

# Combining Regression Quantile Estimators

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## Abstract

Model selection for quantile regression is often a challenging problem. In addition to the well-known general difficulty of model selection uncertainty, when quantiles at multiple probability levels are of interest, typically a single candidate does not serve all of them well simultaneously. In this paper, we propose methods to combine quantile estimators. Oracle inequalities show that at each given probability level, the combined estimators automatically perform nearly as well as the best candidate. Simulation and real examples show that the proposed model combination approach often leads to a substantial gain in accuracy under global measures of performance.

**KEY WORDS:** Model combination; Aggregation of estimators; Adaptive quantile regression

# 1 Introduction

Conditional quantile estimation has been used for a long period of time in various contexts including agriculture, economics and finance. Numerous methods have been proposed under different settings including the classical linear regression, nonlinear regression, time series, and longitudinal experiment (see e.g., He, Ng, and Portnoy 1998; Yu, Lu and Stander 2003; Koenker 2005 for some recent developments and references). In what follows, we first give a brief review of the general problem of conditional quantile estimation and model selection, and then set up the problem for this work.

## 1.1 A background on conditional quantile estimation (CQE)

In regression, besides the conditional mean, we are often interested in other summary measures of the conditional distribution of  $Y$  given the input  $X$ . Quantile regression is used to obtain an estimate of the conditional quantile function at a given probability level  $\tau$  ( $\tau \in (0, 1)$ ). When a range of  $\tau$  values are considered, the quantile profile provides information much beyond the conditional mean. Conditional quantile estimation may also be used to produce confidence bands for the distribution of  $Y$  given  $X$  (see e.g., Zhou and Portnoy 1996; Koenker 2005, for applications). Quantile estimation also gets attention due to its robustness property, compared to conditional mean, in case of strong skewness in the true conditional distribution (see, e.g., Yu, Lu and Stander 2003; Geraci and Bottai 2007).

Koenker and Bassett (1978) introduced regression quantile estimation by minimizing an asymmetric loss function  $L_\tau(\xi) = \{\tau - I_{\{\xi < 0\}}\} \xi$  for  $0 < \tau < 1$ , which is known as the check loss or the pinball loss. It is not hard to verify that the minimizer  $c(x)$  of  $EL_\tau(Y - c(X)|X = x)$  is the lower- $\tau$  conditional quantile of  $Y$  given  $X = x$ . They considered  $c(x)$  of the form  $x'\beta$  and the coefficients  $\beta$  is estimated by minimizing  $\sum_i L_\tau(y_i - x_i'\beta)$ . This method is commonly known as **linear quantile regression** (LQR). A slightly more general loss, called lin-lin loss, was considered in Granger

(1969).

To reduce of the impact of parametric assumptions, nonparametric and semiparametric methods have also been developed for quantile regression. For example, one may assume that the quantile function is of the semi-parametric form  $q_\tau(X, T) = X'\beta + g(T)$ , where both  $X$  and  $T$  are vectors of explanatory variables,  $\beta$  denotes a vector of unknown regression coefficients and  $g$  represents an unparameterized smooth function to be estimated. Analogous to semiparametric mean regression, we can estimate  $\beta$  and  $g$  by minimizing  $\sum_{i=1}^n L_\tau(y_i - x_i'\beta - g(t)) + \alpha \int g''^2 dt$ , where  $\alpha$  is a smoothing parameter to control the amount of penalty on the roughness of  $g$ . Interested readers are referred to Yu *et al.* (2003), Koenker (2005) and references therein for more details.

More recently, Meinshausen (2006) proposed a nonparametric method called **quantile regression forests (QRF)**, which was inspired by the idea of random forests of Breiman (2001). As in the random forests algorithm, for each tree, one selects a random subset of all predictors to split nodes and a large number of (random) trees are grown in this fashion. The conditional quantile of  $Y$  given  $X = x$  is then approximated by the average prediction from the collection of random trees. This method was shown to be consistent and numerical results demonstrated its good performance in problems with high-dimensional predictors, particularly at extreme values of  $\tau$  ( $\tau$  near zero or one).

Regression quantile is also important in areas of application other than the conventional i.i.d. setting. In longitudinal studies, Geraci and Bottai (2007) used the loss function  $L_{0.5}$  to construct Normal-Laplace joint likelihood in a mixed effect model. An interesting quantile autoregression theory is given in Koenker and Xiao (2006). Wei and He (2006) proposed a useful semi-parametric quantile regression method for constructing conditional growth charts based on longitudinal observations.

Besides the check loss, other asymmetric loss functions have also been investigated (see, e.g., Hall, Wolff and Yao 1999), although they are used less often in the statistical literature.

## 1.2 Model selection and combination in CQE

On quantile regression, the issue of model selection has also been studied. Ronchetti (1985) introduced a robust version of AIC, called AICR, which takes the form of the observed check loss plus a multiple of model size (see also Cade, Noon and Flather 2005). Machado (1993) proposed a generalized Schwarz Information Criterion, which is similar to BIC except that the squared error loss is replaced by a more robust loss function. Some other model selection criteria can be found in Burman and Nolan (1995) and Ronchetti, Field and Blanchard (1997).

In an effort to combine different methods, if  $\hat{q}_\tau^A(x)$  and  $\hat{q}_\tau^B(x)$  are two estimates of the conditional lower- $\tau$  quantile of  $Y$  given  $X = x$ , Granger (1989) proposed the use of weights from  $\min_{\alpha, \beta_A, \beta_B} \sum_i L_\tau(y_i - \alpha - \beta_A \hat{q}_\tau^A(x_i) - \beta_B \hat{q}_\tau^B(x_i))$ . Taylor and Bunn (1998) extended this linear combination methodology by considering a number of constraints on the coefficients  $\alpha, \beta_A, \beta_B$ , such as zero intercept, convex coefficients on the predictors, and so on. To our knowledge, theoretical results on combining quantile regression estimators have not been given in the literature.

When the quantile profile is of interest, it is particularly important to consider model combination methods. A main reason is that the different quantile regression estimators typically have distinct relative performances that depend on the value of  $\tau$  (as seen in our numerical results). Integrating the advantages of the candidates and thus globally improving over them is a valuable task for reliable quantile regression.

A recent focus on combining or aggregating models (procedures) is the construction of methods that adaptively share the strengths of a list of arbitrary estimators (see, Nemirovskii 2000; Yang 2001 and 2004b; Catoni 2004; Tsybakov 2003), which allows the integration of powers of different methodologies. See Leung and Barron (2006); Birgé (2006); Bunea and Nobel (2006); Bunea, Tsybakov and Wegkamp (2006); Audibert (2006); Lecué (2006), for some more recent research results in the area. We follow this spirit in the present work and present both theoretical and numerical results on combining quantile estimators.

### 1.3 Problem of interest in this work

We assume that we observe  $(Y_i, X_i)$ ,  $i = 1, \dots, n$ , where  $X_i = (X_{i1}, \dots, X_{ip})$  is a  $p$ -dimensional predictor. Assume the true underlying relationship between  $Y$  and  $X$  is characterized by:

$$Y_i = m(X_i) + \sigma(X_i)\epsilon_i, \quad i = 1, \dots, n,$$

where  $\epsilon_i$  are i.i.d. from a distribution with mean zero and variance one and are independent of the predictors. A time series setting which does not require that  $Y_i, i = 1, \dots, n$  are independent will be considered as well.

Based on the aforementioned data generating model, the conditional quantile of  $Y$  given  $X = x$  has the form

$$q_\tau(x) = m(x) + \sigma(x)F^{-1}(\tau), \quad (1)$$

where  $F$  is the cumulative distribution function of the error. This provides one method for estimating  $q_\tau(x)$ , namely, by first obtaining  $\hat{m}(x)$ ,  $\hat{\sigma}(x)$  and  $\hat{F}^{-1}(\tau)$  (if  $F$  needs to be estimated).

Based on (1), it can be observed that if the  $m(\cdot)$  is a linear function of  $x$  and  $\sigma(\cdot)$  is constant, linear quantile regression (LQR) is expected to perform well. However, if either the mean function is nonlinear or the scale function is non-constant in the predictors, bias will be involved which may lead to poor performance of LQR. Also in real applications, the performance of LQR on extreme quantiles is usually impaired by insufficient extreme observations.

Now suppose we have a pool of  $M$  candidate estimators of the conditional quantile function  $q_\tau(x)$ , denoted by  $\{\hat{q}_{\tau,j}(x)\}_{j=1}^M$ . Our goal is to combine these estimators for an optimal performance. Specifically, at each given  $\tau$ , we hope that the combined estimator performs as well as the best candidate. Since the best candidate often depends on  $\tau$ , our combining approach can improve over all of the candidate procedures in terms of global performance measures over  $\tau$ , as will be seen in our simulation and real examples.

In the context of conditional mean regression, Yang (2001) proposed the adaptive regression by mixing (ARM) method, in which a set of weights is adaptively calculated

from the data under a specified likelihood function such as Gaussian. Alternatively, risk bounds that relate the performance of the combined estimator to that of the best candidate (typically unknown, of course) under certain quadratic-type of loss functions are given in Catoni (2004) and Yang (2004a) without specifying the error distribution. This latter approach is useful when no obvious choice of error density is available and/or when variance estimation is difficult.

In the current context, instead of a quadratic loss, the check loss function is naturally oriented towards quantile estimation and is thus used in our weight construction. However, the distinct natures of the absolute-type and quadratic-type of losses present a non-trivial work to derive an oracle inequality for our quantile regression problem. Risk bounds in terms of the check loss function, under both i.i.d. and a time series settings, without any assumption on the form of the error density nor requiring boundedness of the response variable, are obtained, which indeed show that at each fixed  $\tau$  our combined estimator performs almost as well as the best candidate. A potential application of our method is on conditional growth chart construction (Wei and He, 2006), where different semi-parameter models can be explored.

The rest of the paper is organized as follows. In Section 2, our model combining methods for regression with iid observations are presented and oracle inequalities that show their optimal performance are given. In Section 3, model combination is considered for a time series framework. Simulation and real examples that demonstrate advantages of our methods are presented in Sections 4 and 5 respectively. Concluding remarks are given in Section 6. Proofs of the theoretical results are in an appendix.

## **2 Adaptive quantile regression by mixing (AQRM)**

In this section, we consider the framework in Section 1.3 with iid observations, and take two weighting approaches, one directly based on the cumulative check loss and the other on a mixture of the check and squared losses.

## 2.1 Weighting based on check loss

The AQR algorithm for conditional quantile estimation is as follows. Fix a probability level  $0 < \tau < 1$ . Let  $1 \leq n_0 \leq n - 1$  be an integer (typically  $n_0$  is of the same order as or slightly larger order than  $n - n_0$ ).

1. Randomly partition the data into two parts:  $Z^{(1)} = \{y_l, x_l\}_{l=1}^{n_0}$  for training and  $Z^{(2)} = \{y_l, x_l\}_{l=n_0+1}^n$  for evaluation.

2. Based on  $Z^{(1)}$ , obtain candidate estimates of the conditional quantile function  $q_\tau(x)$  by  $\hat{q}_{\tau,j,n_0}(x) = \hat{q}_{\tau,j,n_0}(x; Z^{(1)})$ . Use  $\hat{q}_{\tau,j,n_0}$  to obtain the predicted quantiles from the  $j^{\text{th}}$  candidate procedure for  $Z^{(2)}$ , for each  $j = 1, \dots, M$ .

3. Compute the candidate weights as follows

$$W_j = \frac{\prod_{l=n_0+1}^n \exp\{-\lambda L_\tau(y_l - \hat{q}_{\tau,j,n_0}(x_l))\}}{\sum_{k=1}^M \prod_{l=n_0+1}^n \exp\{-\lambda L_\tau(y_l - \hat{q}_{\tau,k,n_0}(x_l))\}},$$

where  $\lambda > 0$  is a tuning parameter.

4. Repeat steps 1 – 3 ( $B - 1$ ) more times and average the weights  $W_j$  over  $B$  random permutations. Denote them by  $\tilde{W}_j$ . Our final estimator of the conditional quantile function of  $Y$  at  $X = x$  is  $\hat{q}_{\tau,.,n}(x) = \sum_{j=1}^M \tilde{W}_j \hat{q}_{\tau,j,n}(x)$ .

**Remark:** The tuning parameter  $\lambda$  controls how much the weights rely on the check loss performance. In the extreme case when  $\lambda \downarrow 0$ , simple averaging results; when  $\lambda \rightarrow \infty$ , the candidate with the best historic check loss is selected.

In certain problems such as online estimation/prediction, a sequential updating mechanism is also of interest. Namely, we first obtain  $\hat{q}_{\tau,j,n_0}$  from  $\{(y_l, x_l)\}_{l=1}^{n_0}$  (the initial set of observations) and the weights are updated sequentially once an additional observation is made. In such a setting, we define sequential weight  $W_{j,i}$  as

$$W_{j,i} = \frac{\prod_{l=n_0+1}^{i-1} \exp\{-\lambda L_\tau(y_l - \hat{q}_{\tau,j,l}(x_l))\}}{\sum_{k=1}^M \prod_{l=n_0+1}^{i-1} \exp\{-\lambda L_\tau(y_l - \hat{q}_{\tau,k,l}(x_l))\}},$$

and the combined estimate of  $q_\tau(x)$  at time  $i$  is  $\hat{q}_{\tau,.,i}(x) = \sum_{j=1}^M W_{j,i} \hat{q}_{\tau,j,i}(x)$ . Also of interest is an overall convex combination

$$\hat{q}_{\tau,.,.}(x) = \frac{1}{n - n_0} \sum_{j=1}^M \sum_{i=n_0+1}^n W_{j,i} \hat{q}_{\tau,j,i}(x),$$

which estimates  $q_\tau(x)$  in a way that utilizes the online estimates at different sample sizes and it has a nice risk property as will be seen shortly. Note that in numerical implementation for batch learning, since a sequential updating algorithm can be much more time-consuming when the sample size is not small, we follow the earlier algorithm and the candidate quantile estimators are not updated in the weight construction.

## 2.2 Oracle inequalities on performance

**Condition 0:** The observed vectors  $(Y_i, X_i), i \geq 1$  are iid.

**Condition 1:** The quantile estimators satisfy that  $\sup_{j \geq 1, i \geq 1} |\hat{q}_{\tau, j, i}(x_i) - q_\tau(x_i)| \leq A_\tau$ , for some positive constant  $A_\tau$  with probability one. In what follows, we omit the subscript  $\tau$  to simplify notation.

**Condition 2:** There exist a positive constant  $t_0$  and a monotone function  $0 < H(t) < \infty$  on  $[-t_0, t_0]$  such that for all  $n \geq 1$  and  $-t_0 \leq t \leq t_0$ ,

$$E(|\epsilon_n|^2 + 1) \exp(t|\epsilon_n|) \leq H(t),$$

where  $\epsilon_n$  is the unobservable true error for the  $n^{\text{th}}$  observation.

**Condition 3:** There exist positive constants  $C_1$  (that depends on  $\tau$ ) and  $C_2$  such that  $|m(X) - q_\tau(X)| \leq C_1$  and  $|\sigma^2(X)| \leq C_2$ , with probability one.

Condition 1 requires that the performance of all the candidate estimators are not too far away from the true conditional quantile. This is a mild technical condition, weaker than assuming  $Y$  is bounded, and is typically assumed in the statistics literature of theoretical work on combining estimators. Condition 2 is satisfied by error distributions with moment generating functions well defined, such as normal, shifted gamma and double exponential distribution. These error distributions are considered in our numerical study. Condition 3 requires some regularity of the underlying conditional distribution of  $Y$  given the predictors, but neither constant is required to be known for application.

Let  $B(\lambda) = e^{2\lambda \max(\tau, 1-\tau)(A+1)} (1 + (A+1)^2) H(2\lambda \max(\tau, 1-\tau))$  and in Theorem 1 below, define  $a_\lambda = 2\lambda (\max(\tau, 1-\tau))^2 B(t_0)$ .

**Theorem 1** *Under Conditions 0-3, when the tuning parameter  $\lambda \leq \lambda_0 = \frac{t_0}{2 \max(\tau, 1-\tau)}$ , we have*

$$\frac{1}{n - n_0} \sum_{i=n_0+1}^n EL_\tau(Y_i - \hat{q}_{\tau, \cdot, i}(X_i)) \leq \inf_j \left\{ \frac{1}{n - n_0} \sum_{i=n_0+1}^n EL_\tau(Y_i - \hat{q}_{\tau, j, i}(X_i)) + \frac{\log(M)}{\lambda(n - n_0)} + \frac{a_\lambda(C_2 + C_1^2)}{2} \right\}.$$

*In particular, when  $\lambda = \left( \frac{\log(M)}{(\max(\tau, 1-\tau))^2 B(t_0)(C_2 + C_1^2)(n - n_0)} \right)^{1/2}$ , we have*

$$\frac{1}{n - n_0} \sum_{i=n_0+1}^n EL_\tau(Y_i - \hat{q}_{\tau, \cdot, i}(X_i)) \leq \inf_j \left\{ \frac{1}{n - n_0} \sum_{i=n_0+1}^n EL_\tau(Y_i - \hat{q}_{\tau, j, i}(X_i)) + \tilde{C} \sqrt{\frac{\log(M)}{n - n_0}} \right\}, \quad (2)$$

and

$$EL_\tau(Y - \hat{q}_{\tau, \cdot, \cdot}(X)) \leq \inf_j \left\{ \frac{1}{n - n_0} \sum_{i=n_0+1}^n EL_\tau(Y_i - \hat{q}_{\tau, j, i}(X_i)) + \tilde{C} \sqrt{\frac{\log(M)}{n - n_0}} \right\},$$

where  $\tilde{C}$  is a constant that depends on  $\tau, A, C_1, C_2$ .

### Remarks:

1. For the third display in the theorem, the risk of the combined estimator at sample size  $n$  is upper bounded in terms of the best averaged risk at different sample sizes plus a penalty. Ideally one would want to replace the averaged risk by the risk of the candidate at the full sample size  $n$ , which is not obtained in this work.

2. Note that unboundedness of the response variable makes the derivation of oracle inequalities substantially different from the earlier work on combining predictions in the machine learning literature, which typically requires that  $Y$  has a bounded support (or the loss is bounded). See Bunea and Nobel (2005) for a different approach to address the issue of unbounded response under squared error loss.

The inequalities above say that the risks of the combined prediction are automatically close to the risks of the best individual, with the difference being of order  $(n - n_0)^{-1/2}$  when  $\lambda$  is chosen properly. Note that for  $L_1$  type of risk for regression

estimation, the rate of convergence typically is  $n^{-1/2}$  for parametric cases, and is slower than  $n^{-1/2}$  for nonparametric cases (e.g., Yang and Barron 1999). Therefore, with a choice of  $n_0$  and  $n - n_0$  of the same order, the risk bounds show that the combined quantile predictions adaptively converge at the best rate offered by the candidate procedures for both parametric and nonparametric situations. Furthermore, for nonparametric quantile regression, since the extra term in the risk bound is asymptotically negligible relative to the risk of estimating  $q_\tau(x)$ , under some regularity conditions, AQRM yields combined predictions that perform asymptotically as well as the best procedure among the candidates.

Although at each given probability level  $\tau$ , our approach of combining the quantile estimators does not necessarily lead to performance improvement over the best individual candidate estimator, the results are useful for three reasons. First, for various situations (e.g., one of the candidate procedures is based on the true model), the best individual procedure simply cannot be improved and thus the combined estimator performs optimally or near optimally. Second, since the best procedure is unknown and as is well-known one often pays a high price in trying to find it out (see, e.g., Yuan and Yang 2005, for references and simulation results on reducing model selection uncertainty by model combination), it is important to show that the combining approach, as an alternative, indeed leads to optimal performance. Third, because conditional quantile functions at a range of probability level are often of interest at the same time but the candidate quantile estimators typically have different ranks in performance, the combined estimators have a good potential to beat each of the candidates in terms of global performance, as will be seen later.

### 2.3 Weighting using a mixture of check and squared losses

Define a surrogate loss function  $L_{\tau,a}(\xi) = L_\tau(\xi) + a\xi^2$  for a given  $a > 0$  (see Figure 1) and use it in the construction of weights of the candidate quantile regression

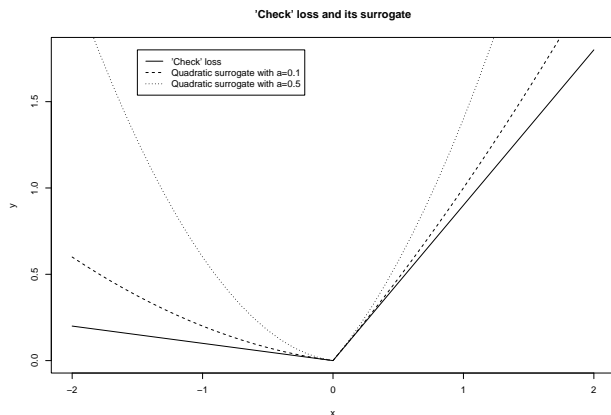


Figure 1: Check loss function with  $\tau = 0.9$  and its surrogates.

procedures. The new weight is

$$W_j^a = \frac{\prod_{l=n_0+1}^n \exp\{-\lambda L_{\tau,a}(y_l - \hat{q}_{\tau,j,n_0}(x_l))\}}{\sum_{k=1}^M \prod_{l=n_0+1}^n \exp\{-\lambda L_{\tau,a}(y_l - \hat{q}_{\tau,k,n_0}(x_l))\}}.$$

We can then derive a similar risk upper bound for the corresponding combined estimator  $\hat{q}_{\tau,\cdot,n_0}^a$ .

**Theorem 2** *Under the same assumptions in Theorem 1, when  $\lambda$  is chosen as for (2) and  $a = 2\lambda (\max(\tau, 1 - \tau))^2 B(t_0)$ , we have*

$$EL_{\tau}(Y - \hat{q}_{\tau,\cdot,n_0}^a(X)) \leq \inf_j \left\{ \frac{1}{n - n_0} \sum_{i=n_0+1}^n EL_{\tau}(Y_i - \hat{q}_{\tau,j,i}^a(X_i)) + C' \sqrt{\frac{\log(M)}{n - n_0}} \right\},$$

where  $C'$  is a constant that depends on  $\tau, A, C_1, C_2$ .

### 3 Combining quantile estimators for time series

For time series data, we typically have autocorrelation between observations. Consider the model

$$Y_t = m_t(X_t) + \sigma_t(X_t)\epsilon_t,$$

where  $X_t$  is the explanatory variable (which may include the past values of the response variable) at time  $t$ . We assume that the errors  $\epsilon_t$  are i.i.d. from a distribution

with mean zero and variance one, and  $\epsilon_t$  is independent of  $\{(Y_s, X_s) : s < t\}$  and  $X_t$ . Our goal is to derive a combined (conditional) quantile estimator  $\hat{q}_{\tau, \cdot, t}(x_t) = \sum_{j=1}^M W_{j,t} \hat{q}_{\tau, j, t}(x_t)$ .

We follow an online setting which means that data come in sequentially and the candidate estimators will also be updated sequentially with each incoming observation. Let  $T$  be the length of the whole series. The combining algorithm AQRM for the time series setting is given below.

1. Start with  $T_0$  observations and let  $t_1 = T_0$ .
2. Denote the first  $t_1$  observations in the series by  $Z^{(1)} = (y_t, x_t)_{t=1}^{t_1}$ .
3. Based on  $Z^{(1)}$ , construct the candidate estimates of the conditional quantile function  $q_\tau(x)$  by  $\hat{q}_{\tau, j, t_1}(x) = \hat{q}_{\tau, j, t_1}(x; Z^{(1)})$ .
4. For each  $j$ , we update the candidate weight sequentially as follows

$$W_{j, t_1+1} = \frac{W_{j, t_1} \exp \{-\lambda L_\tau(y_{t_1} - \hat{q}_{\tau, j, t_1}(x_{t_1}))\}}{\sum_{k=1}^M W_{k, t_1} \exp \{-\lambda L_\tau(y_{t_1} - \hat{q}_{\tau, k, t_1}(x_{t_1}))\}},$$

where  $W_{j, T_0+1} = \frac{1}{M}$ .

5. We increase  $t_1$  by 1 and repeat steps 2 – 4, until  $t_1 = T$ .

Since in the time series setting, the conditional quantiles, conditional means and conditional variances of  $Y_t$  usually depend on both the predictor and time, Conditions 1-3 need to be modified accordingly.

**Theorem 3** *Under Conditions 1-3 on the conditional quantiles, conditional means and conditional variances, when the tuning parameter  $\lambda$  is chosen as for (2),*

$$\sum_{t=T_0+1}^T EL_\tau(Y_t - \hat{q}_{\tau, \cdot, t}(X_t)) \leq \inf_j \left\{ \sum_{t=T_0+1}^T EL_\tau(Y_t - \hat{q}_{\tau, j, t}(X_t)) + \tilde{C} \sqrt{\log(M)} \times \sqrt{T - T_0} \right\},$$

where  $\tilde{C}$  is a constant that depends on  $\tau, A, C_1, C_2$ .

## 4 Simulation results

In this section, four cases are considered to investigate the performance of AQRM. Together with real examples in the next section, we intend to gain some insight on

the differences of behaviors of the methods involved, which may be more helpful than giving one or two favorable examples.

## 4.1 Candidate procedures and performance measures

We consider LQR (Koenker and Bassett 1978) and QRF (Meinshausen 2006), using *R* packages *quantreg* and *quantregForest*.

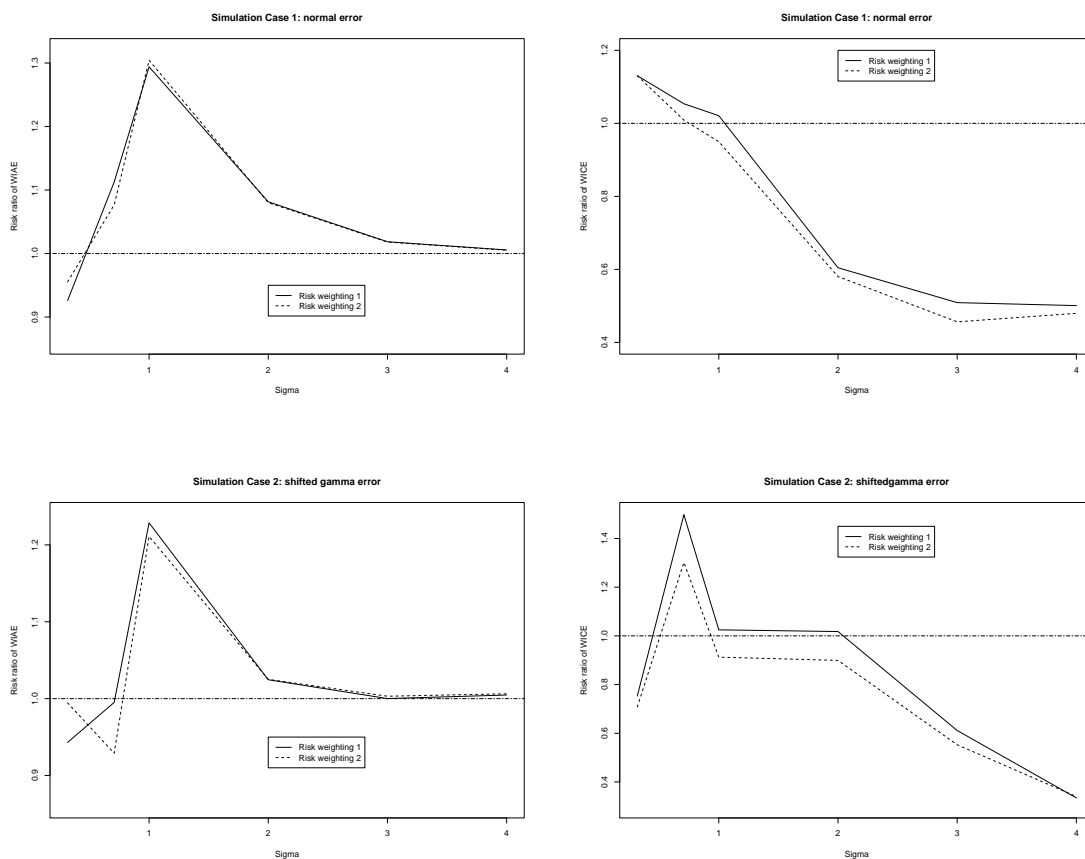
In the literature, performance of quantile regression is usually measured by the coverage probability at some fixed  $\tau$  value(s), such as 90% and 95% (see, e.g., Koenker and Bassett 1978; Taylor and Bunn 1998). For a given quantile estimator at a given  $\tau$ , its empirical coverage probability is defined as the fraction of observations which fall on or below the estimated quantile function in a new (unused) evaluation set.

Here, we focus on the overall performance of a quantile regression procedure over the full range of  $\tau$  in  $(0, 1)$ . One reason is that quantiles at multiple levels are often of interest at the same time (e.g., for growth charts) and global measures over a range of  $\tau$  are naturally relevant. Another related motivation is due to the fact that different regression quantile estimators often have distinct relative performances according to the value of  $\tau$ , and therefore the consideration of a range of  $\tau$  values yields an overall comparison of different methods. We introduce two overall performance measures below.

Let  $g$  denote a weighting function on  $\tau \in (0, 1)$  such that  $g \geq 0$  and  $\int_0^1 g(\tau) d\tau = 1$ , which is used to differentiate the importance of  $\tau$  values in different regions. We choose two different  $g$  functions in this work, one being the uniform weight and the other being the Beta(0.8,0.8) density, which emphasizes extreme  $\tau$ 's.

In simulations, considering the integrated absolute difference between the true  $q_\tau(\cdot)$  and an estimator  $\hat{q}_\tau(\cdot)$ , under a given weight function, we define Weighted Integrated Absolute Error (WIAE) as the expectation of  $\int \int |\hat{q}_\tau(x) - q_\tau(x)| g(\tau) d\tau P(dx)$ . For real data, since we do not know the true conditional quantile function, obviously, we cannot compute WIAE. Instead, we consider the discrepancy between the nominal level  $\tau$  and the empirical coverage probability  $\hat{\tau}$ , and define Weighted Integrated

Coverage Error (WICE) as  $\int_0^1 |\hat{\tau} - \tau|g(\tau)d\tau$ . In implementing this, we use a random data splitting, which reserves part of the given data as an (artificial) evaluation set. This random partition of data is repeated for 100 times and the average performance measures over these repetitions are reported.



(Case 1 and Case 2) The plots give risk ratios of AQRM to that of the best candidate.

Figure 2: Summary plot for Case 1 and Case 2.

To approximate the integrals in the definitions of WIAE and WICE, we select a number of discrete  $\tau$  values,  $\tau \in \{0.01, 0.05 \times k, 0.99\}_{k=1}^{19}$ . We also calculate, for each fixed  $\tau$ , the simulation standard errors of both the candidate methods and AQRM.

In our investigation, we also assess the role of  $\lambda$  on the performance of AQRM (automatic selection of  $\lambda$  will not be addressed in this work). We define the optimal  $\lambda$  as the one that yields the smallest WICE (or WIAE) among all  $\lambda$  considered, and

define the risk ratio of AQRM over the best individual candidate as

$$RR = \frac{\text{WICE (or WIAE) of AQRM under the optimal } \lambda}{\text{WICE (or WIAE) of the best individual candidate}}.$$

The simulation results in this section are based on 100 runs in each case. The sample size is 200, with equal training-testing data splitting randomly done 50 times. To compute the loss of absolute error or coverage error defined above, an independent evaluation set of size 1900 is used.

The tuning parameter  $\lambda$  is taken of the form  $\lambda_\tau = \lambda \times \min(\tau, 1 - \tau)$ , where  $\tau \in \{0.01, 0.05 \times k, 0.99\}_{k=1}^{19}$ . Empirical evidence suggests that our combined estimator performs better with  $\lambda_\tau$  than using a constant value for all  $\tau$ . In what follows, we omit the subscript  $\tau$  in  $\lambda_\tau$  to simplify notation.

## 4.2 Simulation models

We consider 4 cases, the last two with randomly generated coefficients to reduce the reliance of the simulation results on specific choice of parameter values.

**Case 1.** The first model, an example used in the *R*-package *quantreg*, is

$$Y = Z + \log(X) + 0.1 \times (\log(X))^2 + 0.25 \times \log(X) \times \epsilon_2,$$

where  $X \sim \chi_4^2$ ,  $\epsilon_1 \sim N(0, 1)$ ,  $Z = X + \epsilon_1$ ,  $\epsilon_2 \sim f$ , with  $\mu_f = 0$  and  $\sigma_f = \sigma$ , and  $X$ ,  $\epsilon_1$ ,  $\epsilon_2$  are generated independently of each other. Besides  $N(0, 1)$ , shifted gamma distribution is also considered to allow asymmetric error, and the error standard deviation takes 6 values:  $\sigma = 0.316, 0.717, 1, 2, 3, 4$ . In this and other simulations in Section 4.3, unless stated otherwise, shifted gamma errors are generated from a gamma distribution with shape parameter of three and scale parameter of  $\frac{\sigma}{3}$ , then shifted to have zero mean. We observe two predictors  $X$  and  $Z$ , as well as the response  $Y$ . Nine  $\lambda$  values are considered and they are  $\lambda = \min(\tau, 1 - \tau) \times \{0, 0.01, 0.1, 1, 5, 8, 10, 50, 100\}$ .

We consider three candidate procedures: 1) LQR with predictors  $X$  and  $Z$ , 2) QRF with predictors  $X$  and  $Z$ , 3) Linear regression quantile with predictors  $X$ ,  $Z$  and  $\sqrt{X}$ .

**Case 2.** Based on Case 1, to allow more complexity, we modify the scale function from  $\sigma_f(x, z) \equiv \sigma$  to  $\sigma_f(x, z) = \sigma\sqrt{x}$ .

**Case 3.** To avoid “picking the best parameter setting to favor one’s own method”, we now randomly generate coefficients  $\beta = (\beta_1, \dots, \beta_6)$ , with  $\beta_i \stackrel{i.i.d.}{\sim} \text{Unif}[0.5, 2.5]$  for  $i = 1, \dots, 6$  for each data set. The true model is  $Y = \beta'X + \sigma\epsilon$ , where  $X = (X_1, \dots, X_6)$  has independent  $N(0, 1)$  components, and  $\epsilon$  is either from a standard normal distribution or a shifted gamma with mean zero and variance one. Two hundred sets of coefficients are generated this way for each of which the losses of the competing procedures are calculated.

**Case 4.** The model is  $Y = \beta'X + 2 \exp(-0.35X_2 - 1.1X_3) + \sigma\epsilon\sqrt{X_2^2 + 0.8X_4^2}$  and the other aspects are the same as Case 3.

### 4.3 Results

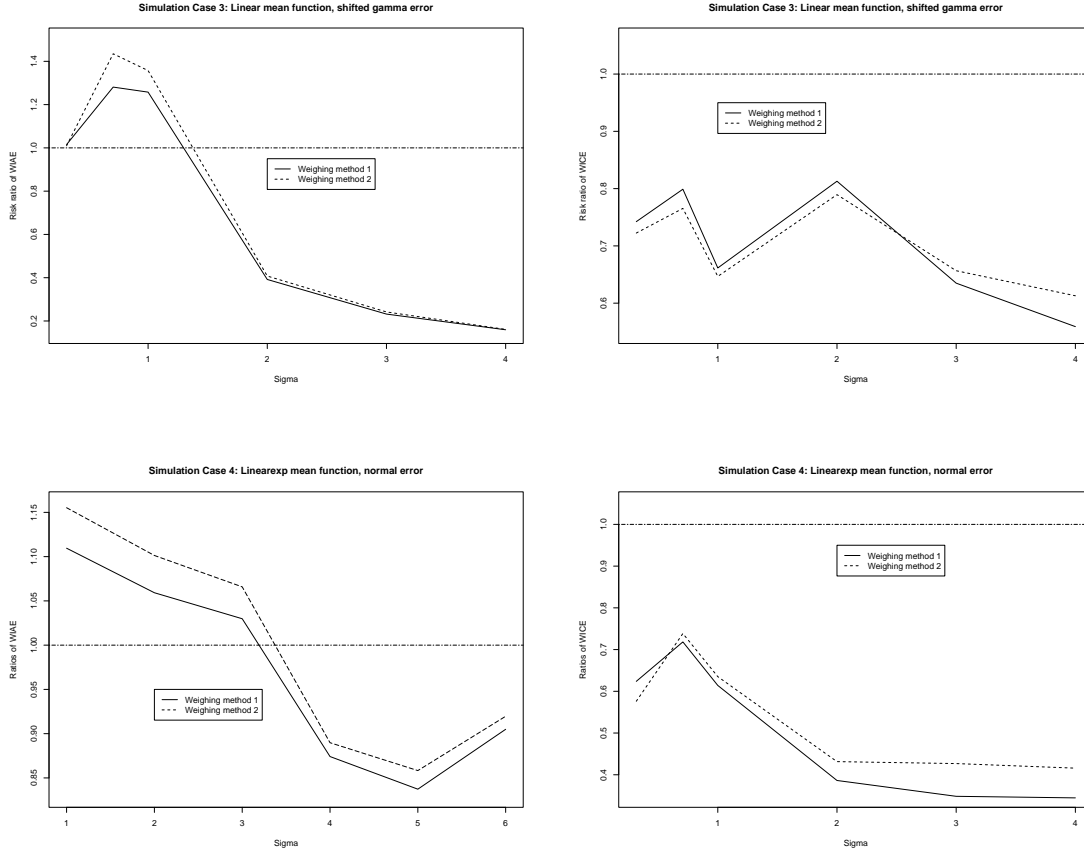
We give graphical summaries of the overall performance of the combined estimator relative to the best candidate under the two loss functions. Here the best candidate refers to the one which has the smallest mean WICE (or WIAE) under the corresponding weighting function. The plots are in Figures 2 and 3.

We are also interested in performance at fixed values of  $\tau$ . We compute the  $L_1$  risks in estimating  $q_\tau$  (along with the simulation standard errors) for several values of  $\tau$ . The results are given in Table 1 for Case 1 (as an example). Note that the optimal  $\lambda$  given in the tables is  $\tau$ -dependent.

The results are summarized as follows (although some are not given in this paper due to space limitation).

1. For the  $\sigma$  and error distributions considered, when  $\tau$  is near either zero or one, QRF has observed coverage probability closer to the nominal level  $\tau$  than LQR. But LQR performed better in the middle range of  $\tau$ .

2. The  $L_1$  risk of QRF for estimating  $q_\tau(x)$  is often the largest, compared to the other two candidates, when  $\sigma$  is small ( $\sigma \leq 0.707$ ) and often the smallest when  $\sigma$  is large ( $\sigma \geq 2$ ). This and the item above indicate that it is unwise to use a single



(Case 3 and Case 4) The plots give the risk ratios of AQRM to that of the best candidate.

Figure 3: Summary plot for Case 3 and Case 4.

quantile regression estimator for all  $\tau$  values.

3. AQRM performed well. For the error distributions considered and almost all  $\tau$ , when  $\sigma \geq 2$ , AQRM can basically tie with or perform better than the best candidate both in terms of observed coverage probability and in  $L_1$  risks (see Table 1).

4. The two performance measures are quite different. For example, the best candidate estimator under the coverage error is not the same as that under the  $L_1$  error. Also, AQRM did not improve over the candidates under the  $L_1$  error, but did so significantly under the coverage error when  $\sigma$  is not small.

5. The random coefficient cases reveal substantial advantages of AQRM. At the noise levels considered, the coverage errors of AQRM are consistently smaller than those of the candidates. Because the coefficients were randomly generated, the rank-

ing of LQR versus QRF can change as well, in which case the combined estimator can be much better than any fixed choice of the candidates.

$\sigma = 2$	$\tau = 0.05$	$\tau = 0.1$	$\tau = 0.5$	$\tau = 0.9$	$\tau = 0.95$
Best candidate	1.1594(0.0136)	0.9566(0.0130)	0.6845(0.0081)	1.0022(0.0153)	1.2619(0.0048)
Combined with optimal $\lambda$	0.8132(0.0055)	0.8077(0.0042)	0.8149(0.0026)	0.8683(0.0041)	0.9216(0.0088)

Table 1:  $L_1$  risks at fixed  $\tau$  for Case 1 under normal errors.

## 5 Real examples

### 5.1 Two regression data sets

The data set *Autoprice* is from UCI machine learning repository. There are  $n = 159$  observations with 14 continuous variables and one nominal variable. After inspecting the data, we decided to take logarithmic transformation on the response variable *price* and removed three outliers: #149, #151 and #153. The two candidate quantile regression methods are the LQR and QRF. In LQR, the best submodel selected by AIC via backward elimination is used.

In Table 2, we choose six distinct values of  $\lambda$  to assess its influence. To compare the quantile regression procedures, we randomly choose 75% of all data for training (including weight construction for combining the procedures), and remaining 25% for final performance evaluation. This is repeated for another 199 times through random partition of the data and the final coverage performance (WICE) in Table 2 is the average over all 200 repetitions.

Method	LQR	QRF	$\lambda = 0$	$\lambda = 0.5$	$\lambda = 1$	$\lambda = 3$	$\lambda = 6$	$\lambda = 10$
Uniform	6.20	1.40	1.03	1.05	1.00	1.11	1.41	1.56
Beta(0.8,0.8)	7.09	1.36	1.09	1.11	1.06	1.15	1.44	1.56

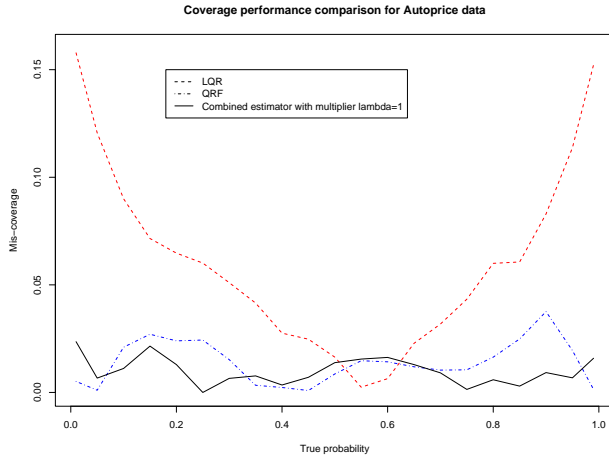
Table 2: Weighted Integrated Coverage Errors ( $\times 10^{-2}$ ) for *Autoprice* data.

$\tau$	0.05	0.1	0.25	0.5	0.75	0.9	0.95
Best candidate	0.001(0.003)	0.021(0.003)	0.024(0.006)	0.009(0.006)	0.011(0.005)	0.038(0.003)	0.019(0.002)
Combined with $\lambda = 1$	0.007(0.003)	0.011(0.004)	0.001(0.006)	0.014(0.006)	0.001(0.005)	0.009(0.003)	0.007(0.003)

Table 3: Mis-coverages at fixed  $\tau$  for *Autoprice* data.

We also report mis-coverages along with permutation standard errors for several  $\tau$  values in Table 3. The numerical results are summarized below.

1. QRF performed better than LQR under both weighting functions, although LQR was slightly more accurate in terms of coverage probability than QRF when  $\tau$  is near 0.5 (see Figure 4).
2. The combined estimators achieved better accuracy under both weighting functions as long as  $\lambda$  is not too large.
3. AQRM had good performance under almost all  $\tau$ .



Mis-coverages for AQRM and candidate estimators at different probability levels for *Autoprice* data.

Figure 4: Summary plot for *Autoprice* data.

Our second data, *Landrent* (see Weisberg 2005) has 67 observations. The response  $Y$  is the average rent per acre planted to alfalfa. There are four predictors.

Besides LQR and QRF, we also included a plug-in estimate (see, e.g., Cai 2002), which is based on linear regression of  $Y$  on  $X_1, \dots, X_4$  with stepwise selection of the variables based on AIC. We use an estimate of the form  $\hat{q}_\tau(x) = \hat{m}(x) + \Phi^{-1}(\tau) \times \hat{\sigma}$ , where both  $\hat{m}(x)$  and  $\hat{\sigma}$  are obtained from the selected model. The graphical

diagnostics on the residuals do not provide strong evidence against the normality assumption. We use 80% of all data for training (including weight construction), and the remaining 20% is reserved for performance evaluation. The final coverage performance (WICE) in Table 4 is the average over 200 repetitions.

Method	LQR	QRF	Plug-in	$\lambda = 0$	$\lambda = 0.5$	$\lambda = 1$	$\lambda = 3$	$\lambda = 6$	$\lambda = 10$
Uniform	2.88	2.44	2.11	2.96	2.03	1.83	1.61	1.62	1.64
Beta(0.8,0.8)	3.32	2.29	2.05	2.78	1.96	1.75	1.53	1.54	1.57

Table 4: Weighted Integrated Coverage Errors ( $\times 10^{-2}$ ) for *Landrent* data.

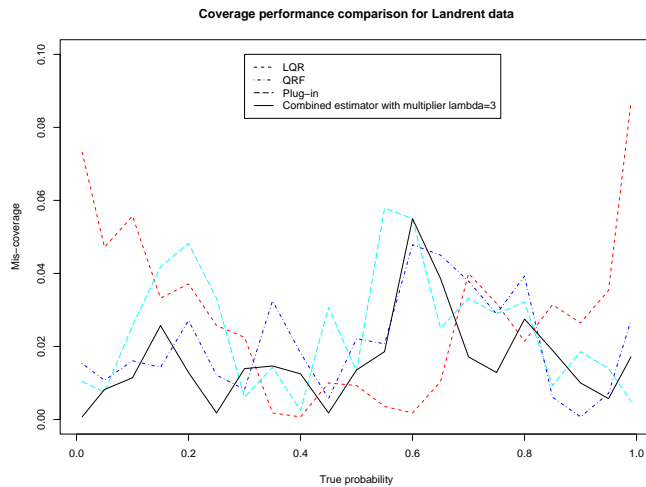
$\tau$	0.05	0.1	0.25	0.5	0.75	0.9	0.95
Best candidate	0.008(0.004)	0.016(0.007)	0.012(0.009)	0.009(0.010)	0.029(0.008)	0.001(0.006)	0.007(0.005)
Combined with $\lambda = 3$	0.008(0.004)	0.011(0.006)	0.002(0.009)	0.014(0.010)	0.013(0.008)	0.010(0.006)	0.005(0.005)

Table 5: Mis-coverage at fixed  $\tau$  for *Landrent* data.

We find that:

1. Among the candidates, the plug-in method performed the best under both weighting functions, possibly because the normal linear model describes the data very well.
2. LQR performed the best only when high weights are put on moderate  $\tau$  values (results are not presented here), where it has an advantage over the other two competing methods.
3. The combined estimators achieved better estimation accuracy under both weighting functions for all  $\lambda$ 's that are not too small.
4. Unlike the previous case, simple averaging did not produce better coverage probability over the best candidate under either weighting function.

Figures 4 and 5 present the coverage performance for each candidate and our combined estimator as a function of  $\tau$  for the two data sets. They show that our method had good performance under most  $\tau$ , especially under very large or very small values of  $\tau$ .



Mis-coverages for AQRM and candidate estimators at different probability levels for *Landrent* data.

Figure 5: Summary plot for *Landrent* data.

## 5.2 A time series

Consider the following methods (for details, see e.g., Allen *et al* 2003).

1. The standard GARCH( $p, q$ ) model with Gaussian innovations:

$$y_t = \beta_0 + \epsilon_t, \quad \epsilon_t = \sigma_t z_t, \quad z_t \stackrel{i.i.d.}{\sim} N(0, 1),$$

$$\sigma_t^2 = \alpha_0 + \sum_{i=1}^q \alpha_i \epsilon_{t-i}^2 + \sum_{j=1}^p \theta_j \sigma_{t-j}^2.$$

2. The “historical simulation” method simply uses the sample quantile of a given number (such as 100 or 500) of the most recent observations. A critique on this method can be found in Pritsker (2001).

In the financial markets, Value-at-Risk (VaR) is defined as the predicted worst-case loss with a specific confidence level (for example, 95%) over a period of time (for example, 1 day). Here we consider VaR estimation of the daily index distribution for S&P500 energy sector with data from January 3, 2000 to November 10, 2006 (available at <http://www.globalfinancialdata.com>). By examining the autocorrelation plot of this series, we decide to apply our candidate procedures to the differenced series. The candidates are GARCH(1,1) model, historical simulation with up to 100 (HS100) and 250 (HS250) most recent observations. Historical simulation method with more than

	GARCH(1,1)	HS(100)	HS(250)	$\lambda = 0$	
$\tau = .01$	0.029	0.012	0.000	0.000	
$\tau = .05$	0.070	0.029	0.000	0.000	
$\tau = .10$	0.116	0.047	0.000	0.006	
$\tau = .90$	0.872	0.761	0.552	0.720	
$\tau = .95$	0.953	0.837	0.727	0.855	
$\tau = .99$	0.988	0.930	0.901	0.971	
Avg mis-coverage	0.0147	0.0645	0.1366	0.0746	
	$\lambda = 0.5$	$\lambda = 1$	$\lambda = 3$	$\lambda = 6$	$\lambda = 10$
$\tau = .01$	0.000	0.000	0.000	0.006	0.006
$\tau = .05$	0.058	0.058	0.064	0.070	0.070
$\tau = .10$	0.110	0.110	0.110	0.116	0.116
$\tau = .90$	0.890	0.883	0.878	0.884	0.884
$\tau = .95$	0.959	0.953	0.953	0.953	0.953
$\tau = .99$	0.971	0.983	0.988	0.988	0.988
Avg mis-coverage	0.0112	<b>0.0093</b>	0.0102	0.0103	0.0103

Table 6: Observed coverage probabilities for *S&P500 energy* series.

250 observations was tried (not shown here) but much worse coverage performance was observed. For constructing the combined estimate, we initialize the estimation of the candidate methods with  $T_0 = 200$  and update both the estimators and weights sequentially. We estimate VaR at  $\tau = 0.01, 0.05, 0.1, 0.9, 0.95, 0.99$ , since VaR with moderate  $\tau$  is of little interest to the market analysts. We use the last 10% of the series for evaluation.

In Table 6, we report the observed coverage probabilities at each chosen  $\tau$  values for all procedures. It is observed that:

1. HS100 performed the best at 1%. But for all other quantiles, GARCH(1,1) performed the best among the candidates.
2. Our combined estimate with tuning parameter of  $\min(\tau, 1 - \tau)$  achieved the best overall performance at the selected  $\tau$  values.
3. Simple average (i.e.,  $\lambda = 0$ ) performed rather poorly, suggesting the need of intelligent model combining methods such as AQRM.

## 6 Concluding remarks

A lot of work has been done on the estimation of conditional quantiles. Although theoretically speaking many of the proposed parametric methods work well asymptotically, for any realized data with a moderate sample size, insufficient extreme observations typically impair their estimation accuracy at high/low quantiles even if the assumed underlying model (e.g., linear quantile functions) is proper. Nonparametric methods can improve in some aspects, especially for extreme quantiles (as is seen in this work).

Choosing a model (or procedure) from a list for quantile regression can be very challenging. Like in other contexts, model selection instability, which can substantially affect estimation/prediction accuracy, is a major issue that should not be ignored. The simulation and real examples in this paper show that the candidate procedures performed very differently (relatively speaking) at moderate and extreme quantiles. Thus selecting a single model based on a traditional model selection criterion is not a good idea for estimating multiple quantiles.

A good approach to address the aforementioned difficulties is the use of model (or procedure) combining as an alternative to choosing a single one. Under mild regularity conditions, we derived a theory which shows that our proposed estimator performs as well as the best individual candidate in terms of the asymmetric linear risk, with a cost that vanishes at  $O\left(n^{-\frac{1}{2}}\right)$  rate. Simulation examples clearly demonstrate that our method yields improved performance in terms of better overall coverage probability when error standard deviation is not small. The example of the financial series *S&P500 energy* demonstrates that our approach can be very useful for Value-at-Risk estimation.

In summary, for the reasons of model selection uncertainty and the typical dependence of the best candidate quantile regression method on the probability level, model combination methods have a great potential for reliable performance. Our proposed method AQRM can integrate the advantages of general candidate procedures that occur at different probability levels and thus globally improve over them.

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