ABSTRACT

Motivation. Accurate knowledge of the genome-wide binding of transcription factors in a particular cell type or under a particular condition is necessary for understanding transcriptional regulation. Using epigenetic data such as histone modification and DNase I accessibility data has been shown to improve motif-based in silico methods for predicting such binding, but this approach has not yet been fully explored.

Results. We describe a probabilistic method for combining one or more tracks of epigenetic data with a standard DNA sequence motif model to improve our ability to identify active transcription factor binding sites (TFBSs). We convert each data type into a position-specific probabilistic prior and combine these priors with a traditional probabilistic motif model to compute a log-posterior odds score. Our experiments, using histone modifications H3K4me1, H3K4me3, H3K9ac and H3K27ac, as well as DNase I sensitivity, show conclusively that the log-posterior odds score consistently outperforms a simple binary filter based on the same data. We also show that our approach performs competitively with a more complex method, CENTIPEDE, and suggest that the relative simplicity of the log-posterior odds scoring method makes it an appealing and very general method for identifying functional TFBSs on the basis of DNA and epigenetic evidence.


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1 INTRODUCTION

Binding of transcription factors (TFs) to DNA is a key step in the regulation of gene expression, strongly influencing the activation or repression of genes. However, accurately locating transcription factor binding sites (TFBSs) in eukaryotic genomes containing gigabases of DNA is challenging because a typical TF binds to relatively short (4–10 bp) DNA sequences that occur at many places throughout the genome, only a small portion of which are actively involved in gene regulation in any particular cell type.

In vivo, several molecular mechanisms limit the binding of TFs to a small fraction of the potential sites. Perhaps most significantly, local chromatin structure, arising from the packaging of DNA, histones and other proteins, occludes access of the TF to many genomic loci. Some aspects of this structure— notably, numerous types of post-translational covalent modifications such as acetylation, methylation, phosphorylation, ubiquitination, ribosylation, glycosylation and sumoylation—affect TF binding in a complex fashion that is not yet completely understood. Empirically, however, numerous studies have shown that TF binding is associated with several types of histone marks (e.g., mono- and tri-methylation of histone H3 lysine 4) (Barski et al., 2007; Cui et al., 2009; Heintzman et al., 2007; Heintzman et al., 2006; Robertson et al., 2007) and with hypersensitivity to cleavage by DNase I (Barnet et al., 2006; Crawford et al., 2006; Hessellberth et al., 2009; Keene et al., 1981; McArthur et al., 2001; Wu, 1980).

In light of these observations, a variety of computational methods have been proposed that aim to improve our ability to identify active TFBSs on the basis of the DNA sequence plus one or more epigenetic experimental assays. Perhaps the simplest such approach scans for occurrences of a specified DNA motif, but deterministically filters out some positions on the basis of the raw signal from a histone modification ChIP-seq assay (Whitington et al., 2009). This type of filter greatly improves the accuracy of TFBS prediction. Several other integrative methods have been proposed for combining DNA motifs with chromatin data (Ernst et al., 2010; Won et al., 2009, 2010). Among these methods, the one most directly comparable to the one proposed here is CENTIPEDE (Pique-Regi et al., 2011), which employs a hierarchical Bayesian mixture model that incorporates information about the DNA sequence, evolutionary conservation, distance to the transcription start site, DNase I accessibility and activating and repressing histone marks.

In this work, we propose a probabilistic method for combining one or more tracks of epigenetic data with a standard DNA sequence motif model to improve our ability to identify active TFBSs. Our approach itself is novel, though it is motivated by similar methods introduced in the context of
motif discovery by the Hartemink group (Gordon et al., 2010; Narlikar et al., 2006), and in the context of TFBS prediction by Lahdesmaki et al. (2008). Critically, our method is applicable to any real-valued data type, requiring the specification of a single hyperparameter with an intuitive interpretation. Second, we describe how to combine a given sequence motif, represented as a position weight matrix (PWM), with the prior to compute a log-posterior odds score. This approach is directly analogous to the most widely used method for scanning for occurrences of TFBSs, and contrasts with a more computationally intensive Markov Chain Monte Carlo-based method proposed previously (Lahdesmaki et al., 2008).

In addition to these methodological contributions, our work suggests three significant conclusions. First, our experiments, using histone modifications H3K4me1, H3K4me3, H3K9ac and H3K27ac, as well as DNase I sensitivity, show conclusively that the log-posterior odds score consistently outperforms a previously described binary filter based on the same data (Whitington et al., 2009). This point is critical, because the simple filter has clear intuitive appeal; we need to begin by demonstrating that our method improves upon this simple baseline. Second, among epigenomic datasets, we find that DNase I sensitivity provides the most value, with only a relatively small additional boost in performance when we also consider histone marks. Third, we evaluate our method using the benchmark dataset described in the CENTIPEDE paper. In this setting, CENTIPEDE performs better on average, but with the relative simplicity of the method, make it an appealing and

2 METHODS

2.1 Creating priors from epigenetic data

We desire a prior distribution on the probability that genomic position \( i \) is bound by any TF in a given tissue. (Throughout the article, we use the terms ‘tissue’ and ‘tissue-specific’ to refer to data from a particular tissue, cell line or condition.) If we have tissue- and position-specific information \( y_i \) for each genomic position, \( 1 \leq i \leq n \), we denote the desired prior distribution by \( \Pr(B_i|y_i) \), where the random variable \( B_i \) is true if and only if any TF is bound at position \( i \). The \( y_i \) can be, for example, tag counts at genomic positions from a histone modification or DNase I assay. We use a heuristic procedure to convert these tag counts to a function that we can use as the desired prior distribution.

To carry out the conversion from tag count to prior, we require a monotonically increasing function \( f(y_i) \), such that \( 0 \leq f(y_i) \leq 1 \) for all values of \( y_i \). Although we use \( f(y_i) \) as a prior distribution, we do not require that it be transformed \( y_i \) values sum to one. Rather, we would like them to sum to any user-chosen estimate, \( \beta > 0 \), of the total number of binding sites of all TFs in the given tissue.

To compute the mapping function \( f(y_i) \), we first map the interval bounded by the minimum and maximum observed values of \( y_i \), \((y_{\min}, y_{\max})\), to the interval \([0,1]\) using the linear function

\[
g(y_i) = \frac{1-a}{y_{\max}-y_{\min}}(y_i-y_{\min}) + a,\]

where \( 0 < a < 1 \). In a moment we will see that the value of \( a \) is not free, but is determined by our choice of \( \beta \). In this work, we employ a linear mapping, but any monotonically increasing function could be used for \( g(y) \).

We then define \( f(y_i) \) in terms of \( g(y_i) \) as

\[
f(y_i) = \frac{\beta}{\sum_j g(y_j)} g(y_i).\]

It is trivial to see that summing \( f(y_i) \) over all genomic positions gives \( \beta \) as desired. Since \( g(y_i) \) is monotonically increasing and \( \beta \) is positive, \( f(y_i) \) is clearly monotonically increasing, as required. We also require that \( f(y_i) \geq 0 \), and we note that this is true because \( g(y_i) \) and \( \beta \) are both positive. To ensure that \( f(y_i) \leq 1 \), we note that the smallest that \( g(y_i) \) can be is \( a \), and the largest it can be is 1, so

\[
f(y_i) = \frac{\beta}{\sum_j g(y_j)} \leq \frac{\beta}{1+\alpha(n-1)}.
\]

Setting the right-hand side above equal to 1 and solving for \( a \) gives

\[
a = \frac{\beta-1}{n-1}.
\]

Using this value for \( a \) ensures that \( f(y_i) \leq 1 \). Note that since it makes no sense to set \( \beta > n \) (because \( \beta \) is the prior estimate of the total number of binding sites in the given tissue, which cannot exceed the number of genomic positions \( n \)), \( a \) will always be \( < 1 \), as required to define the mapping function \( g(y_i) \) above.

2.2 Log-posterior odds motif scoring

2.2.1 Definition

In the previous section, we described a function on epigenetic data, \( f(y_i) \), which we will interpret as \( \Pr(B_i|y_i) \), the prior probability that genomic position \( i \) is occupied by any TF in the tissue of the epigenetic data. Thus, \( \Pr(B_i|y_i) \) represents the ‘general binding propensity’ of genomic region \( x_i \). In this section, we derive a scoring function for predicting the TFBSs of a given TF in a given tissue using a standard PWM TF motif, \( M \), and the position-specific prior function, \( f(y_i) \). This scoring function, which we call the log-posterior odds score, is simply the sum of the traditional log-likelihood ratio motif score, \( S(i) \), and the log-prior odds score, \( P(i) \),

\[
\hat{S}(i) = S(i) + P(i).
\]

The log-prior odds score is defined as

\[
P(i) = \log \left( \frac{f(y_i)}{1-f(y_i)} \right).
\]

The log-likelihood ratio motif score is

\[
S(i) = \sum_j M_{ji},
\]

where \( w \) is the width of the motif and \( b = x_{ij} \) refers to the DNA base in the \( j \)-th position of site \( x_i \).

2.2.2 Derivation

The log-posterior odds score [Equation (3)] is motivated by Bayes decision theory, which states that the best score for discriminating sites from non-sites given an observed sequence and any additional information is the posterior odds (Duda et al., 2001). We will show that the log-posterior odds score defined by Equation (3) is approximately equal to the log-posterior odds that a candidate site for the given TF starts at position \( i \) in the sequence.

\[
\hat{S}(i) \approx \log \left( \frac{\Pr(x_i|y_i, M)}{\Pr(x_i,M)} \right).
\]

Here, the Boolean variable \( x_i \) is true if and only if \( x_i \) is a binding site for the given TF in a given tissue. Note that Boolean variable \( x_i \) differs from \( B_i \), which is true when any TF binds to site \( x_i \) in the given tissue.

We make the simplifying assumption that the sequence data \( x_i \) and epigenetic data \( y_i \) are conditionally independent given the ‘class’ variable
In other words, we assume that once we know whether \( x_i \) is a site or not, knowing \( y_i \) adds no additional information about the particular sequence of bases comprising \( x_i \) and vice versa. Under this assumption, we can write

\[
Pr(x_i,y_i) = Pr(x_i)Pr(y_i|\beta_i) = Pr(x_i,y_i)Pr(\beta_i|y_i)
\]

By applying Bayes' rule twice, and by exploiting Equation (6), we can rewrite the probability of the class being 'true' (e.g. position \( i \) is a site), as

\[
Pr(x_i|y_i) = \frac{Pr(x_i,y_i)Pr(x_i|\beta_i)}{Pr(x_i|\beta_i)} = \frac{Pr(x_i|\beta_i)Pr(y_i|\beta_i)}{Pr(x_i|\beta_i)} = \frac{Pr(x_i|\beta_i)Pr(y_i|\beta_i)}{Pr(x_i|\beta_i)\sum_{\alpha_i}Pr(\alpha_i|\beta_i)Pr(\alpha_i)}
\]

A similar derivation rewrites the denominator of Equation (5), the probability of position \( i \) not being a site, as

\[
Pr(\overline{x}_i|y_i) = \frac{Pr(\overline{x}_i,y_i)Pr(\overline{x}_i|\beta_i)}{Pr(\overline{x}_i|\beta_i)} = \frac{Pr(\overline{x}_i|\beta_i)Pr(y_i|\beta_i)}{Pr(\overline{x}_i|\beta_i)\sum_{\alpha_i}Pr(\alpha_i|\beta_i)Pr(\alpha_i)}
\]

Thus, under the independence assumption above, the log-posterior ratio (the right-hand side of Equation (5)) can be expressed as the sum of a log-likelihood ratio and the log-prior odds

\[
\log \frac{Pr(x_i|\beta_i)}{Pr(\overline{x}_i|\beta_i)} + \log \frac{Pr(y_i|\beta_i)}{Pr(\alpha_i)Pr(\alpha_i|\beta_i)} = \log \frac{Pr(x_i|\beta_i)}{Pr(\overline{x}_i|\beta_i)} + \log \frac{Pr(y_i|\beta_i)}{Pr(\alpha_i)}
\]

by applying Equations (7) and (8) and a bit of algebra. This approach can be easily extended to multiple priors under similar independence assumptions, adding additional log-prior odds terms to the above sum (Section 1 in Supplementary Material).

We estimate the log-likelihood ratio (the first term in Equation (9)) using the standard PWM, \( S(x_i) \). We assume that the binding sites of the given TF are correctly modeled by the position-specific probability matrix \( \Theta_\alpha = [\theta_{ji}] \), where \( \theta_{ji} \) is the probability of DNA base \( c \) at position \( j \) in the motif for motif \( \alpha \). Further assume that non-sites are modeled by a zero-order Markov model \( \Theta_\emptyset = [\rho_i] \), where \( \rho_i \) is the probability of base \( c \) at any non-position. If we let the PWM M have entries \( M_{ji} = \log(\theta_{ji}/\rho_i) \), it is easy to see that

\[
\log \frac{Pr(x_i|\beta_i)}{Pr(\overline{x}_i|\beta_i)} = \sum_{j \in x_i} M_{ji} = S(x_i),
\]

where \( h\cdot x_i \) refers to the DNA base in the \( j \)-th position of site \( x_i \). In our implementation of the log-prior odds score, we actually score both DNA strands independently, and our notation could be easily extended (with some loss of clarity) to include this complication (e.g. \( x^r \) and \( x^i \) to indicate the corresponding sites on the two DNA strands).

We estimate the log-prior odds (second term in Equation (9)) using \( f(y_i) \), our estimate of \( Pr(\beta_i|y_i) \), the prior probability of any TF binding to position \( i \). Recall that the Boolean variable \( b_i \) is true if genomic region \( x_i \) is bound by some TF in a given tissue or cell type or condition. We note that if \( b_i \) is false, then \( x_i \) must also be false, since \( b_i \) being false means that no TF binds region \( x_i \), that is, \( Pr(\overline{x}_i|\beta_i) = 0 \). To estimate \( Pr(\overline{x}_i|\beta_i) \), we make the further simplifying assumption that the epigenetic data \( y_i \) adds no knowledge if we already know the value of the general binding propensity variable \( \beta_i \). That is, we assume that \( Pr(\overline{x}_i|\beta_i) = Pr(\overline{x}_i) \). We further assume that the probability of a position being the site of a given TF, conditional on it being the site of some TF, is a constant. That is, we assume that \( Pr(\overline{x}_i|\beta_i) = \alpha \) for some \( \alpha \geq 1 \).

In any case, with \( \alpha = 1 \) in that equation. In any case, with \( \alpha = 1 \) in that equation.

\[
Pr(y_i|\beta_i) = \frac{Pr(x_i|\beta_i)Pr(y_i|x_i|\beta_i)}{Pr(x_i|\beta_i)} + \frac{Pr(\overline{x}_i|\beta_i)Pr(y_i|\overline{x}_i|\beta_i)}{Pr(\overline{x}_i|\beta_i)} = \frac{Pr(x_i|\beta_i)Pr(y_i|x_i|\beta_i)}{Pr(x_i|\beta_i)} = \frac{Pr(\overline{x}_i|\beta_i)Pr(y_i|\overline{x}_i|\beta_i)}{Pr(\overline{x}_i|\beta_i)} = \alpha Pr(\beta_i)
\]

Typically, the epigenetic data \( y_i \) will not be specific to a given TF, but will rather describe the general binding propensity of genomic region \( x_i \).
2.4 Defining TFBS gold standards based on ChIP-seq data

We used two types of gold standard for evaluating TFBS prediction methods. For a given TF and tissue, each gold standard is created by considering ChIP-seq data for that TF in that tissue as well as a known PWM for the TF. These PWDs are described in Supplementary Table S3.

Our first two experiments are carried out using a ‘peak-centric’ gold standard, which defines a single position in each ChIP-seq peak as a binding site for the TF (a positive), and treats all other genomic positions as negatives. The single positive position assigned to each ChIP-seq peak is the position with the highest PWM score. No positive site is assigned to peaks that do not contain a site with a P < 0.005, as computed by FIMO (Grant et al., 2011). Using this gold standard allows us to compare with the earlier work of Whittington et al. (2009). The ChIP-seq datasets we use for generating the peak-centric gold standards are summarized in Supplementary Table S1.

In our third experiment, to compare our approach against a recently published method, we use the ‘site-centric’ gold standard that was used by Pique-Regi et al. (2011). This gold standard first removes from consideration all genomic positions that are not potential sites, and labels as ‘positive’ and ‘negative’ the sites with strong ChIP-seq evidence for or against occupancy by the TF in the given tissue. Potential sites are those with PWM scores greater than \( \log_2(10000) \) = 13.28 bits. Positive sites are potential sites that occur within a declared ChIP-seq peak. Negative sites are sites that are not within a ChIP-seq peak and the ChIP-seq control signal (tag count) is higher than the ChIP-seq signal. All potential sites that are not clearly positives or negatives according to these two rules are removed from consideration.

2.5 TFBS prediction accuracy metrics

We use two metrics for measuring the accuracy of predictions. Both are based on the true positive rate (TPR) and the false positive rate of a predictor. The TPR, which is often called sensitivity, is the number of true positives divided by the number of known positives at a given score threshold. The FPR is the number of false positives divided by the number of known negatives at a given score.

The first accuracy metric we use is the area under the receiver operating characteristic curve (AUC). This metric gives a combined measure of accuracy at all sensitivity rates, by integrating TPR over all values of FPR (Swets, 1988). The second metric we use is sensitivity at 1% FPR, which allows us to compare the accuracy of prediction methods when a premium is placed on avoiding false positives.

3 RESULTS

3.1 H3K4me3 prior improves TFBS prediction accuracy in mouse ES cells

In our first experiment, we use the log-posterior odds score with the H3K4me3 prior to predict the genome-wide binding of 13 TFs in mouse ES (mES) cells. This experiment tests the hypothesis that our probabilistic prior yields improved TFBS prediction accuracy relative to two competing methods: using no prior at all, or using a ‘hard’ prior, in the form of a site filter based on histone modification ChIP-seq tag counts. We described the latter approach previously (Whittington et al., 2009); accordingly, this experiment uses the same experimental design as that previous work. Specifically, we employ peak-centric gold standards, as described in Section 2, prepared using the TF ChIP-seq datasets listed in Supplementary Table S1. The 13 TFs are CMye, CTCF, E2F4, Eserb, Klf4, Nanog, nMyc, Otx4, Smad1, Sox2, Stat3, Tclcp211 and Zfx.

The results shown in Figure 1 demonstrate that predictions made using the log-posterior odds score are substantially more accurate than those made using the PWM score. The sensitivity at 1% FPR improves from an average of 37% using the PWM score to over 53% using the H3K4me3 prior. As indicated by the small standard error bars, accuracy using the prior is fairly consistent for all 13 TFs, ranging from 77% for Nanog to 98% for CTCF (data not shown). In the Supplementary Material, we show that the log-posterior odds score is also substantially more accurate than the PWM score when we measure the FPR at a given sensitivity (Supplementary Fig. S3 and Table S6). For example, at 20% sensitivity, using the log-posterior odds score reduces the FPR by 68%.

In this experiment, the H3K4me3 log-posterior odds score is also substantially more accurate than our previous method of filtering by histone score and sorting by PWM score (Whittington et al., 2009). The best filtering approach (\( \alpha \geq 2 \)) has a higher sensitivity at only 42%, compared with 53% using the log-posterior odds score (Fig. 1). The H3K4me3 log-posterior odds score is also more accurate than the filtering approach in terms of FPR at sensitivity levels between 20% and 100% (Supplementary Fig. S3).

The H3K4me3 log-posterior odds score also achieves higher overall AUC than either the PWM score or any of the filtering methods tested. (Supplementary Fig. S2). The average AUC for predicting the binding sites of the 13 TFs is 92.4% for using the log-posterior odds score. The best filtering approach we test (\( \alpha \geq 2 \)) has average AUC of 91.5%, and the PWM score achieves 86.6% AUC. The small differences in AUC among the different methods shows that AUC obscures the significant differences at low false positive rates. This is due to the large imbalance in the number of positives and negatives in the peak-centric gold standard causing AUC to emphasize accuracy at high false positive rates.

3.2 Histone modification and DNase I HS priors improve TFBS prediction in human K562 cells

In our second experiment, having established the utility of our log-posterior odds scoring in mice, we perform a systematic evaluation of a variety of priors in human K562 cells. Our aims in this experiment are 2-fold: first, to verify that our proposed method works well with priors derived from a variety of types of histone modifications or from DNase I sensitivity data; second, to examine the extent to which multiple priors can be combined to achieve even better motif search performance. For evaluation, we again
Fig. 2. The log-posterior odds score based on various histone marks and DNase I data improves binding site recognition in human K562 cells. Results are shown for predicting the binding sites of 15 TFs in K562 (human erythroleukaemia) cells. The height of each bar corresponds to the average sensitivity at 1% false positive rate, and error bars indicate standard error. All DNase I hypersensitivity data are from the Stamatoyannopoulos lab at the University of Washington.

use a peak-centric gold standard, this time derived from 15 TFs (Supplementary Table S1). The priors used in this second experiment are based on the histone ChIP-seq and DNase I datasets for K562 cells listed in Supplementary Table S2. The 15 TFs are Atf3, Brg1, cFos, cMyc, Egr1, G4BP, Gata1, Jund, Max, Nfe2, NfyA, NfyB, Sirt6, Usf1 and YY1.

The results of this experiment, summarized in Figure 2, show clearly that log-posterior odds scoring is much more accurate than the PWM score. For example, for 15 TFs the average sensitivity at 1% FPR is 53.7% using the DNase I prior, compared with only 31.3% using PWM scanning. Among the five priors based on single epigenetic datasets used in Figure 2 (DNase I HS, H3K4me1, H3K4me3, H3K9ac and H3K9ac), the DNase I prior is the most informative according to this accuracy metric and according to AUC (Supplementary Fig. S4). Strikingly, at both 20 and 50% sensitivity, the log-posterior odds score with the DNase I HS prior reduces the number of false positives by almost 90% on average when predicting the binding of these 15 TFs (Supplementary Fig. S5). Among the four single-histone priors, H3K27ac and H3K4me3 are the most informative, achieving sensitivities of 49.3 and 48.8%, respectively, at 1% FPR. This result agrees with our previous observation that H3K4me3 is highly predictive of TF binding compared with other histone marks (Whittington et al., 2009). The prior based on H3K4me1 data is much less effective in this experiment, yet still achieves a higher sensitivity rate (38.5%) than PWM scanning (31.5%). These results suggest that DNase I sensitivity data is most useful in predicting TF binding, but that in the absence of DNase I HS data, priors based on H3K27ac, H3K4me3 or even H3K4me1 ChIP-seq data can still greatly increase the accuracy of TFBS prediction when used with the log-posterior odds score.

Finally, we investigate if multiple priors can be combined to achieve even greater prediction accuracy. We observe in Figure 2 that combining DNase I sensitivity data with an informative histone modification (H3K4me3 or H3K27ac) yields an improvement in sensitivity, relative to using any of the three priors alone. For example, the sensitivity of the log-posterior score using a composite prior based on DNase I and H3K4me3 data is 60.3% at 1% FPR, compared with 53.7% using the DNase I prior or 48.8% using the H3K4me3 prior alone.

Somewhat surprisingly, we found that using multiple histone marks to construct priors did not improve accuracy greatly compared to using a single histone prior. For example, a prior constructed from all four of our (non-control) histone marks—H3K4me1, H3K4me3, H3K9ac and H3K27ac—achieves sensitivity of 49.7% at 1% FPR, only marginally better than the accuracy of the prior based on H3K27ac alone (49.3% sensitivity at 1% FPR). However, a prior constructed from H3K9ac and H3K27ac does show moderately improved accuracy (Fig. 2). When accuracy is measured in terms of AUC, the conclusions are the same (Supplementary Fig. S4).

The failure of the combined histone priors to improve over the individual priors may be due to the fact that these three histone marks are highly correlated (Supplementary Fig. S7). On the assumption that the lack of improvement was due to correlated priors not contributing additional information, we also tested a prior combining H3K9ac, which is far less correlated with the other three, with the best histone prior, H3K4me3. However, this prior—H3K4me1 and H3K4me3—performs no better than the H3K4me3 prior alone (Fig. 2).

3.3 DNase I-based scores dramatically improve TFBS prediction in human GM12878 cells

In our third experiment, we compare the log-posterior odds score method to a recently described method called CENTIPEDE (Pique-Regi et al., 2011). Accordingly, this experiment follows the design used by Pique-Regi et al. (2011). Specifically, we use site-centric
Table 1. Performance of five TFBS prediction methods on putative sites with high PWM score

<table>
<thead>
<tr>
<th>TF</th>
<th>Neg</th>
<th>Pos</th>
<th>Area under the ROC curve (%)</th>
<th>Sensitivity at 1% false positive rate</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>PWM</td>
<td>H-p</td>
<td>D-p</td>
<td>Cpd</td>
</tr>
<tr>
<td>CTCF</td>
<td>21320</td>
<td>17093</td>
<td>80.2</td>
<td>82.5</td>
</tr>
<tr>
<td>Galpa</td>
<td>19150</td>
<td>924</td>
<td>61.4</td>
<td>96.4</td>
</tr>
<tr>
<td>Junq</td>
<td>72172</td>
<td>14</td>
<td>82.0</td>
<td>92.7</td>
</tr>
<tr>
<td>Max</td>
<td>13611</td>
<td>72</td>
<td>70.6</td>
<td>93.3</td>
</tr>
<tr>
<td>Nrsf</td>
<td>1642</td>
<td>1193</td>
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<td>88.1</td>
</tr>
<tr>
<td>Srf</td>
<td>1146</td>
<td>133</td>
<td>72.3</td>
<td>90.6</td>
</tr>
</tbody>
</table>

The table shows two measurements of accuracy of tissue-specific TFBS prediction in GM12878 (lymphoblastoid) cells. The first three columns list the transcription factor (TF) and numbers of negative (Neg) and positive (Pos) sites in the site-centric gold standard for that TF. Accuracies (area under the ROC curve and sensitivity at 1% false positive rate) are shown for sorting sites by either the PWM score (PWM), the log-posterior odds using the H3K4me3 prior (H-p), the log-posterior odds using the DNase I prior (D-p), the CENTIPEDE log-posterior odds (Cpd) or the DNase I score (D-s). All DNase I data are from the Crawford lab at Duke University. Best results for each accuracy metric are shown in bold.

gold standards for six TFs in the GM12878 lymphoblastoid cell line. The first three columns of Table 1 list these TFs along with the corresponding numbers of positive and negative sites in the site-centric gold standards. As before, we use as performance metrics the AUC and sensitivity achieved at 1% false positive rate, as used by Pique-Regi et al. (2011). We also explored accuracy in terms of sensitivity at 1% false discovery rate (Supplementary Table S8), with results essentially identical to those described below.

The results of this experiment, summarized in the left half of Table 1, show two very clear trends. First, as in our previous two experiments, using the log-posterior odds score with an epigenetic prior results in higher TFBS prediction accuracy, compared with traditional PWM scoring, with the DNase I prior outperforming the H3K4me3 histone prior. For example, the mean AUC when predicting the binding of these six TFs in GM12878 cells is 75.9% when putative sites are sorted by the PWM score, 90.6% using the H3K4me3 prior and 95.8% when we use the DNase I prior. Using either of these two priors also results in a large increase in the sensitivity achieved at a 1% FPR. On average, the sensitivity more than doubles when we use the DNase I HS prior, compared with using the PWM score (55.3% versus 21.2% sensitivity). Using the H3K4me3 prior, the improvement in sensitivity at 1% FPR is far less (sensitivity = 28.6%). We also tested other histone priors and combinations of histone priors, and the results are very similar (data not shown).

Second, the CENTIPEDE method, which leverages DNase I HS, sequence conservation and distance from TSS data, usually performs better than the log-posterior odds score using only the H3K4me3 or DNase I prior (Table 1). For all six TFs, CENTIPEDE’s AUC is higher than the corresponding AUC for our method using the DNase I prior. Overall, the average difference in AUC is 2.5%, which is relatively large considering that the AUC scores themselves are quite close to 100%. The results with respect to sensitivity at 1% FPR are similar, with CENTIPEDE performing better than the DNase I prior in four out of six cases.

However, while investigating the cause for CENTIPEDE’s strong performance in this experiment, we made a surprising discovery: we are able to achieve performance comparable to CENTIPEDE’s by completely ignoring the PWM and using only the DNase I data.

This strawman method, which we call the ‘DNase I score’, simply sorts all putative sites by the total number of DNase I cuts (tags) in a window of 150 bp around the site. Recall that the site-centric gold standards contain only putative sites with PWM scores greater than log2(10000) = 13.28 bits. Hence, the DNase I score is analogous to the ‘histone filter’ proposed by Whittington et al. (2009), except that rather than filtering by histone tag counts and ranking by PWM score, the DNase I score filters by PWM score and ranks by DNase I tag counts. As can be seen from Table 1 (column ‘D-s’), the DNase I score AUC is higher than CENTIPEDE’s (and than the DNase I prior’s) for five out of six TFs. For the remaining TF (Nrsf), CENTIPEDE achieves a much higher AUC (95.2% compared with 89.3%). In terms of sensitivity at 1% FPR, the DNase I score performs substantially better than CENTIPEDE for three of the six TFs, but is much worse for the other three TFs: Nrsf, CTCF and Jund. These results suggest that the DNase I score method is as accurate as CENTIPEDE in terms of AUC, but less accurate at very low false positive rates.

The above observation—that simply ranking sites based on DNase I sensitivity yields results that are competitive with the best-performing method—suggests that, if we only consider sequences whose PWM scores are relatively high, DNase I sensitivity is more useful than the DNA sequence at discriminating between bound and unbound sequences. It also makes clear a weakness in the site-centric gold standard. Since the site-centric gold standard is based solely on sites with PWM scores exceeding a given threshold, it cannot tell us anything about the accuracy of a prediction method on sites with PWM scores below the threshold. Other types of gold standards, such as the ‘peak-centric’ one used here, must be employed in order to compare the ability of prediction methods on more difficult binding sites with lower PWM scores.

4 DISCUSSION

We have demonstrated that the log-posterior odds score is a powerful method for improving tissue-specific TFBS prediction when genome-wide histone modification ChIP-seq or DNase I hypersensitivity data is available. Compared to a sequence-only scan or to using a hard filter based on histone or DNase I HS tag counts,
the log-posterior odds score offers a large and consistent advantage. Among the datasets that we considered, DNase I sensitivity provides the most useful prior, though some additional benefit can be gained by combining priors from DNase I and an informative histone modification such as H3K4me3 or H3K27ac.

When we compare our proposed method to CENTIPEDE, which leverages additional sources of information such as distance from transcription start sites, the conclusions are less clear. On the one hand, when measured via AUC or sensitivity at 1% FPR, CENTIPEDE appears to provide slightly better results than the log-posterior odds score. On the other, we have shown that, for this particular benchmark, simply ranking sites by their smoothed DNase I tag counts performs comparably to CENTIPEDE. This observation suggests that, among sites with high PWM scores, measurements of DNase I sensitivity are much more useful than the PWM score itself in discriminating between bound and unbound sequences. The experiment also leaves open the possibility that, in a more realistic setting (e.g. using a gold standard that includes all sites in the genome), the relative performance of the methods would be quite different. Finally, it is worth noting that the relative simplicity of the proposed log-posterior odds score makes it an appealing alternative to a more complex approach such as CENTIPEDE.

Although our work has focused on scanning for occurrences of DNA motifs, the log-posterior odds score framework can be trivially extended to protein motifs, given a suitable prior. Such priors might be derived, for example, from the measurements of evolutionary conservation or from the predicted local attributes such as sequence, solvent accessibility, transmembrane topology or propensity to bind to DNA or to other proteins. Similarly, the method described here could be easily incorporated into existing probabilistic approaches for modeling regulatory sequences as cis-regulatory modules composed of multiple motifs (Bailey and Noble, 2003; Sinha et al., 2008; Zhou and Wong, 2004). Although our linear mapping function for constructing epigenetic priors, \( p(y) \), works well in practice, the use of a non-linear function is a promising direction for future research.

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REFERENCES


