς_i of a function g_i is usually unknown and its computation can be expensive, because it involves solving the original optimization problem, which was indicated in [Braverman et al., 2020]. For this reason, we are usually interested to find efficiently computable upper bounds $s_i \geq \varsigma_i$ for the sensitivities and then to draw samples proportional to the upper bounds s_i. As we will see, as long as the sum $S = \sum_{i=1}^n s_i$ of the upper bounds is sufficiently small, the coreset size will be small as well.

The second element of the sensitivity framework, which [Feldman and Langberg, 2011] related to the concept of sensitivity sampling in order to obtain small coresets, is the theory of range spaces and the VC-dimension. Its relevant definitions are given below.

Definition 8 ([Feldman and Langberg, 2011]). *A range space is a pair $\mathfrak{R} = (F, \text{ranges})$, where F is a set and* ranges *is a family (set) of subsets of F.*

Definition 9 ([Feldman and Langberg, 2011]). *The VC-dimension $\Delta(\mathfrak{R})$ of a range space $\mathfrak{R} = (F, \text{ranges})$ is the size $|G|$ of the largest subset $G \subseteq F$ such that*

$$|\{G \cap R \mid R \in \text{ranges}\}| = 2^{|G|},$$

i.e. G is shattered by ranges.

Definition 10 ([Feldman and Langberg, 2011]). *Let F be a finite set of functions mapping from $\mathbb{R}^d$ to $\mathbb{R}^{\geq 0}$. For every $\beta \in \mathbb{R}^d$ and $r \geq 0$, let*

$$\text{range}(F, \beta, r) = \{f \in F \mid f(\beta) \geq r\}$$

and let

$$\text{ranges}(F) = \left\{\text{range}(F, \beta, r) \mid \beta \in \mathbb{R}^d,\ r \geq 0\right\}.$$

Then we call $\mathfrak{R}_F := (F, \text{ranges}(F))$ the range space induced by F.

The following theorem is the basis of the sensitivity framework and combines the theory of range spaces with the concept of sensitivity sampling. Its original version goes back to [Feldman and Langberg, 2011], but it was further improved by [Braverman et al., 2020]. In this work, we will use the following variant by [Feldman et al., 2020]:

Theorem 5 ([Feldman et al., 2020]). *Let $F = \{g_1, ..., g_n\}$ be a set of functions, $g_i : \mathbb{R}^d \to \mathbb{R}_{\geq 0},\ i \in [n]$ and let $w \in \mathbb{R}^n_{>0}$ be a vector of positive weights. Let $\epsilon, \delta \in (0, \frac{1}{2})$.*

Let $s_i \geq \varsigma_i$ be upper bounds of the sensitivities and let $S = \sum_{i=1}^{n} s_i$. Given s_i, one can compute in time $O(|F|)$ a set $R \subseteq F$ of

$$|R| \in O\left(\frac{S}{\epsilon^2}\left(\Delta \log S + \log\left(\frac{1}{\delta}\right)\right)\right)$$

weighted functions, such that with probability $1-\delta$ we have for all $\beta \in \mathbb{R}^d$ simultaneously

$$(1-\epsilon)\sum_{g_i \in F} w_i g_i(\beta) \leq \sum_{g_i \in R} u_i g_i(\beta) \leq (1+\epsilon)\sum_{g_i \in F} w_i g_i(\beta).$$

Each element of R is sampled independently with probability $p_j = \frac{s_j}{S}$ from F, $u_i = \frac{Sw_j}{s_j|R|}$ denotes the weight of a function $g_i \in R$ that corresponds to $g_j \in F$ and Δ is an upper bound on the VC-dimension of the range space $\mathfrak{R}_{F^*}$ induced by F^*, where F^* is the set of functions $g_i \in F$ scaled by $\frac{Sw_i}{s_i|R|}$, i.e. $F^* = \left\{\frac{Sw_i}{s_i|R|} g_i(\beta) \mid i \in [n]\right\}$.

From this theorem, it follows that there are two things that have to be done in order to find a small coreset for probit regression.

The first one is to find small and efficiently computable upper bounds on the sensitivities and the second thing is to find a small upper bound on the VC-dimension of the range space induced by F^*. We will do both in the following section.

3.3 Constructing the Coreset

3.3.1 Bounding the Sensitivity

The first thing we need to do in order to find upper bounds on the sensitivities is to find bounds on the function g, which we do in the following two lemmas.

Lemma 2. *Let $g(x) = \ln\left(\frac{1}{1-\Phi(x)}\right)$. Then, for all $x \geq 0$, it holds that:*

$$\frac{1}{2}x^2 \leq g(x).$$

Proof. We first show the claim for all $x \geq 1$, by using the following inequality:

$$\begin{aligned}
\Phi(-x) &= \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{-x} \exp\left(-\frac{1}{2}z^2\right) dz \\
&\leq \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{-x} -z \exp\left(-\frac{1}{2}z^2\right) dz \\
&= \frac{1}{\sqrt{2\pi}} \exp\left(-\frac{1}{2}x^2\right) \\
&\leq \exp\left(-\frac{1}{2}x^2\right).
\end{aligned}$$

In the next step, we use this inequality to show that for $x \geq 1$:

$$e^{g(x)} = e^{\ln\left(\frac{1}{1-\Phi(x)}\right)} = \frac{1}{\Phi(-x)} \geq e^{\frac{1}{2}x^2}$$
$$\iff$$
$$g(x) \geq \frac{1}{2}x^2,$$

which proves the theorem for $x \geq 1$.

Let us now turn to the case when $0 \leq x \leq 1$. Both $g(x)$ and $\frac{1}{2}x^2$ are monotonically increasing and continuous functions for $0 \leq x \leq 1$. Making use of the fact that $g(0) > \frac{1}{2}$, it follows for all $0 \leq x \leq 1$, that

$$g(x) \geq g(0) > \frac{1}{2} = \max_{0 \leq x \leq 1} \frac{1}{2}x^2 \geq \frac{1}{2}x^2,$$

which concludes the proof. □

Lemma 3. *Let $g(x) = \ln\left(\frac{1}{1-\Phi(x)}\right)$. Then, for all $x \geq 2$, it holds that:*

$$g(x) \leq x^2.$$

Proof. In [Gordon, 1941], it was shown that the following inequality holds for all $x \geq 0$:

$$\Phi(-x) \geq \frac{1}{\sqrt{2\pi}} \frac{x}{x^2+1} e^{-\frac{1}{2}x^2}.$$

We can use this equality to establish that for all $x \geq 2$ it holds that:

$$\begin{aligned}
e^{x^2} \cdot \Phi(-x) &\geq e^{x^2} \frac{1}{\sqrt{2\pi}} \frac{x}{x^2+1} e^{-\frac{1}{2}x^2} \\
&= e^{\frac{1}{2}x^2} \frac{1}{\sqrt{2\pi}} \frac{x}{x^2+1} \\
&= e^{\frac{1}{2}x^2} \frac{1}{\frac{4}{3}(x^2+1)} \frac{\frac{4}{3}x}{\sqrt{2\pi}} \\
&\geq \frac{e^{\frac{1}{2}x^2}}{\frac{4}{3}(x^2+1)} \\
&\geq \frac{e^{\frac{1}{2}x^2}}{e^{\frac{1}{2}x^2}} \\
&= 1 \\
\iff & \\
e^{x^2} &\geq \frac{1}{1-\Phi(x)} \\
\iff & \\
x^2 &\geq \ln\left(\frac{1}{1-\Phi(x)}\right) = g(x),
\end{aligned}$$

which completes the proof. □

Having established upper and lower bounds for g, we can now turn to bounding the sensitivity.

Lemma 4. *Let $\mathcal{D}$ be a d-dimensional and μ-complex dataset of size $|\mathcal{D}| = n$ with scaled model matrix $Z \in \mathbb{R}^{n\times d}$ and let $w \in \mathbb{R}^n_{>0}$ be a vector of positive weights. Let $F = \{g_1, ..., g_n\}$ be a set of functions with $g_i(\beta) = g(z_i^T\beta)$ and let $f_w(\beta) = \sum_{i=1}^n w_i g_i(\beta)$. Then, it holds that*

$$w_j g_j(\beta) \leq 2\|U_j\|_2^2(1+\mu)f_w(\beta) \quad \forall\ j \in \{i \in [n] :\ z_i^T\beta \geq 2\},$$

where $U \in \mathbb{R}^{n\times d}$ is an orthonormal basis for the columnspace of $\sqrt{D_w}Z$ and $\sqrt{D_w} \in \mathbb{R}^{n\times n}$ is a diagonal matrix, where the i-th diagonal element is equal to $\sqrt{w_i}$ and $U_j \in \mathbb{R}^d$ is the j-th row of U.

Proof. Let $\sqrt{D_w}Z = UR$, where U is an orthonormal basis for the columnspace of $\sqrt{D_w}Z$. Then, for all $j \in \{i \in [n] :\ z_i^T\beta \geq 2\}$:

$$w_j g_j(\beta) = w_j g(z_j^T\beta) = w_j g\left(\frac{\sqrt{w_j}z_j^T\beta}{\sqrt{w_j}}\right) = w_j g\left(\frac{U_j^T R\beta}{\sqrt{w_j}}\right) \leq w_j g\left(\frac{\|U_j\|_2\|R\beta\|_2}{\sqrt{w_j}}\right),$$

where $U_j \in \mathbb{R}^d$ is the vector that constitutes the j'th row of U and the inequality is true due to the *Cauchy-Schwarz inequality*. We continue the proof as follows:

$$\begin{aligned} w_j g\left(\frac{\|U_j\|_2\|R\beta\|_2}{\sqrt{w_j}}\right) &= w_j g\left(\frac{\|U_j\|_2\|UR\beta\|_2}{\sqrt{w_j}}\right) \\ &= w_j g\left(\frac{\|U_j\|_2\|\sqrt{D_w}Z\beta\|_2}{\sqrt{w_j}}\right) \\ &\leq \|U_j\|_2^2\|\sqrt{D_w}Z\beta\|_2^2 \\ &= \|U_j\|_2^2 \sum_{i=1}^n w_i(z_i^T\beta)^2. \end{aligned}$$

Here, the first equality follows from the fact that U is orthonormal, i.e. multiplying by U doesn't change the norm of a vector. The inequality follows from the bound $g(x) \leq x^2$ that holds for all $x \geq 2$, which was shown in lemma 3.

Now, let $I_\beta^+ = \{i \in [n] :\ w_i z_i^T\beta > 0\}$ and let $I_\beta^- = \{i \in [n] :\ w_i z_i^T\beta < 0\}$ like in definition 6, the definition of μ-complexity. We continue the proof by making use of the

relationship that was shown in lemma 1:

$$\begin{aligned}
\|U_j\|_2^2 \sum_{i=1}^{n} w_i (z_i^T \beta)^2 &= \|U_j\|_2^2 \left(\sum_{i \in I_\beta^+} w_i (z_i^T \beta)^2 + \sum_{i \in I_\beta^-} w_i (z_i^T \beta)^2 \right) \\
&\leq \|U_j\|_2^2 \left(\sum_{i \in I_\beta^+} w_i (z_i^T \beta)^2 + \mu \sum_{i \in I_\beta^+} w_i (z_i^T \beta)^2 \right) \\
&= \|U_j\|_2^2 (1+\mu) \sum_{i \in I_\beta^+} w_i (z_i^T \beta)^2 \\
&\leq 2\|U_j\|_2^2 (1+\mu) \sum_{i \in I_\beta^+} w_i g(z_i^T \beta),
\end{aligned}$$

where the last inequality follows from the bound $g(x) \geq \frac{1}{2}x^2$, that holds for all $x \geq 0$, which we proved in lemma 2.

From here, we can use the fact that g is a strictly positive function to complete the proof:

$$2\|U_j\|_2^2 (1+\mu) \sum_{i \in I_\beta^+} w_i g(z_i^T \beta) \leq 2\|U_j\|_2^2 (1+\mu) \sum_{i=1}^{n} w_i g(z_i^T \beta) = 2\|U_j\|_2^2 (1+\mu) f_w(\beta)$$

□

Lemma 5. *Let $\mathcal{D}$ be a d-dimensional and μ-complex dataset of size $|\mathcal{D}| = n$ with scaled model matrix $Z \in \mathbb{R}^{n \times d}$ and let $w \in \mathbb{R}^n_{>0}$ be a vector of positive weights. Let $F = \{g_1, ..., g_n\}$ be a set of functions with $g_i(\beta) = g(z_i^T \beta)$ and let $f_w(\beta) = \sum_{i=1}^n w_i g_i(\beta)$. Then, it holds that*

$$w_j g_j(\beta) \leq \frac{w_j}{\mathcal{W}} (80 + 16\mu) f_w(\beta) \quad \forall\ j \in \{i \in [n] :\ z_i^T \beta \leq 2\},$$

where $\mathcal{W} = \sum_{i=1}^n w_i$ is the sum of all weights.

Proof. We first start by noting that $g(-1) > \frac{1}{10}$ and that $g(2) < 4$. Now, we partition the indices into two sets as follows:

$$\begin{aligned}
K_\beta^- &= \{i \in [n] \mid z_i^T \beta \leq -1\} \\
K_\beta^+ &= \{i \in [n] \mid z_i^T \beta > -1\}.
\end{aligned}$$

In the case that $\sum_{j \in K_\beta^+} w_j \geq \frac{1}{2}\mathcal{W}$, the following relationship holds:

$$f_w(\beta) = \sum_{i=1}^{n} w_i g(z_i^T \beta) \geq \sum_{i \in K_\beta^+} w_i g(z_i^T \beta) \geq \frac{\sum_{i \in K_\beta^+} w_i}{10} \geq \frac{\mathcal{W}}{20} = \frac{\mathcal{W}}{20 w_j} w_j \geq \frac{\mathcal{W}}{80 w_j} w_j g(z_j^T \beta),$$

where $j \in \{i \in [n] : \ z_i^T \beta \leq 2\}$. Thus, we have in this case:

$$w_j g(z_j^T \beta) \leq \frac{80 w_j}{\mathcal{W}} f_w(\beta).$$

If on the other hand $\sum_{j \in K_\beta^-} w_j \geq \frac{1}{2}\mathcal{W}$, we have that

$$f_w(\beta) = \sum_{i=1}^{n} w_i g(z_i^T \beta) \geq \sum_{i \in I_\beta^+} w_i g(z_i^T \beta) \geq \frac{1}{2} \sum_{i \in I_\beta^+} w_i (z_i^T \beta)^2 \geq \frac{1}{2\mu} \sum_{i \in I_\beta^-} w_i (z_i^T \beta)^2,$$

where $I_\beta^+ = \{i \in [n] : \ w_i z_i^T \beta > 0\}$ and $I_\beta^- = \{i \in [n] : \ w_i z_i^T \beta < 0\}$ like in definition 6 (μ-complexity). The second inequality is true due to the lower bound $g(x) \geq \frac{1}{2}x^2$ that holds for all $x \geq 0$ (see lemma 2) and the third inequality is true due to a property of μ that was proved in lemma 1.

We continue the proof as follows:

$$\frac{1}{2\mu} \sum_{i \in I_\beta^-} w_i (z_i^T \beta)^2 \geq \frac{1}{2\mu} \sum_{i \in K_\beta^-} w_i (z_i^T \beta)^2 \geq \frac{1}{2\mu} \sum_{i \in K_\beta^-} w_i \geq \frac{\mathcal{W}}{4\mu} \geq \frac{\mathcal{W}}{16\mu w_j} w_j g(z_j^T \beta),$$

which leads us to the upper bound for the second case:

$$w_j g(z_j^T \beta) \leq \frac{16\mu w_j}{\mathcal{W}} f_w(\beta).$$

We can conclude the proof by adding both upper bounds:

$$w_j g_j(\beta) = w_j g(z_j^T \beta) \leq \frac{80 w_j}{\mathcal{W}} f_w(\beta) + \frac{16\mu w_j}{\mathcal{W}} f_w(\beta) = \frac{w_j}{\mathcal{W}} (80 + 16\mu) f_w(\beta).$$

□

Lemma 6. *Let $\mathcal{D}$ be a d-dimensional and μ-complex dataset of size $|\mathcal{D}| = n$ with scaled model matrix $Z \in \mathbb{R}^{n \times d}$, let $w \in \mathbb{R}^n_{>0}$ be a vector of positive weights and let $U \in \mathbb{R}^{n \times d}$ be an orthonormal basis for the columnspace of $\sqrt{D_w} Z$. Let $F = \{g_1, ..., g_n\}$ be a set of functions with $g_i(\beta) = g(z_i^T \beta)$ and let $f_w(\beta) = \sum_{i=1}^n w_i g_i(\beta)$. Then, the sensitivity ς_i of g_i (see definition 7) is upper bounded by*

$$\varsigma_i \leq s_i = (80 + 16\mu)(\|U_i\|_2^2 + \frac{w_i}{\mathcal{W}}),$$

and the total sensitivity is bounded by

$$\mathfrak{S} = \sum_{i=1}^{n} \varsigma_i \leq 192\mu d.$$

Proof. We can use the bounds that we derived in lemma 4 and lemma 5 to bound the sensitivities:

$$\begin{aligned}
\varsigma_i &= \sup_{\beta \in \mathbb{R}^d,\ f_w(\beta)>0} \frac{w_i g(z_i\beta)}{f_w(\beta)} \\
&\leq \sup_{\beta \in \mathbb{R}^d,\ f_w(\beta)>0} \frac{2\|U_i\|_2^2(1+\mu)f_w(\beta) + \frac{w_i}{\mathcal{W}}(80+16\mu)f_w(\beta)}{f_w(\beta)} \\
&= 2\|U_i\|_2^2(1+\mu) + \frac{w_i}{\mathcal{W}}(80+16\mu) \\
&\leq \|U_i\|_2^2(80+16\mu) + \frac{w_i}{\mathcal{W}}(80+16\mu) \\
&= (80+16\mu)(\|U_i\|_2^2 + \frac{w_i}{\mathcal{W}}),
\end{aligned}$$

which completes the first part of the proof. For the next part, we use that U is an orthonormal matrix. The Frobenius norm $\|U\|_F$ (see for example [Golub and van Loan, 2013]) of an orthonormal matrix is equal to $\sqrt{d}$, as can easily be verified:

$$\|U\|_F = \sqrt{\sum_{k=1}^{d}\sum_{l=1}^{n}|u_{lk}|^2} = \sqrt{\sum_{k=1}^{d} 1} = \sqrt{d},$$

where the second equality follows from the fact that the columns of U have unit norm due to its orthonormality. We can now conclude the proof as follows:

$$\begin{aligned}
\mathfrak{S} = \sum_{i=1}^{n}\varsigma_i &\leq (80+16\mu)\sum_{i=1}^{n}\|U_i\|_2^2 + \frac{w_i}{\mathcal{W}} \\
&= (80+16\mu)(\|U\|_F^2 + 1) \\
&= (80+16\mu)(d+1) \\
&\leq 96\mu(d+1) \\
&\leq 192\mu d.
\end{aligned}$$

□

3.3.2 Bounding the VC dimension

Lemma 7. *Let $Z \in \mathbb{R}^{n\times d}$, let $z_i \in \mathbb{R}^d$ be the i-th row of Z and let $c \in \mathbb{R}_{>0}$. Let $F = \{g_1, ..., g_n\}$ be a set of functions with $g_i(\beta) = g(z_i^T\beta)$. The VC-dimension of the range space induced by*

$$\mathcal{F}^c = \{cg_i(\beta) \mid i \in [n]\}$$

is bounded by $\Delta(\mathfrak{R}_{\mathcal{F}^c}) \leq d+1$.

Proof. The idea of this proof is closely related to a proof in [Huggins et al., 2016], where a similar theorem was proven in the context of logistic regression.

We start by noting that for all $G \subseteq \mathcal{F}^c$ we have

$$|\{G \cap R \mid R \in \text{ranges}(\mathcal{F}^c)\}| = \left|\left\{\text{range}(G, \beta, r) \mid \beta \in \mathbb{R}^d,\ r \geq 0\right\}\right|.$$

Since g is invertible and monotone, we have for all $\beta \in \mathbb{R}^d$ and $r \geq 0$ that

$$\begin{aligned}\text{range}(G, \beta, r) &= \{g_i \in G \mid g_i(\beta) \geq r\} \\ &= \left\{g_i \in G \mid cg(z_i^T\beta) \geq r\right\} \\ &= \left\{g_i \in G \mid z_i^T\beta \geq g^{-1}\left(\frac{r}{c}\right)\right\}.\end{aligned}$$

Note, that $\left\{g_i \in G \mid z_i^T\beta \geq g^{-1}\left(\frac{r}{c}\right)\right\}$ corresponds to the positively classified points of the affine hyperplane classifier $x \mapsto \text{sign}\left(x^T\beta - g^{-1}\left(\frac{r}{c}\right)\right)$. We thus have for all $G \subseteq \mathcal{F}^c$, that

$$|\{G \cap R \mid R \in \text{ranges}(\mathcal{F}^c)\}| = \left|\left\{\left\{g_i \in G \mid z_i^T\beta - s \geq 0\right\} \mid \beta \in \mathbb{R}^d,\ s \in \mathbb{R}\right\}\right|.$$

The VC dimension of the set of affine hyperplane classifiers is $d + 1$ (see for example [Kearns and Vazirani, 1994]), so it follows that $\Delta(\mathfrak{R}_{\mathcal{F}^c}) \leq d + 1$, which concludes the proof. □

Lemma 8. *Let $Z \in \mathbb{R}^{n\times d}$, let $z_i \in \mathbb{R}^d$ be the i-th row of Z and let $w \in \mathbb{R}^n_{>0}$ be a vector of positive weights, where $w_i \in \{v_1, ..., v_t\}$ for all $i \in [n]$. Let $F = \{g_1, ..., g_n\}$ be a set of functions with $g_i(\beta) = g(z_i^T\beta)$. The VC-dimension of the range space induced by*

$$\mathcal{F}^w = \{w_i g_i(\beta) \mid i \in [n]\}$$

is bounded by $\Delta(\mathfrak{R}_{\mathcal{F}_{probit}}) \leq t \cdot (d + 1)$.

Proof. This proof follows the same line of argumentation as a similar proof in [Munteanu et al., 2018], where the authors derived a similar bound in the context of logistic regression.

We start by partition the functions in $\mathcal{F}^w$ into t disjoint classes

$$F_j = \{w_i g(z_i\beta) \in \mathcal{F}_{probit} \mid w_i = v_j\}, \quad j \in [t].$$

The functions in each of these classes have an equal weight, wich means that by lemma 7, each of their induced range spaces has a VC-dimension of at most $d + 1$.

For the sake of contradiction, assume that $\Delta(\mathfrak{R}_{\mathcal{F}^w}) > t \cdot (d + 1)$ and let G be the corresponding set of size $|G| > t \cdot (d + 1)$ that is shattered by ranges$(\mathcal{F}^w)$. Since the sets F_j are disjoint, each intersection $F_j \cap G$ must be shattered by ranges(F_j) as well. Further, at least one of the intersections must have at minimum $\frac{|G|}{t}$ elements, which means that for at least one $j \in [t]$ it holds that $|F_j \cap G| \geq \frac{|G|}{t} > \frac{t\cdot(d+1)}{t} = d + 1$. This is a contradiction to lemma 7, which concludes the proof. □

3.3.3 A simple two-pass algorithm

Theorem 6. *Let $\mathcal{D}$ be a d-dimensional and μ-complex dataset of size $|\mathcal{D}| = n$ with scaled model matrix $Z \in \mathbb{R}^{n \times d}$, let $w \in \mathbb{R}^n_{>0}$ be a vector of positive weights, with $\omega = \frac{w_{max}}{w_{min}}$ being the ratio of the largest and smallest weight, $\mathcal{W} = \sum_{i=1}^n w_i$ being the sum of all weights, and let $U \in \mathbb{R}^{n \times d}$ be an orthonormal basis for the columnspace of $\sqrt{D_w}Z$, where $U_i \in \mathbb{R}^d$ is the vector that constitutes the i-th row of U. Let $\epsilon \in (0, \frac{1}{2})$.*

If $\mathcal{C} \subseteq \mathcal{D}$ is a subset of $\mathcal{D}$ of size $|\mathcal{C}| = k$, that was obtained by independently sampling

$$k \in O\left(\frac{\mu d}{\epsilon^2} \log(\omega n)\right)$$

elements from $\mathcal{D}$ proportional to

$$q_i = \min\left\{2^l \mid l \in \mathbb{Z},\ 2^l \geq \|U_i\|_2^2 + \frac{w_i}{\mathcal{W}}\right\},$$

i.e. with sampling probability $p_i = \frac{q_i}{\sum_{i=1}^n q_i}$ for all $i \in [n]$ and $u \in \mathbb{R}^k_{>0}$ is a new weight vector, where $u_j = \frac{w_i \sum_{l=1}^n p_l}{k p_i}$ is the new weight for an element in $\mathcal{C}$ that corresponds to the i-th element of $\mathcal{D}$, then with probability $1 - \log^{-c}(n)$, $\mathcal{C}$ with weights u is a $(1 \pm \epsilon)$-coreset of $\mathcal{D}$ for probit regression for any absolute constant $c > 1$.

Proof. TODO. □

4 Data Streams

Content.

5 Experiments

Content.

6 Concluding Remarks

Content.

7 Notes

7.1 VC Dimension

An alternative approach is to write down the VC dimension by using an instance space and a concept class as given in [Kearns and Vazirani, 1994].

Lemma 9. *Let $X = \{x_1, ..., x_n\} \subset \mathbb{R}^d \times \mathbb{R}_{>0}$ be the instance space consisting of n points with their last coordinate being positive. The concept class of interest, $\mathcal{C}$ over X, is given as follows:*

$$\mathcal{C} = \left\{ \{x \in X : \ f_{\beta,r}(x) \geq 0\} \mid \beta \in \mathbb{R}^d, r \geq 0 \right\},$$

with

$$f_{\beta,r}(x) = x_{d+1} \cdot g\left(\sum_{i=1}^{d} x_i \beta_i\right) - r$$

and

$$g(x) = -\log \Phi(-x).$$

The VC dimension of $\mathcal{C}$ is equal to the VC dimension of the range space induced by $\mathcal{F}^w_{probit} = \{w_i g(z_i \beta) \mid i \in [n]\}$, $Z \in \mathbb{R}^{n \times d}$, $w \in \mathbb{R}^n_{>0}$.

There a few different strategies that can be used to find an upper bound on the VC dimension of $\mathcal{C}$, as shown by the following lemmas. The first one is a simple upper bound for finite concept classes:

Lemma 10. *Let X be an instance space and $\mathcal{C}$ be a concept class over X. If the cardinality of $\mathcal{C}$ can be bounded by m, i.e. $|\mathcal{C}| \leq m$, then $VCdim(\mathcal{C}) \leq log(m)$.*

The next lemma partitions the concept class into smaller classes, for each of which the VC dimension can be bounded:

Lemma 11. *Let X be an instance space and $\mathcal{C}$ be a concept class over X. Let $\mathcal{C}_1, ..., \mathcal{C}_k$ be a partition of $\mathcal{C}$ into k disjoint subsets, i.e. $\mathcal{C} = \bigcup_{i=1}^{k} \mathcal{C}_i$ and $\mathcal{C}_i \cap \mathcal{C}_j = \emptyset \ \forall i \neq j$. Then, $VCdim(\mathcal{C}) \leq \sum_{i=1}^{k} VCdim(\mathcal{C}_i)$.*

Proof. For the sake of contradiction, assume there was a set $S \subseteq X$ of size $|S| > \sum_{i=1}^{k} VCdim(\mathcal{C}_i)$ that is shattered by $\mathcal{C}$. If S is shattered by C, every subset of S must also be shattered by C. Consider the intersections $T_i = \bigcup_{c \in \mathcal{C}_i} S \cap c$. Every T_i is a subset of S and $S = \bigcup_{i=1}^{k} T_i$. Since S is shattered by $\mathcal{C}$, every T_i must be shattered by $\mathcal{C}_i$. We assumed that $|S| > \sum_{i=1}^{k} VCdim(\mathcal{C}_i)$. It follows that there exists a T_j with $T_j > VCdim(\mathcal{C}_j)$. Since T_j is also shattered by $\mathcal{C}_j$, this is a contradiction, which concludes the proof. □

A result in [Linial et al., 1991] suggests an even smaller upper bound:

Lemma 12 ([Linial et al., 1991])**.** *Let X be an instance space and $\mathcal{C}$ be a concept class over X. Let $\mathcal{C} = \bigcup_{i=1}^{k} \mathcal{C}_i$ and $VCdim(\mathcal{C}_i) \leq m$. If k is bounded by a polynomial function of m, then $VCdim(\mathcal{C}) \leq 3m$.*

Instead of partitioning the concept class, we could also partition the instance space and obtain a similar bound:

Lemma 13. *Let X be an instance space and $\mathcal{C}$ be a concept class over X. Let $X_1, ..., X_k$ be a partition of X into k disjoint subsets, i.e. $X = \bigcup_{i=1}^{k} X_i$ and $X_i \cap X_j = \emptyset \ \forall i \neq j$. Let $\mathcal{C}_i = \{X_i \cap c \mid c \in \mathcal{C}\}$ be a concept class over X_i for all $i \in [k]$.*
Then, $VCdim(\mathcal{C}) \leq \sum_{i=1}^{k} VCdim(\mathcal{C}_i)$.

Proof. Again, assume there existed a set $S \subseteq X$ of size $|S| > \sum_{i=1}^{k} VCdim(\mathcal{C}_i)$ that is shattered by $\mathcal{C}$. S can be partitioned into disjoined subsets $T_i = S \cap X_i$, with $\bigcup_{i=1}^{k} T_i = S$. Every T_i must be shattered by $\mathcal{C}_i$. Since we assumed that $|S| > \sum_{i=1}^{k} VCdim(\mathcal{C}_i)$, there exists a T_j with $|T_j| > VCdim(C_j)$ which is also shattered by C_j. This contradiction concludes the proof. □

7.2 New idea for VC dimension proof

Lemma 14. *Let*

$$h_{\beta,r}(x) = \begin{cases} 1 \text{ if } & x_{d+1} \cdot g\left(\sum_{i=1}^{d} x_i \beta_i\right) - r \geq 0 \\ 0 \text{ else} & \end{cases}$$

Be a function from $\mathbb{R}^{d+1}$ *to* $\{0,1\}$ *with parameters* $\beta \in \mathbb{R}^d$ *and* $r \in \mathbb{R}_{\geq 0}$ *with*

$$g(x) = \log\left(\frac{1}{1-\Phi(x)}\right),$$

$$\Phi(x) = \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{x} e^{-\frac{1}{2}z^2} dz \ .$$

Let

$$H = \left\{x \mapsto h_{\beta,r}(x) \mid \beta \in \mathbb{R}^d,\ r \in \mathbb{R}_{\geq 0}\right\}$$

be the hypothesis class determined by h*. Then, the VC dimension of* H *is ...*

Proof. Let $S = \sum_{i=1}^{d} x_i \beta_i$. We show that h can be computed in t steps as follows:

$$\begin{aligned}
& x_{d+1} \cdot g\left(S\right) - r \geq 0 \\
\iff & \log\left(\frac{1}{1-\Phi(S)}\right) \geq \frac{r}{x_{d+1}} \\
\iff & \frac{1}{1-\Phi(S)} \geq \exp\left(\frac{r}{x_{d+1}}\right) \\
\iff & 1-\Phi(S) \leq \exp\left(-\frac{r}{x_{d+1}}\right) \\
\iff & \Phi(S) \geq 1-\exp\left(-\frac{r}{x_{d+1}}\right) \\
\iff & \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{S} e^{-\frac{1}{2}z^2} dz \geq 1-\exp\left(-\frac{r}{x_{d+1}}\right) \\
\iff & \int_{-\infty}^{S} e^{-\frac{1}{2}z^2} dz \geq \sqrt{2\pi}\left(1-\exp\left(-\frac{r}{x_{d+1}}\right)\right)
\end{aligned}$$

□

7.3 Online Leverage Scores

The leverage scores of a matrix $A \in \mathbb{R}^{n \times d}$ are given by $l_i = a_i^T (A^T A)^{-1} a_i$ [Cohen et al., 2020]. According to [Cohen et al., 2020], we can obtain overestimates of these scores by using only a subset of the rows in A to compute them.

Let A_j be a matrix that contains only the first j rows of A. It follows that the the estimated leverage score $\tilde{l}_j = a_j^T (A_j^T A_j)^{-1} a_j$ is an overestimate of l_j. In a recent paper by [Chhaya et al., 2020], it was shown that the sum of these overestimates can be bounded regardless of how the rows in A are ordered:

Lemma 15 ([Chhaya et al., 2020])**.**

$$\sum_{i=1}^{n} \tilde{l}_j \in O(d + d \log \|A\| - \min_{i \in [n]} \|a_i\|)$$

Next, we show how a simple algorithm that computes $\tilde{l}_j$ in an online manner (passing row by row over the data stream) can be constructed requiring only $\mathcal{O}(d^2)$ of working memory. The idea is to only keep the matrix $A_j^T A_j \in \mathbb{R}^{d \times d}$ in memory and update it for every new row a_{j+1} using a rank one update $A_{j+1}^T A_{j+1} = A_j^T A_j + a_{j+1} \cdot a_{j+1}^T$. See [Golub and van Loan, 2013] for more on matrix multiplication using outer products. The algorithm is given in algorithm 2.

Algorithm 2: Online Leverage Scores

Input: Matrix $A \in \mathbb{R}^{n \times d}$

Output: Online leverage scores $\tilde{l}_i$ for all $i \in [n]$

Initialize $M_0 = 0^{d \times d}$

foreach $a_i := i$'th row vector of A, $a_i \in \mathbb{R}^d$ **do**

  $M_i = M_{i-1} + a_i \cdot a_i^T$

  $\tilde{l}_i = a_i^T M_i^{\dagger} a_i$

return $\tilde{l}_i,\ i \in [n]$

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