

Structures and Assumptions: Strategies to Harness Gene \times Gene and Gene \times Environment Interactions in GWAS

Charles Kooperberg, Michael LeBlanc, James Y. Dai and Indika Rajapakse

Abstract. Genome-wide association studies, in which as many as a million single nucleotide polymorphisms (SNP) are measured on several thousand samples, are quickly becoming a common type of study for identifying genetic factors associated with many phenotypes. There is a strong assumption that interactions between SNPs or genes and interactions between genes and environmental factors substantially contribute to the genetic risk of a disease. Identification of such interactions could potentially lead to increased understanding about disease mechanisms; drug \times gene interactions could have profound applications for personalized medicine; strong interaction effects could be beneficial for risk prediction models. In this paper we provide an overview of different approaches to model interactions, emphasizing approaches that make specific use of the structure of genetic data, and those that make specific modeling assumptions that may (or may not) be reasonable to make. We conclude that to identify interactions it is often necessary to do some selection of SNPs, for example, based on prior hypothesis or marginal significance, but that to identify SNPs that are marginally associated with a disease it may also be useful to consider larger numbers of interactions.

Key words and phrases: Adaptive regression modeling, high dimensional, disease risk, association study.

1. INTRODUCTION

Genome-wide association studies (GWAS), in which as many as a million single nucleotide polymorphisms (SNP) are measured on several thousand samples, are

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quickly becoming common for identifying genetic factors associated with many phenotypes. Until now most analyses of GWAS have taken a one-SNP-at-a-time approach, some analyses are employing haplotypes, but this is mostly as surrogates for unmeasured SNPs. There is, however, the strong assumption that interactions between SNPs or genes (gene \times gene interactions) and interactions between genes and environmental factors (gene \times environment interactions) substantially contribute to the genetic risk of a disease (e.g., Frankel and Stork, 1996; Philips, 2008). Identification of such interactions could potentially lead to increased understanding about disease mechanisms; drug \times gene interactions could have profound applications for personalized medicine; strong interaction effects could be beneficial for risk prediction models.

The GWAS that have been carried out to date have not lead to the identification of many such interactions, other than that some of the SNPs, that were found to

be marginally associated with a disease, are sometimes also tested for differences in subgroups (e.g., Gudbjartsson et al., 2007, Table 2). There are several reasons for this lack of identification of interactions:

1. The odds ratios of SNPs that have main effects are usually small, and there is no reason to expect that interaction effects are any bigger. Increased degrees of freedom for models involving both main effects and interaction effects result in even lower power to specifically identify the interactions than the already low power to identify main effects.
 2. The potential number of gene \times gene interactions is very large, for example, $\sim 5 \times 10^{11}$ two-SNP interactions for the 1 million SNP chip, making it computationally impossible to do anything more than the simplest model for all interactions, and reducing the already limited power further because of the required multiple comparisons correction.
 3. The potential number of gene \times environment interactions is smaller; however, when there are several environmental factors of interest the number of multiple comparisons for which we have to correct is still considerably larger than the number of SNPs in the initial scan. In addition, some environmental factors may have measurement error, further reducing power to identify interactions involving those factors.
 4. With single SNP models there are a variety of (genetic) models to choose from: for example, additive, recessive, dominant and codominant models. For interaction models there are many more, yielding an unwieldy array of models and approaches to choose from, some of which probably could be dismissed as not having the power to identify interactions from the start. An example of such a model with little power is one for k th order interactions, with $k > 2$, where all 3^k possible combinations of SNPs are modeled separately.
 5. Imputation methods for marginal SNP effects work fairly well in identifying disease associated SNPs that are unmeasured, although the power for these unmeasured SNPs is somewhat lower than for measured SNPs; for interactions it becomes infeasible to check all possible interactions involving unmeasured SNPs.
- *If a main effect fits the data well enough, don't use an interaction.* This may seem obvious, and many established function estimation methods adhere to this principle. However, not all models for genetic interactions adhere to this keep-it-simple principle.
 - *Think about which model to use; think how the method scales up.* This is another reason to keep-it-simple. If the model is too computer intensive, it may not be feasible to fit it many millions of times.
 - *If you need to divide the cake, give a slightly larger crumb to the people who will enjoy it—that is, spend your power on the most likely interactions.* Good candidates for SNPs involved in gene \times gene or gene \times environment interactions are prior hypothesized effects (if there are any) and SNPs that show main effects. (Partly) ignoring other possible interactions will eliminate the possibility of identifying these ignored ones as significant; since the power is very small to start out, we may as well use it wisely and at least identify the more likely interactions.
 - *Possibly insignificant interactions may help us to identify disease associated genes.* If there is some difference in genetic effects, looking wisely in subgroups may help you find groups where the effect is strongest.
 - *Be willing to exploit the genetic structure [e.g., linkage (dis)equilibrium, SNPs taking only three values], be willing to make some assumptions (e.g., gene environment independence), but be very aware what you loose if these assumptions are wrong.*

Obviously, many of these principles are not that different from studying interactions in smaller problems: the main difference is that the size of the problem for logistical reasons forces us to make the right choice immediately—we may never get a second chance to correct ourselves.

Furthermore, the nature of potential models of interactions has implications both for interpretation and statistical power. One class of methods uses models leading to traditional odds ratio estimates of main effects and interactions. These methods are based on multiplicative interactions in a logistic regression and can be represented more generally as tensor product regression spline models. On the other hand, another class of strategies address the potential for strong associations within subgroups of subjects; these models include Logic Regression, tree-based regression, haplotype analysis (we describe an adaptive regression strategy called SHARE in Section 3) or adaptively

In this paper we will review a number of approaches that can be used for finding interactions in GWAS. We will use a few guiding principles when using these methods:

weighted subgroup analysis. We note that many methods which focus on subgroup effects should not be interpreted as definitively describing an interaction in the multiplicative sense but rather as tools to increase the potential of finding any association between gene and outcome or as tools for better risk prediction. We make the case, that the choices of interaction strategies which are potentially most powerful depend on the setting, whether it is gene \times gene interaction studies, gene \times environment studies, dimensionality of the SNP or environmental data and hypothesized genetic structure.

The next section starts by giving some general background about methods for interaction modeling in data analysis. In Section 3 we discuss some methods for interactions that have been developed for genetic studies. In Section 3.4 we see how we can do some (limited) identification of interactions in GWAS, and in Section 4 we see how we can use interactions in GWAS to identify SNPs that are marginally associated with a disease. We end with a brief discussion.

2. INTERACTIONS IN STATISTICAL MODELING

Interactions are often characterized by departures from a simple additive combination of effects in the context of some regression model. Such models are of interest in a genetic association study, since one may like to describe instances where a gene is associated with a disease only in the presence of another gene or in combination with an environmental factor. Alternatively, modeling more complex models including interactions can improve function approximations to derived better risk prediction models.

The “Simplest” Interaction Model

To start the discussion, consider a simple interaction model involving two variables. A logit model for a binary disease phenotype $Y \in \{0, 1\}$ is

$$(1) \quad \begin{aligned} \text{logit}[P(Y = 1|\mathbf{X})] \\ = \beta_0 + \beta_1 X_1 + \beta_2 X_2 + \beta_3 X_1 X_2, \end{aligned}$$

where X_1 and X_2 may code genes (SNPs or haplotypes). The final term in the model expresses a departure from a simple additive model, at least on the logit scale. In the setting of genomic association studies, interaction models can also include two classes of variables, for instance, genetic and environmental factors. In that case the model can be represented as

$$\text{logit}[P(Y = 1|\mathbf{X}, \mathbf{Z})] = \beta_0 + \beta_1 Z + \beta_2 X_1 + \beta_3 Z X_1,$$

where Z denotes an environmental variable or some other subject exposure variable, such as assigned treatment. The classic schema of a two variable interaction is given in Figure 1 which shows the odds of disease as a function of two variables. It shows two cases, one where the effect of the second variable is only evident (or largely evident) within a subset of levels of the first variable and a second case, where the effect is actually reversed at one level of the second variable. Sometimes this second case is called a qualitative interaction, because unlike the first one, it does not depend on the choice of the link function to the outcome. Again, the figures would apply to both gene \times gene and gene \times environment interactions. Most would expect the effects shown in the left panel to be the most plausible for

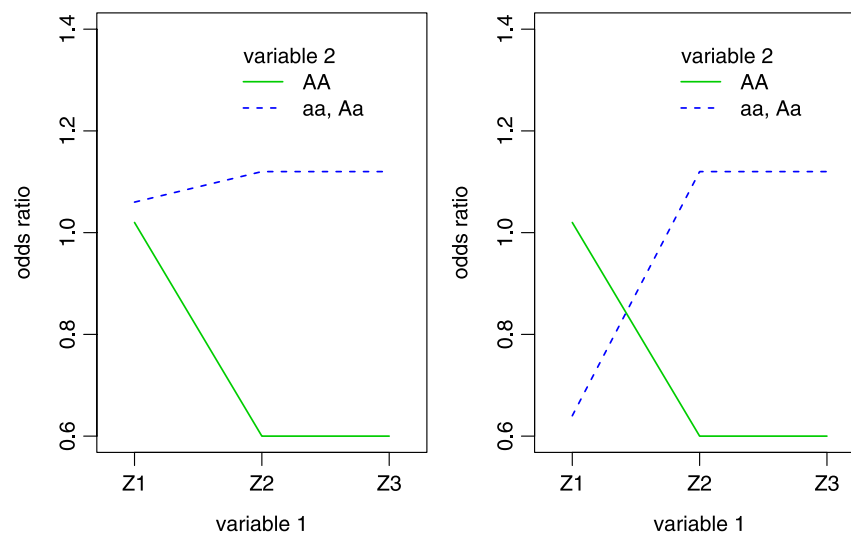


FIG. 1. Two different scenarios of interaction effects. In the left panel disease risk is always higher in the $\{aa, Aa\}$ group; in the right panel the disease risk actually is reversed for variable level Z1 versus Z2 and Z3.

gene by environmental effects in GWAS; that is, effect modification, but not effect reversal. A modest twist for GWAS is that one could envisage 100s of thousands to millions of such interaction plots.

However, the simple model and Figure 1 hide important aspects to the complexity of actually constructing statistical interaction models in several ways: (1) the variables that are involved in the interaction may only be observable as a derived function of several variables; that is, in the models given above, the terms X_i or Z may represent parametric or nonparametric functions of several other variables, for instance, several SNPs representing a gene (2) the system of predictors could be very high dimensional with respect to variable X (for instance, the many SNPs or genes in GWAS) and (3) even after selecting variables, there are other problem specific issues like phasing, measurement error and missingness. For point 1 and more recently for point 2, there is extensive statistical literature addressing function approximation and prediction modeling to draw upon and potentially use, at least in concept, for the analysis of GWAS.

Additive Expansion Interaction Models

The simple interaction model above can be extended to a broader class of models for nonlinear, nonadditive, multivariate regression methods. Assume that the disease model indexed by regression function $\eta(\mathbf{X})$ is in some K -dimensional linear space $\mathcal{B}(\mathcal{X})$,

$$(2) \quad \eta(\mathbf{X}) = \sum_{i=1}^K \beta_i g_i(\mathbf{X})$$

for a given set of basis functions $g_1(\mathbf{X}), \dots, g_p(\mathbf{X})$. Several nonparametric multivariate regression methodologies use this additive expansion (often called a basis function expansion). We review three common function approximation methods that express the basis functions as tensor products of individual covariates. While these methods should probably not be directly applied to whole genome data for both computational and statistical reasons, they can be useful after selecting smaller subsets of variables. More importantly, they follow the common and important paradigm, useful for modeling interactions from GWAS, which involves searching for models sequentially by *first identifying main or (locally) marginal effects before fitting higher order terms*.

(a) *Logistic Regression*. Presented with more than two predictor combinations, several variables may

yield better predictions of the outcome variable. Assuming modeling disease probability is a goal, the expansion model (2) would have component functions $g_i(\mathbf{X})$ which include products of two or more predictors, for example, $g_i(\mathbf{X}) = X_j X_k$. An example model with at most 2-way interactions is

$$\eta(\mathbf{X}) = \beta_0 + \sum_j^p \beta_j X_j + \sum_{j \leq k}^p \beta_{jk} X_j X_k,$$

where $\eta(\mathbf{X})$ may represent the conditional logit of the probability of disease. Even with this simple model form, the number of potential interaction models is order p^2 . If one does not limit the interactions to only involve pairs of variables, there are order 2^p models. The numbers quickly rise with high order interactions and numbers of variables. Therefore, even if we keep the model sparse with only a few β_j and β_{jk} not equal to zero, one would anticipate advantages both in computation and variance control by constructing a reasonable pathway though the model space. A standard approach is to use forward variable selection and consider adding interaction terms if one or more of the variables is already included in the model as a main effect term. The above model is often not sufficiently flexible for prediction modeling especially if X_j are multilevel or continuous. We consider some alternatives below.

(b) *Regression Spline Methods*. One possibility is that the $g_i(\mathbf{X})$ in the expansion model (2) are tensor products of piecewise linear splines. Some examples of methods using such an expansion are Multivariate Adaptive Regression Splines (MARS, Friedman, 1991) and related spline methods Hazard Regression HARE and PolyMARS (e.g., Kooperberg et al., 1997).

Regression spline algorithms exploit lower marginal or additive structure to guide the search for interactions. In addition, these methods use basis functions that are tensor products of basis functions in one dimension. For example, if $g_1(\mathbf{X}) = b_1(X_k)$ and $g_2(\mathbf{X}) = b_2(X_l)$ are two basis functions that depend on a single predictor, the $g_3(\mathbf{X}) = b_1(X_k)b_2(X_l)$ is a tensor product basis function that depends on two predictors. Truncated linear basis splines are used to deal with continuous or ordered covariates $(X_i - t_{k_i})_+ = (X_i - t_{k_i})I\{X_i > t_{k_i}\}$. Given that SNP data only has only 3 categories, piecewise linear components are probably most useful with respect to environmental factors in gene \times environment interaction models. Typically interactions of variables are included only if one or both of the variables are already identified as single variable terms

[e.g., $b_1(X_k)$ or $b_2(X_l)$] described above. This strategy yields more interpretable models since the models contain main effects and it also limits the search over the number of possible models, which better controls variance compared to a search that evaluates all tensor products. The exact restrictions on when tensor product basis functions are allowed in spline models differs from one methodology to the other: for example, MARS (Friedman, 1991) has fewer restrictions than HARE (Koooperberg et al., 1995) and Polyclass and PolyMARS (Koooperberg et al., 1997). All these methods identify lower order effects first to control the search for higher order interaction terms and hence control the variability of the search.

When building more complex models, one may also introduce regularization or penalization to reduce the impact of including too many parameters or cells with very small counts of observations. Park and Hastie (2008) develop an algorithm that constructs penalized regression models for detecting gene environmental interaction models which only includes interactions if one of the main effects are in the model. Three level SNP variables are coded as dummy variables and it uses quadratic penalization to stabilize the estimation.

(c) *Tree Regression.* Regression trees are flexible methods capable of capturing interactions by recursively selecting and partitioning data based on low order associations. For tree based methods, the $g_i(\mathbf{X})$ in the expansion model (2) are indicator functions corresponding to rectangular regions, R_h , of the predictor space, $g_h(X) = I\{X \in R_h\}$. The best known example in the statistical literature is Classification and Regression Trees (CART, Breiman et al., 1984). Therefore, tree models can be represented as a binary tree T , where the set of terminal nodes \tilde{T} corresponds to the partition of the covariate space into disjoint subsets. A tree model can also be expressed by a basis function representation

$$\eta(\mathbf{X}) = \sum_{h \in \tilde{T}} \eta_h g_h(\mathbf{X}),$$

where $g_h(\mathbf{X})$ is the region corresponding to a terminal node h . This function is a tensor product $g_h(X) = I\{X_i \in S_1\} \cdots I\{X_p \in S_p\}$. To control the amount of computation, and to construct a limited path through the large class of potential tensor product interaction models of this form, the model is grown in a forward stepwise fashion, similar to stepwise regression. The assumption used by tree regression is that effects can be found by searching for a *local marginal association* with outcome. The method is applied to the entire

data set and predictor space, each variable and potential split point is evaluated. Of course if the predictors were just SNPs with coding $\{0, 1, 2\}$, then only two splits are possible: one corresponding to a recessive effect and one corresponding to a dominant effect. The split point and variable that leads to the “best” split (as described below) is chosen. The data and the predictor space are partitioned into two groups. The same algorithm is then recursively applied to each of the resulting groups. Therefore, at any point on the regression tree, a split at a node h yields two nodes which can also be represented with the pair of basis functions

$$b_{hj}^+(\mathbf{X}) = I\{X_{h(j)} \in S_{h(j)}\} \quad \text{and} \\ b_{hj}^-(\mathbf{X}) = I\{X_{h(j)} \notin S_{h(j)}\},$$

where $S_{h(j)}$ is a subset of the values of $X_{h(j)}$, leading to terminal nodes basis functions $g_h(X)$ which products of such indicator functions built up at each step. Typically a large tree is grown to avoid missing structure and then pruned back: model complexity is reduced by constructing a backward sequence of models using the cost-complexity pruning algorithm.

(d) *Other Expansion Models.* Of course, other interaction outside the tensor product form of individual predictors can be expressed in this general form using parametric or nonparametric smooth based component functions. For instance, multilayer neural networks construct the $g_i(\mathbf{X})$ as composite function of functions of a linear combinations of subsets of the predictors $g_i(\mathbf{X}) = \phi(\sum_{i=1}^p \alpha_i X_i)$. The regression tree methods above can also be extended to indicator functions based on linear combinations $\{\sum \alpha_i X_i > c\}$. One implementation, FlexTree (Huang et al., 2004), uses this model form.

It is clear that with more than a modest number of predictors the potential number of interaction models is huge and, hence, variance control is critical in the model search.

Bias/Variance Trade-Off: Picking Model Complexity

Stepwise logistic regression, tree-based methods and adaptive regression splines use a forward selection strategy. A final model can be selected to minimize a penalized measure of error,

$$-l_\alpha = -l(\mathcal{M}, \beta; Y_i, \mathbf{X}_i, i = 1, \dots, n) + \alpha |\mathcal{M}|,$$

where $l(\mathcal{M}, \beta; Y_i, \mathbf{X}_i, i = 1, \dots, n)$ is the fitted log-likelihood for a model (of dimension $|\mathcal{M}|$) that was considered, and α is a penalty parameter. Small penalty parameters would lead to large models with limited

bias, but potentially high variance; larger penalty parameters lead to the selection of models biased toward the null model, but with less variance. Given sufficient computation time, a model that minimizes the negative of the cross-validated likelihood (or some resampling analog) may be the preferred method to address the bias versus variance trade-off.

The above penalty only involves the number of parameters; often other common penalties can be useful. Additional terms that penalize either the L^1 norm of the coefficients (e.g., LASSO, Tibshirani, 1996) or the L^2 norm of the coefficients (Ridge Regression, Hoerl and Kennard, 1970) or a linear combination of the two penalty terms (e.g., Elastic Net, Zou and Hastie, 2005) can lead to additional effective ways to control variance.

While these are flexible statistical strategies for interaction model building and regularization, they do not directly incorporate any genomic structure of the predictors. As we describe in the next section, improved inferences, including improved power for testing, can be obtained by incorporating the special form of trinary SNP data 0, 1, 2, the nature of dependence and/or independence between SNPs and environmental variables, and haplotype structures.

3. MODELS FOR INTERACTIONS IN GENETIC STUDIES

Genetic data have a number of special features that can be exploited in modeling interactions. In this section we discuss some of those special features of genetic data, and how they have been used in modeling interactions. Most of these methods cannot deal with all SNPs as are commonly measured in GWAS simultaneously. However, they can be directly used for targeted regions based on either prior biological hypothesis or top hits from an initial single-SNP filtering, as discussed in Section 3.4. It is hard to put the size restrictions of the various methods on one scale. For example, the SHARE method, discussed in Section 3.2, is intended to find interactions *within* a block of SNPs in linkage disequilibrium, and would thus be applied to a fairly small number of SNPs, for instance, 50 tag SNPs between recombination hotspots. Nevertheless, for SHARE it is straightforward to apply it to a GWAS using a “sliding window” approach, where the method is applied to overlapping blocks of SNPs that are close to each other in the genome.

On the other hand, methods like Multifactor Dimensionality Reduction and Logic Regression, discussed

in Section 3.1, are intended to find long-range interactions between a smaller number of SNPs. These methods do not scale up to complete GWAS. However, they could be applied to subsets of SNPs, like candidate gene studies, SNPs in a particular pathway or SNPs that attain a certain (marginal) significance level. The number of SNPs that these methods can deal with is typically up to a few hundred, or maybe a thousand, obviously depending on the sample size (so that the methods have enough power) and the available computing resources (so that sufficiently many models can be examined in a reasonable time). These limitations are quite understandable if we just consider the number of SNPs measured in a GWAS. With, say, one million SNPs, there are 5×10^{11} two-SNP combinations. So, even examining the simplest interaction model for each combination of SNPs is expensive.

These size restrictions are much less severe for identifying gene \times environment interactions. Again, this becomes clear from examining the scale of the problem. Typically we will only be interested in a few environmental factors, thus, the number of potential single SNP \times environment interactions is smaller than the potential number of SNP \times SNP interactions. Thus, some of the ideas on how to identify gene \times environment interactions discussed in Section 3.3 are directly applicable to GWAS.

To find gene \times gene interactions in a GWAS, we have to take a much simpler approach, for example, the two-stage approach discussed in Section 3.4.

3.1 Genetic Data Is “Almost Binary”

Humans carry two copies of each chromosome, and most genetic data comes from typing of Single Nucleotide Polymorphisms (SNPs). SNP data is commonly coded as 0/1/2, indicating the number of minor alleles at a particular locus. If this locus has a dominant effect on a disease phenotype, the genetic factor X can be coded 1 if the SNP is 1 or 2, and 0 otherwise, and if this locus has a recessive effect on a disease phenotype, the genetic factor X can be coded 1 if the SNP is 2, and 0 otherwise. Dealing with binary data is attractive, as models are often easier to interpret, and many computations can be done more efficiently.

This binary coding of genetic data is especially exploited in Logic Regression (Ruczinski et al., 2003). The logic regression model is

$$g[E(Y|\mathbf{X}, \mathbf{Z})] = \beta_0 + \sum_{j=1}^m \beta_j L_j + \sum_k \beta_{k+m} Z_k,$$

where Y is the disease response, \mathbf{X} a vector of recoded SNPs, \mathbf{Z} a vector of other (environmental) covariates, $g(\cdot)$ is a link function, such as the logit, and the L_j are binary combinations of the \mathbf{X} , such as

$$((X_1 \text{ and } X_7^c) \text{ or } X_3).$$

The L_j can be interpreted as risk factors. In Logic Regression model selection is carried out using permutation tests and cross-validation. In particular, conditional permutation tests are used to select simpler models when those fit the data. An alternative approach is to sample Logic Regression models using Markov chain Monte Carlo (Kooperberg and Ruczinski, 2005) or bagging (Schwender and Ickstadt, 2007). The search among candidate models is carried out using a stochastic simulated annealing algorithm, though if the number of SNPs (\mathbf{X}) is limited and the maximum number of SNPs in each L_j were limited to, say, 3, all models could be enumerated.

Logic Regression can be used to find gene \times gene interactions, and gene \times environment interactions for binary environmental variables. For example, in Kooperberg et al. (2007), for an analysis of a candidate gene study of cardiovascular disease among hypertensive, Logic Regression was used to identify the model

$$\begin{aligned} & \text{logit}[P(\text{myocardial infarction}) \\ & \quad \text{AGTR2 SNPs, hypertensive drugs}] \\ & = -0.90 - 0.72[(\geq 1 \text{ A allele at rs171231429}) \\ & \quad \text{and (no calcium channel} \\ & \quad \quad \quad \text{blockers)}]. \end{aligned}$$

As most methods described in this section, Logic Regression does not scale up well to the size of GWAS and needs some selection of SNPs. The stochastic search algorithm could not possibly examine a sufficiently large number of models when there are hundreds of thousands of SNPs. Permutation tests or cross-validation are even more prohibitive.

While Logic Regression reduces SNPs from a 0/1/2 variable to a binary variable, multifactor dimensionality reduction (MDR, Ritchie et al., 2001) makes use of the 0/1/2 nature of the SNP data. For two specific SNPs MDR divides the nine combinations into those that are associated with high and low risk for a particular outcome. Thus, an MDR model for two SNPs may be

		SNP A		
		0	1	2
SNP B	0	H	H	L
	1	L	H	H
	2	L	H	L

where H and L refer to high and low risk, respectively. Three-level interactions are modeled using all 27 possible combinations of three SNPs, and so on. Among all interactions up to a certain level, MDR chooses the best combination by cross-validation. Dividing the SNP combinations in high-risk and low-risk clearly has a close connection to classification trees (see Section 2), but the nonmonotonicity for some SNP combinations (e.g., for SNP B = 2 in the example above) makes the method less regularized and maybe harder to interpret. The cross-validation for MDR does not specifically prefer lower order models over interactions. Thus, for example, a two-level MDR may be chosen, by chance, if in fact there are two main effects. This leads to potentially increasing the type 1 error of incorrectly identifying an interaction (but not increasing the type 1 error of incorrectly identifying a genetic effect). Similarly, three SNP interactions may be identified when a model with two SNP interactions fits the model well. MDR has been applied to a substantial number of candidate gene studies.

3.2 Linkage Disequilibrium

SNPs that are close to each other on the chromosome are typically highly correlated, because of the shared ancestral history. The extent of this correlation or linkage disequilibrium (LD) is known from databases such as the HapMap (HapMap Consortium, 2007) and the 1000 genomes project (Kaiser, 2008) which is currently underway. Known LD structure can be used to impute SNPs that are not measured (e.g., Servin and Stephens, 2007; Marchini et al., 2007). LD can also be used to develop multilocus association methods, often based on haplotype reconstructions (e.g., Browning and Browning, 2007; Epstein et al., 2003; Lin and Zeng, 2006). Statistically, the main effect of a haplotype can be deemed as a combination of main effects and interactions in a locus (SNP) regression model (Schaid, 2001). Using haplotypes can be an effective way to model interactions between multiple mutations within a gene. The latter perspective may receive an increasing appreciation as genome-wide sequencing technologies hold promise to directly capture the rare variants in the near future. Several recent candidate gene studies suggested that accumulation of multiple rare alleles may contribute to the risk of some common diseases (Vermeire et al., 2002; Cohen et al., 2004; Nejentsev et al., 2009). Theoretic population genetic studies also support the presence of multiple deleterious susceptibility loci with low frequencies (Pritchard, 2001, Kryukov, Pennacchio and Sunyaev, 2007). To

this end, a haplotype analysis can be useful to assemble the cumulative, possibly interactive, effect of rare variants within a gene.

Certainly locus-based models, such as the stepwise regression method of Cordell and Clayton (2002), can be used to model the local interactions directly without having to deal with haplotype phase ambiguity. However, with the high density of SNP data currently being collected, haplotype phase ambiguity is less a concern for power. Furthermore, the main effect of a haplotype already contains interactions when casted in locus (SNP)-based models, and hence tend to be more parsimonious than locus-based models that may require high-order interaction terms. In situations where the LD is strong, additional power gain can be achieved by fewer parameters used to characterize the genetic risk profile. For instance, consider a situation with 2 SNPs in complete LD and hence 3 haplotypes (00, 10, 11). A logistic regression with additive main effects and interactions uses 3 parameters, whereas a logistic regression with additive haplotype main effects uses one less parameter. With more rare, recent variants (with less opportunity of recombination) discovered through sequencing, it is anticipated that haplotype-based models will be more cost-effective than locus-based models in modeling such local interactions.

When many SNPs in a region (gene) are under investigation, the number of haplotypes constructed from all SNPs can be excessively large. A haplotype scan using a moving window of 5 to 10 SNPs may miss the interactions between SNPs separated in the region, since the 3-D structure of a protein can bring amino acids further apart into one functional domain. Strategies to perform model selection seem inevitable in order to characterize relevant genetic variants. Dai et al. (2009) propose the SNP-Haplotype Adaptive REgression (SHARE) algorithm that seeks the most informative set of SNPs for genetic association in a targeted region by growing and shrinking haplotypes with one more or less SNP in a stepwise fashion. Though it is not always “optimal,” the stepwise selection is a rational choice given the computational demand facing a large number of SNPs and the fact that haplotypes we observe today were formed by sequential (stepwise) mutations in history. It is hard to imagine that there is no marginal effect for a haplotype carrying disease risk. Contrary to the popular haplotype clustering approaches (e.g., Seltman et al., 2001; Durrant et al., 2004; Browning and Browning, 2007), in the SHARE algorithm both the trait and the genotypes guide the model selection process, and the SNP selection is irrespective of the order of the SNPs in the region (gene). Cross-validation

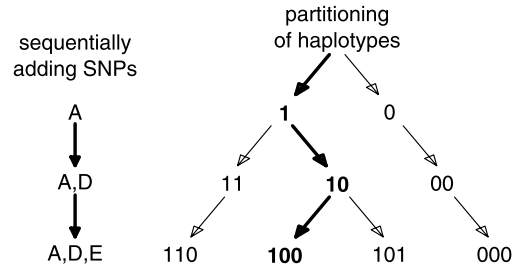


FIG. 2. Tree illustration of the sequential partitioning of haplotypes when 5 SNPs A–E are present. The left panel shows the growing set of SNPs used in the analysis and the right panel shows the partitions resulted from the haplotypes based on the current set of SNPs. The minimal set of SNPs that captures the genetic association is (A, D, E), with the disease risk concentrated on the haplotype “100”. The path leading to discovering it is 1 → 10 → 100. The order of SNPs in the haplotypes follows A, AD and ADE respectively.

is used to select the best set of SNPs for a haplotype analysis, and phase ambiguity is accounted for by treating haplotype estimation as a part of the procedure.

Despite the resemblance to the stepwise logistic regression of Cordell and Clayton (2002), SHARE is actually more closely related to the Classification and Regression Tree (CART, Breiman et al., 1984) algorithm (see Section 2) in that it recursively partitions the haplotype sample space. Figure 2 shows a simple example of the pathway of partition. Among 5 SNPs in a region, we first find the SNP A as the most significant SNP by a single-locus scan, so haplotypes are partitioned by “0” and “1” at the A locus. Next, haplotypes constructed by A and D are found to be the best 2-SNP haplotypes, and finally, the algorithm reaches the most informative set {A, D, E} so that one 3-SNP haplotype concentrates the disease risk. To increase the chance of finding the most informative set, we grow longer haplotypes and then prune back one SNP at a time. We use cross-validation to select the partition with the minimal prediction error. While CART is effective to dissect high-order interactions, growing haplotypes is essentially refining high-order interactions between loci. The difference between SHARE and CART is that the recursive partitioning of CART is binary, while SHARE potentially creates multiple splits when adding one SNP, if there is recombination. Moreover, every subject has two sets of haplotypes and we need some genetic model to describe the combined haplotype effect, such as additive, dominant or recessive genetic models.

3.3 Independence Assumptions

When it is known that pairs of SNPs or SNPs and environmental factors are independent of each other

in the population, this knowledge can be effectively used to identify interactions. Noticing that these independent quantities are not independent in a particular group of subjects (e.g., cases of a particular disease) now implies an interaction effect. This approach has been used to develop methods for SNPs that are (assumed to be) in linkage equilibrium, and for non-genetic environmental factors.

In randomized clinical trials treatment assignment is made independent of the genetic status of the participants. In some other situations it may also be reasonable to assume gene \times environment independence for some environmental exposures. If for a case-control study this independence holds for the controls (e.g., under the usual rare disease assumption), it is immediate that a significant correlation between a genetic effect and the environmental factor among cases implies a gene \times environment interaction on the disease outcome. This is the basic idea behind the case-only analysis (Albert et al., 2001; Piegorsch, Weinberg and Taylor, 1994; Umbach and Weinberg, 1997): there is a simple relation between the odds-ratio among the cases and the parameter for the interaction in a linear logistic regression model. Recently a number of methods have been proposed to exploit this independence also for estimating main effects, and to avoid having to make the rare disease assumption (Chatterjee and Carroll, 2005; Dai et al., 2009).

It has been pointed out that violation of the gene-environment independence assumption can seriously increase the type 1 error (Albert et al., 2001). An empirical Bayes approach that “averages” the analysis assuming independence and a traditional case-control analysis has been proposed as a save alternative (Mukherjee and Chatterjee, 2008) that maintains some of the advantage of the independence assumption, when that assumption cannot be fully confirmed. In this approach the effect estimate under the (unbiased) case-control design is averaged with the (more efficient) estimate using the case only analysis, with weights balancing the variance of the case-control estimator, and an empirical Bayes estimate of the uncertainty of independence assumption.

Clearly, testing whether two factors are independent (e.g., gene and environment among the controls) and then using a test with or without assuming independence, depending on the result of the independence test, will seriously inflate the type 1 error. On the other hand, if a preliminary test is independent of the final test for interaction in the data set, such a test can be used to prioritize potential models that are tested,

thereby alleviating the multiple comparisons problem. This is the approach taken by Millstein et al. (2006). In this paper the authors first test for (gene-gene) independence in the complete data set of cases and controls. Situations where there is substantial deviation of independence are prioritized for testing for interaction effects. The motivation seems to be that if two genes are dependent, this dependence may very well be different between cases and controls, for example, but not necessary, because the genes are independent among the controls but not among the cases. As case-control status is not used in this preliminary analysis, the eventual analysis for interactions only needs to be corrected for the interactions actually tested, which increases power considerably.

Linkage equilibrium. In a homogeneous population SNPs that are on different chromosomes or far apart on the same chromosome are (approximately) independent. This suggests an alternative way to identify interactions. For “rare diseases,” this population-wise independence implies that for controls the SNPs should be independent. Simple properties of log-linear models show that dependence between two SNPs among the cases now implies an interaction effect of those SNPs on case-control status. Zhao et al. (2006) used this property to develop a test for interactions by reconstructing haplotypes (implicitly assuming Hardy-Weinberg equilibrium) between two unlinked loci (SNPs) to get a measure of LD between these SNPs. Assuming no LD among the controls, they developed a (asymptotically χ^2) test for an interaction effect of these two SNPs on a disease outcome.

In practice, it is unlikely that a population is completely homogeneous. Thus, it may be dangerous to assume that SNPs are indeed in linkage equilibrium among the controls. Given the large number of SNPs that are typically tested, a small amount of correlation will already inflate the type 1 error rate. It is, however, a valid test of interaction in genetic association studies to test whether the correlation between two SNPs is the same among the cases and the controls. In fact, Zhao et al. (2006) also provide a test for interaction using this approach. However, the examples in their paper that compare their approach to logistic regression (which does not use an independence assumption) make this independence assumption.

Rajapakse, Perlman and Kooperberg (2009) generalizes the approach by Zhao et al. (2006) by constructing a correlation matrix between groups of SNPs using the generalized or composite LD of Weir (1996), separately for cases and controls. The advantage of using

TABLE 1
Comparison of *P*-values for testing gene \times gene interactions

Method	No independence assumed			Independence assumed	
	Logistic regression	Zhao et al. (2006)	Rajapakse, Perlman and Kooperberg (2009)	Zhao et al. (2006)	Rajapakse, Perlman and Kooperberg (2009)
TP53 \times CD14	0.0418	0.0325	0.0364	0.1062	0.0210
TNFR1 \times APOC3	0.0001	0.0005	0.0005	0.0009	0.0000
TP53 \times MDM2	0.0123	0.0715	0.0767	0.8607	0.0836

the generalized LD over other measures of LD is that, since phase information is not required, no haplotype reconstruction is needed, and under certain conditions the generalized LD between two SNPs reduces to the correlation between these SNPs when coded as 0/1/2. Using this approach, a simple test for identity between the correlation between SNPs for the cases and controls becomes a test of the interaction effect of these SNPs on case-control status. There are a number of advantages for this approach. (i) This test of independence can easily be extended to blocks of SNPs. An interaction between two different blocks of SNPs on case-control status might suggest that a haplotype, that may be a surrogate for an unmeasured SNP in the first block, and a haplotype, that may be a surrogate for an unmeasured SNP in the second block, have an interaction effect on case-control status. Therefore, this method is an alternative to the method of Chatterjee et al. (2006) discussed in Section 4. (ii) A test for identity between the complete correlation matrix for the cases and the controls is a global test for interactions among the SNPs considered on case-control status. (iii) Assumed independence between selected SNPs among the controls can easily be incorporated into this procedure by setting elements of the correlation matrix for the controls equal to 0. When only two SNPs are examined, and no independence assumption is made, we would expect the methods of Zhao et al. (2006) and Rajapakse, Perlman and Kooperberg (2009) to give similar results, and that the results would be similar to the four degree-of-freedom test of comparing

$$\begin{aligned} & \text{logit}[P(Y = 1|X_i, X_j)] \\ &= \beta_0 + \beta_1 I(X_i = 1) + \beta_2 I(X_i = 2) \\ &+ \beta_3 I(X_j = 1) + \beta_4 I(X_j = 2) \\ &+ \beta_5 I(X_i = 1)I(X_j = 1) \\ &+ \beta_6 I(X_i = 2)I(X_j = 2) \\ &+ \beta_7 I(X_i = 1)I(X_j = 1) \\ &+ \beta_8 I(X_i = 2)I(X_j = 2) \end{aligned}$$

and

$$\begin{aligned} & \text{logit}[P(Y = 1|X_i, X_j)] \\ &= \beta_0 + \beta_1 I(X_i = 1) + \beta_2 I(X_i = 2) \\ &+ \beta_3 I(X_j = 1) + \beta_4 I(X_j = 2). \end{aligned}$$

In the implementation of Rajapakse, Perlman and Kooperberg (2009) these tests use the Kullback–Leibler distance between two matrices. In Table 1 we provide results on a previously analyzed case-control study consisting of 779 heart disease patients, 342 of whom showed restenosis, and 437 who did not (Hoh et al., 2001; Kooperberg and Ruczinski, 2005). All individuals were genotyped for 89 SNPs/variants in 62 genes that were previously associated with heart disease. We show results for three of the two-SNP interactions that were identified in Table III of Kooperberg and Ruczinski (2005) (the other four interactions in this table involved a variant that had no homozygotic minor allele subjects). The significance level for the methods of Zhao et al. (2006) and Rajapakse, Perlman and Kooperberg (2009) are based on 10,000 permutations. We note that, as expected, the three methods give similar results when there is no independence assumption. When we do make an independence assumption, the approach of Zhao et al. (2006) appears less powerful than the one of Rajapakse, Perlman and Kooperberg (2009), which does not require a haplotype reconstruction. The approach of Rajapakse, Perlman and Kooperberg (2009) offers the additional advantage of potential for extension to tests of interaction effects between blocks of SNPs.

3.4 Using Main Effects to Find Interactions in GWAS

The methods discussed above exploit the genetic structure of the data to be analyzed. However, mostly those approaches do not scale up to GWAS, both because methods become computationally too demanding, and because the number of multiple comparisons

becomes so large that the power to identify significant interactions for anything other than the strongest effects is missing. Interestingly, testing all models that include an interaction for the combined effect of a particular SNP on a disease outcome can increase the power to identify an individual SNP as being associated with a disease outcome (Marchini et al., 2005; Evans et al., 2006), but it does not increase the power to identify an interaction. We discuss this approach further in Section 4.

If only a few environmental factors are examined, the problem to identify simple multiplicative gene \times environment interactions is essentially the same as studying marginal effects. Thus, while power is limited, just like for any GWAS study, computationally studying gene \times environment interactions is straightforward. However, the filtering procedures that we suggest below for gene \times gene interactions can increase the power to identify gene \times environment interactions in GWAS as well.

While it is probably obvious that enumerating all interactions in a GWAS will be computationally too expensive, except for the simplest possible models involving just 2-SNP interactions, adaptive search algorithms do not circumvent these problems. Adaptive algorithms can be roughly divided in those using stochastic search algorithms and greedy search algorithms. Stochastic search algorithms, like the simulated annealing algorithm used by Logic Regression, search a stochastically selected set of models, selecting the best fitting model(s) among those examined. For well structured model classes these algorithms can avoid looking at many poorly fitting models. However, to have a reasonable opportunity to find good fitting models, the number of models examined needs to increase considerably. In particular, since in humans the extent of linkage disequilibrium is small compared to the length of the genome (r^2 is typically much smaller than 0.8 after less than 50 kb; Pritchard and Przeworski, 2001), the number of models that are examined needs to go up with close to the number of SNPs, and likely even more if higher order interactions are studied to find good interactions. Greedy search algorithms, for example, the stepwise and tree algorithms described in Section 2, are much more likely to end up in local optimal models, that are globally not very good, when the number of predictors increases.

For these reasons, the most viable solutions to extend the methods discussed above to GWAS would be to select SNPs based on some marginal criterion, and only search for interactions among the selected SNPs. It is clear that this approach reduces the computational

requirements. The two main questions are, however, as follows:

- does a filtering procedure alleviate the multiple comparisons problem?
- are we able to identify the “important” interactions?

As Marchini et al. (2005) point out, the multiple comparisons correction for testing interactions after marginal filtering needs to take the filtering into account. The “safe” approach is to correct the number of tests, for example, using a family-wide error rate (FWER) or a false discovery rate (FDR) approach for the number of interactions that *could* have been examined. Clearly, with such an approach the power to identify significant interactions cannot be larger than when all possible interactions *would* have been examined (but at reduced computational cost). This was part of what was found by Marchini et al. (2005) and Evans et al. (2006). We should note here that, besides that this is computationally infeasible, there is no simple permutation test for (the strongest) interaction effect, as a simple permutation of case-control status not only removes the interaction effect, but also removes all main effects. Such a permutation test would thus be a test of main effect combined with interaction—a topic which we discuss in Section 4.

Kooperberg and LeBlanc (2008) establish that if the marginal testing of SNPs is carried out using regression models of the form

$$\gamma_0 + \gamma_1 X_i,$$

for some coding X_i of a SNP i and the interaction model examined is of the form

$$(3) \quad \beta_0 + \beta_1 X_i + \beta_2 X_j + \beta_3 X_i X_j,$$

then the least squares estimates of γ_1 and β_3 are independent. While the estimation in case-control studies is typically carried out using logistic regression, this independence result gives some justification of only correcting for the number of tests that actually were examined, for example, using a Bonferoni approach. Kooperberg and LeBlanc (2008) also develop a resampling procedure based on scores, extending a technique proposed by Lin (2006), that offers an alternative to permutation tests, and is applicable to two stage studies in which only SNPs that are marginally significant are tested for interactions. In this approach a sample from the efficient score for the interaction model (3) is generated under the null hypothesis that $\beta_3 = 0$.

In particular, the efficient score for the addition of an interaction $X_i X_j$, conditional on X_i and X_j already being in the model, is $U_{ij} = \sum_k U_{ijk}$, where

$U_{ijk} = (Y_k - p_{ijk})(X_{ik}X_{jk} - \mu_{ijk})$. Here Y_k is case-control status for subject k , p_{ijk} is the fitted probability of subject k being a case in a logistic regression model using X_i and X_j , but not X_iX_j , as predictor for Y , and μ_{ijk} is the fitted value for subject k in the linear regression model using X_i and X_j as predictor for X_iX_j . Under the null hypothesis of no association, U_{ij} is approximately normal with mean 0 and variance $V_{ij} = \sum_k U_{ijk}^2$. In computing the significance level, we compare $T = \max_{ij} U_{ij}^2/V_{ij}$ with $T^* = \max_{ij} (\sum_k U_{ijk} Z_k)^2/V_{ij}$, where the Z_k are independent standard normal random variables. This approach does not assume independence of the stage one and two tests, as the Bonferoni approach does, but rather the “permutations” are carried out conditional on the results of the first stage.

A simple routine for computing power calculations for two stage tests of interaction is implemented in the CRAN package `powerGWASinteraction`. Kooperberg and LeBlanc (2008) contain extensive simulation studies establishing that the two-stage approach indeed maintains the correct type 1 error rate, that the power in most reasonable situations is vastly improved over a one-stage analysis, and that this power is well approximated by the routines from `powerGWASinteraction`.

Clearly, this approach can also be applied to identify gene \times environment interactions: now only one SNP needs to be marginally significant to be tested as part for a gene \times environment interaction. In Table 2 we present power calculations for identifying a gene \times environment interaction. We assume model (3) with X_i a binary SNP with $P(X_i = 1) = 0.4375$, which corresponds to a dominant SNP effect for a SNP with mi-

nor allele frequency 0.25, a binary environmental factor X_j with $P(X_j = 1) = 0.5$, a case-control study with 5000 cases and controls, 500,000 SNPs, $\beta_0 = -2$ (not a rare disease), $\beta_1 = 0$ (no genetic effect when $X_j = 0$), $\beta_2 = 0.5$ (a moderate environmental effect), and an overall multiple-comparisons controlled significance level $\alpha = 0.05$. We show results for several gene \times environment interaction effects β_3 , and several levels for the marginal level of significance α_1 that a SNP has to satisfy before it is tested for the gene \times environment interaction. Besides power to identify the interaction using a regular analysis, we also show power for an analysis that assumes that the gene and the environmental factor are independent.

We see from Table 2 that a two-stage procedure increases the power considerably. If only the top 500 SNPs are tested for gene \times environment interactions ($\alpha_1 = 0.0001$), the power to identify an interaction with odds ratio 1.4 is over 70%. If the gene and environmental factor are assumed to be independent, and we analyze the data using, for example, a case-only analysis or the approach of Dai et al. (2009), the power increases to about 90%.

We applied the two-stage approach to the WTCCC Crohn’s disease data (WTCCC, 2007). We identified 211 SNPs that had minor allele frequency ≥ 0.1 , less than 5% missing data and marginal significance level $p < 0.0001$, and did not grossly violate Hardy Weinberg Equilibrium. Among the $\binom{211}{2} = 22155$ two-SNP interactions among these SNPs, three had a q -value < 0.05 Storey and Tibshirani (FDR, 2003), and are thus plausible. Two of these interactions involve SNPs on different chromosomes, the third one involves two SNPs relatively close together on the same chromosome. Nineteen more possible interactions have q -values < 0.25 , suggesting that more than ten of those

TABLE 2
Sample power calculations for gene \times environment interactions. The numbers 0.00001, ..., 1 are the first stage significance α_1

β_3	OR	0.00001	0.0001	0.001	0.01	0.1	1
Not assuming gene–environment independence							
0.262	1.3	0.21	0.26	0.18	0.10	0.04	0.02
0.336	1.4	0.66	0.72	0.56	0.37	0.21	0.11
0.405	1.5	0.94	0.94	0.85	0.69	0.51	0.34
Assuming gene–environment independence							
0.262	1.3	0.29	0.45	0.52	0.47	0.33	0.19
0.336	1.4	0.73	0.87	0.93	0.89	0.79	0.65
0.405	1.5	0.95	0.99	1.00	0.99	0.97	0.94

show some reproducible association with Crohn's disease.

If the effect of a SNP goes in the opposite direction for two levels of another SNP or an environmental variable, a two-stage analysis may have less power than a one-stage analysis (if these opposite effects just cancel each other out). We believe that some of the more unusual interaction effects considered in Evans et al. (2006) are less likely, and we believe that using the power for more likely scenarios is a potentially more fruitful way of α -spending.

4. USING INTERACTIONS TO FIND MAIN EFFECTS IN GWAS

Modeling gene \times environment or gene \times gene interactions is useful even if the goal of the analysis is primarily the identification of simple gene-disease associations. For instance, if not acknowledged in the analysis method, interactions can lead to attenuation of the marginal effect size and reduce the power to detect true associations. Several authors have incorporated interactions into their search of marginal genetic association. A common thread of successful methods is that they allow model flexibility, but not so much model flexibility as to substantially increase variance.

For instance, in the simple two variable model (1) in Section 2, one can test the overall association of X_1 with disease outcome by testing the null hypothesis $H_0 : \beta_1 = \beta_3 = 0$ using a 2 degree of freedom test. Exploiting potential interactions to detect genetic association with this simple testing technique has been explored by Kraft et al. (2007). In a less directed fashion, Marchini et al. (2005) tested both the main effects and interactions to explain the 3×3 table of two SNPs to assess individual associations.

Extensions to multiple predictors can substantially increase the potential number of parameters. For instance, consider two sets of variables X_{1i} and X_{2j} , which could represent two sets of SNPs or SNPs and environmental variables. For assessing the association of a given X_{1i} with outcome, one can simultaneously test the $1 + q$ terms β_{1i} and β_{ij} (if there are multiple X_{1i} , $i = 1, \dots, p$, testing involves $p + pq$ terms) in the model

$$\eta(\mathbf{X}) = \beta_0 + \sum_{i=1}^p \beta_{1i} X_{1i} + \sum_{j=1}^q \beta_{2j} X_{2j} + \sum_{i,j} \beta_{ij} X_{1i} X_{2j},$$

where the components X_{2j} could represent other SNPs or environmental factors. The difficulty is that as q

increases, the potential power of the test may significantly decrease due to the increased number of parameters. One way to limit model complexity is to specify a restricted form for the interaction model such as

$$\eta(\mathbf{X}) = \beta_0 + \sum_{i=1}^p \beta_{1i} X_{1i} + \sum_{j=1}^q \beta_{2j} X_{2j} + \theta \sum_{i,j} \beta_{1i} \beta_{2j} X_{1i} X_{2j}.$$

This class of models is used by Chatterjee et al. (2006), and relates to the idea of a one-degree of freedom interaction test, dating back to Tukey (1949). Suppose a gene is identified by p SNPs X_{1i} , then a hypothesis of no association with outcome for this set X_{1i} could be phrased as $H_0 : \beta_{1i} = 0$, $i = 1, \dots, p$, where this indicates no association through main effects or interactions since the β_{1i} also appear in the interaction term. The strategy assesses overall variable importance of a gene (potentially represented by several X_{1i}) in the context of a more general model which could include interactions. For instance, one could use any regularized and or stepwise model building strategy (e.g., regression trees or regression splines described in Section 2, or ensembles of such models) and evaluate the impact of removing the gene of interest from the model. One technique to measure the importance of the variable is to evaluate the difference in the fit or log-likelihood compared to the fit with the gene permuted with respect to all other variables (e.g., Breiman, 2001). However, this last strategy is likely not computationally feasible in the context of GWAS.

An alternative idea is to modify simple gene association test statistics by weighting them to take advantage of an interaction and increase power of the test statistic. For instance, one could focus on subgroups of subjects to test for genetic association. In addition, if the procedure was computational efficient, it could be an alternative to fitting full interactions with maximum likelihood methods.

Since many useful association tests are score type statistics, one can outline the method in terms of weighted score test statistics. Let Z denote an environment or treatment variable and X_j a genetic factor. For example, in the case of binary outcome data, let X_{ji} be the gene j value and Z_{ki} environmental factor k for individual i , the score component would be $U_{ji} = X_{ji}(Y_i - \exp(\alpha + \beta Z_{ki})) / (1 + (\alpha + \beta Z_{ki}))$. If the association is thought to be stronger in a subgroup of subjects (e.g., heavier smokers) based on some environmental factors, then a subgroup weighted marginal

test statistic

$$U_W(Z, X_j; \theta) = \sum_{i=1}^n h(Z_i, \theta) U_{ji},$$

where $h(Z; \theta) = I\{Z > \theta\}$ and Z is an ordered environmental variable, may have more power. The generalization to multiple or alternative basis functions

$$U_W(Z, X_j; \theta) = \sum_{i=1}^n \sum_{k=1}^q \alpha_k h(Z_{ik}, \theta_k) U_{ji}$$

allows more flexibility. The basis functions, $h(Z_{ik}, \theta_k)$, could be simple subset functions, $h(Z_{ik}; \theta) = I\{Z_{ik} > c_k\}$, or piecewise linear functions, $h(Z_{ik}; \theta) = \{Z_{ik} - c_k\}^+$, similar to those used in tree-based models and regression spline models described in Section 2. As the direction of association is usually unknown, the weights α_k can be derived from the data. LeBlanc and Kooperberg (2009) obtain weights using stage-wise regression based on the Least Angle Regression algorithm (LARS, Efron et al., 2004), where the score components are the outcome variable. The intent is to focus the test statistic on the environmental combination which leads to maximal genetic association. They show that weighting the association test statistics can significantly increase power in many simulated situations where gene \times environment interactions exist. As an example, Figure 3 shows simulations for 2000 cases and 2000 controls generated from the logistic interaction regression model

$$\eta(Z, X) = \beta_0 + \beta_1 X + \beta_2 h(Z) + \beta_3 X h(Z),$$

where we assume a binary SNP, X with frequency equal to 0.2, and $h(Z)$ a function of several environmental variables. There were five environmental variables corresponding to the basis functions $(1, \{Z_j < c_{0.25}\}, \{Z_j < c_{0.50}\}, \{Z_j < c_{0.75}\})_{j=1, \dots, 5}$, available for modeling and $h(Z)$ depended on the linear combination of two of those variables: $h(Z) = \{Z_1 < c_{0.50}\} - \{Z_2 < c_{0.50}\} + 0.25$. We evaluated a marginal test of the $1 + k$, $k = 3 \times 5$ parameters (including main gene effect and the modifying variables), and a regularized stage-wise test. The type I error was controlled to approximately 0.00001. The results are presented in Figure 3. The parameter values main effects are $\beta_1 = 0.05$ and $\beta_2 = 0$ so that the small genetic effect increases as $h(Z)$ and β_3 increase. This would be a plausible scenario for effects in a GWAS, a genetic effect that is more apparent within a subgroup of subjects exposed to a set of environmental conditions. This model is similar to the hypothetical effects shown in the left panel of Figure 1.

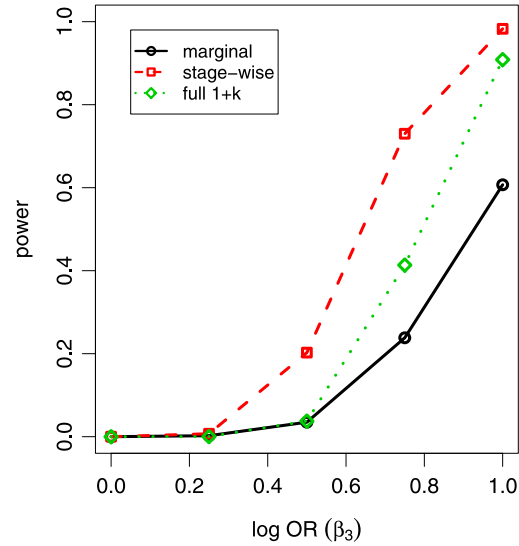


FIG. 3. Using interactions in tests of association: Power for marginal test, full $1 + k$ parameter association test, and stage-wise weighted test with 2000 cases and 2000 controls and $\alpha = 0.00001$. The full $1 + k$ parameter test is based on testing the main effect and all interaction terms in a logistic regression model and the stage-wise test is based on using weighted score test statistics where the weights are derived from the Least Angle Regression (LARS) algorithm. The main effect parameters in the data generating model are $\beta_1 = 0.05$ and $\beta_2 = 0$.

As the magnitude of the interaction effect increases, using the more complex model and joint association testing substantially increases power over marginal testing. The full model weighting, depending on 16 basis functions and parameters, suffers somewhat from increased variance. However, the stage-wise test statistic performs the best of the three methods by controlling the overall variance. The tuning parameter for the stage-wise method was set to correspond to approximately 2.5 degrees of freedom.

Therefore, if there are one or a small number of well characterized environmental factors that are substantially modifying the association of a gene with disease, statistical strategies which incorporate interactions, and jointly test the main effect and the interaction, are useful for improving power over marginal association tests.

While the adaptive weighting strategy is more computationally demanding than calculating traditional score test statistics, it is feasible to conduct the analysis on the GWAS because each SNP calculation is independent and the Least Angle Regression algorithm using a small number of environmental variables is very efficient, if the tuning parameter is set a priori as we suggest. However, the impact on variance is potentially a greater concern, there is still likely be some

power advantage of filtering on main effects to, say, a small number of 100–1000s, before applying the adaptive method analogous strategy described in Section 3.

5. DISCUSSION

Identifying interactions is typically not the main goal of a GWAS analysis. Interaction effects may teach us things about the biology behind a disease, or they may be beneficial in constructing prediction models. However, at least as important, interaction effects or differences in the gene (SNP) effect between different subgroups may actually help us in identifying the significant SNPs. Therefore, we believe that it is important to pay some attention to the identification of interactions.

Because of the number of SNPs under consideration in a typical GWAS, it is virtually impossible to identify gene \times gene interaction effects, unless additional assumptions are being made. We believe that the most fruitful approach is to first identify SNPs that are (marginally) associated with a disease, and then examine interactions involving those SNPs. Not only does this seem reasonable because SNPs that have an interaction effect typically will also show some modest main effect, it also adheres to a basic premise in statistical modeling which reduces the variance in model building: don't model interactions without main effects.

After such initial filtering, there is a substantial number of approaches that can be used to identify interactions that make use of the specific form of genetic data. This is a reasonable two-stage approach if the methods to identify interactions are used to independent data than what was used to identify the marginal significant SNPs. If the same data is used, however, care has to be taken that the initial selection of SNPs does not bias the inference about the interactions. We have shown that for a simple model this is possible—but this is certainly not generally true.

The story for gene \times environment interactions is similar. The problem of identifying such interactions is “smaller,” but it is still so large that some filtering will often increase the power.

Exploiting the genetic structures and making additional assumptions, like gene \times gene independence among genes on different chromosomes, among controls, or gene \times environment independence, can substantially increase the power to identify interactions. Clearly, however, if the assumptions are not true, making those assumptions can substantially increase the type 1 error. Thus, if those assumptions are uncertain, an empirical Bayes approach like the one by Mukherjee and Chatterjee (2008) (see Section 3) may be safer.

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