How many faces can be recognized? Performance extrapolation for multi-class classification

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Abstract

The difficulty of multi-class classification generally increases with the number of classes. Using data from a subset of the classes, can we predict how well a classifier will scale with an increased number of classes? Under the assumption that the classes are sampled exchangeably, and under the assumption that the classifier is generative (e.g. QDA or Naive Bayes), we show that the expected accuracy when the classifier is trained on k classes is the $k-1$st moment of a *conditional accuracy distribution*, which can be estimated from data. This provides the theoretical foundation for performance extrapolation based on pseudolikelihood, unbiased estimation, and high-dimensional asymptotics. We find empirically that some of the methods work well even for non-generative classifiers.

1 Introduction

In multi-class classification, one observes pairs (z, y) where $y \in \mathcal{Y} \subset \mathbb{R}^p$ are feature vectors, and z are unknown labels, which lie in a countable label set $\mathcal{Z}$. The goal is to construct a classification rule for predicting the label of a new data point; generally, the classification rule $h : \mathcal{Y} \to \mathcal{Z}$ is learned from previously observed data points. In many applications of multi-class classification, such as face recognition or image recognition, the space of potential labels is practically infinite. In such a setting, one might consider a sequence of classification problems on finite label subsets $\mathcal{Z}_1 \subset \cdots \subset \mathcal{Z}_K$, where in the i-th problem, one constructs the classification rule $h^{(i)} : \mathcal{Y} \to \mathcal{Z}_i$. Supposing that (Z, Y) have a joint distribution, define the accuracy for the i-th problem as

$$\mathrm{acc}^{(i)} = \Pr[h^{(i)}(Y) = Z | Z \in \mathcal{Z}_i].$$

Using data from only $\mathcal{Z}_k$, can one predict the accuracy achieved on the larger label set $\mathcal{Z}_K$, with $K > k$? This is the problem of *performance extrapolation*.

A practical instance of performance extrapolation occurs in neuroimaging studies, where the number of classes k is limited by experimental considerations. Kay et al. [1] obtained fMRI brain scans which record how a single subject's visual cortex responds to natural images. The label set $\mathcal{Z}$ corresponds to the space of all grayscale photographs of natural images, and the set $\mathcal{Z}_1$ is a subset of 1750 photographs used in the experiment. They construct a classifier which achieves over 0.75 accuracy for classifying the 1750 photographs; based on exponential extrapolation, they estimate that it would take on the order of $10^{9.5}$ photographs before the accuracy of the model drops below 0.10! Directly validating this estimate would take immense resources, so it would be useful to develop the theory needed to understand how to compute such extrapolations in a principled way.

However, in the fully general setting, it is impossible on construct non-trivial bounds on the accuracy achieved on the new classes $\mathcal{Z}_K \setminus \mathcal{Z}_k$ based only on knowledge of $\mathcal{Z}_k$: after all, $\mathcal{Z}_k$ could consist entirely of well-separated classes while the new classes $\mathcal{Z}_K \setminus \mathcal{Z}_k$ consist entirely of highly inseparable

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classes, or vice-versa. Thus, the most important assumption for our theory is that of *exchangeable sampling*. The labels in $\mathcal{Z}_i$ are assumed to be an exchangeable sample from $\mathcal{Z}$. The condition of exchangeability ensures that the separability of random subsets of $\mathcal{Z}$ can be inferred by looking at the empirical distributions in $\mathcal{Z}_k$, and therefore that some estimate of the achievable accuracy on $\mathcal{Z}_K$ can be obtained.

The assumption of exchangeability greatly limits the scope of application for our methods. Many multi-class classification problems have a hierarchical structure [2], or have classes distributed according to non-uniform discrete distributions, e.g. power laws [3]; in either case, exchangeability is violated. It would be interesting to extend our theory to the hierarchical setting, or to handle non-hierarchical settings with non-uniform prior class probabilities, but again we leave the subject for future work.

In addition to the assumption of exchangeability, we consider a restricted set of classifiers. We focus on *generative classifiers*, which are classifiers that work by training a model separately on each class. This convenient property allows us to characterize the accuracy of the classifier by selectively conditioning on one class at a time. In section 3, we use this technique to reveal an equivalence between the expected accuracies of $\mathcal{Z}_k$ to moments of a common distribution. This moment equivalence result allows standard approaches in statistics, such as U-statistics and nonparametric pseudolikelilood, to be directly applied to the extrapolation problem, as we discuss in section 4. In non-generative classifiers, the classification rule has a joint dependence on the entire set of classes, and cannot be analyzed by conditioning on individual classes. We confirmed that our methods work well for generative classifiers in simulations, which we have omitted from the paper. Since generative classifiers are rarely used in large classification problems, we felt that a more interesting demonstration would be to apply our methods to extrapolate the performance of multinomial logistic regression, support vector machines, and k-nearest neighbors–three commonly used non-generative classifiers. In Section 5, we see that our methods achieve accurate extrapolation for these three non-generative classifiers. A possible reason for this fact is that since the Bayes classifier (the optimal classifier as implemented by an oracle) is generative, our theory must also apply to any classifier which adequately approximates the Bayes classifier. We discuss this idea further in Section 6.

To our knowledge, we are the first to formalize the problem of prediction extrapolation. We introduce three methods for prediction extrapolation: the method of extended unbiased estimation and the constrained pseudolikelihood method are novel. The third method, based on asymptotics, is a new application of a recently proposed method for estimating mutual information [4].

2 Setting

Having motivated the problem of performance extrapolation, we now reformulate the problem for notational and theoretical convenience. Instead of requiring $\mathcal{Z}_k$ to be a random subset of $\mathcal{Z}$ as we did in section 1, take $\mathcal{Z} = \mathbb{N}$ and $\mathcal{Z}_k = \{1, \dots, k\}$. We fix the size of $\mathcal{Z}_k$ without losing generality, since any monotonic sequence of finite subsets can be embedded in a sequence with $|\mathcal{Z}_k| = k$. In addition, rather than randomizing the labels, we will randomize the marginal distribution of each label; Towards that end, let $\mathcal{Y} \subset \mathbb{R}^p$ be a space of feature vectors, and let $\mathcal{P}(\mathcal{Y})$ be a measurable space of probability distributions on $\mathcal{Y}$. Let $\mathcal{F}$ be a probability measure on $\mathcal{P}$, and let $F_1, F_2, \dots$ be an infinite sequence of i.i.d. draws from $\mathbb{F}$. We refer to $\mathbb{F}$, a probability measure on probability measures, as a *meta-distribution*. The distributions $F_1, \dots, F_k$ are the marginal distributions of the first k classes. We therefore rewrite the accuracy as

$$\text{acc}^{(t)} = \frac{1}{t} \sum_{i=1}^{t} \Pr_{F_i}[h^{(t)}(Y) = i].$$

where the probabilities are taken over $Y \sim F_i$.

In order to construct the classification rule $h^{(t)}$, we need data from the classes $F_1, \dots, F_t$. In most instances of multi-class classification, one observes independent observations from each F_i which are used to construct the classifier. Since the order of the observations does not generally matter, a sufficient statistic for the training data for the t-th classification problem is the collection of empirical distributions $\hat{F}_1^{(t)}, \dots, \hat{F}_t^{(t)}$ for each class. Henceforth, we make the simplifying assumption that the training data for the i-th class remains fixed from $t = i, i+1, \dots$, so we drop the superscript on

$\hat{F}_i^{(t)}$. Write $\hat{\mathbb{F}}(F)$ for the conditional distribution of $\hat{F}_i$ given $F_i = F$; also write $\hat{\mathbb{F}}$ for the marginal distribution of $\hat{F}$ when $F \sim \mathbb{F}$. As an example, suppose every class has the number of training examples $r \in \mathbb{N}$; then $\hat{F}$ is the empirical distribution of r i.i.d. observations from F, and $\hat{\mathbb{F}}(F)$ is the *empirical meta-distribution* of $\hat{F}$. Meanwhile, $\hat{\mathbb{F}}$ is the meta-distribution of the empirical distribution of r i.i.d. draws from a random $F \sim \mathbb{F}$.

2.1 Multiclass classification

Extending the formalism of Tewari and Bartlett [5][1], we define a classifier as a collection of mappings $\mathcal{M}_i : \mathcal{P}(\mathcal{Y})^k \times \mathcal{Y} \to \mathbb{R}$ called *margin functions.* Intuitively speaking, each margin function *learns a model* from the first k arguments, which are the empirical marginals of the k classes, $\hat{F}_1, \ldots, \hat{F}_k$; for each class, the classifier assigns a *margin* or *score* to the *query point* $y \in \mathcal{Y}$. A higher score $\mathcal{M}_i(\hat{F}_1, \ldots, \hat{F}_k, y)$ indicates a higher estimated probability that y belongs to the k-th class. Therefore, the classification rule corresponding to a classifier $\mathcal{M}_i$ assigns a class with maximum margin to y:

$$h(y) = \text{argmax}_{i \in \{1,\ldots,k\}} \mathcal{M}_i(y).$$

For some classifiers, the margin function $\mathcal{M}_i$ is especially simple in that $\mathcal{M}_i$ is only a function of $\hat{F}_i$ and y. Furthermore, due to symmetry, in such cases one can write

$$\mathcal{M}_i(\hat{F}_1, \ldots, \hat{F}_k, y) = \mathcal{Q}(\hat{F}_i, y),$$

where $\mathcal{Q}$ is called a *single-class margin* (or simply *margin*), and we say that $\mathcal{M}$ is a *generative classifier*. Quadratic discriminant analysis and Naive Bayes [6] are two examples of generative classifiers[2].

For notational convenience, we assume that ties occur with probability zero: that is, Note that the tie-breaking property implies that $\mathbb{F}$ contains no atoms. The *generative* property allows us to prove strong results about the accuracy of the classifier under the exchangeable sampling assumption, as we see in Section 3.

3 Performance extrapolation for generative classifiers

Let us specialize to the case of a generative classifier, with scoring rule $\mathcal{Q}$. Consider estimating the expected accuracy for the t-th classification problem,

$$p_t \stackrel{def}{=} \mathbf{E}[\text{acc}^{(t)}]. \tag{1}$$

In the case of a generative classifier, we have

$$p_k = \mathbf{E}[acc^{(k)}] = \mathbf{E}\left[\frac{1}{k}\sum_{i=1}^k \Pr_{Y \sim F_i}[\mathcal{Q}(\hat{F}_i, Y) > \max_{j \neq i} \mathcal{Q}(\hat{F}_j, Y)]\right].$$

Define the *conditional accuracy* function $u(\hat{F}, y)$ which maps a distribution $\hat{F}$ on $\mathcal{Y}$ and a *test* observation y to a real number in $[0, 1]$. The conditional accuracy gives the probability that for independently drawn $\hat{F}'$ from $\hat{\mathbb{F}}$, that $\mathcal{Q}(\hat{F}, y)$ will be greater than $\mathcal{Q}(\hat{F}', y)$:

$$u(\hat{F}, y) = \Pr_{\hat{F} \sim \hat{\mathbb{F}}}[\mathcal{Q}(\hat{F}, y) > \mathcal{Q}(\hat{F}', y)].$$

[1]We borrow their terminology of *margin functions*, but introduce the notion of a classifier as a multiple-argument functional on empirical distributions. The functional formulation of a classifier echoes the functional formulation of estimators common in the statistical literature.

[2]For QDA, the margin is given by

$$\mathcal{Q}_{QDA}(\hat{F}, y) = -(y - \mu(\hat{F}))^T \Sigma(\hat{F})^{-1}(y - \mu(\hat{F})) - \log\det(\Sigma(\hat{F})),$$

where $\mu(F) = \int y dF(y)$ and $\Sigma(F) = \int (y - \mu(F))(y - \mu(F))^T dF(y)$. In Naive Bayes, the margin is

$$\mathcal{Q}_{NB}(\hat{F}, y) = \sum_{i=1}^n \log \hat{f}_i(y_i),$$

where $\hat{f}_i$ is a density estimate for the i-th component of $\hat{F}$.

Define the *conditional accuracy* distribution ν as the law of $u(\hat{F}, Y)$ where $\hat{F}$ and Y are generated as follows: (i) a true distribution F is drawn from $\mathbb{F}$; (ii) the query Y is drawn from F, and (iii) the empirical distribution $\hat{F}$ is drawn from $\hat{\mathbb{F}}(F)$ (e.g., the distribution of the empirical distribution of r i.i.d. observations drawn from F), with Y independent of $\hat{F}$. The significance of the conditional accuracy distribution is that the expected generalization error p_t can be written in terms of its moments.

Theorem 3.1. *Let $\mathcal{Q}$ be a single-distribution margin, and let $\mathbb{F}$, $\hat{F}(F)$ be a distribution on $\mathcal{P}(\mathcal{Y})$. Further assume that $\hat{\mathbb{F}}$ and $\mathcal{Q}$ jointly satisfy the* tie-breaking *property:*

$$\Pr[\mathcal{Q}(\hat{F}, y) = \mathcal{Q}(\hat{F}', y)] = 0 \tag{2}$$

for all $y \in \mathcal{Y}$, where $\mathbb{F}, \mathbb{F}' \stackrel{iid}{\sim} \hat{\mathbb{F}}$. Let U be defined as the random variable

$$U = u(\hat{F}, Y)$$

for $F \sim \mathbb{F}$, $Y \sim F$, and $\hat{F} \sim \hat{\mathbb{F}}(F)$ with $Y \perp \hat{F}$. Then

$$p_k = \boldsymbol{E}[U^{k-1}],$$

where p_k is the expected accuracy as defined by (1)*.*

Proof. Write $q^{(i)}(y) = \mathcal{Q}(\hat{F}_i, y)$. By using conditioning and conditional independence, p_k can be written

$$\begin{aligned}
p_k &= \mathbf{E}\left[\frac{1}{k}\sum_{i=1}^{k} \Pr_{F_i}[q^{(i)}(Y) > \max_{j\neq i} q^{(j)}(Y)]\right] \\
&= \mathbf{E}\left[\Pr_{F_1}[q^{(1)}(Y) > \max_{j\neq 1} q^{(j)}(Y)]\right] \\
&= \mathbf{E}_{F_1}[\Pr[q^{(1)}(Y) > \max_{j\neq 1} q^{(j)}(Y) | \hat{F}_1, Y]] \\
&= \mathbf{E}_{F_1}[\Pr[\cap_{j>1} q^{(1)}(Y) > q^{(j)}(Y) | \hat{F}_1, Y]] \\
&= \mathbf{E}_{F_1}[\prod_{j>1} \Pr[q^{(1)}(Y) > q^{(j)}(Y) | \hat{F}_1, Y]] \\
&= \mathbf{E}_{F_1}[\Pr[q^{(1)}(Y) > q^{(2)}(Y) | \hat{F}_1, Y]^{k-1}] \\
&= \mathbf{E}_{F_1}[u(\hat{F}_1, Y)^{k-1}] = \mathbf{E}[U^{k-1}].
\end{aligned}$$

□

Theorem 3.1 tells us that the problem of extrapolation can be approached by attempting to estimate the conditional accuracy distribution. The $(t-1)$-th moment of U gives us p_t, which will in turn be a good estimate of $\text{acc}^{(t)}$.

While $U = u(\hat{F}, Y)$ is not directly observed, we can obtain unbiased estimates of $u(\hat{F}_i, y)$ by using test data. For any $\hat{F}_1, \ldots, \hat{F}_k$, and independent test point $Y \sim F_i$, define

$$\hat{u}(\hat{F}_i, Y) = \frac{1}{k-1}\sum_{j\neq i} I(\mathcal{Q}(\hat{F}_i, Y) > \mathcal{Q}(\hat{F}_j, Y)). \tag{3}$$

Then $\hat{u}(\hat{F}_i, Y)$ is an unbiased estimate of $u(\hat{F}_i, Y)$, as stated in the following theorem.

Theorem 3.2. *Assume the conditions of theorem 3.1. Then defining*

$$V = (k-1)\hat{u}(\hat{F}_i, y), \tag{4}$$

we have

$$V \sim \text{Binomial}(k-1, u(\hat{F}_i, y)).$$

Hence,

$$\boldsymbol{E}[\hat{u}(\hat{F}_i, y)] = u(\hat{F}_i, y).$$

In section 4, we will use this result to estimate the moments of U. Meanwhile, since U is a random variable on $[0,1]$, we also conclude that p_t follows a *mixed exponential decay*. Let α be the law of $-\log(U)$. Then from change-of-variables $\kappa = -\log(u)$, we get

$$\mathbf{E}[\text{acc}^{(t)}] = \mathbf{E}[U^{t-1}] = \int_0^1 u^{t-1} d\nu(u) = \int_0^1 e^{t\log(u)} \frac{1}{u} d\nu(u) = \int_{\mathbb{R}^+} e^{-\kappa t} d\alpha(\kappa).$$

This fact immediately suggests the technique of fitting an mixture of exponentials to the test error at $t = 2, 3, \ldots, k$: we explore this idea further in Section 4.1.

3.1 Properties of the conditional accuracy distribution

The conditional error distribution ν is determined by $\mathbb{F}$ and $\mathcal{Q}$. What can we say about the the conditional accuracy distribution without making any assumptions on either $\mathbb{F}$ or $\mathcal{Q}$? The answer is: not much–for an arbitrary probability measure ν' on $[0,1]$, one can construct $\mathbb{F}$ and $\mathcal{Q}$ such that $\nu = \nu'$, even if one makes the *perfect sampling assumption* that $\hat{F} = F$.

Theorem 3.3. *Let U be defined as in Theorem 3.1, and let ν denote the law of U. Then, for any probability distribution ν' on $[0,1]$, one can construct a meta-distribution $\mathbb{F}$ and a scoring rule $\mathcal{Q}$ such that $\nu = \nu'$ under perfect sampling (that is, $\hat{F} = F$.)*

Proof. Let G be the cdf of ν, $G(x) = \int_0^x d\nu(x)$, and let $H(u) = \sup_x\{G(x) \leq u\}$. Define $\mathcal{Q}$ by

$$\mathcal{Q}(\hat{F}, y) = \begin{cases} 0 & \text{if } \mu(\hat{F}) > y + H(y) \\ 0 & \text{if } y + H(y) > 1 \text{ and } \mu(\hat{F}) \in [H(y) - y, y] \\ 1 + \mu(\hat{F}) - y & \text{if } \mu(\hat{F}) \in [y, y + H(y)] \\ 1 + y + \mu(\hat{F}) & \text{if } \mu(\hat{F}) + H(y) > 1 \text{ and } \mu(\hat{F}) \in [0, H(y) - y]. \end{cases}$$

Let $\theta \sim \text{Uniform}[0,1]$, and define $F \sim \mathbb{F}$ by $F = \delta_\theta$, and also $\hat{F} = F$. A straightforward calculation yields that $\nu = \nu'$. $\square$

On the other hand, we can obtain a positive result if we assume that the classifier approximates a *Bayes classifier.* Assuming that F is absolutely continuous with respect to Lebesgue measure Λ with probability one, a Bayes classifier results from assuming perfect sampling ($\hat{F} = F$) and taking $\mathcal{Q}(\hat{F}, y) = \frac{dF}{d\Lambda}(y)$. Theorem 3.4. states that for a Bayes classifier, ν has a density $\eta(u)$ which is monotonically increasing. Since a 'good' classifier approximates the Bayes classifier, we intuitively expect that a monotonically increasing density η is a good model for the conditional accuracy distribution of a 'good' classifier.

Theorem 3.4. *Assume the conditions of theorem 3.1, and further suppose that $\hat{F} = F$, F is absolutely continuous with respect to Λ with probability one, that $\mathcal{Q}(\hat{F}, y) = \frac{dF}{d\Lambda}(y)$, and that $F|Y$ has a regular conditional probability distribution. Let ν denote the law of U. Then ν has a density $\eta(u)$ on $[0,1]$ which is monotonic in u.*

Proof. It suffices to prove that

$$\nu([u, u+\delta]) < \nu([v, v+\delta])$$

for all $0 < u < v < 1$ and $0 < \delta < 1 - v$. Let $\mathcal{P}_{ac}(\mathcal{Y})$ denote the space of distributions supported on $\mathcal{Y}$ which are absolutely continuous with respect to p-dimensional Lebesgue measure Λ. Let $\mathbb{Y}$ denote the marginal distribution of Y for $Y \sim F$ with $F \sim \mathbb{F}$. Define the set

$$J_y(A) = \{F \in \mathcal{P}_{ac}(\mathcal{Y}) : u(F, y) \in A\}.$$

for all $A \subset [0,1]$. One can verify that for all $y \in \mathcal{Y}$,

$$\Pr_{\mathbb{F}}[J_y([u, u+\delta])|Y = y] \leq \Pr_{\mathbb{F}}[J_y([v, v+\delta])|Y = y],$$

using the fact that $\mathbb{F}$ has no atoms. Hence, we obtain

$$\Pr[U \in [u-\delta, u+\delta]] = \mathbf{E}_{\mathbb{Y}}[\Pr_{\mathbb{F}}[J_Y([u, u+\delta])|Y]] \leq \mathbf{E}_{\mathbb{Y}}[\Pr_{\mathbb{F}}[J_Y([v, v+\delta])|Y]] = Pr[U \in [v-\delta, v+\delta]].$$

Taking $\delta \to 0$, we conclude the theorem. $\square$

4 Estimation

Suppose we have m independent test repeats per class, $y^{(i),1} \ldots, y^{(i),m}$. Let us define

$$V_{i,j} = \sum_{\ell \neq i} I(\mathcal{M}_i(\hat{F}_1, \ldots, \hat{F}_k, y^{(i,j)}) > \mathcal{M}_\ell(\hat{F}_1, \ldots, \hat{F}_k, y^{(i,j)})),$$

which coincides with the definition (4) in the special case that $\mathcal{M}$ is generative.

At a high level, we have a hierarchical model where U is drawn from a distribution ν on $[0,1]$ and then $V_{i,j} \sim \text{Binomial}(k, U)$. Let us assume that U has a density $\eta(u)$: then the marginal distribution of $V_{i,j}$ can be written

$$\Pr[V_{i,j} = \ell] = \binom{k}{\ell} \int_0^1 u^\ell (1-u)^{k-\ell} \eta(u) du.$$

However, the observed $\{V_{i,j}\}$ do *not* comprise an i.i.d. sample.

We discuss the following three approaches for estimating $p_t = \mathbf{E}[U^{t-1}]$ based on $V_{i,j}$. The first is an extension of *unbiased estimation* based on binomial U-statistics, which is discussed in Section 4.1. The second is the *pseudolikelihood* approach. In problems where the marginal distributions are known, but the dependence structure between variables is unknown, the *pseudolikelihood* is defined as the product of the marginal distributions. For certain problems in time series analysis and spatial statistics, the maximum psuedolikelihood estimator (MPLE) is proved to be consistent [7]. We discuss psuedolikelihood-based approaches in Section 4.2. Thirdly, we note that the high-dimensional theory of Anon 2006 can be applied for prediction accuracy, which we discuss in Section 4.3.

4.1 Extensions of unbiased estimation

If $V \sim \text{Binomial}(k, U)$, then an unbiased estimator of U^t exists if and only if $0 \leq t \leq k$.

The theory of U-statistics [8] provides the minimal variance unbiased estimator for U^t:

$$U^t = \mathbf{E}\left[\binom{V}{t}\binom{k}{t}^{-1}\right].$$

This result can be immediately applied to yield an unbiased estimator of p_t, when $t \leq k$:

$$\hat{p}_t^{UN} = \frac{1}{km} \sum_{i=1}^k \sum_{j=1}^m \binom{V_{i,j}}{t-1}\binom{k}{t-1}^{-1}. \tag{5}$$

However, since $\hat{p}_t^{UN}$ is undefined for $k \geq t$, we can use exponential extrapolation to define an extended estimator $\hat{p}_t^{EXP}$ for $k > t$. Let $\hat{\alpha}$ be a measure defined by solving the optimization problem

$$\text{minimize} \sum_{t=2}^k \left(\hat{p}_t^{UN} - \int_0^\infty \exp[-t\kappa] d\alpha(\kappa)\right)^2.$$

After discretizing the measure $\hat{\alpha}$, we obtain a convex optimization problem which can be solved using non-negative least squares [9]. Then define

$$\hat{p}_t^{EXP} = \begin{cases} \hat{p}_t^{UN} & \text{for } t \leq k, \\ \int_0^\infty \exp[-t\kappa] d\hat{\alpha}(\kappa)) & \text{for } t > k. \end{cases}$$

4.2 Maximum pseudolikelihood

The pseudolikelihood is defined as

$$\ell(\eta) = \sum_{i=1}^k \sum_{j=1}^m \log\left(\int u^{V_{i,j}} (1-u)^{k-V_{i,j}} \eta(u) du\right), \tag{6}$$

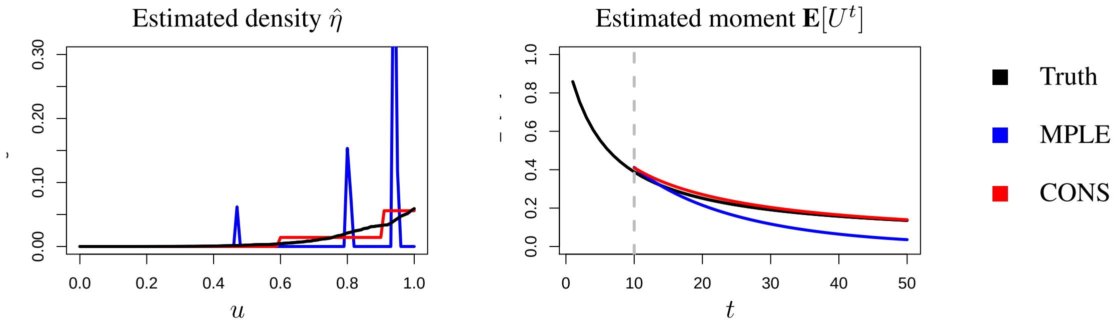

Figure 1: Maximum pseudolikelihood (MPLE) versus constrained pseudolikelihood (CONS). Adding constraints improves the estimation of the density $\eta(u)$, as well as moment estimation.

and a maximum pseudolikelihood estimator (MPLE) is defined as any density $\hat{\eta}$ such that

$$\ell(\hat{\eta}_{MPLE}) = \sup_{\eta} \ell(\eta).$$

The motivation for $\hat{\eta}_{MPLE}$ is that it consistently estimates η in the limit where $k \to \infty$. However, in finite samples, $\hat{\eta}_{MPLE}$ is not uniquely defined, and if we define the plug-in estimator

$$\hat{p}_t^{MPLE} = \int u^{t-1} \hat{\eta}_{MPLE}(u) du,$$

$\hat{p}_t^{MPLE}$ can vary over a large range, depending on which $\hat{\eta} \in \text{argmax}_{\eta} \ell_t(\eta)$ is selected. These shortcomings motivate the adoption of additional constraints on the estimator $\hat{\eta}$.

Theorem 3.4. motivates the *monotonicity constraint* that $\frac{d\hat{\eta}}{du} > 0$. A second constraint is to restrict the k-th moment of $\hat{\eta}$ to match the unbiased estimate. The addition of these constraints yields the constrained PMLE $\hat{\eta}_{CON}$, which is obtained by solving

$$\text{maximize } \ell(\eta) \text{ subject to } \int u^{k-1} \eta(u) du = \hat{p}_k^{UN} \text{ and } \frac{d\hat{\eta}}{du} > 0.$$

By discretizing η, all of the above maximization problems can be solved using a general-purpose convex solver[3]. As seen in Figure 1, the added constraints can improve estimation of η and thus improve moment estimation.

4.3 High-dimensional asymptotics

Under a number of conditions on the distribution $\mathbb{F}$, including (but not limited to) having a large dimension p, Anon et al. [4] relate the accuracy p_t of the Bayes classifier to the mutual information between the label z and the response y:

$$p_t = \bar{\pi}_t(\sqrt{2I(Z;Y)}).$$

where

$$\bar{\pi}_k(c) = \int_{\mathbb{R}} \phi(z-c)\Phi(z)^{k-1} dz.$$

While our goal is not to estimate the mutual information, we note that the results of Anon 2016 imply a relationship between p_k and p_K for the Bayes error under the high-dimensional regime:

$$p_K = \bar{\pi}_K \left(\bar{\pi}_k^{-1}(p_k) \right).$$

Therefore, under the high-dimensional conditions of [4] and assuming that the classifier approximates the Bayes classifier, we naturally obtain the following estimator

$$\hat{p}_t^{HD} = \bar{\pi}_K \left(\bar{\pi}_k^{-1}(\hat{p}_k^{UN}) \right).$$

[3] We found that the CVX discipline convex programming language, using the ECOS second-order cone programming solver, succeeds in optimizing the problems where the dimension of the discretized η is as large as 10,000 [10, 11].

Classifier	Test acc$^{(20)}$	Test acc$^{(400)}$	$\hat{p}_{400}^{EXP}$	$\hat{p}_{400}^{CON}$	$\hat{p}_{400}^{HD}$
Logistic	0.922	0.711	0.844	**0.721**	0.686
SVM	0.860	0.545	0.737	0.575	**0.546**
ϵ-NN	0.880	0.591	0.903	**0.608**	0.839

Figure 2: Performance extrapolation: predicting the error on 400 classes using data from 20 classes on a Telugu character dataset. $\epsilon = 0.002$ for ϵ-nearest neighbors.

5 Results

We applied the methods described in Section 4 to predict the 400-class accuracy of multinomial logistic regression, SVM [6], and ϵ-nearest neighbors[4] on a Telegu character classification task [12], using 20-class data with 100 examples per class. The results are displayed in Figure 2.

Taking the test accuracy on 400 classes (using 50 test examples per class) as a proxy for acc$^{(400)}$, we compare the performance of the three extrapolation methods. The exponential extrapolation method makes use of the fewest theoretical assumptions, but performs badly on all three problems. Meanwhile, constrained PMLE makes an extra assumption in the monotonicity of $\eta(u)$, which is true if the classifier is sufficiently close to the Bayes classifier, and achieves the best and most consistent results. The high-dimensional estimator $\hat{p}^{HD}$ is the most assumption-heavy; in addition to assuming approximation to the Bayes classifier, it also requires Y to be high-dimensional, and to satisfy a number of other technical conditions (Anon 2016). Nevertheless, it performs well on the multinomial logistic and SVM classifiers. That said, all three classifiers studied are non-generative, hence violating an assumption common to both the $\hat{p}^{EXP}$ and $\hat{p}^{CON}$ estimators: therefore, it is doubtful if the theory developed so far can do much to explain the relative performance of these methods.

6 Discussion

We have developed a theory of prediction extrapolation for generative classifiers, under the assumption of exchangeable classes. The equivalence between the expected t-class accuracy and the $t-1$-th moment of the conditional accuracy distribution allows a variety of methods to be applied to the problem. We develop two novel extrapolation methods, and also propose a new application of the mutual information estimator [4] as a third method for prediction extrapolation.

Empirical results indicate that our methods generalize beyond generative classifiers. A possible explanation is that since the Bayes classifier is generative, any classifier which approximates the Bayes classifier is also 'approximately generative.' However, an important caveat is that the classifier must already attain close to the Bayes accuracy on the smaller subset of classes. If the classifier is initially far from the Bayes classifier, and then becomes more accurate as more classes are added, our theory could underestimate the accuracy on the larger subset. This is a non-issue for generative classifiers when the training data per class is fixed, since a generative classifier approximates the Bayes rule if and only if the single-class margin approximates the Bayes optimal single-class margin. On the other hand, for classifiers with built-in *model selection* or *representation learning*, it is expected that the individual class margins become more accurate, in the sense that they better approximate a monotonic function of the class margins of the Bayes classifier, as data from more classes is added.

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