Bias-corrected GEE estimation and smooth-threshold GEE variable selection for single-index models with clustered data

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Abstract

In this paper, we present a generalized estimating equations based estimation approach and a variable selection procedure for single-index models when the observed data are clustered. Unlike the case of independent observations, bias-correction is necessary when general working correlation matrices are used in the estimating equations. Our variable selection procedure based on smooth-threshold estimating equations (Ueki, 2009) can automatically eliminate irrelevant parameters by setting them as zeros and is computationally simpler than alternative approaches based on shrinkage penalty. The resulting estimator consistently identifies the significant variables in the index, even when the working correlation matrix is misspecified. The asymptotic property of the estimator is the same whether or not the nonzero parameters are known (in both cases we use the same estimating equations), thus achieving the oracle property in the sense of Fan and Li (2001). The finite sample properties of the estimator are illustrated by some simulation examples, as well as a real data application.

Keywords and phrases: Generalized estimating equation; Longitudinal data; Oracle property; Single-index model; Variable selection.

Short title: Variable selection for SIM
1 Introduction

Many data sets nowadays are characterized by two properties that make their statistical analysis complicated, high-dimensionality and dependence of observations. In fact, clustered data with a medium to large number of covariates are often produced in fields such as biology, engineering, or medicine. For different clusters $1 \leq i \leq n$, let $Y_i = (Y_{i1}, \ldots, Y_{imi})^T$ denote the vector of outcome values, which depends on a $p \times m_i$ covariate matrix $X_i = (X_{i1}, \ldots, X_{imi})$, $X_{ij} = (X_{ij1}, \ldots, X_{ijp})^T$. When the dimension of $X_{ij}$ is high, it is worthwhile to spend efforts in seeking a more parsimonious representation of the regression function in the hope of making estimation feasible for moderate sample size. Dimension reduction is one way towards this goal.

As a popular instantiation of dimensional reduction idea, the single-index model for the clustered data is defined by

$$Y_i = g(X_i^T \beta) + \varepsilon_i, i = 1, 2, \ldots, n,$$

where

$$g(X_i^T \beta) = \begin{pmatrix} g(X_{i1}^T \beta) \\ \vdots \\ g(X_{imi}^T \beta) \end{pmatrix}, \varepsilon_i = \begin{pmatrix} \varepsilon_{i1} \\ \vdots \\ \varepsilon_{imi} \end{pmatrix}.$$  

Here $g(\cdot)$ is an unknown link function and $\varepsilon_i$ is mean-zero random error with covariance matrix $Var(\varepsilon_i) = \Sigma_i$ for the $i$th subject, and $\beta = (\beta_1, \ldots, \beta_p)^T$ is the unknown parameters for the index associated with covariates. Since both $g$ and $\beta$ are unknown, it is commonly assumed that $\| \beta \| = 1$ for identifiability, where $\| \cdot \|$ is the Euclidean norm. The true value of $\beta$ will be denoted by $\beta_0$. Throughout this paper we assume that the total sample size $N = \sum_{i=1}^n m_i$ is large (diverges to $\infty$ in our theoretical investigations) while $\{m_i, i = 1, \ldots, n\}$ are uniformly bounded.

The popularity of the semiparametric single-index model presented above can be attributed to its ability to address the so-called “curse of dimensionality” problem in multi-dimensional nonparametric regression by making use of a combination of predictors as univariate index, which hopefully can still capture some im-
portant relationships between the covariates and the responses. As a dimension
reduction method, single-index models have been studied extensively. See for ex-
ample, Ichimura (1993); Härdle et al. (1993); Carroll et al. (1997); Xia et al. (1999);
Naik and Tsai (2000, 2001, 2004); Yu and Ruppert (2002); Delecroix et al. (2003);
Zhu and Xue (2006); Xia and Härdle (2006); Kong and Xia (2007); Wong et al. (2008).
More recently, Bai et al. (2009) studied the single-index model for longitudinal data,
and proposed to use splines to estimate $\beta$ and the unknown link function based on
quadratic inference functions. Our study here is different from that work in many
respects. Bai et al. (2009) considered asymptotic analysis with a fixed number of
knots and thus their analysis is not appropriate when the true link function is not
inside the spline space. In particular, their asymptotic analysis is only for a para-
metric model since the number of unknown parameters does not diverge with sample
size. Our estimation method and asymptotic analysis does not pose this constraint,
and treat the unknown link function as a truly nonparametric component. Further-
more, we will consider variable selection problem which was not investigated before
for single-index models on longitudinal data.

Even though single-index models avoid the problem of “curse of dimensionality”
to some extent, in practice, one would still want to investigate which covariates
are relevant for prediction, both for better interpretation of the model, and for
better efficiency of the estimator. In recent years, penalization or shrinkage based
variable selection methods have attracted lots of attention, due to their computa-
tional efficiency for high-dimensional problems, and their statistical stability com-
pared to information criterion based methods (Fan and Li, 2001; Zou, 2006). Ex-
amples of shrinkage estimation methods include LASSO (Tibshirani, 1996), SCAD
(Fan and Li, 2001), Adaptive Lasso (Zou, 2006), Dantzig selector (Candes and Tao,
2007), and many others. For single-index models, Naik and Tsai (2001) considers
variable selection using sliced inverse regression, Kong and Xia (2007) uses cross-
validation to select the significant variables, but these estimators are not expected
to have the oracle property (Fan and Li, 2001).

In this paper, we build on the estimating equations based approach for single-index models (Chang et al., 2010), which was shown to result in a more efficient estimator for the index vector, and extend it to the case where data are clustered. The bias-corrected estimating equations we use here were proposed in Li et al. (2010), which focused on the construction of confidence regions of partially linear single-index models for longitudinal data through the empirical likelihood method. Furthermore, variable selection is achieved by extending the smooth-threshold estimating equations proposed in Ueki (2009). Compared to shrinkage methods reviewed above, this approach dispenses with convex optimization and is thus computationally simpler. We will theoretically demonstrate the oracle property of the estimator as well as empirically illustrate its performance. We also note that recently Cui et al. (2011) has extended the estimating equations approach to generalized single-index models which do not involve clustered data. We expect that this can also be extended to the case with variable selection for clustered data, although this is outside the scope of the current paper.

The rest of the paper is organized as follows. In Section 2 we present our estimation approach for single-index models with clustered data, and in Section 3 a variable selection procedure based on smooth-threshold generalized estimating equations is presented. The oracle property for the proposed estimator is also discussed. In Section 4, we report some simulation studies as well as an application to a real data set. Our simulations show the advantage of incorporating the intra-cluster correlation in estimation. The proofs of theoretic results are presented in the Appendix.

2 Bias-corrected GEE estimation

In model (1.1), we imposed \( \| \beta \| = 1 \) for identifiability, which implies that the parameter is not an interior point of the \( p \)-dimensional space, causing some difficulty in inference. We use the “remove one component” method used previ-
ously in Yu and Ruppert (2002); Zhu and Xue (2006); Chang et al. (2010). Without loss of generality, we assume that for some $1 \leq r \leq p$, $\beta_r > 0$. Let $\beta^{(r)} = (\beta_1, \ldots, \beta_{r-1}, \beta_{r+1}, \ldots, \beta_p)^T$ be the $(p - 1)$-dimensional parameter vector after removing the $r$th component $\beta_r$ of $\beta$. Then, we may write

$\beta(\beta^{(r)}) = (\beta_1, \ldots, \beta_{r-1}, (1 - \|\beta^{(r)}\|^2)^{1/2}, \beta_{r+1}, \ldots, \beta_p)^T$.

Since $\|\beta^{(r)}_0\| < 1$, $\beta(\cdot)$ is infinitely differentiable in a neighborhood of $\beta^{(r)}_0$, and the Jacobian is

$J_{\beta^{(r)}} = \frac{\partial \beta}{\partial \beta^{(r)}} = (b_1, \ldots, b_p)^T$,

where $b_s$ is a $(p - 1)$-dimensional unit vector with $s$th component 1 for $s \neq r$, and $b_r = -(1 - \|\beta^{(r)}\|^2)^{-1/2}\beta^{(r)}$.

Based on these notations, we construct the generalized estimating equation (GEE) for the single-index model with clustered data as

$$\sum_{j=1}^{n} Z_j^T R_j^{-1} (Y_j - g(X_j^T \beta)) = 0,$$

(2.1)

where

$Z_j = \begin{pmatrix} g'(X_{j1}^T \beta)(J_{\beta^{(r)}_j}^T X_{j1})^T \\ \vdots \\ g'(X_{jm_j}^T \beta)(J_{\beta^{(r)}_j}^T X_{jm_j})^T \end{pmatrix}$, $j = 1, \ldots, n$,

and $R_j, j = 1, \ldots, n$ are the working covariance matrices, possibly depending on some unknown parameter $\alpha$, which can be estimated by the method of Liang and Zeger (1986). From the estimating equations, we can see that if $R_j = I_{m_j}$, with $I_{m_j}$ the $m_j \times m_j$ identity matrix, we just ignore the dependence of the data within a cluster, that is, assume working independence (Lin and Carroll, 2000). For the following theoretical results, we do not require $R_j$ to be the same as the true covariance $\Sigma_j$, although $R_j = \Sigma_j$ results in the most efficient estimator.

The estimating equation (2.1) contains the unknown functions $g(\cdot)$ and $g'(\cdot)$. To solve this problem, we need to plug in some estimates for these two unknown
functions. Here we use the local linear regression (Fan and Gijbels, 1996). Similar to Chang et al. (2010), for any given $\beta$, we can estimate $g(t)$ and $g'(t)$ by minimizing

$$
\min_{a,b} \sum_{i=1}^{n} \sum_{j=1}^{m_i} \{Y_{ij} - a - b(X_{ij}^T \beta - t)\}^2 K_h(X_{ij}^T \beta - t),
$$

where $K$ is a kernel function, $K_h(\cdot) = K(\cdot/h)/h$ and $h$ is the bandwidth. Let $(\hat{a}, \hat{b})$ be the minimizers and set $\hat{g}(t, \beta) = \hat{a}$ and $\hat{g}'(t, \beta) = \hat{b}$. Simple and standard calculations yield the closed form expression

$$
\hat{g}(t, \beta) = \sum_{i=1}^{n} \sum_{j=1}^{m_i} W_{nij}(t, \beta) Y_{ij} \quad \text{and} \quad \hat{g}'(t, \beta) = \sum_{i=1}^{n} \sum_{j=1}^{m_i} \tilde{W}_{nij}(t, \beta) Y_{ij},
$$

(2.2)

where

$$
W_{nij}(t, \beta) = \frac{U_{nij}(t, \beta)}{\sum_{i=1}^{n} \sum_{j=1}^{m_i} U_{nij}(t, \beta)}, \quad \tilde{W}_{nij}(t, \beta) = \frac{\tilde{U}_{nij}(t, \beta)}{\sum_{i=1}^{n} \sum_{j=1}^{m_i} U_{nij}(t, \beta)},
$$

$$
U_{nij}(t, \beta) = K_h(X_{ij}^T \beta - t)\{S_{n,2}(t, \beta) - (X_{ij}^T \beta - t)S_{n,1}(t, \beta)\}, \quad \tilde{U}_{nij}(t, \beta) = K_h(X_{ij}^T \beta - t)\{(X_{ij}^T \beta - t)S_{n,0}(t, \beta) - S_{n,1}(t, \beta)\},
$$

and

$$
S_{n,l}(t, \beta) = \frac{1}{N} \sum_{i=1}^{n} \sum_{j=1}^{m_i} (X_{ij}^T \beta - t)^l K_h(X_{ij}^T \beta - t), \quad l = 0, 1, 2.
$$

Plugging these estimators into (2.1), we obtain the estimating equations

$$
\sum_{j=1}^{n} \hat{Z}_j^T R_j^{-1} (Y_j - \hat{g}(X_j^T \beta)) = 0,
$$

(2.3)

where

$$
\hat{Z}_j = \begin{bmatrix}
\hat{g}'(X_{j1}^T \beta)(J_{\beta(c)}^T X_{j1})^T \\
\vdots \\
\hat{g}'(X_{jm_j}^T \beta)(J_{\beta(c)}^T X_{jm_j})^T
\end{bmatrix}, \quad j = 1, \ldots, n.
$$

We can also obtain an initial estimator of $\beta$, denoted by $\tilde{\beta}$, by assuming working independence. When assuming working independence, the results in Wang et al.
apply with few changes, and in particular, \( \tilde{\beta} \) is \( \sqrt{n} \)-consistent under standard assumptions.

For our theoretical analysis, we will assume that \( R_1, \ldots, R_n \) are prespecified and known. We briefly discuss the more general case where \( R_i \) must be estimated in Remark 1 below. However, when we do not assume that \( R_j, j = 1, \ldots, n \), are all equal, similar to Wang et al. (2010), (2.3) leads to

\[
\sum_{j=1}^{n} \hat{Z}_j^T R_j^{-1} (Y_j - \hat{g}(X_j^T \beta)) = U_1(\beta_0^{(r)}) - nV_1(\tilde{\beta}^{(r)}) - U_2(\tilde{\beta}^{(r)}) + o_p(\sqrt{n}),
\]

where

\[
U_1(\beta_0^{(r)}) = \sum_{k=1}^{n} (g'(X_k^T \beta_0) J_{\beta_0}^T(\beta_0) X_k - E[X_k^T \beta_0]) \right)^T R_k^{-1} \varepsilon_k
\]

\[
V_1 = \lim_{n \to \infty} \frac{1}{n} \sum_{k=1}^{n} E \left[ (g'(X_k^T \beta_0) J_{\beta_0}^T(\beta_0) X_k \right)^T R_k^{-1} (g'(X_k^T \beta_0) J_{\beta_0}^T(\beta_0) X_k)]
\]

and

\[
U_{2s}(\beta_0^{(r)}) = \sum_{k=1}^{n} \sum_{i=1}^{m_k} \sum_{j=1}^{m_k} \left( \sum_{l_1=1}^{m_{i1}} \sum_{l_2=1}^{m_{i2}} W_{nkj}(X_{l_1 l_2 k}^T, \beta_0) g'(X_{l_1 l_2 k}^T, \beta_0) X_{l_1 l_2 k}^T R_k^{-1} \varepsilon_{kj} \right)
\]

\[
\text{if } R_k, k = 1, \ldots, n \text{ are not equal to each other, the arguments contained in Wang et al. (2010) that show the term } U_{2s}(\beta_0^{(r)}) \text{ is asymptotically negligible do not apply, and thus we cannot show the asymptotic normality of } \tilde{\beta}. \text{ Therefore, instead of GEE (2.3), we incorporate bias correction which was previously used in Li et al. (2010), leading to the bias-corrected GEE}
\]

\[
\sum_{j=1}^{n} \hat{Z}_j^T R_j^{-1} (Y_j - \hat{g}(X_j^T \beta)) = 0, \quad (2.4)
\]
where
\[ \hat{\beta}_j = \left( \begin{array}{c} g'(X_{j1}^T\beta)(J_{\beta}^T(X_{j1} - \hat{\beta}[X_{j1}|X_{j1}^T\beta]))^T \\ \vdots \\ g'(X_{jm_j}^T\beta)(J_{\beta}^T(X_{jm_j} - \hat{\beta}[X_{j1}|X_{j1}^T\beta]))^T \end{array} \right), j = 1, \ldots, n, \]

and \( \hat{E}(X_{jk}|X_{jk}^T\hat{\beta}) \) is a nonparametric estimate of \( E(X_{jk}|X_{jk}^T\beta_0) \) with \( \beta_0 \) replaced by the initial estimator \( \hat{\beta} \), that is
\[ \hat{E}(X_{jk}|X_{jk}^T\beta) = \sum_{l_1=1}^{m_j} \sum_{l_2=1}^{m_j} W_{nl_1l_2}(X_{jk}^T\beta)X_{jlq}. \]

In the following, and also in the proofs in the Appendix, with misuse of notation but for simplicity in writing, we will write the matrix such as
\[ \left( \begin{array}{c} g'(X_{j1}^T\beta)(J_{\beta}^T(X_{j1} - E[X_{j1}|X_{j1}^T\beta]))^T \\ \vdots \\ g'(X_{jm_j}^T\beta)(J_{\beta}^T(X_{jm_j} - E[X_{jm_j}|X_{jm_j}^T\beta]))^T \end{array} \right) \]
simply as \( g'(X_{j}^T\beta)J_{\beta}^T(X_j - E[X_j|X_j^T\beta]) \) and take \( X_j - E(X_j|X_j^T\beta) \) to denote the \( m_j \times p \) matrix with entries \( X_{jlq} - E[X_{jlq}|X_{jlq}^T\beta], 1 \leq l \leq m_j, 1 \leq q \leq p. \)

Denote the solution of (2.4) by \( \hat{\beta}_s^{(r)} \) (the notations \( \hat{\beta} \) and \( \hat{\beta}_s^{(r)} \) are reserved for the estimator based on smooth-threshold generalized estimating equations later when we deal with variable selection), thus our final estimator for \( \beta \) is \( \hat{\beta}_s = \beta(\hat{\beta}_s^{(r)}). \) We have the following asymptotic property for \( \hat{\beta}_s. \)

**Theorem 1.** Under the regularity conditions given in the Appendix, and suppose the initial estimator \( \hat{\beta} \) is \( \sqrt{n} \)-consistent, then there exists a solution \( \hat{\beta}_s \) of (2.4) inside the ball \( B = \{ \| \beta - \beta_0 \| \leq Cn^{-1/2} \} \) for \( C \) sufficiently large. Furthermore,
\[ \sqrt{n}(\hat{\beta}_* - \beta_0) \xrightarrow{d} N(0, \Sigma_a), \]
where
\[ \Sigma_a = J_{\hat{\beta}_0} V^{-1} \Omega V^{-1} J_{\hat{\beta}_0}^T. \]
The matrices \( V \) and \( \Omega \) are defined in condition C7 of the Appendix.
Remark 1. We have assumed that \( R_i \) are prespecified and known in the above. However, from the proof, one easily sees that when \( R_i \) is replaced by a consistent estimator \( \hat{R}_i \), the theorem still holds. When \( \Sigma_1 = \cdots = \Sigma_n \), a consistent estimator of \( \Sigma_i \) is \( \sum_{j=1}^{n} \hat{\varepsilon}_j \hat{\varepsilon}_j^T / n \), where

\[
\hat{\varepsilon}_j = (Y_{j1} - \tilde{g}(X_{j1}^T \tilde{\beta}), \ldots, Y_{jm_j} - \tilde{g}(X_{jm_j}^T \tilde{\beta}))^T,
\]

with \( \tilde{g} \) and \( \tilde{\beta} \) obtained from the working independence assumption \( \text{Balan and Schiopu-Kratina, 2003} \). Alternatively, when \( R_i, 1 \leq i \leq n \) depend on some fixed parameter \( \alpha \), moments-based method can be used to estimate \( \alpha \) consistently, resulting in consistent estimator of \( R_i \) \( \text{Liang and Zeger, 1986} \).

Remark 2. When \( R_1 = \cdots = R_n = R \), it is not necessary to use bias-corrected GEE \( (2.4) \). In particular, when using GEE \( (2.3) \), Lemma A.7 in \( \text{Wang et al., 2016} \) can be followed line by line (with the extra simplification that we are dealing with single-index models instead of partially linear single-index models in that paper) to show that \( \hat{\beta}_* \) is asymptotically normal with covariance matrix \( \Sigma_b = J_{\beta_0^{(r)}} V_1^{-1} \Omega V_1^{-1} J_{\beta_0^{(r)}}^T \), where

\[
V_1 = \lim_{n \to \infty} \frac{1}{n} \sum_{k=1}^{n} E \left[ (g'((X_k^T \beta_0) J_{\beta_0^{(r)}}^T X_k))^T R_k^{-1} (g'((X_k^T \beta_0) J_{\beta_0^{(r)}}^T X_k))^T \right].
\]

It is obvious that \( V_1 \geq V \) (i.e. \( V_1 - V \) is nonnegative definite) and thus \( \Sigma_b \leq \Sigma_a \), which means estimator obtained from \( (2.3) \) is more efficient than that obtained from \( (2.4) \). However, theoretically, using bias-correction leads to simpler assumptions on the bandwidth. In particular, unlike the theoretical results presented in \( \text{Chang et al., 2016} \), we do not need to use different bandwidths when estimating \( g \) and \( g' \) if \( (2.4) \) is used. In our simulation results, our experience is that empirically the difference between using \( (2.3) \) and \( (2.4) \) is very small and thus we only report the simulation results based on bias-corrected GEE only. When \( R_i \) are not all equal, the original proof in \( \text{Wang et al., 2016} \) fall through and this is the reason for proposing \( (2.4) \) to make our presentation much more general and work in all cases.
3 Variable selection and the oracle property

So far in our discussions, all the covariates are assumed to be important for predicting \( Y \). However, in many practical situations, some covariate variables are independent of or have negligible correlations with the response variable. As mentioned in the introduction, many shrinkage based approaches have been proposed in the literature to solve this variable selection problem, most of which are based on penalty functions with a singularity at zero. As an alternative method, Ueki (2009) proposed smooth-threshold estimating equations (SEE). This method is easily implemented with Newton-Raphson type algorithms, which is almost the same as solving the original estimating equations under the full model.

Let \( A = \{1, 2, \ldots, p\} \) be the index set for the components of \( \beta \). We make the sparsity assumption that some components of \( \beta_0 \) are zeros and without loss of generality assume the first \( p_0 \) components are nonzero and let \( A_0 = \{1, 2, \ldots, p_0\} \), and thus \( A_0^c \) contains all the indices of the zero components. Following Ueki (2009), we propose the following smooth-threshold generalized estimating equations (SGEE) for simultaneous variable selection and estimation,

\[
(I_{p-1} - \hat{D}) \sum_{j=1}^{n} Z_j^0 R_j^{-1} (Y_j - \hat{g}(X_j^T \beta)) + \hat{D} \beta^{(r)} = 0, \tag{3.1}
\]

where

\[
\hat{Z}_j^0 = \begin{pmatrix} 
\hat{g}'(X_{j1}^T \beta)(J_{\beta^{(r)}}^T[j_{\beta^{(r)}}][X_{j1} - \hat{E}(X_{j1}|X_{j1}^T \beta)])^T \\
\vdots \\
\hat{g}'(X_{jm}^T \beta)(J_{\beta^{(r)}}^T[j_{\beta^{(r)}}][X_{jm} - \hat{E}(X_{jm}|X_{jm}^T \beta)])^T 
\end{pmatrix}, j = 1, \ldots, n;
\]

\[
\hat{D} = \begin{pmatrix} 
\hat{\delta}_1 & & \\
& \ddots & \\
& & \hat{\delta}_{r-1} & 0 \\
& & \hat{\delta}_r & \ddots \\
& & 0 & \ddots \\
& & & 0 & \ddots \\
& & & & \hat{\delta}_p 
\end{pmatrix},
\]
with $\hat{\delta}_i = \min(1, \frac{\lambda}{|\tilde{\beta}_i|^{1+\gamma}}), i = 1, \ldots, p, i \neq r$; and $\tilde{\beta}$ is the initial $\sqrt{n}$-consistent estimator as before. The estimate of $\beta$ obtained from SGEE (3.1) is denoted by $\hat{\beta}$ and the set of estimated nonzero indices is $\hat{A} = \{i : \hat{\beta}_i \neq 0\}$.

From (3.1), we see that $\hat{\delta}_i = 1$ implies $\hat{\beta}_i = 0$, while if $\hat{\delta}_i$ is negligibly close to zero, then (3.1) is similar to (2.4). The choice $\hat{\delta}_i = \min(1, \frac{\lambda}{|\tilde{\beta}_i|^{1+\gamma}})$, proposed in Ueki (2009), satisfies the desired property that $\hat{\delta}_i = 1$ for insignificant variables and negligible for significant variables, if the parameter $\lambda > 0$ is appropriately chosen.

**Theorem 2.** Suppose the conditions C1-C7 in the Appendix hold, and $r \leq p_0$. For any positive $\lambda$ and $\gamma$ such that $n^{1/2}\lambda \to 0$ and $n^{(1+\gamma)/2}\lambda \to \infty$ as $n \to \infty$, we have: (i) variable selection consistency, i.e. $P(\hat{A} = A_0) \to 1$; (ii) asymptotic normality, i.e. $n^{1/2}(\hat{\beta}_{A_0} - \beta_{0,A_0})$ is asymptotically normal with mean zero and covariance matrix the same as when $A_0$ is known.

We note that in the statement of the theorem, we need to assume $r \leq p_0$, that is, the removed component is significant. In practice, we select this component based on the initial estimator under the full model, and choose the component that has the largest absolute value.

To use the SGEE in practice, we need to choose appropriately the tuning parameters $(\lambda, \gamma)$. Following Ueki (2009), we use BIC-type criterion to choose these two parameters. That is, we choose $(\lambda, \gamma)$ as the minimizer of

$$BIC_{\lambda,\gamma} = \sum_{i=1}^n (Y_i - \hat{g}(X_i^T \tilde{\beta}_{\lambda,\gamma}, \tilde{\beta}_{\lambda,\gamma}))R_i^{-1}(Y_i - \hat{g}(X_i^T \tilde{\beta}_{\lambda,\gamma}, \tilde{\beta}_{\lambda,\gamma})) + df_{\lambda,\gamma} \log(n),$$

where $\tilde{\beta}_{\lambda,\gamma}$ is the estimator for given $(\lambda, \gamma)$, $df_{\lambda,\gamma}$ is the number of estimated nonzero parameters.
4 Numerical studies

4.1 Simulations

In this section, we carry out some simulations to evaluate the finite sample performance of our proposed method. For each example below, we generate 200 data sets, each consisting of \( n = 50 \) or 100 subjects. For Examples 1-3, we have \( m_k \equiv m = 3 \) observations per subject. Within a cluster, the covariance of the error is specified by \( \text{Cov}(\varepsilon_{m'}, \varepsilon_{m''}) = 0.5^{|m' - m''|}, m', m'' = 1, \ldots, m \). For Example 4, we have \( m_k = 1, 2, 3 \) for \( k \leq n/3, n/3 < k \leq 2n/3 \), and \( k > 2n/3 \), respectively. Within a cluster, the covariance of the error is specified by \( \text{Cov}(\varepsilon_{km'}, \varepsilon_{km''}) = 0.5^{|m' - m''|}, 1 \leq m', m'' \leq m_k \).

The kernel function is taken to be \( K(x) = \frac{3}{4}(1 - x^2) \) if \(|x| \leq 1\), 0 otherwise, and the bandwidth \( h \) is selected by leave-one-out cross validation. We compare the proposed estimator \( \hat{\beta} \) with the oracle estimator (when the zero coefficients are known), the estimator \( \hat{\beta}_s \), and also with \( \hat{\beta}_I \), which is the solution of SGEE (3.1) using identity matrices as working covariance matrices. The following criterions are considered.

- The square of the R statistic: \( R^2 = \frac{[\hat{\beta}^T \beta_0]^2}{[\beta_0^T \beta_0]^2} \);
- The number of zero coefficients and nonzero coefficients obtained by different methods: “TN” is the average number of zero coefficients correctly estimated as zero, and “TP” is the number of nonzero coefficients identified as nonzero.

In these simulations, for SGEE estimator, the common intra-cluster covariance matrix is estimated nonparametrically from the residuals based on the initial estimator assuming working independence.

Example 1. Consider the single-index model for longitudinal data

\[
Y_{ij} = \exp(X_{ij}^T \beta_0) + \varepsilon_{ij}, \quad i = 1, \ldots, n, j = 1, \ldots, 3,
\]

where \( X_{ij} = (X_{ij1}, \ldots, X_{ij6})^T \) was generated from multivariate normal distribution with identity covariance matrix. The true parameter is \( \beta_0 = \frac{1}{\sqrt{2}}(1, 1, 0, 0, 0, 0)^T \). The numerical results are reported in Table 1.
| n  | Method | $R^2$ | TN | TP  |
|----|--------|-------|----|-----|
|    | Oracle | 0.9982 | 4  | 2   |
| 50 | $\hat{\beta}_s$ | 0.9895 | 0  | 2   |
| 50 | $\hat{\beta}$   | 0.9935 | 3.955 | 2 |
|    | $\hat{\beta}_I$  | 0.9817 | 3.525 | 2 |
|    | Oracle | 0.9994 | 4  | 2   |
| 100| $\hat{\beta}_s$  | 0.9962 | 0  | 2   |
| 100| $\hat{\beta}$    | 0.9960 | 3.985 | 2 |
|    | $\hat{\beta}_I$  | 0.9854 | 3.75 | 2 |

Table 1: Simulation results for Example 1.

Example 2. Similarly to Example 1 except that we let $\beta_0 = \frac{1}{\sqrt{4}}(1, 0.6, 0.2, 0, 0, 0)^T$. The numerical results are reported in Table 2.

Example 3. Similar to Example 1, except we use a different link function $g(X^T\beta_0) = \sin(X^T\beta_0)$ which is nonmonotone. The numerical results are reported in Table 3.

Example 4. Similar to Example 1, except that $m_k, k = 1, \ldots, n$, are different. The numerical results are reported in Table 4.

Tables 1-4 show that for our three examples, SGEE can satisfactorily identify the true model. Besides, it is advantageous to take into account the correlation of the observations.

4.2 Real data

We now apply the proposed procedure to the CD4 data from the Multi-Center AIDS Cohort Study. This data set has been studied in Kaslow et al. (1987); Fan and Li (2004); Fan et al. (2007); Li et al. (2010). The data set contains the human immunodeficiency virus (HIV) status of 283 homosexual men who were infected with HIV during the follow-up period between 1984 and 1991. Details of the study design, methods, and medical implications can be found in Kaslow et al. (1987). All individuals were scheduled to have their measurements made during semiannual visits.
| n  | Method | $R^2$ | TN | TP |
|-----|--------|-------|----|----|
| 50  | Oracle | 0.9907 | 3  | 3  |
|     | $\hat{\beta}_*$ | 0.9840 | 0  | 3  |
|     | $\hat{\beta}$ | 0.9329 | 2.78 | 2.255 |
|     | $\hat{\beta}_I$ | 0.9128 | 2.615 | 2.47 |
| 100 | Oracle | 0.9986 | 3  | 3  |
|     | $\hat{\beta}_*$ | 0.9925 | 0  | 3  |
|     | $\hat{\beta}$ | 0.9527 | 2.94 | 2.3  |
|     | $\hat{\beta}_I$ | 0.9411 | 2.8  | 2.395 |

Table 2: Simulation results for Example 2.

| n  | Method | $R^2$ | TN | TP |
|-----|--------|-------|----|----|
| 50  | Oracle | 0.9832 | 4  | 2  |
|     | $\hat{\beta}_*$ | 0.9171 | 0  | 2  |
|     | $\hat{\beta}$ | 0.9756 | 3.885 | 2 |
|     | $\hat{\beta}_I$ | 0.9558 | 3.775 | 2 |
| 100 | Oracle | 0.9928 | 4  | 2  |
|     | $\hat{\beta}_*$ | 0.9648 | 0  | 2  |
|     | $\hat{\beta}$ | 0.9912 | 3.94 | 2 |
|     | $\hat{\beta}_I$ | 0.9889 | 3.88 | 2 |

Table 3: Simulation results for Example 3.

| n  | Method | $R^2$ | TN | TP |
|-----|--------|-------|----|----|
| 50  | Oracle | 0.9980 | 4  | 2  |
|     | $\hat{\beta}_*$ | 0.9806 | 0  | 2  |
|     | $\hat{\beta}$ | 0.9972 | 3.965 | 2 |
|     | $\hat{\beta}_I$ | 0.9966 | 3.85 | 2 |
| 100 | Oracle | 0.9996 | 4  | 2  |
|     | $\hat{\beta}_*$ | 0.9931 | 0  | 2  |
|     | $\hat{\beta}$ | 0.9990 | 3.995 | 2 |
|     | $\hat{\beta}_I$ | 0.9986 | 3.99 | 2 |

Table 4: Simulation results for Example 4.
However, many participants missed some of their scheduled visits resulting in different measurement time points and unequal number of measurements per individual. In our analysis, we let $y_{ij}$ be the CD4 cell counts for individual $i$ at the $j$th visit, $x_{ij}\, 1$ be the smoking status with 1 for a smoker and 0 for a non-smoker, $x_{ij}\, 2$ be the person’s age, and $x_{ij}\, 3$ be last measured CD4 level before HIV infection. For exploratory purposes, we also consider possible interactions of the covariates and also squares of $x_{ij}\, 2$ and $x_{ij}\, 3$, resulting in the following model:

$$y_{ij} = g(x_{ij}\, 1\beta_1 + x_{ij}\, 2\beta_2 + x_{ij}\, 3\beta_3 + x_{ij}\, 2^2\beta_4 + x_{ij}\, 3^2\beta_5 + x_{ij}\, 1x_{ij}\, 2\beta_6 + x_{ij}\, 1x_{ij}\, 3\beta_7 + x_{ij}\, 2x_{ij}\, 3\beta_8) + \varepsilon_{ij}.$$

We apply the SGEE approach to this data set to select significant variables and estimate the effects. The tuning parameters $\lambda$ and $\gamma$ are selected by the BIC-type criterion. For any individual, we assume the correlation between visits at time $t_{j1}$ and $t_{j2}$ is $\alpha^{|t_{j1} - t_{j2}|}$. The fitted model is

$$y_{ij} \sim g(0.4531x_{ij}\, 1 - 0.6744x_{ij}\, 2 + 0.5829x_{ij}\, 3)$$

By our variable selection procedure, we can see that only the linear terms are significant.

**APPENDIX**

In order to study the asymptotic behavior of the estimator, the following standard assumptions are imposed (Li et al., 2010).

- **C1.** The density function $f_{ij}(t)$ of $X_{ij}^T\beta$ is bounded away from zero and continuously differentiable on $\{t : t = X_{ij}^T\beta, X_{ij} \in A, i = 1, \ldots, n; j = 1, \ldots, m_i\}$ and $A$ is the support of $X_{ij}$ which is assumed to be compact.

- **C2.** The function $g(\cdot)$ is twice continuously differentiable, and $E(X_{klq}|X_{klq}^T\beta = x), 1 \leq l \leq m_k, 1 \leq k \leq n, 1 \leq q \leq p$ as a function of $x$ is Lipschitz continuous.
• C3. The kernel $K$ is a bounded, continuous and symmetric probability density function, satisfying
\[ \int_{-\infty}^{\infty} u^2 K(u) du < \infty. \]

• C4. There exists a positive constant $M$, such that $\max_{1 \leq k \leq n, 1 \leq j \leq m_k} E(\varepsilon^4_{kj}) \leq M < \infty$.

• C5. The bandwidth $h$ satisfies $nh^3 \to \infty, nh^8 \to 0$.

• C6. The eigenvalues of $R_i$ and $\Sigma_i$ are uniformly bounded and bounded away from zero.

• C7. $\Omega = \lim_{n \to \infty} \frac{1}{n} \sum_{k=1}^{n} E \left\{ \left( g'(X^T_k \hat{\beta}_0) J^T_{\hat{\beta}_0} \tilde{X}_0^T - R_k^{-1} \varepsilon_k \right)^{\otimes 2} \right\}$ is positive definite, where we use the notation $\tilde{X}_0^T = X_k - E(X_k | X^T_k \hat{\beta}_0)$ is $m_k \times p$ matrix with entries $X_{klq} - E[X_{klq} | X^T_{klq}]$, $1 \leq l \leq m_k, 1 \leq q \leq p$, and $E(A)^{\otimes 2} = E(AA^T)$ for any matrix $A$.

\[ V = \lim_{n \to \infty} \frac{1}{n} \sum_{k=1}^{n} E \left[ \left( g'(X^T_k \hat{\beta}_0) J^T_{\hat{\beta}_0} \tilde{X}_0^T R_k^{-1} (g'(X^T_k \hat{\beta}_0) J^T_{\hat{\beta}_0} \tilde{X}_0^T \tilde{X}_0^T) \right)^{\otimes 2} \right] \] is also positive definite.

**Remark.** Note that in condition C1 we allow the distributions of $X^T_{ij} \beta$ to be different for different $i, j$, and in particular $m_i, i = 1, \ldots, n$, are not required to be the same.

**Proof of Theorem 1.** Proof of existence of $\sqrt{n}$-consistent solution to (2.4) is almost same as in Wang et al. (2010) and omitted here. Thus we proceed to consider asymptotic normality. By (2.4), since

\[ \sum_{k=1}^{n} \tilde{Z}_k^T R_k^{-1} (Y_k - \hat{g}(X_k \hat{\beta}_*)) = 0, \]
it follows
\[
\begin{align*}
\sum_{k=1}^{n} Z_k^0 T R_k^{-1} (Y_k - \hat{g}(X_k^T \hat{\beta}_*)) \\
= \sum_{k=1}^{n} (g'(X_k^T \hat{\beta}_*) J_{\hat{\beta}_*}^T [X_k - E(X_k|X_k^T \hat{\beta}_*)])^T R_k^{-1} \varepsilon_k \\
+ \sum_{k=1}^{n} (g'(X_k^T \hat{\beta}_*) J_{\hat{\beta}_*}^T [X_k - \hat{E}(X_k|X_k^T \hat{\beta})])^T R_k^{-1} \varepsilon_k \\
+ \sum_{k=1}^{n} (g'(X_k^T \hat{\beta}_*) J_{\hat{\beta}_*}^T [X_k - \hat{E}(X_k|X_k^T \hat{\beta})])^T R_k^{-1} (g(X_k^T \hat{\beta}_0) - \hat{g}(X_k^T \hat{\beta}_*)) \\
+ \sum_{k=1}^{n} ([g'(X_k^T \hat{\beta}_*) - g'(X_k^T \beta_0)] J_{\beta_0}^T [X_k - \hat{E}(X_k|X_k^T \hat{\beta})])^T R_k^{-1} (g(X_k^T \hat{\beta}_0) - \hat{g}(X_k^T \hat{\beta}_*)) \\
: = Q_1(\hat{\beta}_*) + Q_2(\hat{\beta}_*) + Q_3(\hat{\beta}_*) + Q_4(\hat{\beta}_*) + Q_5(\hat{\beta}_*). \quad (A.1)
\end{align*}
\]
Noting that \(J_{\hat{\beta}_*} - J_{\beta_0} = O_p(n^{-1/2})\), we have
\[
Q_1(\hat{\beta}_*) - U(\beta_0) = o_p(\sqrt{n}), \quad (A.2)
\]
where
\[
U(\beta_0) = \sum_{k=1}^{n} (g'(X_k^T \beta_0) J_{\beta_0}^T [X_k - E(X_k|X_k^T \beta_0)])^T R_k^{-1} \varepsilon_k.
\]
For \(Q_2(\hat{\beta}_*)\), denote
\[
R_k^{-1} = \begin{pmatrix}
R_{k11}^{-1} & \cdots & R_{k1m_k}^{-1} \\
\vdots & \ddots & \vdots \\
R_{km_k1}^{-1} & \cdots & R_{km_km_k}^{-1}
\end{pmatrix},
\]
then
\[
Q_2(\hat{\beta}_*) = \sum_{k=1}^{n} \sum_{j=1}^{m_k} \sum_{i=1}^{m_k} \varepsilon_{kj} \sum_{i=1}^{m_k} R_{ki}^{-1} [g'(X_{ki}^T \hat{\beta}_*) - g'(X_{ki}^T \beta_0)] J_{\beta_0}^T [X_{ki} - \hat{E}(X_{ki}|X_k^T \hat{\beta})]. \quad (A.3)
\]
Note that \( \hat{\beta}_s, \tilde{\beta} \in \mathbf{B} \), together with conditions C2 and C3, we have

\[
Q_2(\hat{\beta}_s^{(r)}) = \sum_{k=1}^{n} \sum_{j=1}^{m_k} \varepsilon_{kj} R^{-1}_{kij} \left[ g'(X_{ki}^T \beta_0) - g'(X_{ki}^T \hat{\beta}_s^{(r)}) + \frac{\partial}{\partial \hat{\beta}_s^{(r)}} g'(X_{ki}^T \hat{\beta}_s^{(r)}) \right]
\]

\[
\times J_{\hat{\beta}_s^{(r)}}' \left[ \hat{X}_{ki} - \hat{E}(X_{ki}|X_k^T \hat{\beta}_2) \right] + o_p(\sqrt{n}).
\]

where \( \hat{\beta}_1, \tilde{\beta}_2 \) are the intermediate values between \( \beta_0, \hat{\beta}_s \). Thus,

\[
Q_2(\hat{\beta}_s^{(r)}) = \sum_{k=1}^{n} \sum_{j=1}^{m_k} \varepsilon_{kj} R^{-1}_{kij} \left[ g'(X_{ki}^T \beta_0) - g'(X_{ki}^T \hat{\beta}_s^{(r)}) \right] J_{\hat{\beta}_s^{(r)}}' \left[ \hat{X}_{ki} - \hat{E}(X_{ki}|X_k^T \beta_0) \right] + o_p(\sqrt{n}).
\]

Let \( Q_2(\hat{\beta}_s^{(r)}) = J_{\hat{\beta}_s^{(r)}}^T Q_2(\hat{\beta}_s^{(r)})^* \), where the sth component of \( Q_2(\hat{\beta}_s^{(r)})^* \) is

\[
Q_2(\hat{\beta}_s^{(r)})_{s}^* = \sum_{k=1}^{n} \sum_{j=1}^{m_k} \epsilon_{kj} R^{-1}_{kij} \left[ g'(X_{ki}^T \beta_0) - g'(X_{ki}^T \hat{\beta}_s^{(r)}) \right] [\hat{X}_{kis} - \hat{E}(X_{kisi}|X_k^T \beta_0)].
\]

By (2.2), let \( \tilde{X}_{kis} = [\hat{X}_{kis} - \hat{E}(X_{kisi}|X_k^T \beta_0)] \) be the sth component of \( X_{kisi} - \hat{E}(X_{kisi}|X_k^T \beta_0) \), we have

\[
Q_2(\hat{\beta}_s^{(r)})_{s}^* = \sum_{k=1}^{n} \sum_{j=1}^{m_k} \varepsilon_{kj} R^{-1}_{kij} \left[ \sum_{l_1=1}^{m_{k_1}} \bar{W}_{nl_1l_2} (X_{ki}^T \beta_0, \hat{\beta}_s^{(r)}) g(X_{ki}^T \beta_0) - g'(X_{ki}^T \beta_0) \right]
\]

\[
+ \sum_{k=1}^{n} \sum_{j=1}^{m_k} \tilde{W}_{nkj} (X_{kj}^T \beta_0, \hat{\beta}_s^{(r)}) \tilde{X}_{kj} \varepsilon_{kj} R^{-1}_{kij} + \sum_{k=1}^{n} \sum_{j \neq i}^{m_k} \tilde{W}_{nkj} (X_{kj}^T \beta_0, \hat{\beta}_s^{(r)}) \tilde{X}_{kj} \varepsilon_{kj} R^{-1}_{kij}
\]

\[
+ \sum_{k=1}^{n} \sum_{j=1}^{m_k} \sum_{l_1 \neq k} \sum_{l_2 \neq i} \tilde{W}_{nl_1l_2} (X_{ki}^T \beta_0, \hat{\beta}_s^{(r)}) \tilde{X}_{kis} \varepsilon_{kj} R^{-1}_{kij} + \sum_{k=1}^{n} \sum_{j \neq i}^{m_k} \tilde{W}_{nkj} (X_{kj}^T \beta_0, \hat{\beta}_s^{(r)}) \tilde{X}_{kj} \varepsilon_{kj} R^{-1}_{kij}
\]

\[
:= Q_{21s} + Q_{22s} + Q_{23s} + Q_{24s}.
\]

Similar to the proof of Lemma A.4 in Li et al. (2010), utilizing also Lemmas A.1-A.3 there, we can show that \( Q_2(\hat{\beta}_s^{(r)})_{s}^* = o_p(\sqrt{n}) \) and thus

\[
Q_2(\hat{\beta}_s^{(r)}) = o_p(\sqrt{n}).
\]
Similarly, we can obtain
\[ Q_3(\beta_*(r)) = o_p(\sqrt{n}), \quad Q_5(\beta_*(r)) = o_p(\sqrt{n}). \] (A.5)

For \( Q_4(\beta_*(r)) \), simple calculations yield
\[
Q_4(\beta_*(r)) = \sum_{k=1}^{n} \sum_{i=1}^{m_k} \sum_{j=1}^{m_k} g'(X^T_{ki} \hat{\beta}_*) J^T_{\beta_0(r)} (X_{ki} - E(X_{ki} | X_{ki}^T \hat{\beta}_*)) R^{-1}_{kij} (g(X^T_{kj} \hat{\beta}_*) - g(X^T_{kj} \hat{\beta}_*)) \\
- \sum_{k=1}^{n} \sum_{i=1}^{m_k} \sum_{j=1}^{m_k} g'(X^T_{ki} \beta_0) J^T_{\beta_0(r)} (X_{ki} - E(X_{ki} | X_{ki}^T \beta_0)) R^{-1}_{kij} \\
\times g'(X^T_{kj} \beta_0) \{ J^T_{\beta_0(r)} (X_{ki} - E(X_{ki} | X_{kj}^T \beta_0)) \}^T (\hat{\beta}_*(r) - \beta_0(r)) + o_p(\sqrt{n}) \\
= Q_{41}(\beta_*(r)) + Q_{42}(\beta_*(r)) + o_p(\sqrt{n}).
\]

It is easy to show that \( Q_{41}(\beta_*(r)) = o_p(\sqrt{n}) \) and that
\[ Q_{42}(\beta_*(r)) - nV(\beta_*(r) - \beta_0(r)) = o_p(\sqrt{n}), \] (A.6)

where
\[
V = \lim_{n \to \infty} \frac{1}{n} \sum_{k=1}^{n} E \left[ (g'(X^T_k \beta_0) J^T_{\beta_0(r)} \bar{X}_k^r)^T R^{-1}_k (g'(X^T_k \beta_0) J^T_{\beta_0(r)} \bar{X}_k^r) \right].
\] (A.7)

is a positive definite matrix. Thus
\[ Q_4(\beta_*(r)) - nV(\beta_*(r) - \beta_0(r)) = o_p(\sqrt{n}), \] (A.8)

In summary, by estimating equation (2.4), together with (A.2), (A.4), (A.5) and (A.8), it follows
\[
0 = Q_1(\beta_*(r)) + Q_2(\beta_*(r)) + Q_3(\beta_*(r)) + Q_4(\beta_*(r)) + Q_5(\beta_*(r)) \\
= U(\beta_0(r)) + o_p(\sqrt{n}) - nV(\beta_*(r) - \beta_0(r)) \\
\Rightarrow \sqrt{n}(\beta_*(r) - \beta_0(r)) = V^{-1} n^{-1/2} U(\beta_0(r)) + o_p(1). \] (A.9)

Thus, we have
\[ \sqrt{n}(\beta_0 - \beta_0) = J_{\beta_0(r)} V^{-1} n^{-1/2} U(\beta_0(r)) + o_p(1). \] (A.10)
The asymptotic normality of $\hat{\beta}_*$ directly follows from this representation and the central limit theorem. □

Proof of Theorem 2. First, for $j \in A_0^c$, we have $|\tilde{\beta}^{(r)}_j| = O(n^{-1/2})$ by the assumption of $\sqrt{n}$-consistency of the initial estimator. Using the condition on $\lambda$ in the statement of the theorem, we get

$$P(\lambda/|\tilde{\beta}^{(r)}_j|^{1+\gamma} < 1) \to 0, j \in A_0^c,$$

and thus

$$P(\hat{\delta}_j = 1 \text{ for all } j \in A_0^c) \to 1.$$

On the other hand, we have for any $\epsilon > 0$ and $j \in A_0 - \{r\}$,

$$P(\hat{\delta}_j > n^{-1/2} \epsilon) = P(\lambda n^{1/2}/\epsilon > |\tilde{\beta}^{(r)}_j|^{1+\gamma}) \to 0$$

using that $\lambda n^{1/2} \to 0$ and that $|\tilde{\beta}^{(r)}_j|$ is bounded away from zero. Thus $\hat{\delta}_j = o_p(n^{-1/2})$ for each $j \in A_0 - \{r\}$, implying trivially $P(\hat{\delta}_j < 1 \text{ for all } j \in A_0 - \{r\}) \to 1$, and (i) is proved.

Next, we prove (ii). From (i) and the assumption that the $r$th component of $\beta_0$ is nonzero, the SGEE coincide with

$$(1 - \hat{\delta}_j) u_j(\hat{\beta}^{(r)}) + \hat{\delta}_j \hat{\beta}^{(r)}_j = 0, \text{ for } j \in A_0 - \{r\}$$

(A.13)

and $\hat{\beta}_j = 0$ for $j \in A_0^c$, with probability tending to one, where $u_j(\hat{\beta}^{(r)})$ is the $j$th component of $\sum_{k=1}^n \hat{Z}^T_k R_k^{-1} (Y_k - \hat{g}(X_k^T \hat{\beta}))$, $j \in A_0 - \{r\}$. Using that $\hat{\delta}_j = o_p(n^{-1/2})$ for $j \in A_0 - \{r\}$, it is easy to show that (A.13) is asymptotically equivalent to $u_j(\hat{\beta}^{(r)}) = 0$ and the asymptotic normality follows the same way as in the proof of Theorem 1.

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References

Bai, Y., Fung, W., Zhu, Z., 2009. Penalized quadratic inference functions for single-index models with longitudinal data. Journal of Multivariate Analysis 100 (1), 152–161.

Balan, R., Schiopu-Kratina, I., 2005. Asymptotic results with generalized estimating equations for longitudinal data. Annals of Statistics 33 (2), 522–541.

Candes, E., Tao, T., 2007. The dantzig selector: statistical estimation when p is much larger than n. Annals of Statistics 35 (6), 2313–2351.

Carroll, R., Fan, J., Gijbels, I., Wand, M., 1997. Generalized partially linear single-index models. Journal of the American Statistical Association 92 (438), 477–489.

Chang, Z., Xue, L., Zhu, L., 2010. On an asymptotically more efficient estimation of the single-index model. Journal of Multivariate Analysis 101 (8), 1898–1901.

Cui, X., Härdle, W. K., Zhu, L. X., 2011. The EFM approach for single-index models. Annals of Statistics 39 (3), 1658–1688.

Delecroix, M., Härdle, W., Hristache, M., 2003. Efficient estimation in conditional single-index regression. Journal of Multivariate Analysis 86 (2), 213–226.

Fan, J., Gijbels, I., 1996. Local polynomial modelling and its applications. Chapman & Hall/CRC.
Fan, J., Huang, T., Li, R., 2007. Analysis of longitudinal data with semiparametric estimation of covariance function. Journal of the American Statistical Association 102 (478), 632–641.

Fan, J., Li, R., 2001. Variable selection via nonconcave penalized likelihood and its oracle properties. Journal of the American Statistical Association 96 (456), 1348–1360.

Fan, J., Li, R., 2004. New estimation and model selection procedures for semiparametric modeling in longitudinal data analysis. Journal of the American Statistical Association 99 (467), 710–723.

Härdle, W., Hall, P., Ichimura, H., 1993. Optimal smoothing in single-index models. The annals of Statistics 21 (1), 157–178.

Ichimura, H., 1993. Semiparametric least squares (SLS) and weighted SLS estimation of single-index models. Journal of Econometrics 58 (1-2), 71–120.

Kaslow, R., Ostrow, D., Detels, R., Phair, J., Polk, B., Rinaldo Jr, C., 1987. The Multicenter AIDS Cohort Study: rationale, organization, and selected characteristics of the participants. American Journal of Epidemiology 126 (2), 310–318.

Kong, E., Xia, Y., 2007. Variable selection for the single-index model. Biometrika 94 (1), 217–229.

Li, G., Zhu, L., Xue, L., Feng, S., 2010. Empirical likelihood inference in partially linear single-index models for longitudinal data. Journal of Multivariate Analysis 101 (3), 718–732.

Liang, K., Zeger, S., 1986. Longitudinal data analysis using generalized linear models. Biometrika 73 (1), 13–22.
Lin, X., Carroll, R., 2000. Nonparametric Function Estimation for Clustered Data When the Predictor Is Measured Without/With Error. Journal of the American Statistical Association 95 (450), 520–534.

Naik, P., Tsai, C., 2000. Partial least squares estimator for single-index models. Journal of the Royal Statistical Society: Series B (Statistical Methodology) 62 (4), 763–771.

Naik, P., Tsai, C., 2001. Single-index model selections. Biometrika 88 (3), 821–832.

Naik, P., Tsai, C., 2004. Residual information criterion for single-index model selections. Journal of Nonparametric Statistics 16 (1), 187–195.

Tibshirani, R., 1996. Regression shrinkage and selection via the lasso. Journal of the Royal Statistical Society. Series B (Methodological) 58 (1), 267–288.

Ueki, M., 2009. A note on automatic variable selection using smooth-threshold estimating equations. Biometrika 96 (4), 1005–1011.

Wang, J., Xue, L., Zhu, L., Chong, Y., 2010. Estimation for a partial-linear single-index model. The Annals of Statistics 38 (1), 246–274.

Wong, H., Ip, W., Zhang, R., 2008. Varying-coefficient single-index model. Computational statistics & data analysis 52 (3), 1458–1476.

Xia, Y., Härdle, W., 2006. Semi-parametric estimation of partially linear single-index models. Journal of Multivariate Analysis 97 (5), 1162–1184.

Xia, Y., Tong, H., Li, W., 1999. On extended partially linear single-index models. Biometrika 86 (4), 831–842.

Yu, Y., Ruppert, D., 2002. Penalized spline estimation for partially linear single-index models. Journal of the American Statistical Association 97 (460), 1042–1054.
Zhu, L., Xue, L., 2006. Empirical likelihood confidence regions in a partially linear single-index model. Journal of the Royal Statistical Society: Series B (Statistical Methodology) 68 (3), 549–570.

Zou, H., 2006. The adaptive lasso and its oracle properties. Journal of the American Statistical Association 101 (476), 1418–1429.