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FUNCTIONAL ADDITIVE QUANTILE REGRESSION

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Abstract: We investigate a functional additive quantile regression that models the conditional quantile of a scalar response based on the nonparametric effects of a functional predictor. We model the nonparametric effects of the principal component scores as additive components, which are approximated by B-splines. We select the relevant components using a nonconvex smoothly clipped absolute deviation (SCAD) penalty. We establish that, when the relevant components are known, the convergence rate of the estimator using the estimated principal component scores is the same as that using the true scores. We also show that the estimator based on relevant components is a local solution of the SCAD penalized quantile regression problem. The practical performance of the proposed method is illustrated using simulation studies and an empirical application to corn yield data.

Key words and phrases: Additive quantile regression, functional data, principal component analysis, splines.

1. Introduction

A functional quantile regression provides an overall picture of the predictive distribution of the scalar response, given the function-valued covariates, rather than focusing simply on the mean response. This model is used, for instance, by Cardot, Crambes and Sarda (2005), Chen and Müller (2012), and Kato (2012), who construct the τ th conditional quantile of a scalar response y from a functional predictor X using the following linear operation:

$$Q_y(\tau|X) = \alpha(\tau) + \int_{\mathcal{T}} X(t)\beta(t, \tau)dt. \quad (1.1)$$

Here, $X(t)$ is a predictor process that is a square integrable random function defined on a compact interval \mathcal{T} , and $\beta(t, \tau)$ is the square integrable coefficient function for a given τ . Yao, Sue-Chee and Wang (2017) and Ma et al. (2019) extended the model to accommodate high-dimensional scalar predictors. To deal with functional data that are infinite-dimensional objects, the most widely used

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approach is to project the functional data onto a space spanned by a finite number of basis functions. The basis can be fixed in advance (e.g, B-splines, Fourier basis; see Crambes, Gannoun and Henchiri (2013) and Crambes, Gannoun and Henchiri (2014)) or data-driven. A convenient choice for the data-driven option is to use the eigenbasis of the covariance operator of $X(t)$, which often provides a parsimonious and efficient representation. More specifically, the covariance kernel is defined as $G(s, t) = \text{Cov}(X(s), X(t))$. By the well-known functional principal component analysis (FPCA) (Yao, Müller and Wang (2005), Hall, Müller and Wang (2006), Li and Hsing (2010)), we have the spectral expansion $G(s, t) = \sum_{k=1}^{\infty} \lambda_k \phi_k(s) \phi_k(t)$ and the Karhunen–Loeve expansion $X(t) = \sum_{k=1}^{\infty} \xi_k \phi_k(t)$, where $\lambda_1 \geq \lambda_2 \geq \dots \geq 0$ are ordered eigenvalues, $\{\phi_k\}_{k=1}^{\infty}$ are eigenfunctions making up an orthonormal basis of $L_2(\mathcal{T})$, and $\xi_k = \int_{\mathcal{T}} X(t) \phi_k(t) dt$ are called the principal component scores for $X(t)$. Using the expansion $\beta(t, \tau) = \sum_{k=1}^{\infty} b_k(\tau) \phi_k(t)$, where $b_k(\tau) = \int_{\mathcal{T}} \beta(t, \tau) \phi_k(t) dt$, the model (1.1) is transformed into a quantile regression model with an infinite number of “regressors”:

$$Q_y(\tau|X) = \alpha(\tau) + \sum_{k=1}^{\infty} b_k(\tau) \xi_k, \quad (1.2)$$

and thus regularization is necessary. In Kato (2012), this regularization is achieved by truncating the eigensequence to the first K leading terms, where K is chosen such that it retains most of the variation in predictor $X(t)$.

There is an obvious limitation to model (1.2) in that the linear relationship can be restrictive for general applications. To make it more flexible, we propose a functional additive quantile regression in which the linear components are replaced by the sum of nonlinear functional components; that is,

$$Q_y(\tau|X) = \alpha(\tau) + \sum_{k=1}^{\infty} f_{k,\tau}(\xi_k),$$

where $\{f_{k,\tau}(\cdot)\}$ are unknown smooth functions. To make the estimation feasible, we assume all useful information is contained in the first s components. That is, we assume $f_{k,\tau} \equiv 0$, $k > s$, for some sufficiently large s . Furthermore, to avoid possible scaling issues, we use the standardized version $\zeta_{ik} = \Phi(\lambda_k^{-1/2} \xi_{ik})$, where $\Phi(\cdot)$ is a continuously differentiable map from \mathbb{R} to $[0, 1]$. Then, the model becomes

$$Q_y(\tau|X) = \alpha(\tau) + \sum_{k=1}^s f_{k,\tau}(\zeta_k). \quad (1.3)$$

To ensure that we can identify $f_{k,\tau}$, we assume that $E f_{k,\tau}(\zeta_k) = 0$, for $k =$

$1, 2, \dots, s$. We assume the significant components are contained in the first s components, but that the s components are not all significant.

We approximate the component functions using the B-splines method, which is computationally convenient. To automatically select the significant components, we impose a smoothly clipped absolute deviation (SCAD) penalty (Sherwood and Wang (2016)) on the l_1 norm of each coefficient group. With the l_1 penalty, the minimization can be solved using linear programming. However, the computation becomes more involved if the l_2 penalty is used. This choice is also made for theoretical convenience. If we penalize the l_2 norm, the sufficient condition for the local minimizer of the convex difference problem in Wang, Wu and Li (2012) is too restrictive and difficult to satisfy. This is because the subgradients of $\|\boldsymbol{\theta}\|_1$ involves only the signs of $\boldsymbol{\theta}_k$, whereas the gradients of $\|\boldsymbol{\theta}\|_2$ are not restricted to $\{-1, 0, 1\}$. Empirically, we find that when we penalize $\|\boldsymbol{\theta}\|_1$ inside a SCAD penalty, the individual components of $\boldsymbol{\theta}$ are not sparse, as long as $\boldsymbol{\theta} \neq 0$. Other penalties, such as the group LASSO or adaptive group LASSO, could also be used instead of the SCAD penalty. Our choice is mainly for convenience; several alternatives in the literature exhibit similar performance.

There is a rich body of literature on functional additive models when the conditional mean of the scalar response is of interest. Müller and Yao (2008) considered the model $E(y|X) = \alpha + \sum_{k=1}^{\infty} f_k(\xi_k)$. They estimated $\{f_k\}$ using local polynomial smoothing, and regularized the model using truncation, as we do here. Zhu, Yao and Zhang (2014) estimated and selected the additive components in a reproducing kernel Hilbert space (RKHS) framework, adopting the component selection and smoothing operator (COSSO) penalty. Furthermore, Wong, Li and Zhu (2018) extended the model to include partial linear functional additive regressions with multivariate functional predictors. However, fewer works examine conditional quantile modeling with functional predictors. Kato (2012) pioneered this field by investigating the functional linear quantile regression. His analysis is based on fully observed $X(t)$. In contrast, we allow observation errors in the predictor, and apply a B-spline approximation, which is computationally more convenient than the kernel method in terms of implementation. The relevant components are then selected using a SCAD penalty on the l_1 norm of each coefficient group.

A related line of research is that on additive or sparse additive quantile regression models; see Horowitz and Lee (2005), Koenker (2011), Kato (2011), Lian (2012) and Lv et al. (2018). Our work differs from these studies in that we estimate the scores, which serve as pseudo-predictors, using an FPCA. Thus, theoretically, we need to deal with the error caused by the estimated predictors.

This requires new bounds throughout the proof and new conditions constraining the number of components, number of knots in the splines, and tuning parameter in the penalty.

The rest of the paper is organized as follows. In Section 2, we present the proposed approach and the computational algorithm. In Section 3, we investigate the asymptotic properties of the proposed estimator. We illustrate the method by means of simulation studies in Section 4, and apply it to a real data set in section 5. Section 6 concludes the paper. All proofs are contained in the online Supplementary Material.

2. Proposed Methodology

Let $\{y_i, X_i(t)\}_{i=1}^n$ be independent and identically distributed (i.i.d.) realizations of the pair $\{y, X(t)\}$. The trajectories $\{X_i(t) : t \in \mathcal{T}\}$ are observed intermittently on possibly irregular grids $\mathbf{t}_i = (t_{i1}, \dots, t_{iN_i})^T$. We further assume that the predictor trajectories are subject to i.i.d. measurement errors; that is, $X_{ij} = X_i(t_{ij}) + \varepsilon_{ij}$, with $E\varepsilon_{ij} = 0$ and $\text{var}(\varepsilon_{ij}) = \sigma^2$, for $j = 1, \dots, N_i$. The sequence of functional principal component (FPC) scores of $X_i(t)$ is denoted by $\boldsymbol{\xi}_{i,\infty} = (\xi_{i1}, \xi_{i2}, \dots)^T$. Denote the s truncated FPC scores as $\boldsymbol{\xi}_i = (\xi_{i1}, \dots, \xi_{is})^T$. Similarly, write $\boldsymbol{\zeta}_{i,\infty} = (\zeta_{i1}, \zeta_{i2}, \dots)^T$ and $\boldsymbol{\zeta}_i = (\zeta_{i1}, \dots, \zeta_{is})^T$. The transformed FPC scores $\{\boldsymbol{\zeta}_i\}$ cannot be observed, and need to be estimated from discrete observations of $X_i(t)$.

When $X_i(t)$ are fully observed, the mean and covariance of $X(t)$ can be estimated using the sample counterparts, $\hat{\mu}(t)$ and $\hat{G}(s, t)$, respectively. The spectral decomposition on the estimated covariance function, $\hat{G}(s, t) = \sum_{k=1}^{n-1} \hat{\lambda}_k \hat{\phi}_k(s) \hat{\phi}_k(t)$, provides the estimated eigenvalues and eigenfunctions. The FPC scores are estimated by projecting $X_i(t)$ onto the eigenfunctions, $\hat{\xi}_{ik} = \int_{\mathcal{T}} X_i(t) \hat{\phi}_k(t) dt$, and we set $\hat{\zeta}_{ik} = \Phi(\hat{\lambda}_k^{-1/2} \hat{\xi}_{ik})$ for the transformed FPC scores. When only discrete observations are available, we focus on the case in which dense measurements are made, such that each $X_i(t)$ can be recovered effectively using smoothing. The eigenvalues, eigenfunctions, and FPC scores are estimated by replacing $X_i(t)$ with the estimated $\hat{X}_i(t)$. For the detailed algorithm, refer to Section 3.1 in Wong, Li and Zhu (2018). The estimated transformed FPC scores are denoted by $\hat{\boldsymbol{\zeta}}_i$, which serve as the predictors in the following.

The additive components $f_{k,\tau}(\cdot)$, for $k = 1, \dots, s$ are approximated using B-spline basis functions. Let $(b_1(t), \dots, b_{K_n+l+1}(t))^T$ be the vector of normalized B-spline basis functions of order l , with K_n quasi-uniform internal knots on $[0, 1]$. Refer to Schumaker (2007) for details of the B-spline construction. For

ease of notation and to simplify the proofs, we use the same number of basis functions for all nonlinear components. To accommodate the identifying restriction $E f_{k,\tau}(\zeta_k) = 0$, for $1 \leq k \leq s$, we use the centered B-spline basis functions $B_m(\zeta_k) = b_{m+1}(\zeta_k) - (1/n) \sum_{i=1}^n b_{m+1}(\hat{\zeta}_{ik})$ for $m = 1, \dots, K_n + l$, as in Huang, Horowitz and Wei (2010), and denote $\mathbf{w}(\zeta_k) = (B_1(\zeta_k), \dots, B_{K_n+l}(\zeta_k))^T$. We can approximate each $f_{k,\tau}(t)$ by $f_{k,\tau}(t) \approx \mathbf{w}(t)' \boldsymbol{\theta}_k$. Let $\boldsymbol{\theta} = (\theta_0, \boldsymbol{\theta}_1^T, \dots, \boldsymbol{\theta}_s^T)^T$ be the spline coefficient for the estimation of the component functions, and define $\mathbf{W}(\zeta_i) = (K_n^{-1/2}, \mathbf{w}(\zeta_{i1})^T, \dots, \mathbf{w}(\zeta_{is})^T)^T$, where $K_n^{-1/2}$ is used to make the scale of the intercept comparable with that of the B-spline basis functions. Now, we can apply the SCAD penalty on the l_1 norms of $\boldsymbol{\theta}_k$, for $1 \leq k \leq s$, to select the significant components. We estimate $\boldsymbol{\theta}$ by $\hat{\boldsymbol{\theta}}$, which minimizes

$$S_n(\boldsymbol{\theta}) = n^{-1} \sum_{i=1}^n \rho_\tau(y_i - \mathbf{W}(\hat{\zeta}_i)^T \boldsymbol{\theta}) + \sum_{k=1}^s p_\lambda(\|\boldsymbol{\theta}_k\|_1), \tag{2.1}$$

where $\rho_\tau(u) = |u| + (2\tau - 1)u$ is the quantile loss function, $\|\boldsymbol{\theta}_k\|_1 = \sum_{j=1}^{K_n+l} |\theta_{kj}|$, and $p_\lambda(t)$ is the SCAD penalty function, defined as

$$p_\lambda(t) = \lambda t I(0 \leq t < \lambda) + \frac{a\lambda t - (t^2 + \lambda^2)/2}{a - 1} I(\lambda \leq t \leq a\lambda) + \frac{(a + 1)\lambda^2}{2} I(t > a\lambda),$$

for some $a > 2$. Then, we can estimate the parameters in model (1.3) by $\hat{\alpha}(\tau) = K_n^{-1/2} \hat{\theta}_0$ and $\hat{f}_{k,\tau}(t) = \mathbf{w}(t)^T \hat{\boldsymbol{\theta}}_k$.

Owing to the l_1 norm in the penalization, the above penalized problem can be solved using the local linear approximation (LLA) proposed by Zou and Li (2008). Specifically, for each step t , we update the estimator as follows:

$$\hat{\boldsymbol{\theta}}^t = \underset{\boldsymbol{\theta}}{\operatorname{argmin}} n^{-1} \sum_{i=1}^n \rho_\tau(y_i - \mathbf{W}(\hat{\zeta}_i)^T \boldsymbol{\theta}) + \sum_{k=1}^s p'_\lambda(\|\hat{\boldsymbol{\theta}}_k^{t-1}\|_1) \|\boldsymbol{\theta}_k\|_1.$$

This problem can be transformed to an unpenalized weighted quantile regression problem based on the observation $|\theta_{kj}| = \rho_\tau(\theta_{kj}) + \rho_\tau(-\theta_{kj})$ and an augmented data set. A similar algorithm was used in Sherwood and Wang (2016).

In practice, the selection of the tuning parameter λ is important. We choose λ to minimize the Bayesian information criterion (BIC; see Section 3.1 in Lee, Noh and Park (2014)), defined as

$$\text{BIC}(\lambda) = \log \sum_{i=1}^n \rho_\tau(y_i - \mathbf{W}(\hat{\zeta}_i)^T \hat{\boldsymbol{\theta}}_\lambda) + J_\lambda \frac{\log n}{2n} C_n, \tag{2.2}$$

where $\hat{\boldsymbol{\theta}}_\lambda$ is the SCAD penalized estimator given λ , $J_\lambda = 1 + (K_n + l)|\mathcal{S}_\lambda|$, $|\mathcal{S}_\lambda|$ is the number of selected components, and C_n is defined in the asymptotic theory.

3. Asymptotic Properties

In this section, we first study the asymptotic properties of the oracle estimator when the important components are known a priori. Then, we show that the oracle estimator is a local minimizer of $S_n(\boldsymbol{\theta})$.

3.1. Oracle estimator

Assume there are q nonzero components in $\{f_{k,\tau}(\cdot), 1 \leq k \leq s\}$. In particular, we denote by $\mathcal{S}^* = \{k_1, k_2, \dots, k_q\} \subseteq \{1, 2, \dots, s\}$ the index set of the important components, with $|\mathcal{S}^*| = q$, where $|\cdot|$ denotes the cardinality of a set. Denote the corresponding transformed scores as $\boldsymbol{\zeta}_{i,\mathcal{S}^*} = (\zeta_{i,k_1}, \dots, \zeta_{i,k_q})^T$. Then, $\boldsymbol{\theta}$ can be divided as $\boldsymbol{\theta}_{\mathcal{S}^*} = (\theta_0, \boldsymbol{\theta}_{k_1}^T, \dots, \boldsymbol{\theta}_{k_q}^T)^T$ and $\boldsymbol{\theta}_{\mathcal{S}^{*c}}$, which is the complement vector of $\boldsymbol{\theta}_{\mathcal{S}^*}$ in $\boldsymbol{\theta}$. Similarly, define the B-spline basis $\mathbf{W}(\boldsymbol{\zeta}_{i,\mathcal{S}^*}) = (K_n^{-1/2}, \mathbf{w}(\zeta_{ik_1})^T, \dots, \mathbf{w}(\zeta_{ik_q})^T)^T$. Then, we can obtain the oracle estimator $\boldsymbol{\theta}^*$ as $\boldsymbol{\theta}_{\mathcal{S}^{*c}}^* = \mathbf{0}$ and

$$\boldsymbol{\theta}_{\mathcal{S}^*}^* = \underset{\boldsymbol{\theta}_{\mathcal{S}^*}}{\operatorname{argmin}} n^{-1} \sum_{i=1}^n \rho_\tau(y_i - \mathbf{W}(\hat{\boldsymbol{\zeta}}_{i,\mathcal{S}^*})^T \boldsymbol{\theta}_{\mathcal{S}^*}). \quad (3.1)$$

The oracle estimations for the component functions are $\alpha^*(\tau) = K_n^{-1/2} \theta_0^*$, $f_{k_j,\tau}^*(t) = \mathbf{w}(t)^T \boldsymbol{\theta}_{k_j}^*$, for $j = 1, \dots, q$, and $f_{k,\tau}^*(t) = 0$, for $k \notin \mathcal{S}^*$. The following technical conditions are imposed in order to analyze the asymptotic behavior of $\boldsymbol{\theta}^*$ and $f_{k,\tau}^*(t)$.

(C1) Condition on the functional predictor: $E(\|X(t)\|^4) < \infty$, and there exists a constant $C_\xi > 0$, such that $E(\xi_k^2 \xi_{k'}^2) \leq C_\xi \lambda_k \lambda_{k'}$ and $E(\xi_k^2 - \lambda_k)^2 < C_\xi \lambda_k^2$, for all k and $k' \neq k$. In addition, $C_\lambda^{-1} k^{-\beta} \leq \lambda_k \leq C_\lambda k^{-\beta}$ and $\lambda_k - \lambda_{k+1} \geq C_\lambda^{-1} k^{-1-\beta}$, for some constant C_λ and $\beta > 1$.

This condition is used in Wong, Li and Zhu (2018). It imposes a weak moment condition on the functional predictor, and is satisfied if $X(t)$ is a Gaussian process. In addition, it assumes the eigenvalues decay at a polynomial rate. Under this condition, we have the following lemma (Proposition 1 in Wong, Li and Zhu (2018)).

Lemma 1. *Suppose the transformation function $\Phi(\cdot)$ has a bounded derivative. Under condition (C1) and $\min_i N_i > C_1 n^{1/4}$, for some positive constant C_1 , there exists a constant C such that $E(\hat{\zeta}_{ik} - \zeta_{ik})^2 \leq Ck^2/n$ uniformly for $k \leq G_n$, where $G_n = C_2 n^{1/(2+2\beta)}$, for some constant $C_2 > 0$.*

The dense condition $\min_i N_i > C_1 n^{1/4}$ is justified in Hall, Müller and Wang (2006) to ensure that the smoothed function estimators $\hat{X}(t)$ are as good as the true functions $X(t)$, in the sense that the estimators of λ_k and ϕ_k are both first-order equivalent to those from applying a conventional principal component analysis to the true curves $X(t)$.

(C2) Condition on the random error: The random error $\epsilon_i = y_i - \alpha(\tau) - \sum_{k=1}^s f_{k,\tau}(\zeta_{ik})$ has the conditional distribution function F_i and conditional density function f_i , given $X_i(t)$. In addition, f_i is uniformly bounded away from zero and infinity in a neighborhood of zero. Its first derivative f'_i has a uniform upper bound in a neighborhood of zero, for $i = 1, \dots, n$.

(C3) Condition on the component functions: $E f_{k_j,\tau}(\zeta_{ik}) = 0$ and $f_{k_j,\tau} \in \mathcal{H}_r$, for all $1 \leq j \leq q$, where q is fixed and \mathcal{H}_r is the collection of functions f on $[0, 1]$, such that the v th-order derivative satisfies the Hölder condition of order m with $r = m + v > 3/2$, v is a positive integer, and $m \in (0, 1]$; that is, $|f^{(v)}(t) - f^{(v)}(t')| \leq C_h |t - t'|^m$, for all $t, t' \in [0, 1]$, for some positive constant C_h . In addition, $f_{k,\tau}(t) = 0$, for all $k \notin \mathcal{S}^*$.

Condition 2 is more relaxed than that usually imposed on the random error of a mean regression, which often requires independence from the predictors or homoscedasticity. Conditions 3 and 4 are typical for applications involving B-splines. From Schumaker (2007), under Condition 3, there exists $\theta_{\mathcal{S}^*}^0$, such that

$$\sup_{\mathbf{t} \in [0,1]^q} |\mathbf{W}(\mathbf{t})^T \theta_{\mathcal{S}^*}^0 - \alpha(\tau) - \sum_{j=1}^q f_{k_j,\tau}(t_j)| = O(K_n^{-r}),$$

where $\mathbf{W}(\mathbf{t}) = (K_n^{-1/2}, \mathbf{w}(t_1)^T, \dots, \mathbf{w}(t_q)^T)^T$ and $\mathbf{t} = (t_1, \dots, t_q)^T$. The following theorem summarizes the asymptotic properties of the oracle estimator.

Theorem 1. *Assume the conditions in Lemma 1 and (C2)–(C3) hold. The number of oracle predictors q is fixed and $q \leq s \leq G_n = C_2 n^{1/(2+2\beta)}$. If $K_n^3 s^2 \ll n$ and $\max\{K_n, s^2, K_n^{-2r} n\} \gg K_n^2 \{s/\sqrt{n} + K_n^{-r}\} \log n$, then we have*

$$\|\theta_{\mathcal{S}^*}^* - \theta_{\mathcal{S}^*}^0\|_2 = O_p \left(\frac{K_n}{\sqrt{n}} + \sqrt{\frac{K_n}{n}} s + K_n^{-r+1/2} \right), \tag{3.2}$$

$$n^{-1} \sum_{i=1}^n (g^*(\hat{\zeta}_{i,\mathcal{S}^*}) - g(\zeta_{i,\mathcal{S}^*}))^2 = O_p \left(\frac{K_n}{n} + \frac{s^2}{n} + K_n^{-2r} \right), \tag{3.3}$$

where $g^*(\hat{\zeta}_{i,\mathcal{S}^*}) = \alpha^*(\tau) + \sum_{j=1}^q f_{k_j,\tau}^*(\hat{\zeta}_{i,k_j})$ and $g(\zeta_{i,\mathcal{S}^*}) = \alpha(\tau) + \sum_{j=1}^q f_{k_j,\tau}(\zeta_{i,k_j})$.

Remark 1. In addition to the two terms in the convergence rate also found in nonparametric regression, we have a third term s^2/n caused by the error-contaminated predictors $\hat{\zeta}_{i,S^*}$. The technical condition $K_n^3 s^2 \ll n$ guarantees that the B-spline design matrix with estimated scores behaves as well as that with true scores. Furthermore, $\max\{K_n, s^2, K_n^{-2r} n\} \gg K_n^2 \{s/\sqrt{n} + K_n^{-r}\} \log n$ follows after applying Bernstein's inequality in the proof.

Remark 2. The optimal choice of K_n is $K_n = O(n^{1/(2r+1)})$, which satisfies the order condition $\max\{K_n, s^2, K_n^{-2r} n\} \gg K_n^2 \{s/\sqrt{n} + K_n^{-r}\} \log n$, for any s . If s is fixed, the convergence rate for g^* is the traditional nonparametric convergence rate $O_p(K_n/n + K_n^{-2r})$. This implies that $\hat{\zeta}_{i,S^*}$ converges sufficiently quickly and does not influence the global convergence rate. If $s = O(n^{1/(2+2\beta)})$ under the framework laid out in Lemma 1, the condition $K_n^3 s^2 \ll n$ becomes $3/(2r+1) + 1/(1+\beta) < 1$. In this case, there is a trade-off between r and β . A smaller β allows us to consistently estimate a higher number of scores. However, we require a stronger smoothness assumption on the functions to be estimated; that is, $f_{k_j, \tau}$ must satisfy the Hölder condition with a larger r .

3.2. SCAD penalized estimator

To investigate the asymptotic properties of the SCAD penalized estimator, we need an additional condition on how quickly the nonzero signal can decay.

(C4) Condition on the signal strength: The minimal signal $\min_{k \in S^*} \|\theta_k^0\|_2 \geq C\sqrt{K_n/n}(K_n^{1/2} + s)n^\alpha$, for some positive constants C and α .

Owing to the nonsmoothness and nonconvexity of the penalized objective function, in our proof, we use a sufficient condition for the local minimizer of a convex difference problem, as in Wang, Wu and Li (2012) and Sherwood and Wang (2016). Specifically, the penalized objective function can be represented as the difference of two convex functions. By verifying that the oracle estimator meets the sufficient condition, we obtain our main theorem. Let $\mathcal{E}(\lambda)$ be the set of local minima of $S_n(\theta)$. The oracle estimator belongs to the set $\mathcal{E}(\lambda)$, with probability approaching one.

Theorem 2. *Assume the conditions in Theorem 1 and (C4) are satisfied. The optimal rate $K_n = O(n^{1/(2r+1)})$ is used. Let θ^* be the oracle estimator defined in (3.1). If $\max\{1/\sqrt{n}, s/\sqrt{K_n n}\} \ll \lambda \ll \sqrt{K_n/n}(K_n^{1/2} + s)n^\alpha$, then $P(\theta^* \in \mathcal{E}(\lambda)) \rightarrow 1$ as $n \rightarrow \infty$.*

3.3. Consistency of the BIC in model selection

In this section, we investigate the consistency of the BIC in terms of model selection. The generic notation $\mathcal{S} \subseteq \{1, \dots, s\}$ denotes an arbitrary candidate model. We define the BIC for model \mathcal{S} as

$$\text{BIC}(\mathcal{S}) = \log \sum_{i=1}^n \rho_\tau(y_i - \mathbf{W}(\hat{\boldsymbol{\zeta}}_{i,\mathcal{S}})^T \hat{\boldsymbol{\theta}}_{\mathcal{S}}) + J_{\mathcal{S}} \frac{\log n}{2n} C_n, \tag{3.4}$$

where $\hat{\boldsymbol{\theta}}_{\mathcal{S}}$ is the estimator under model \mathcal{S} , $J_{\mathcal{S}} = (1 + (K_n + l)|\mathcal{S}|)$, and C_n diverges to infinity. The diverging order of C_n is specified later. The main challenge is that the number of candidate models increases exponentially with s . We use additional conditions to establish the model selection consistency of the BIC for our model.

(C5) Condition on the spline design covariates:

- (i) $\max_{1 \leq k \leq s} \|W(\zeta_{i,k})\| = O_p(1)$,
- (ii) $b_2/K_n \leq \inf_{\mathcal{S}} E\lambda_{\min}(W(\zeta_{i,\mathcal{S}})W(\zeta_{i,\mathcal{S}})^T) \leq \sup_{\mathcal{S}} E\lambda_{\max}(W(\zeta_{i,\mathcal{S}})W(\zeta_{i,\mathcal{S}})^T) \leq b_2^*/K_n$, for some positive constants b_2 and b_2^* .

Condition (C5) is not strong, and is satisfied in most nonparametric estimations based on a B-spline basis approximation (see Lee, Noh and Park (2014)). It is well known that $\|W(\zeta_{i,k})\| = O_p(1)$, for every $1 \leq k \leq s$. In addition, we have $b_2/K_n \leq E\lambda_{\min}(W(\zeta_{i,\mathcal{S}})W(\zeta_{i,\mathcal{S}})^T) \leq \sup_{\mathcal{S}} E\lambda_{\max}(W(\zeta_{i,\mathcal{S}})W(\zeta_{i,\mathcal{S}})^T) \leq b_2^*/K_n$, for every \mathcal{S} (see the proof of Lemma S1.1(2)). Condition (C5) specifies a uniform version. Using these conditions, we obtain the model selection consistency of the BIC when s is diverging under the framework laid out in Lemma 1.

Theorem 3. *Assume the conditions in Theorem 2 and (C5) are satisfied. The number of candidate components $s = O(n^{1/2(1+\beta)})$ and $\min_{k \in \mathcal{S}^*} \|\boldsymbol{\theta}_k^0\| \geq C\sqrt{K_n}$. For any sequence $C_n \rightarrow \infty$ satisfying $K_n(\log n/(2n))C_n \rightarrow 0$ and $s/(K_n C_n) \rightarrow 0$, we have*

$$P\left(\inf_{\mathcal{S} \neq \mathcal{S}^*} \text{BIC}(\mathcal{S}) > \text{BIC}(\mathcal{S}^*)\right) \rightarrow 1.$$

Let S_λ denote the model selected using the penalized estimation with λ . From the definitions, $\text{BIC}(\lambda) \geq \text{BIC}(S_\lambda)$ because the latter uses an unpenalized estimator. In addition, Theorem 1 implies that, with high probability, the oracle estimator can be produced by some λ^* on the solution path; thus, $\text{BIC}(\lambda^*) = \text{BIC}(S_{\lambda^*})$. Therefore, by Theorem 3, for any λ not inducing the oracle model, we have $\text{BIC}(\lambda) \geq \text{BIC}(S_\lambda) > \text{BIC}(S_{\lambda^*}) = \text{BIC}(\lambda^*)$. This suggests that the BIC is consistent for the tuning parameter selection.

4. Simulation Studies

We conduct simulation studies to illustrate the empirical performance of the proposed method. We generate predictor trajectories $X_i(t)$ over a grid with 100 equally spaced points over $0 \leq t \leq 10$ from

$$X_i(t) = t + \sin(t) + \sum_{k=1}^{10} \xi_{ik} \phi_k(t),$$

where $\xi_{ik} \sim N(0, \lambda_k)$, $\lambda_k = 30k^{-2}$, $\text{corr}(\xi_{ik}, \xi_{ik'}) = 0$, for $k \neq k'$, and $\phi_k(t) = (1/\sqrt{10})\sin(\pi kt/10 + \pi/4)$. This is the Karhunen–Loève expansion, and the scores are independent. The measurement errors are generated independently from $N(0, 0.2^2)$. For the regression function, we set $f_1(\zeta_1) = 3\zeta_1 - 3/2$, $f_2(\zeta_2) = \sin(2\pi(\zeta_2 - 1/2))$, $f_3(\zeta_3) = 8(\zeta_3 - 1/3)^2$, and $f_4(\zeta_4) = 8(\zeta_4 - 1/3)^2 - 8/9$. Then, we generate responses using five models:

Model 1: $y_i = 1.4 + f_1(\zeta_{i1}) + f_2(\zeta_{i2}) + f_4(\zeta_{i4}) + \epsilon_i$, $\epsilon_i \sim N(0, 1)$;

Model 2: $y_i = 1.4 + f_1(\zeta_{i1}) + f_2(\zeta_{i2}) + f_4(\zeta_{i4}) + \epsilon_i$, $\epsilon_i \sim t(5)$;

Model 3: $y_i = 1.4 + f_1(\zeta_{i1}) + f_2(\zeta_{i2}) + f_4(\zeta_{i4}) + f_3(\zeta_{i3})\epsilon_i$, $\epsilon_i \sim N(0, 1)$;

Model 4: $y_i = 1.4 + f_1(\zeta_{i1}) + f_2(\zeta_{i2}) + f_4(\zeta_{i10}) + \epsilon_i$, where ϵ_i is $N(0.5, 0.2)$ when $u < 0.3$, and $N(-0.5, 1)$ when $u > 0.7$ with $u \sim \text{Uniform}(0, 1)$;

Model 5: $y_i = 1.4 + f_1(\zeta_{i1}) + f_2(\zeta_{i2}) + f_{10}(\zeta_{i10}) + \epsilon_i$, where $\epsilon_i \sim N(0, 1)$ and $f_{10}(x) = 8(x - 1/3)^2 - 8/9$.

Model 1 is homoscedastic and the τ th conditional quantile is $Q_{y_i}(\tau|X_i) = 1.4 + \Phi^{-1}(\tau) + f_1(\zeta_{i1}) + f_2(\zeta_{i2}) + f_4(\zeta_{i4})$. For identification, we fit the demeaned functions $f_1(\zeta_1) - E(f_1(\zeta_1))$, $f_2(\zeta_2) - E(f_2(\zeta_2))$, and $f_4(\zeta_4) - E(f_4(\zeta_4))$. With a slightly abuse of notation, we still denote these demeaned functions as $f_1(\zeta_1)$, $f_2(\zeta_2)$, and $f_4(\zeta_4)$ respectively. Model 2 mimics Model 1 with a heavy-tailed error. Model 3 is heteroscedastic and the τ th conditional quantile is $Q_{y_i}(\tau|X_i) = 1.4 + f_3(\zeta_{i3})\Phi^{-1}(\tau) + f_1(\zeta_{i1}) + f_2(\zeta_{i2}) + f_4(\zeta_{i4})$. Here, ζ_3 plays a role in the conditional distribution of y given $X(t)$, but does not directly influence the center (mean or median) of the conditional distribution. Model 4 mimics Model 1 with a bimodal distribution. Model 5 uses the 10th score instead of the fourth score in Model 1. Note that the estimation of the 10th score is more challenging than that of the fourth score.

We consider sample sizes $n = 200$ and $n = 400$ with three quantiles $\tau = 0.1, 0.5$, and 0.9 . We choose s to recover at least 99.9% of the total variation in

$X(t)$, use cubic splines with two knots, and select the tuning parameters using the BIC with $C_n = 1.5$, which is shown by Lee, Noh and Park (2014) to perform well when s is small relative to n . One can select the number of knots using cross-validation or some other information criterion, but this would increase the computational burden, without offering appreciable numerical advantages. In the literature, using a pre-fixed number of basis functions is not uncommon; see for example, Huang, Horowitz and Wei (2010) and Fan, Feng and Song (2011). This choice of the number of knots is small enough to avoid overfitting in typical problems in which the sample size is not too small, and big enough to flexibly approximate many smooth functions accurately.

Based on 500 repetitions, Tables 1–5 report the component-selection results for Models 1–5 respectively. We show the selection percentages for the first six component functions from our functional additive modeling and quantile linear regression modeling (also with a SCAD penalty for variable selection) for Models 1–5 (for Model 5, we show the results for the first five and the 10th component functions). The column “correct set” corresponds to the percentages of exact selection, and the column “super set” gives the percentages of fittings that include all nonzero functions. For the linear regression modeling, we also use the SCAD penalty to select important components, and use the ordinary BIC to select the tuning parameter. One can see that the functional additive model performs slightly better than the functional linear model. In addition, the linear model tends to select slightly larger model sizes.

For the estimation accuracy of f_k , we calculate the average integrated squared errors (AISEs) of the estimated component functions. The integrated squared errors are defined as

$$\text{ISE}(\hat{f}_k) = \int_0^1 \{\hat{f}_k(t) - f_k(t)\}^2 dt.$$

Tables 6–10 show the AISEs of the six component functions for Models 1–5. The AISE for \hat{g} is the summation of those of the first six component functions. The last column shows the prediction errors, computed as $(1/n) \sum_{i=1}^n \{\hat{Q}_{y_i}(\tau|X_i) - Q_{y_i}(\tau|X_i)\}^2$, on 200 newly generated testing points. To compute the prediction, we first estimate the transformed FPC scores of $X_i(t)$ in the test set using the estimates of the mean function, eigenvalues, and eigenfunctions from the training data. Then, we substitute these scores into the estimated component functions. Note that the functional additive model performs significantly better than the functional linear model, yielding smaller estimation and prediction errors.

Table 1. Model 1: Selection percentages of the first six component functions and the mean of the selected model size (last column).

Model 1			1	2	3	4	5	6	correct set	super set	model size
$n = 200$	$\tau = 0.1$	QFAM	1.00	0.84	0.03	0.85	0.07	0.05	0.52	0.73	2.98
		QFLM	1.00	0.84	0.20	0.90	0.27	0.14	0.21	0.76	4.00
	$\tau = 0.5$	QFAM	1.00	0.80	0.01	0.99	0.01	0.00	0.78	0.80	2.82
		QFLM	1.00	0.86	0.12	1.00	0.20	0.05	0.52	0.86	3.33
	$\tau = 0.9$	QFAM	1.00	0.85	0.06	1.00	0.08	0.02	0.67	0.85	3.07
		QFLM	1.00	0.79	0.19	1.00	0.22	0.10	0.27	0.79	3.85
$n = 400$	$\tau = 0.1$	QFAM	1.00	0.99	0.01	1.00	0.05	0.01	0.88	0.99	3.10
		QFLM	1.00	0.98	0.18	0.99	0.34	0.09	0.36	0.98	4.00
	$\tau = 0.5$	QFAM	1.00	0.99	0.00	1.00	0.01	0.00	0.99	0.99	3.00
		QFLM	1.00	0.99	0.14	1.00	0.28	0.04	0.56	0.99	3.55
	$\tau = 0.9$	QFAM	1.00	0.99	0.04	1.00	0.06	0.00	0.88	0.99	3.12
		QFLM	1.00	0.97	0.16	1.00	0.28	0.07	0.38	0.97	3.91

QFAM: quantile functional additive model; QFLM: quantile functional linear model.

Table 2. Model 2: Selection percentages of the first six component functions and the mean of the selected model size (last column).

Model 2			1	2	3	4	5	6	correct set	super set	model size
$n = 200$	$\tau = 0.1$	QFAM	1.00	0.60	0.01	0.64	0.04	0.02	0.34	0.46	2.43
		QFLM	1.00	0.72	0.17	0.83	0.24	0.13	0.17	0.61	3.69
	$\tau = 0.5$	QFAM	1.00	0.64	0.00	0.96	0.01	0.00	0.63	0.64	2.61
		QFLM	1.00	0.83	0.11	0.99	0.17	0.03	0.51	0.82	3.26
	$\tau = 0.9$	QFAM	1.00	0.70	0.04	0.97	0.04	0.02	0.57	0.69	2.84
		QFLM	1.00	0.71	0.16	0.99	0.22	0.10	0.29	0.70	3.68
$n = 400$	$\tau = 0.1$	QFAM	1.00	0.90	0.02	0.96	0.03	0.01	0.78	0.88	2.98
		QFLM	1.00	0.88	0.14	0.97	0.26	0.09	0.32	0.86	3.80
	$\tau = 0.5$	QFAM	1.00	0.98	0.00	1.00	0.01	0.00	0.97	0.98	2.99
		QFLM	1.00	0.97	0.09	1.00	0.23	0.02	0.61	0.97	3.41
	$\tau = 0.9$	QFAM	1.00	0.89	0.03	1.00	0.03	0.01	0.81	0.89	2.97
		QFLM	1.00	0.87	0.12	1.00	0.17	0.09	0.42	0.87	3.64

QFAM: quantile functional additive model; QFLM: quantile functional linear model.

Table 3. Model 3: Selection percentages of the first six component functions and the mean of the selected model size (last column).

Model 3			1	2	3	4	5	6	correct set	super set	model size
$n = 200$	$\tau = 0.1$	QFAM	1.00	0.70	0.86	0.91	0.06	0.02	0.52	0.63	3.65
		QFLM	1.00	0.58	0.89	0.84	0.24	0.10	0.16	0.43	4.16
	$\tau = 0.5$	QFAM	1.00	0.96	0.02	0.99	0.01	0.00	0.94	0.96	2.98
		QFLM	1.00	0.90	0.12	1.00	0.17	0.02	0.60	0.90	3.27
	$\tau = 0.9$	QFAM	1.00	0.93	0.95	1.00	0.05	0.01	0.80	0.90	4.00
		QFLM	1.00	0.80	0.92	0.99	0.17	0.08	0.37	0.74	4.46
$n = 400$	$\tau = 0.1$	QFAM	1.00	0.98	1.00	1.00	0.08	0.00	0.88	0.98	4.09
		QFLM	1.00	0.83	0.98	0.98	0.30	0.08	0.38	0.80	4.54
	$\tau = 0.5$	QFAM	1.00	1.00	0.01	1.00	0.01	0.00	0.98	1.00	3.02
		QFLM	1.00	0.99	0.15	1.00	0.29	0.02	0.55	0.99	3.53
	$\tau = 0.9$	QFAM	1.00	1.00	1.00	1.00	0.05	0.00	0.94	1.00	4.05
		QFLM	1.00	0.96	0.99	1.00	0.17	0.07	0.54	0.96	4.55

QFAM: quantile functional additive model; QFLM: quantile functional linear model.

Table 4. Model 4: Selection percentages of the first six component functions and the mean of the selected model size (last column).

Model 4			1	2	3	4	5	6	correct set	super set	model size
$n = 200$	$\tau = 0.1$	QFAM	1.00	0.78	0.04	0.82	0.09	0.04	0.45	0.68	2.95
		QFLM	1.00	0.80	0.21	0.90	0.26	0.14	0.20	0.72	3.98
	$\tau = 0.5$	QFAM	1.00	0.82	0.00	0.98	0.02	0.00	0.79	0.81	2.83
		QFLM	1.00	0.91	0.16	1.00	0.20	0.05	0.53	0.90	3.43
	$\tau = 0.9$	QFAM	1.00	0.96	0.10	1.00	0.06	0.00	0.80	0.95	3.13
		QFLM	1.00	0.90	0.19	1.00	0.22	0.07	0.37	0.89	3.82
$n = 400$	$\tau = 0.1$	QFAM	1.00	0.97	0.03	0.99	0.04	0.01	0.85	0.97	3.10
		QFLM	1.00	0.94	0.18	1.00	0.29	0.12	0.32	0.94	4.13
	$\tau = 0.5$	QFAM	1.00	1.00	0.00	1.00	0.01	0.00	0.98	1.00	3.01
		QFLM	1.00	0.99	0.16	1.00	0.30	0.02	0.54	0.99	3.55
	$\tau = 0.9$	QFAM	1.00	1.00	0.08	1.00	0.09	0.00	0.84	1.00	3.17
		QFLM	1.00	0.98	0.13	1.00	0.25	0.09	0.43	0.98	3.84

QFAM: quantile functional additive model; QFLM: quantile functional linear model.

Table 5. Model 5: Selection percentages of the first six component functions and the mean of the selected model size (last column).

Model 5			1	2	3	4	5	10	correct set	super set	model size
$n = 200$	$\tau = 0.1$	QFAM	1.00	0.74	0.02	0.01	0.03	0.61	0.39	0.51	2.52
		QFLM	1.00	0.85	0.12	0.15	0.15	0.82	0.20	0.70	3.92
	$\tau = 0.5$	QFAM	1.00	0.82	0.00	0.00	0.00	0.95	0.78	0.80	2.80
		QFLM	1.00	0.88	0.06	0.04	0.04	1.00	0.45	0.88	3.52
	$\tau = 0.9$	QFAM	1.00	0.80	0.01	0.00	0.02	0.99	0.60	0.79	3.04
		QFLM	1.00	0.83	0.14	0.09	0.11	1.00	0.24	0.83	4.05
$n = 400$	$\tau = 0.1$	QFAM	1.00	0.98	0.02	0.01	0.01	0.97	0.82	0.95	3.10
		QFLM	1.00	0.98	0.15	0.09	0.11	0.98	0.33	0.97	4.08
	$\tau = 0.5$	QFAM	1.00	1.00	0.00	0.00	0.00	1.00	0.97	1.00	3.02
		QFLM	1.00	1.00	0.06	0.03	0.02	1.00	0.53	1.00	3.58
	$\tau = 0.9$	QFAM	1.00	0.99	0.01	0.00	0.00	1.00	0.88	0.99	3.11
		QFLM	1.00	0.96	0.11	0.09	0.09	1.00	0.36	0.96	4.01

QFAM: quantile functional additive model; QFLM: quantile functional linear model.

5. Empirical Application

Weather has a significant effect on crop yield, as a result, many studies have developed statistical models of this relation (Cadson, Todey and Taylor (1996), Prasad et al. (2006), Lobell and Burke (2010)). In this section, we apply our proposed method to the corn yield data set of Wong, Li and Zhu (2018). We have yield-related variables for corn from 105 counties in Kansas for the period 1999 to 2011. The data set contains the average corn yield per acre for a specific year and county, which is the scalar outcome of interest. The functional predictor is $X(t) = (X_1(t) + X_2(t))/2$, where $X_1(t)$ is the daily maximum temperature trajectory, and $X_2(t)$ is the daily minimum temperature trajectory with the time domain $\mathcal{T} = [0, 365]$. These are gathered from 1,123 weather stations and aggregated at the county level. After deleting missing data, we use the remaining 857 observations

Table 6. Model 1: Average integrated squared errors (AISEs) and standard deviations (in parentheses).

Model 1			\hat{f}_1	\hat{f}_2	\hat{f}_3	\hat{f}_4	\hat{f}_5	\hat{f}_6	\hat{g}	prediction error
$n = 200$	$\tau = 0.1$	QFAM	0.11	0.18	0.01	0.27	0.01	0.01	0.58	0.88
			(0.07)	(0.16)	(0.04)	(0.34)	(0.05)	(0.04)	(0.40)	(0.42)
		QFLM	2.80	0.79	0.09	1.65	0.13	0.04	5.50	1.02
			(1.02)	(0.36)	(0.26)	(0.76)	(0.26)	(0.13)	(1.47)	(0.28)
	$\tau = 0.5$	QFAM	0.07	0.16	0.00	0.13	0.00	0.00	0.36	0.58
			(0.04)	(0.17)	(0.03)	(0.30)	(0.04)	(0.01)	(0.38)	(0.22)
	QFLM	2.67	0.76	0.05	1.76	0.09	0.02	5.35	0.74	
		(0.79)	(0.27)	(0.17)	(0.53)	(0.20)	(0.08)	(0.99)	(0.11)	
$\tau = 0.9$	QFAM	0.11	0.18	0.02	0.17	0.02	0.00	0.51	0.77	
		(0.07)	(0.15)	(0.07)	(0.32)	(0.08)	(0.03)	(0.41)	(0.27)	
	QFLM	2.62	0.77	0.08	1.93	0.09	0.03	5.52	0.96	
		(0.98)	(0.33)	(0.20)	(0.67)	(0.20)	(0.12)	(1.25)	(0.22)	
$n = 400$	$\tau = 0.1$	QFAM	0.05	0.08	0.00	0.08	0.01	0.00	0.22	0.46
			(0.03)	(0.06)	(0.02)	(0.14)	(0.03)	(0.01)	(0.17)	(0.16)
		QFLM	2.70	0.75	0.05	1.69	0.10	0.02	5.30	0.85
			(0.67)	(0.25)	(0.13)	(0.57)	(0.18)	(0.06)	(0.99)	(0.16)
	$\tau = 0.5$	QFAM	0.03	0.06	0.00	0.05	0.00	0.00	0.14	0.36
			(0.02)	(0.05)	(0.00)	(0.13)	(0.01)	(0.00)	(0.14)	(0.11)
	QFLM	2.68	0.73	0.03	1.78	0.07	0.01	5.30	0.68	
		(0.57)	(0.20)	(0.10)	(0.37)	(0.14)	(0.04)	(0.69)	(0.08)	
$\tau = 0.9$	QFAM	0.05	0.08	0.01	0.06	0.01	0.00	0.21	0.45	
		(0.03)	(0.06)	(0.03)	(0.13)	(0.05)	(0.01)	(0.16)	(0.13)	
	QFLM	2.62	0.72	0.03	1.87	0.07	0.01	5.32	0.85	
		(0.68)	(0.24)	(0.09)	(0.40)	(0.15)	(0.06)	(0.84)	(0.16)	

QFAM: quantile functional additive model; QFLM: quantile functional linear model.

Table 7. Model 2: Average integrated squared errors (AISEs) and standard deviations (in parentheses).

Model 2			\hat{f}_1	\hat{f}_2	\hat{f}_3	\hat{f}_4	\hat{f}_5	\hat{f}_6	\hat{g}	prediction error
$n = 200$	$\tau = 0.1$	QFAM	0.16	0.31	0.00	0.44	0.01	0.00	0.92	1.26
			(0.11)	(0.21)	(0.03)	(0.39)	(0.05)	(0.04)	(0.49)	(0.57)
		QFLM	2.65	0.80	0.10	1.68	0.14	0.05	5.43	1.11
			(1.10)	(0.43)	(0.27)	(0.86)	(0.35)	(0.17)	(1.60)	(0.34)
	$\tau = 0.5$	QFAM	0.08	0.23	0.00	0.15	0.00	0.00	0.46	0.66
			(0.05)	(0.23)	(0.00)	(0.31)	(0.02)	(0.00)	(0.42)	(0.27)
	QFLM	2.72	0.79	0.05	1.80	0.08	0.01	5.44	0.76	
		(0.84)	(0.31)	(0.16)	(0.60)	(0.19)	(0.07)	(1.10)	(0.12)	
$\tau = 0.9$	QFAM	0.18	0.28	0.02	0.23	0.01	0.00	0.72	1.07	
		(0.13)	(0.22)	(0.09)	(0.35)	(0.06)	(0.03)	(0.47)	(0.45)	
	QFLM	2.76	0.85	0.09	1.94	0.12	0.04	5.80	1.06	
		(1.20)	(0.45)	(0.25)	(0.71)	(0.30)	(0.15)	(1.49)	(0.29)	
$n = 400$	$\tau = 0.1$	QFAM	0.08	0.14	0.00	0.12	0.00	0.00	0.35	0.61
			(0.05)	(0.13)	(0.02)	(0.16)	(0.03)	(0.01)	(0.23)	(0.27)
		QFLM	2.69	0.76	0.05	1.68	0.10	0.02	5.30	0.90
			(0.79)	(0.32)	(0.14)	(0.64)	(0.22)	(0.09)	(1.13)	(0.22)
	$\tau = 0.5$	QFAM	0.04	0.07	0.00	0.04	0.00	0.00	0.15	0.37
			(0.02)	(0.07)	(0.01)	(0.08)	(0.02)	(0.00)	(0.11)	(0.12)
	QFLM	2.70	0.74	0.03	1.79	0.07	0.00	5.32	0.68	
		(0.57)	(0.22)	(0.09)	(0.35)	(0.14)	(0.02)	(0.72)	(0.08)	
$\tau = 0.9$	QFAM	0.08	0.15	0.01	0.09	0.00	0.00	0.33	0.60	
		(0.05)	(0.14)	(0.04)	(0.08)	(0.03)	(0.02)	(0.19)	(0.22)	
	QFLM	2.66	0.76	0.04	1.90	0.05	0.02	5.44	0.90	
		(0.83)	(0.30)	(0.12)	(0.44)	(0.14)	(0.09)	(0.93)	(0.18)	

QFAM: quantile functional additive model; QFLM: quantile functional linear model.

Table 8. Model 3: Average integrated squared errors (AISEs) and standard deviations (in parentheses).

Model 3			\hat{f}_1	\hat{f}_2	\hat{f}_3	\hat{f}_4	\hat{f}_5	\hat{f}_6	\hat{g}	prediction error	
$n = 200$	$\tau = 0.1$	QFAM	0.06 (0.05)	0.21 (0.20)	0.35 (0.33)	0.18 (0.25)	0.01 (0.04)	0.00 (0.01)	0.81 (0.59)	1.52 (0.74)	
		QFLM	2.45 (0.98)	0.60 (0.21)	1.64 (0.66)	1.44 (0.90)	0.10 (0.22)	0.03 (0.10)	6.26 (1.60)	2.10 (0.45)	
	$\tau = 0.5$	QFAM	0.02 (0.02)	0.07 (0.09)	0.00 (0.02)	0.09 (0.21)	0.00 (0.01)	0.00 (0.00)	0.18 (0.25)	0.46 (0.19)	
		QFLM	2.45 (0.65)	0.70 (0.23)	0.05 (0.16)	1.28 (0.49)	0.06 (0.14)	0.00 (0.03)	4.54 (0.83)	0.73 (0.11)	
	$\tau = 0.9$	QFAM	0.06 (0.05)	0.11 (0.12)	0.25 (0.24)	0.12 (0.24)	0.01 (0.03)	0.00 (0.01)	0.55 (0.48)	1.20 (0.57)	
		QFLM	2.51 (0.98)	0.90 (0.49)	1.90 (0.83)	1.14 (0.53)	0.07 (0.18)	0.03 (0.13)	6.56 (1.39)	2.03 (0.37)	
	$n = 400$	$\tau = 0.1$	QFAM	0.02 (0.02)	0.07 (0.07)	0.15 (0.10)	0.07 (0.06)	0.00 (0.02)	0.00 (0.00)	0.32 (0.15)	0.85 (0.31)
			QFLM	2.43 (0.62)	0.55 (0.12)	1.56 (0.52)	1.34 (0.59)	0.08 (0.15)	0.01 (0.06)	5.97 (1.06)	1.89 (0.31)
		$\tau = 0.5$	QFAM	0.01 (0.01)	0.03 (0.02)	0.00 (0.02)	0.04 (0.06)	0.00 (0.01)	0.00 (0.00)	0.09 (0.07)	0.32 (0.11)
			QFLM	2.43 (0.46)	0.68 (0.17)	0.04 (0.10)	1.28 (0.29)	0.06 (0.10)	0.00 (0.02)	4.49 (0.56)	0.68 (0.08)
		$\tau = 0.9$	QFAM	0.03 (0.02)	0.04 (0.03)	0.13 (0.09)	0.05 (0.09)	0.00 (0.01)	0.00 (0.00)	0.25 (0.13)	0.69 (0.24)
			QFLM	2.49 (0.68)	0.84 (0.32)	1.81 (0.57)	1.11 (0.34)	0.04 (0.11)	0.02 (0.07)	6.30 (0.93)	1.84 (0.26)

QFAM: quantile functional additive model; QFLM: quantile functional linear model.

Table 9. Model 4: Average integrated squared errors (AISEs) and standard deviations (in parentheses).

Model 4			\hat{f}_1	\hat{f}_2	\hat{f}_3	\hat{f}_4	\hat{f}_5	\hat{f}_6	\hat{g}	prediction error	
$n = 200$	$\tau = 0.1$	QFAM	0.13 (0.08)	0.21 (0.17)	0.01 (0.04)	0.28 (0.32)	0.02 (0.08)	0.01 (0.04)	0.65 (0.43)	0.96 (0.42)	
		QFLM	2.75 (1.04)	0.80 (0.38)	0.10 (0.24)	1.66 (0.77)	0.14 (0.30)	0.04 (0.14)	5.49 (1.44)	1.01 (0.27)	
	$\tau = 0.5$	QFAM	0.06 (0.04)	0.15 (0.17)	0.00 (0.01)	0.11 (0.22)	0.00 (0.04)	0.00 (0.00)	0.33 (0.30)	0.60 (0.24)	
		QFLM	2.73 (0.78)	0.78 (0.31)	0.06 (0.16)	1.75 (0.53)	0.09 (0.21)	0.01 (0.06)	5.43 (1.02)	0.77 (0.13)	
	$\tau = 0.9$	QFAM	0.07 (0.05)	0.11 (0.10)	0.02 (0.06)	0.10 (0.23)	0.01 (0.06)	0.00 (0.01)	0.31 (0.29)	0.63 (0.24)	
		QFLM	2.72 (0.89)	0.77 (0.33)	0.06 (0.14)	1.88 (0.52)	0.08 (0.19)	0.02 (0.07)	5.52 (1.07)	1.11 (0.27)	
	$n = 400$	$\tau = 0.1$	QFAM	0.06 (0.04)	0.10 (0.09)	0.00 (0.03)	0.09 (0.14)	0.01 (0.03)	0.00 (0.01)	0.26 (0.18)	0.51 (0.18)
			QFLM	2.68 (0.75)	0.75 (0.29)	0.05 (0.14)	1.71 (0.59)	0.09 (0.18)	0.02 (0.08)	5.31 (1.07)	0.85 (0.17)
		$\tau = 0.5$	QFAM	0.03 (0.02)	0.05 (0.04)	0.00 (0.01)	0.05 (0.13)	0.00 (0.01)	0.00 (0.00)	0.14 (0.14)	0.37 (0.12)
			QFLM	2.66 (0.55)	0.76 (0.22)	0.04 (0.09)	1.78 (0.40)	0.08 (0.13)	0.00 (0.02)	5.31 (0.73)	0.71 (0.10)
		$\tau = 0.9$	QFAM	0.03 (0.02)	0.06 (0.04)	0.01 (0.03)	0.05 (0.16)	0.01 (0.03)	0.00 (0.00)	0.15 (0.18)	0.43 (0.16)
			QFLM	2.63 (0.60)	0.74 (0.22)	0.02 (0.08)	1.91 (0.36)	0.06 (0.12)	0.01 (0.05)	5.37 (0.73)	1.03 (0.21)

QFAM: quantile functional additive model; QFLM: quantile functional linear model.

Table 10. Model 5: Average integrated squared errors (AISEs) and standard deviations (in parentheses).

Model 5			\hat{f}_1	\hat{f}_2	\hat{f}_3	\hat{f}_4	\hat{f}_5	\hat{f}_{10}	\hat{g}	prediction error	
$n = 200$	$\tau = 0.1$	QFAM	0.11 (0.07)	0.21 (0.18)	0.00 (0.02)	0.00 (0.02)	0.00 (0.02)	0.45 (0.38)	0.80 (0.47)	1.22 (0.46)	
		QFLM	2.62 (0.94)	0.77 (0.35)	0.04 (0.12)	0.04 (0.13)	0.05 (0.14)	1.33 (0.61)	5.11 (1.31)	1.16 (0.32)	
	$\tau = 0.5$	QFAM	0.07 (0.05)	0.15 (0.17)	0.00 (0.01)	0.00 (0.00)	0.00 (0.00)	0.18 (0.31)	0.41 (0.41)	0.79 (0.27)	
		QFLM	2.56 (0.76)	0.77 (0.29)	0.02 (0.09)	0.01 (0.07)	0.01 (0.06)	1.59 (0.51)	5.17 (0.99)	0.84 (0.12)	
	$\tau = 0.9$	QFAM	0.12 (0.08)	0.21 (0.17)	0.00 (0.02)	0.00 (0.01)	0.00 (0.04)	0.21 (0.34)	0.61 (0.45)	1.05 (0.36)	
		QFLM	2.58 (0.98)	0.79 (0.36)	0.06 (0.17)	0.03 (0.11)	0.04 (0.13)	1.88 (0.68)	5.69 (1.26)	1.08 (0.24)	
	$n = 400$	$\tau = 0.1$	QFAM	0.05 (0.03)	0.08 (0.07)	0.00 (0.01)	0.00 (0.01)	0.00 (0.01)	0.14 (0.17)	0.29 (0.19)	0.66 (0.21)
			QFLM	2.59 (0.66)	0.75 (0.25)	0.03 (0.09)	0.02 (0.06)	0.02 (0.06)	1.30 (0.40)	4.87 (0.84)	0.96 (0.19)
		$\tau = 0.5$	QFAM	0.03 (0.02)	0.05 (0.04)	0.00 (0.00)	0.00 (0.00)	0.00 (0.00)	0.06 (0.09)	0.15 (0.11)	0.51 (0.12)
			QFLM	2.57 (0.56)	0.75 (0.20)	0.02 (0.07)	0.01 (0.04)	0.00 (0.03)	1.62 (0.32)	5.08 (0.70)	0.75 (0.08)
		$\tau = 0.9$	QFAM	0.05 (0.04)	0.08 (0.06)	0.00 (0.01)	0.00 (0.01)	0.00 (0.01)	0.08 (0.12)	0.24 (0.16)	0.64 (0.17)
			QFLM	2.59 (0.70)	0.74 (0.26)	0.02 (0.07)	0.02 (0.07)	0.02 (0.06)	1.90 (0.47)	5.44 (0.88)	0.96 (0.17)

QFAM: quantile functional additive model; QFLM: quantile functional linear model.

Table 11. Average model size across 500 bootstrap samples and the average prediction error (standard deviation in parentheses) for five-fold cross-validation across 100 random partitions for corn yield data.

	$\tau = 0.1$		$\tau = 0.5$		$\tau = 0.9$	
	QFAM	QFLM	QFAM	QFLM	QFAM	QFLM
MS	7.42	39.89	14.95	40.72	11.69	42.91
PE-BIC	5.77 (0.60)	5.42 (0.30)	14.70 (0.90)	13.13 (0.64)	5.77 (0.56)	5.34 (0.33)
PE-CV	5.77 (0.60)	5.37 (0.34)	13.94 (0.71)	12.89 (0.63)	5.77 (0.56)	5.25 (0.32)

MS: model size; PE-BIC: prediction error with BIC tuning; PE-CV: prediction error with CV tuning; QFAM: quantile functional additive model; QFLM: quantile functional linear model.

in our analysis.

For the data exploration, we include 45 principal components in the regression model, which account for 99.9% of the variation in the daily temperature trajectories. Then, we fit the proposed functional additive quantile regression and linear quantile regression for $\tau = 0.1, 0.5, 0.9$, separately. The linear model is also equipped with a SCAD penalty for the variable selection. We choose the tuning parameter using both the BIC and five-fold cross-validation (CV). The results in Table 11 show that the tuning parameters chosen by CV produce bet-

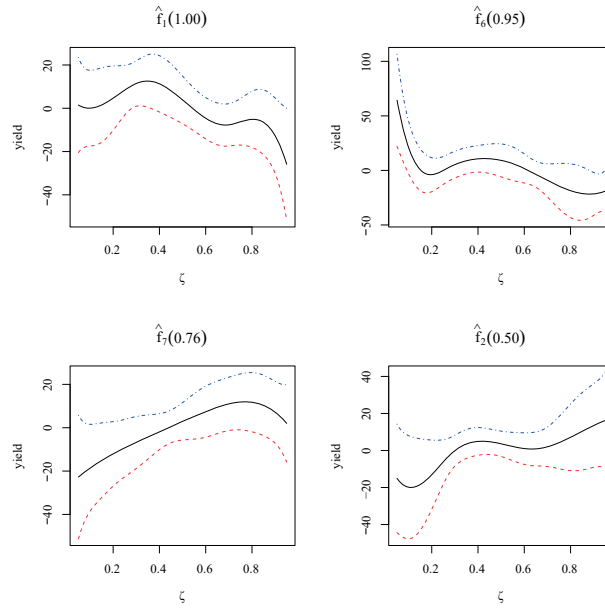


Figure 1. Additive component functions $\hat{f}_k(\zeta)$ selected by FAQR with $\tau = 0.1$, sorted by selection frequencies (in parentheses) across 500 bootstrap samples.

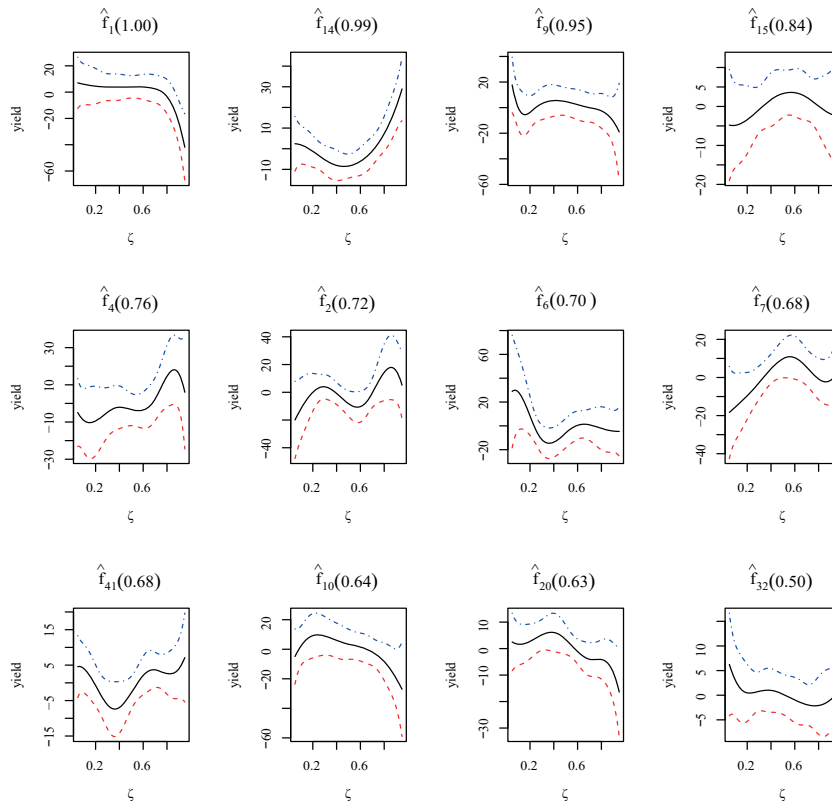


Figure 2. Additive component functions $\hat{f}_k(\zeta)$ selected by FAQR with $\tau = 0.5$, sorted by selection frequencies (in parentheses) across 500 bootstrap samples.

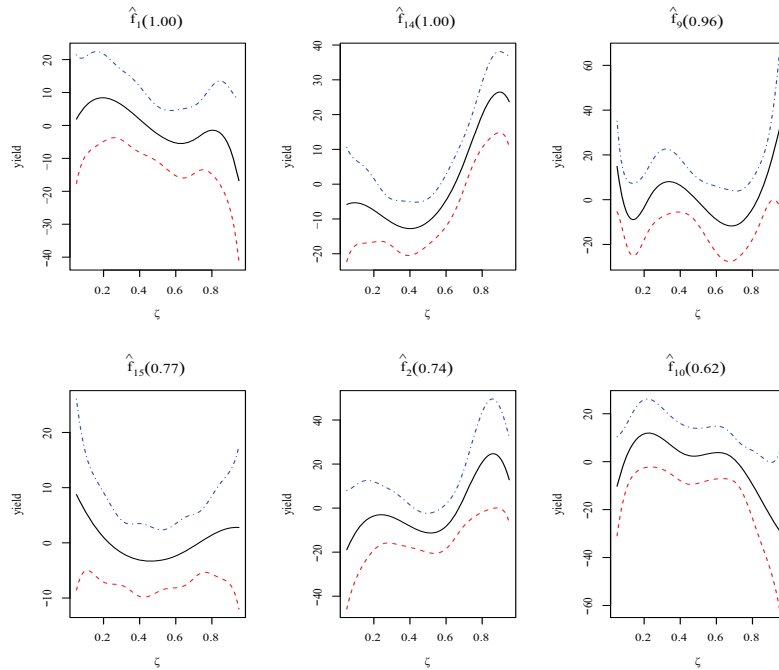


Figure 3. Additive component functions $\hat{f}_k(\zeta)$ selected by FAQR with $\tau = 0.9$, sorted by selection frequencies (in parentheses) across 500 bootstrap samples.

ter predictions. Thus, in the following, we focus on the results with the tuning parameter chosen by CV. In Figures 1–3, we show the corresponding additive component functions $\hat{f}_k(\zeta)$ selected for $\tau = 0.1, 0.5, 0.9$, respectively. The dashed curves are the pointwise confidence bands $\hat{f}_k(\zeta) \pm 2 \times se\{\hat{f}_k(\zeta)\}$. The standard errors are estimated as $se\{\hat{f}_k(\zeta)\} = (1/(B - 1)) \sum_{b=1}^B (\hat{f}_k^{(b)}(\zeta) - \bar{f}_k^{(b)}(\zeta))^2$, where $\bar{f}_k^{(b)}(\zeta) = (1/B) \sum_{b=1}^B \hat{f}_k^{(b)}(\zeta)$ and $\hat{f}_k^{(b)}(\zeta)$ is the estimation from a pair bootstrap sample $\{y_i^{(b)}, X_i^{(b)}\}$, for $b = 1, \dots, B$. We report the average model size over different bootstrap samples in Table 11. Note that the additive regression modeling produces notably smaller models than those of the linear regression, without sacrificing much prediction accuracy. Only ζ_1 and ζ_2 are shared between the three conditional quantile regressions. Figures 1–3 show that the effect of the temperature trajectory on the corn yield varies, because different components are selected.

6. Conclusion

Many researchers have used a functional linear quantile regression to study the relationship between a functional predictor and the conditional quantile of a response. We model the effects of a functional covariate nonparametrically to increase the model flexibility. We consider a nonconvex penalized estimation

of the functional additive quantile regression model. Under mild conditions, we derive the oracle convergence rate when true relevant components are used.

A problem of important practical interest is to extend our model to take into account scalar covariates and functional covariates simultaneously, as Wong, Li and Zhu (2018) have done for the mean regression. Another relevant problem is to estimate the conditional quantile function in the RKHS framework, and then to compare its performance with that of our proposed method. Finally, because we fit each quantile level separately, quantile crossing may occur. Methods proposed in the literature, such as those of Koenker and Ng (2005), Bondell, Reich and Wang (2010), Chernozhukov, Fernandez-Val and Galichon (2010), and Qu and Yoon (2015), may be used to address this problem. These topics are left to future research.

Supplementary Material

The online Supplementary Material contains proofs of the technical results.

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