

Instrumental Variables Estimation and Weak-Identification-Robust Inference Based on a Conditional Quantile Restriction

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Extending the \mathcal{L}_1 -IV approach proposed by Sakata (1997, 2007), we develop a new method, named the ρ_τ -IV estimation, to estimate structural equations based on the conditional quantile restriction imposed on the error terms. We study the asymptotic behavior of the proposed estimator and show how to make statistical inferences on the regression parameters. Given practical importance of weak identification, a highlight of the paper is a proposal of a test robust to the weak identification. The statistics used in our method can be viewed as a natural counterpart of the Anderson and Rubin's (1949) statistic in the ρ_τ -IV estimation.

1 Introduction

In this paper, we develop a new method, named the ρ_τ -IV estimation, to estimate structural equations based on the conditional quantile restriction imposed on the error terms, extending the \mathcal{L}_1 -IV approach proposed by Sakata (1997, 2007). We study the large sample behavior of the new estimator and show how to make statistical inferences on the regression parameters. In particular, we pay attention to the statistical inference under weak identification, as the weak identification is as important a possibility in the regression based on a conditional quantile restriction as in that based on the conditional mean restriction. We propose a weak-identification-robust test that can be viewed as a natural counterpart of the Anderson and Rubin's (1949) statistic in ρ_τ -IV estimation.

The conventional instrumental variables (IV) estimator is based on the identification of the structural parameters through the conditional mean restriction that the mean of the structural error term conditional on a set of instrumental variables is zero. The conditional mean restriction may look appealing, because, unlike the independence between the error term and the instruments, it does not impose restrictions on other features of the conditional distribution of the error term such as the variance of it.

Nevertheless, the conditional mean restriction is considered unsuitable in some applications. The conditional mean of a random variable critically depends on the tails of the conditional distribution of the variable. A small change in the tails can cause a large change in the conditional mean. In many applications, on the other hand, we know little about the part of the population distribution that correspond to the tails of the error distribution. This often makes it difficult to justify the conditional mean restriction.

The conditional mean restriction is not the only natural way to identify the parameters of structural equations. In many applications, the conditional mean restriction comes from an informal intuition that the "location" of the conditional distribution of the error term given a suitably chosen set of instruments should be constant. When we are faced by the above-mentioned concern about the conditional mean restriction, one would desire to capture the location of the conditional distribution of the error term by a measure that does not depend on tails. The conditional quantiles of the error term are examples of such location measures.

Sakata (1997, 2007) proposes identifying and estimating the regression parameters based on the conditional median restrictions. Chernozhukov and Hansen (2001, 2006) also consider identification of the regression parameters based on the conditional quantile restrictions and propose an estimation method, tak-

ing an approach related to but different from Sakata's. In the current paper, we extend the estimator of Sakata (1997, 2007) to propose a method called the ρ_τ -IV method to estimate regression models with the conditional τ -quantile restriction.

Being based on the same identification condition, our estimator is closely related to Chernozhukov and Hansen's estimator. The computation burden of the two estimators are also comparable, as should be clear from the discussion in Section 3 of the current paper. A benefit of our approach is that the objective function to be maximized in the ρ_τ -IV estimation takes a form similar to the "variance ratio" in the (normal) limited information maximum likelihood (LIML) estimation. This allows us to formulate a statistic analogous to the Anderson-Rubin (AR) statistic, with which we can make weak-identification-robust inference on the regression parameters of interest.

In the IV regression literature, many researchers have been paying attention to possible identification issues. Sargan (1983) points out that "near violation" of identifiability is problematic. The analysis of Phillips (1984, 1985) on the exact finite sample distribution of LIML clearly shows that lack of identifiability in structural equation estimation keeps the LIML estimator from consistently estimating the coefficient of the structural equation. Hillier (1990) also shows analogous results in considering the directional estimation of the coefficients of structural equations. Choi and Phillips (1992) further explores the behavior of the IV estimator under lack of identifiability.

When instrumental variables are poorly correlated with endogenous explanatory variables in linear regression, the asymptotic distribution of the IV estimator is quite different from what the standard large sample theory suggests, as demonstrated by Nelson and Startz (1990b, 1990a) and Bound, Jaeger, and Baker (1995). Staiger and Stock (1997) propose an alternative way to approximate the distribution of the IV estimator with weak instruments. Stock and Wright (2000) then establish a way to approximate the distribution of generalized-method-moments (GMM) estimators under weak identification. The proposed approximation methods are useful in theoretically studying the nature of the IV and GMM estimators under weak identification. Nevertheless, they do not offer a way to approximate the distribution of the estimators based on data, involving some unidentifiable nuisance parameters.

Given the absence of a convenient and reliable approximation to the distribution of the IV estimator with weak instruments, it is difficult to perform tests of hypotheses on regression parameters in the usual style (i.e., the t -test, the Wald test, etc.). On the other hand, the AR test originally proposed in Anderson

and Rubin (1949) is not affected by weakness of instruments. For this reason, Staiger and Stock (1997) and Dufour (1997) recommend the use of the AR test. The AR test even has nice power properties if the number of instruments is equal to the number of endogenous explanatory variables (Moreira 2001, Andrews, Moreira, and Stock 2004).

The weak identification is also an important possibility in regression based on a conditional quantile restriction. To this end, we propose a test that has asymptotically correct size regardless of whether the identification is strong or not. The hypothesis we consider is that some regression parameters are equal to prespecified values. If we apply the ρ_τ -IV method imposing the constraints of the null hypothesis, the objective function in the ρ_τ -IV estimation maximized subject to the parameter constraints of the null hypothesis tends to be close to one under the null. The constraint maximum of the objective function is similar to the Anderson and Rubin (1949) statistic in the sense that it captures how much of the fitted structural error can be explained by the instruments. It, ranging between zero and one, is closed to one if the fitted structural residuals cannot be fitted by the instruments in the sample. Its value far from one is thus taken as an evidence against the null in our test. If the conditions in the null hypothesis include the coefficients of all regressors potentially weakly related to the instruments excluded from the regression function, then the proposed test involves no weak identification problem, so that our test is robust to weak-identification.

Our test is closely related to Chernozhukov and Hansen (2008). They formulate a test in a way convenient in the estimation framework of Chernozhukov and Hansen (2001, 2006), while we propose a test convenient in the ρ_τ -IV estimation. Another paper related to our test is Jun (2008). Jun formulates a test adapting the approach of Kleibergen (2005).

The rest of the paper is organized as follows. We first describe the basic setup and define the ρ_τ -IV estimator in Section 2. Then, after briefly discussing the computation of the ρ_τ -IV estimator in Section 3, we establish the consistency and asymptotic normality of the ρ_τ -IV estimator and explains how to consistently estimate the asymptotic covariance matrix of ρ_τ -IV estimator in Section 4. In Section 5, we develop a weak-identification-robust method to test hypotheses on the regression parameters. Throughout the paper, $|\cdot|$ denotes the Euclidean norm for vectors and the Frobenius norm for matrices, and limits are taken along the sequence of sample sizes growing to infinity, unless otherwise indicated.

2 ρ_τ -IV Estimator

ASSUMPTION 1: Let (Ω, \mathcal{F}, P) be a probability space. The data are a realization of an independently and identically distributed stochastic process $\{X_t \equiv (y_t, Y_t', Z_t')' : \Omega \rightarrow \mathbb{R} \times \mathbb{R}^g \times \mathbb{R}^k\}_{t \in \mathbb{N}}$ such that $E[|X_1|] < \infty$, and for each $c \in \mathbb{R}^{1+g+k} \setminus \{0\}$, $P[c'X_t = 0] < 1$.

Partition Z_t as $Z_t = (Z_{t,1}', Z_{t,2}')'$, where $Z_{t,1}$ is $k_1 \times 1$, and $Z_{t,2}$ is $k_2 \times 1$ (so that $k_1 + k_2 = k$). The parameter of interest is the coefficients in regression of y_t on Y_t and $Z_{t,1}$ described in the next assumption.

ASSUMPTION 2: The subset \mathbf{B} of \mathbb{R}^g is nonempty and compact. There exists a unique $\theta_0 \in (\beta_0', \alpha_0')' \in \mathbf{B} \times \mathbb{R}^{k_1}$ such that the conditional τ -quantile of

$$U_1 \equiv y_1 - Y_1' \beta_0 - Z_{1,1}' \alpha_0$$

given Z_1 is zero, where τ is a known real constant in $(0, 1)$.

If instead the conditional τ -quantile of U_1 given Y_1 and $Z_{1,1}$ is known to be zero, β_0 and α_0 could be consistently estimated by the estimator of Koenker and Bassett (1978). In our current setup, Koenker and Bassett's estimator is inconsistent in general.

We here propose an estimator of the structural regression coefficients, following the approach described in Section 11 of Sakata (2007). Define $\rho_\tau : \mathbb{R} \rightarrow \mathbb{R}$ by

$$\rho_\tau(v) \equiv (\tau - 1(v < 0))v, \quad v \in \mathbb{R},$$

where $1(A)$ is the indicator function that becomes one if and only if the condition A is true. Also define functions $R : \mathbf{B} \times \mathbb{R}^{k_1} \times \mathbb{R}^k \rightarrow \mathbb{R}$ and $Q : \mathbf{B} \times \mathbb{R}^{k_1} \rightarrow \mathbb{R}$ by

$$R(\beta, \alpha, \gamma) \equiv E[\rho_\tau(y_1 - Y_1' \beta - Z_{1,1}' \alpha - Z_{1,2}' \gamma)], \quad (\beta, \alpha, \gamma) \in \mathbf{B} \times \mathbb{R}^{k_1} \times \mathbb{R}^k$$

and

$$Q(\beta, \alpha) \equiv \frac{\inf_{\gamma \in \mathbb{R}^k} R(\beta, \alpha, \gamma)}{R(\beta, \alpha, 0)}, \quad (\beta, \alpha) \in \mathbf{B} \times \mathbb{R}^{k_1},$$

where $R(\beta, \alpha, 0) > 0$ by the linear independence of the elements of $X_t = (y_t, Y_t', Z_t')'$ required by Assumption 1. Because of the conditional τ -quantile restriction imposed on U_1 in Assumption 2, we have that $Q(\beta_0, \alpha_0) = 1$, so that for each $(\beta, \alpha) \in \mathbf{B} \times \mathbb{R}^{k_1}$

$$0 \leq Q(\beta, \alpha) \leq Q(\beta_0, \alpha_0) = 1. \tag{1}$$

It follows that $\theta_0 \in (\beta'_0, \alpha'_0)'$ is the maximizer of Q over $\Theta \equiv \mathbf{B} \times \mathbb{R}^{k_1}$.

Our estimator is the maximizer of the sample counterpart of Q , which is given by a sequence of random functions $\{\hat{Q}_n : \mathbf{B} \times \mathbb{R}^{k_1} \times \Omega \rightarrow \mathbb{R}\}_{n \in \mathbb{N}}$ defined by

$$\hat{Q}_n(\beta, \alpha, \omega) \equiv \begin{cases} \frac{\inf_{\gamma \in \mathbb{R}^{k_2}} \hat{R}_n(\beta, \alpha, \gamma, \omega)}{\hat{R}_n(\beta, \alpha, 0, \omega)}, & \text{if } \inf_{(b, a) \in \mathbf{B} \times \mathbb{R}^{k_1}} \hat{R}_n(b, a, 0, \omega) > 0, \\ 1, & \text{otherwise,} \end{cases}$$

$$(\beta, \alpha) \in \mathbf{B} \times \mathbb{R}^{k_1}, \omega \in \Omega, n \in \mathbb{N},$$

where

$$\hat{R}_n(\beta, \alpha, \gamma, \omega) \equiv n^{-1} \sum_{t=1}^n \rho_\tau(y_t(\omega) - Y_t(\omega)' \beta - Z_{t,1}(\omega)' \alpha - Z_t(\omega)' \gamma),$$

$$(\beta, \alpha, \gamma) \in \mathbf{B} \times \mathbb{R}^{k_1} \times \mathbb{R}^k, \omega \in \Omega, n \in \mathbb{N}.$$

We now define our estimator.

DEFINITION 1 (THE ρ_τ -IV ESTIMATOR): *Given Assumption 1, a sequence of random vectors $\{\hat{\theta} \equiv (\hat{\beta}'_n, \hat{\alpha}'_n)'\} : \Omega \rightarrow \mathbf{B} \times \mathbb{R}^{k_1}\}_{n \in \mathbb{N}}$ is called the ρ_τ -IV estimator if for each $n \in \mathbb{N}$, $\hat{Q}_n(\hat{\beta}_n, \hat{\alpha}_n, \cdot) = \sup_{(\beta, \alpha) \in \mathbf{B} \times \mathbb{R}^{k_1}} \hat{Q}_n(\beta, \alpha, \cdot)$.*

For each $(\beta, \alpha) \in \mathbf{B} \times \mathbb{R}^{k_1}$, we have that

$$\begin{aligned} \inf_{\gamma \in \mathbb{R}^k} \hat{R}_n(\beta, \alpha, \gamma, \cdot) &= \inf_{(\gamma_1, \gamma_2) \in \mathbb{R}^{k_1} \times \mathbb{R}^{k_2}} n^{-1} \sum_{t=1}^n \rho_\tau(y_t - Y_t' \beta - Z_{t,1}'(\alpha + \gamma_1) - Z_{t,2}' \gamma_2) \\ &= \inf_{(\gamma_1, \gamma_2) \in \mathbb{R}^{k_1} \times \mathbb{R}^{k_2}} n^{-1} \sum_{t=1}^n \rho_\tau(y_t - Y_t' \beta - Z_{t,1}' \gamma_1 - Z_{t,2}' \gamma_2) = \inf_{\gamma \in \mathbb{R}^k} \hat{R}_n(\beta, 0, \gamma, \cdot). \end{aligned}$$

Given this fact, it holds that whenever $\hat{R}_n(\beta, \alpha, 0, \cdot) > 0$ for every $(\beta, \alpha) \in \mathbf{B} \times \mathbb{R}^{k_1}$,

$$\sup_{\alpha \in \mathbb{R}^{k_1}} \hat{Q}_n(\beta, \alpha, \cdot) = \sup_{\alpha \in \mathbb{R}^{k_1}} \frac{\inf_{\gamma \in \mathbb{R}^k} \hat{R}_n(\beta, 0, \gamma, \cdot)}{\hat{R}_n(\beta, \alpha, 0, \cdot)} = \frac{\inf_{\gamma \in \mathbb{R}^k} \hat{R}_n(\beta, 0, \gamma, \cdot)}{\inf_{\alpha \in \mathbb{R}^{k_1}} \hat{R}_n(\beta, \alpha, 0, \cdot)}, \quad \beta \in \mathbf{B}. \quad (2)$$

Because the numerator and denominator of the ratio on the right-hand side of (2) are continuous in β , $\sup_{\alpha \in \mathbb{R}^{k_1}} \hat{Q}_n(\beta, \alpha, \cdot)$ is continuous in β in all realizations, whenever $\hat{R}_n(\beta, \alpha, 0, \cdot) > 0$ for every $(\beta, \alpha) \in \mathbf{B} \times \mathbb{R}^{k_1}$. The continuity of $\sup_{\alpha \in \mathbb{R}^{k_1}} \hat{Q}_n(\beta, \alpha, \cdot)$ in β is also satisfied when $\hat{R}_n(\beta, \alpha, 0, \cdot)$ can touches zero, because $\hat{Q}_n(\beta, \alpha, \cdot) = 1$ in such case. Thus, given the compactness of \mathbf{B} , ρ_τ -IV estimator $\hat{\beta}_n$ of β_0 exists by the standard result on the existence of extremum estimators such as Gallant and White (1988, Theorem 2.2).

Further, $\hat{\alpha}_n$ is the solution of

$$\inf_{\alpha \in \mathbb{R}^{k_1}} \hat{R}_n(\hat{\beta}_n, \alpha, 0, \cdot) = \inf_{\alpha \in \mathbb{R}^{k_1}} n^{-1} \sum_{t=1}^n \rho_\tau((y_t - Y_t' \hat{\beta}_n) - Z_{t,1}' \alpha).$$

That is, it is the Koenker and Bassett's (1978) quantile regression estimator taking $(y_t - Y_t' \hat{\beta}_n)$ for the dependent variable and $Z_{t,1}$ for the regressors, which surely exists.

THEOREM 2.1: *Given Assumption 1, the ρ_τ -IV estimator exists.*

Remark. We could avoid the compactness requirement of \mathbf{B} by first defining the ρ_τ -IV directional estimator, as Sakata (2007) does, and then deriving the slope estimator in Definition 1 from it. We, however, directly define the slope estimator by imposing compactness on \mathbf{B} for saving space in this paper.

3 Computation of the ρ_τ -IV Estimator

We could calculate the ρ_τ -IV estimator, adapting the algorithm described in Sakata (2007) for the case $\tau = 0.5$ in the straightforward manner. Sakata's algorithm is, however, slow if k_1 is large, because it uses a global search algorithm to minimize \hat{Q}_n over $\mathbf{B} \times \mathbb{R}^{k_1}$.

Given a β , however, the ratio on the right-hand side of (2) can be quickly calculated, because the minimization problems appearing in both the numerator and denominator of the ratio can be rewritten as linear programming problems, as Koenker and Bassett (1978) explains. Thus, the ρ_τ -IV estimator can be calculated by maximizing the ratio in terms of β over \mathbf{B} . Because the ratio may have local maximum, it is advisable to use a global search algorithm such as the simulated annealing algorithm in calculating $\hat{\beta}_n$, while $\hat{\alpha}_n$ is the solution of the minimization problem in the denominator calculated with $\hat{\beta}_n$.

4 Large Sample Properties of the ρ_τ -IV Estimator

In investigating the consistency of the ρ_τ -IV estimator, it is convenient to consider the population counterpart of (2), i.e.,

$$\sup_{\alpha \in \mathbb{R}^{k_1}} Q(\beta, \alpha) = \sup_{\alpha \in \mathbb{R}^{k_1}} \frac{\inf_{\gamma \in \mathbb{R}^k} R(\beta, 0, \gamma)}{R(\beta, \alpha, 0)} = \frac{\inf_{\gamma \in \mathbb{R}^k} R(\beta, 0, \gamma)}{\inf_{\alpha \in \mathbb{R}^{k_1}} R(\beta, \alpha, 0)}, \quad \beta \in \mathbf{B}. \quad (3)$$

By Assumption 2, $\beta \mapsto \sup_{\alpha \in \mathbb{R}^{k_1}} Q(\beta, \alpha) : \mathbf{B} \rightarrow \mathbb{R}$ is a continuous function uniquely maximized at β_0 . We can also show that $\{\inf_{\alpha \in \mathbb{R}^{k_1}} \hat{Q}_n(\beta, \alpha, \cdot)\}_{n \in \mathbb{N}}$ converges to $\inf_{\alpha \in \mathbb{R}^{k_1}} Q(\beta, \alpha)$ uniformly in $\beta \in \mathbf{B}$ a.s.- P (Lemma A.3). By a standard result on consistency of extremum estimators (e.g., Pötscher and Prucha 1991, Lemma 4.2), we can establish the consistency of $\{\hat{\beta}_n\}_{n \in \mathbb{N}}$ for β_0 .

The estimator $\hat{\alpha}_n$, on the other hand, minimizes $\hat{R}_n(\hat{\beta}_n, \alpha, 0, \cdot)$ with respect to α over \mathbb{R}^{k_1} . Given the strong consistency $\{\hat{\beta}_n\}$ for β_0 , we can verify the a.s.- P convergence of $\hat{R}_n(\hat{\beta}_n, \alpha, 0, \cdot)$ to $R(\beta_0, \alpha, 0)$ for each $\alpha \in \mathbb{R}^{k_1}$ and utilize the convexity of $\hat{R}_n(\hat{\beta}_n, \alpha, 0, \cdot)$ in α to establish the strong consistency of $\hat{\alpha}_n$ for α_0 .

THEOREM 4.1: *Under Assumptions 1 and 2, $\{\hat{\theta}_n = (\hat{\beta}'_n, \hat{\alpha}'_n)'\}_{n \in \mathbb{N}}$ converges to $\theta_0 = (\beta'_0, \alpha'_0)'$.*

In establishing the asymptotic normality of the ρ_τ -IV estimator, we impose the additional conditions stated in the next theorem.

ASSUMPTION 3: (a) *The minimizer of $R(\beta_0, \alpha_0, \cdot) : \mathbb{R}^k \rightarrow \mathbb{R}$ over \mathbb{R} is unique (hence, it is uniquely minimized at the origin).*

(b) *The vector β_0 is interior to \mathbf{B} . Also, a neighborhood $B_0 \subset \mathbf{B}$ of β_0 , a neighborhood $A_0 \subset \mathbb{R}^{k_1}$ of α_0 , and a neighborhood $\Gamma_{2,0} \subset \mathbb{R}^{k_2}$ of the origin satisfy the following conditions:*

- (i) *The conditional distribution y_1 given Y_1 and Z_1 has a continuous probability density function (pdf) $f(\cdot | Y_1, Z_1)$ at $Y'_1\beta + Z'_{1,1}\alpha + Z'_{1,2}\gamma_2$ for each $(\beta, \alpha, \gamma_2) \in B_0 \times A_0 \times \Gamma_{2,0}$ a.s.- P .*
- (ii) *There exists a random variable $D : \Omega \rightarrow \mathbb{R}$ with a finite absolute moment such that for each $\beta \in B_0$, each $\alpha \in A_0$, and each $\gamma_2 \in \Gamma_{2,0}$,*

$$f(Y'_1\beta + Z'_{1,1}\alpha + Z'_{1,2}\gamma_2 | Y_1, Z_1) (|Y_1|^2 + |Z_1|^2) < D. \quad (4)$$

(c) *Let J be the Hessian of R at $(\beta'_0, \alpha'_0, 0'_{k \times 1})'$ and partition it as*

$$J \equiv \begin{pmatrix} J_{\beta\beta} & J_{\beta\alpha} & J_{\beta\gamma} \\ J_{\alpha\beta} & J_{\alpha\alpha} & J_{\alpha\gamma} \\ J_{\gamma\beta} & J_{\gamma\alpha} & J_{\gamma\gamma} \end{pmatrix},$$

where $J_{\beta\beta}$ is $g \times g$, $J_{\alpha\alpha}$ is $k_1 \times k_1$, and $J_{\gamma\gamma}$ is $k \times k$. Then the matrix

$$J_{\theta\theta} \equiv \begin{pmatrix} J_{\beta\beta} & J_{\beta\alpha} \\ J_{\alpha\beta} & J_{\alpha\alpha} \end{pmatrix}$$

is positive definite, and $J_{\gamma\theta} \equiv (J_{\gamma\beta}, J_{\gamma\alpha})$ is of full column rank.

(d) $E[|Y_1|^2 + |Z_1|^2] < \infty$.

Assumption 3(b) ensures the twice continuous differentiability of R in a neighborhood of $(\beta_0, \alpha_0, 0)$ in $\mathbb{R}^g \times \mathbb{R}^{k_1} \times \mathbb{R}^k$, which then implies the twice continuous differentiability of Q in a neighborhood of $\theta_0 = (\beta'_0, \alpha'_0)'$. The first condition in Assumption 3(c) ensures that the Hessian of $R(\beta_0, \alpha_0, \cdot) : \mathbb{R}^k \rightarrow \mathbb{R}$ at its minimum is negative definite. Under these conditions, the Hessian of $(\beta, \alpha) \mapsto \log Q(\beta, \alpha) : \mathbf{B} \times \mathbb{R}^{k_1} \rightarrow \mathbb{R}$ at $(\beta'_0, \alpha'_0)'$ is guaranteed to be positive definite, being equal to $-K$, where $K \equiv R(\beta_0, \alpha_0, 0)^{-1} J_{\theta\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\theta}$. The full column rankness of $J_{\gamma\beta}$ means that within a neighborhood of $(\beta'_0, \alpha'_0)'$, moving $(\beta', \alpha)'$ away from $(\beta'_0, \alpha'_0)'$ causes the gradient of $R(\beta, \alpha, \cdot) : \mathbb{R}^k \rightarrow \mathbb{R}$ to be bounded away from zero uniformly in all directions, so that we can choose γ to make $R(\beta, \alpha, \gamma)$ smaller than $R(\beta, \alpha, 0)$, once $(\beta', \alpha)'$ deviates from $(\beta'_0, \alpha'_0)'$. Assumption 3(d) ensures that the Lindeberg-Levy Central Limit Theorem (Rao 1973, p. 127) applies to the generalized score of the ρ_τ -IV estimator. The moment requirements in Assumptions 3(b,d) are mild. If $f(y_1 - Y'_1\beta - Z'_{1,1}\alpha - Z'_1\gamma | Y_1, Z_1)$ is bounded, they merely require that each element of Y_1 and Z_1 has a finite second moment, while the asymptotic normality of the conventional IV estimator is typically established under the assumption that the fourth moments of the dependent variable, the regressors, and the instruments are finite.

LEMMA 4.2: *Suppose that Assumptions 1–3 hold. Let $\{b_n\}_{n \in \mathbb{N}}$ and $\{a_n\}_{n \in \mathbb{N}}$ be sequences of \mathbf{B} - and \mathbb{R}^{k_1} -valued random vectors, respectively. Then there exists a sequence of $k \times 1$ random vectors c_n such that for each $n \in \mathbb{N}$, $\hat{R}_n(b_n, a_n, c_n, \cdot) = \inf_{\gamma \in \mathbb{R}^k} \hat{R}_n(b_n, a_n, \gamma, \cdot)$. If, in addition, Assumptions 3 hold, and $b_n \rightarrow \beta_0$ and $a_n \rightarrow \alpha_0$ in probability- P , then*

$$n^{1/2}c_n = Cn^{1/2} \begin{pmatrix} b_n - \beta_0 \\ a_n - \alpha_0 \end{pmatrix} + J_{\gamma\gamma}^{-1} n^{-1/2} \sum_{t=1}^n (\tau - 1)(U_t < 0) Z_t + o_P(n^{1/2}|b_n - \beta_0| + n^{1/2}|a_n - \alpha_0| + 1),$$

and $C \equiv -J_{\gamma\gamma}^{-1} J_{\gamma\theta}$.

Using this lemma, we can now approximate $\log \hat{Q}_n$.

LEMMA 4.3: *Suppose that Assumptions 1–3 hold and let $\{b_n\}_{n \in \mathbb{N}}$ and $\{a_n\}_{n \in \mathbb{N}}$ be sequences of \mathbf{B} - and*

\mathbb{R}^{k_1} -valued random vectors, respectively, that converge to β_0 and α_0 . Write $\tilde{\theta}_n \equiv (b'_n, a'_n)'$. Then

$$\begin{aligned} n \log \hat{Q}_n(b_n, a_n, \cdot) &= -\frac{1}{2R(\beta_0, \alpha_0, 0)} \left(n^{-1/2} \sum_{t=1}^n (\tau - 1(U_t < 0)) Z_t \right)' J_{\gamma\gamma}^{-1} \left(n^{-1/2} \sum_{t=1}^n (\tau - 1(U_t < 0)) Z_t \right) \\ &\quad - \frac{1}{R(\beta_0, \alpha_0, 0)} n^{-1/2} \sum_{t=1}^n (\tau - 1(U_t < 0)) Z_t' C n^{1/2} (\tilde{\theta}_n - \theta_0) - \frac{1}{2} n^{1/2} (\tilde{\theta}_n - \theta_0)' K n^{1/2} (\tilde{\theta}_n - \theta_0) \\ &\quad + o_P(n^{1/2} |b_n - \beta_0| + n |b_n - \beta_0|^2 + 1). \end{aligned} \quad (5)$$

Given this lemma, it is natural to expect that the minimizer of the the second and third terms on the right-hand side of (5) approximates $\hat{\theta} = (\hat{\beta}'_n, \hat{\alpha}'_n)'$. The next theorem confirms that such approximation bears an $o_P(1)$ approximation error, and derives the asymptotic distribution of $\{\hat{\theta}_n\}_{n \in \mathbb{N}}$ based on the approximation.

THEOREM 4.4: *Suppose that Assumptions 1–3 hold. Then*

$$n^{1/2}(\hat{\theta}_n - \theta_0) = -\frac{1}{R(\beta_0, \alpha_0, 0)} K^{-1} C' n^{-1/2} \sum_{t=1}^n (1(U_t < 0) - \tau) Z_t + o_P(1),$$

and

$$D^{-1/2} n^{1/2} (\hat{\theta}_n - \theta_0) \overset{A}{\approx} N(0, I_1),$$

where $D \equiv K^{-1} C' V C K^{-1}$, $K = R(\beta_0, \alpha_0, 0)^{-1} J_{\theta\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\theta}$ (as introduced earlier), and

$$V \equiv \tau(1 - \tau) R(\beta_0, \alpha_0, 0)^{-2} \mathbf{E}[Z_1 Z_1'].$$

To estimate the asymptotic covariance matrix D consistently, we need to estimate V , K , and C consistently. For consistent estimation of V , we can use its sample analogue,

$$\hat{V}_n \equiv \tau(1 - \tau) \hat{R}_n(\hat{\beta}_n, \hat{\alpha}_n, 0, \cdot)^{-2} n^{-1} \sum_{t=1}^n Z_t Z_t'.$$

On the other hand, K and C are more complicated, depending on J , the Hessian of R . The Hessian of \hat{R}_n is zero at each point in $\mathbf{B} \times \mathbb{R}^{k_1} \times \mathbb{R}^k$, at which it is differentiable. This rules out estimation of J by using of the Hessian of \hat{R}_n .

A way to overcome the difficulty in estimation of K and C is to employ the numerical differentiation approach described in Newey and McFadden (1994, Section 7.3). Because $-K$ is the Hessian of $(\beta, \alpha) \mapsto \log Q(\beta, \alpha) : \mathbf{B} \times \mathbb{R}^{k_1} \rightarrow \mathbb{R}$ at $(\beta'_0, \alpha'_0)'$, -1 times a second-order numerical derivative of $\log \hat{Q}(\beta, \alpha, \cdot)$ at $(\hat{\beta}'_n, \hat{\alpha}'_n)'$ is our estimator of K . Let e_i^m denote the unit vector along the i th axis of the Cartesian coordinate system in \mathbb{R}^m . Assume:

ASSUMPTION 4: *The sequence $\{h_n\}_{n \in \mathbb{N}}$ consists of positive (possibly random) numbers such that $h_n \rightarrow 0$ and $n^{1/2}h_n \rightarrow \infty$ in probability- P .*

Then our estimator of K is \hat{K}_n , whose (i, j) -element is given by

$$\begin{aligned} \hat{K}_{nij} \equiv & -\frac{1}{4h_n^2} (\log \hat{Q}_n(\hat{\theta}_n + h_n e_i + h_n e_j, \cdot) - \log \hat{Q}_n(\hat{\theta}_n - h_n e_i + h_n e_j, \cdot) \\ & - \log \hat{Q}_n(\hat{\theta}_n + h_n e_i - h_n e_j, \cdot) + \log \hat{Q}_n(\hat{\theta}_n - h_n e_i - h_n e_j, \cdot)), \quad (i, j) \in \{1, \dots, (g + k_1)\}^2, n \in \mathbb{N}. \end{aligned}$$

For C , we utilize the result of Lemma 4.2, which suggests that perturbation in $\hat{\theta}_n = (\hat{\beta}'_n, \hat{\alpha}'_n)'$ would change $\hat{\gamma}_n$ approximately by C times the change in $\hat{\theta}_n$. Let $\hat{\gamma}_{ni}(\theta)$ denote the i th element in the usual quantile regression estimator in regression of $y_t - (Y'_t, Z'_{t,1})\theta$ on Z_t ($i \in \{1, 2, \dots, k\}$). Then our estimator of C is \hat{C}_n whose (i, j) -element is given by

$$\hat{C}_{nij} \equiv \frac{1}{2h_n} (\hat{\gamma}_{ni}(\hat{\theta}_n + h_n e_j) - \hat{\gamma}_{ni}(\hat{\theta}_n - h_n e_j)), \quad i \in \{1, \dots, k\}, j \in \{1, \dots, (g + k_1)\}.$$

Given the estimators of K and C , we estimate D by $\hat{D}_n \equiv \hat{K}_n^+ \hat{C}'_n \hat{V}_n \hat{C}_n \hat{K}_n^+$, $n \in \mathbb{N}$, where \hat{K}_n^+ is the Moore-Penrose (MP) inverse of \hat{K}_n (we use the MP inverse instead of the regular inverse to ensure that this estimator is well defined for every realization).

THEOREM 4.5: *Suppose that Assumptions 1–4 hold. Then:*

- (a) $\{\hat{K}_n\}_{n \in \mathbb{N}}$ *is weakly consistent for K .*
- (b) $\{\hat{C}_n\}_{n \in \mathbb{N}}$ *is weakly consistent for C .*
- (c) $\{\hat{V}_n\}_{n \in \mathbb{N}}$ *is weakly consistent for V .*
- (d) $\{\hat{D}_n\}_{n \in \mathbb{N}}$ *is weakly consistent for D .*

Remark. The same step size h_n is used in each element of \hat{K}_n and \hat{C}_n just for simplicity. One could use a different step size for each element in \hat{K}_n and \hat{C}_n without affecting the consistency results in Theorem 4.5, as long as the step size satisfies the requirements in Assumption 4.

5 Testing on the Regression Coefficients under Possible Weak Identification

When $(\beta, \alpha) \mapsto \log Q(\beta, \alpha)$ is flat in some directions from (β_0, α_0) , compared with the size of the error in approximating $\log Q$ by $\log \hat{Q}_n$, the large sample distribution of the ρ_τ -IV estimator established in Section 4 can be unreliable, because the estimator can easily go astray. In other words, we may experience the so-called weak identification problem in the ρ_τ -IV estimation.

The flatness of $(\beta, \alpha) \mapsto \log Q(\beta, \alpha)$ described above implies near singularity of K , which is -1 times the Hessian of $\log Q(\beta, \alpha)$. Because the large sample analysis in Section 4 involves the inverse of K , the near singularity of K makes the results in Section 4 unreliable unless the sample size is extremely large. To verify that the nearly singular K can arise in practice, suppose that Y_t is related to Z_t through

$$Y_t = \Pi_0 Z_t + V_t, \quad t \in \mathbb{N},$$

where Π_0 is a $g \times k$ constant matrix, and V_t is a $g \times 1$ zero-mean random vector independent from Z_t . Let $f_U(\cdot | Z_1)$ denote the conditional pdf of U_1 given Z_1 . Then, under our current assumption, we have that

$$J_{\theta_\gamma} = 2\mathbb{E} \left[f_U(0|Z) \begin{pmatrix} Y_1 \\ Z_{1,1} \end{pmatrix} Z_1' \right] = 2\mathbb{E} \left[f_U(0|Z_1) \begin{pmatrix} \Pi_0 Z_1 \\ Z_{1,1} \end{pmatrix} Z_1' \right]$$

If the last k_2 columns of Π_0 is close to zero, each of the first g rows of J_{θ_γ} can be well approximated by a linear combination of the last k_1 rows of J_{θ_γ} ; i.e., the columns of J_{θ_γ} becomes nearly dependent. This causes $K = R(\beta_0, \alpha_0, 0)^{-1} J_{\theta_\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\theta}$ to be nearly singular and raises concerns about inference on β_0 and α_0 , relying on the asymptotics in Section 4.

Suppose that we are interested in the hypothesis that $H_0: \beta_0 = \bar{\beta}$, where $\bar{\beta}$ is a known $g \times 1$ constant vector. In the usual IV regression based on the zero-conditional mean restriction imposed on the error term, the AR test is known to be robust to weakness of instruments. Given the structural equation estimated under the constraint of the null hypothesis, the AR test regresses the null-restricted fitted structural error term on all instruments and checks if R^2 is close to zero. If R^2 is high enough, it rejects the null hypothesis.

Because the AR test rejects the null when $1 - R^2$ is close to zero, we can view the AR test as rejecting the null hypothesis when the null-restricted fitted structural error term can be well explained by the instruments. Note that $1 - R^2$ is equal to the ratio of the two sample second moments. The denominator in the ratio

is the sample second moment of the fitted structural error term, while the numerator is the sample second moment of the residuals in regression of the structural error term on the instruments. This view gives us a way to adapt Anderson and Rubin's (1949) approach in our problem setup. Namely, we replace the sample second moment in $1 - R^2$ with the corresponding average check functions. The resulting statistic is $\hat{Q}_n(\bar{\beta}, \hat{\alpha}_n^0, \cdot)$, where $\hat{\alpha}_n^0$ is the ρ_τ -IV estimator obtained imposing the constraint of H_0 , which is exactly equal to the Koenker and Bassett's (1978) estimator in regression of $y_t - Y_t' \bar{\beta}$ on $Z_{1,t}$. For convenience, we take the logarithm of it and multiply it by $-2n$ to define a test statistic \mathcal{J}_n .

$$\mathcal{J}_n \equiv -2n \log \hat{Q}_n(\bar{\beta}, \hat{\alpha}_n^0, \cdot) = -2n \log \frac{\inf_{\gamma \in \mathbb{R}^k} \hat{R}_n(\bar{\beta}, 0, \gamma, \cdot)}{\inf_{\alpha \in \mathbb{R}^{k_1}} \hat{R}_n(\bar{\beta}, \alpha, 0, \cdot)}, \quad n \in \mathbb{N}. \quad (6)$$

Let $\bar{\alpha}$ be a $k_1 \times 1$ vector such that $Z'_{1,1} \bar{\alpha}$ be the ρ_τ -metric projection of $y_1 - Y_1' \bar{\beta}$ on the linear space spanned by the elements of $Z_{1,1}$. Then the standard large sample analysis on extremum estimation shows that

$$n^{-1} \mathcal{J}_n \rightarrow -2 \sup_{\alpha \in \mathbb{R}^{k_1}} \log Q(\bar{\beta}, \alpha) = -2 \log Q(\bar{\beta}, \bar{\alpha}) \text{ in probability-}P.$$

Under H_0 , the right-hand side of this equality is zero, because

$$\sup_{\alpha \in \mathbb{R}^{k_1}} Q(\bar{\beta}, \alpha) = \sup_{\alpha \in \mathbb{R}^{k_1}} Q(\beta_0, \alpha) = Q(\beta_0, \alpha_0) = 1.$$

Under the alternative, on the other hand, the limit of $\{n^{-1} \mathcal{J}_n\}_{n \in \mathbb{N}}$ is strictly positive, because

$$Q(\bar{\beta}, \bar{\alpha}) < Q(\beta_0, \alpha_0) = 1.$$

Thus, a test based on \mathcal{J}_n should reject H_0 if \mathcal{J}_n exceeds a suitably chosen critical value. We will discuss how to find the critical value below.

$$\text{Define } C^0 \equiv -J_{\gamma\gamma}^{-1} J_{\gamma\alpha} \text{ and } L \equiv R(\beta_0, \alpha_0, 0)^{-1} J_{\alpha\alpha}.$$

LEMMA 5.1: *Suppose that Assumptions 1–3 hold. If in addition H_0 is true, then*

$$\mathcal{J}_n \overset{A}{\rightsquigarrow} \eta' (L^{-1} - C^0 (C^{0'} L C^0)^{-1} C^{0'}) \eta,$$

where η is a $k \times 1$ random vector distributed with $N(0, V)$,

Thus, $\{\mathcal{J}_n\}_{n \in \mathbb{N}}$ has a non-degenerate limiting distribution, though it is not asymptotically pivotal. Among the unknown parameters in the formula for the asymptotic distribution of $\{\mathcal{J}_n\}_{n \in \mathbb{N}}$, C^0 can be consistently

estimated by applying $\{\hat{C}_n\}_{n \in \mathbb{N}}$ “under the null” (Theorem 4.5). Write $\hat{\theta}_n^0 \equiv (\bar{\beta}', \hat{\alpha}_n^{0'})'$. Then our estimator of C^0 is \hat{C}_n^0 whose (i, j) -element is equal to

$$\hat{C}_{nij}^0 \equiv \frac{1}{2h_n} (\hat{\gamma}_{ni}(\hat{\theta}_n^0 + h_n e_{g+j}) - \hat{\gamma}_{ni}(\hat{\theta}_n^0 - h_n e_{g+j})), \quad i \in \{1, \dots, k\}, j \in \{1, \dots, k_1\}.$$

Analogously, V can be estimated by $\hat{V}_n^0 \equiv \hat{R}_n(\hat{\beta}_n^0, \hat{\alpha}_n^0, 0, \cdot)^{-2} n^{-1} \sum_{t=1}^n Z_t Z_t'$.

The matrix L is the Hessian of $\gamma \mapsto \log R(\beta_0, \alpha_0, \gamma) : \mathbb{R}^k \rightarrow \mathbb{R}$ at the origin. We take a second-order numerical derivative of the sample counterpart of this function to estimate L . The resulting estimator \hat{L}_n of L is the $k \times k$ matrix with (i, j) -element equal to

$$\begin{aligned} \hat{L}_{nij} \equiv & \frac{1}{4h_n^2} (\log \hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0 + h_n e_i + h_n e_j, \cdot) - \log \hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0 - h_n e_i + h_n e_j, \cdot) \\ & - \log \hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0 + h_n e_i - h_n e_j, \cdot) + \log \hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0 - h_n e_i - h_n e_j, \cdot)), \quad i, j = 1, 2, \dots, k, \end{aligned}$$

where $\hat{\gamma}_n^0$ is the τ -quantile regression estimator in regressing $y_t - Y_t' \bar{\beta} - Z_{1,t}' \hat{\alpha}_n^0$ on Z_t .

LEMMA 5.2: *Suppose that Assumptions 1–3 and 4 hold. If in addition H_0 holds, $\{\hat{L}_n^0\}_{n \in \mathbb{N}}$ is consistent for L .*

The limiting distribution of $\{\mathcal{J}_n\}_{n \in \mathbb{N}}$ is that of a positive random variable whose distribution function is positively sloped at each positive point. Let $c(p, \tilde{C}, \tilde{L}, \tilde{V})$ denote the $(1-\alpha)$ -quantile of $\tilde{\eta}'(\tilde{L}^+ - \tilde{C}(\tilde{C}'\tilde{L}\tilde{C}) + \tilde{C}')\tilde{\eta}$ for each $k \times l$ matrix \tilde{C} , each $k \times k$ symmetric matrix \tilde{L} , and each $k \times k$ symmetric matrix \tilde{V} , where $\tilde{\eta}$ is a $k \times 1$ random vector distributed with $N(0, \tilde{V})$, and $p \in (0, 1)$, where (a, b) denotes the open interval between real numbers a and b . We here propose a test that rejects H_0 if and only if \mathcal{J}_n exceeds $c(p, \hat{C}_n, \hat{L}_n, \hat{V}_n)$, where p is the desired size of the test. This test has the correct asymptotic size and it is consistent, as stated in the next theorem.

THEOREM 5.3: *Suppose that Assumptions 1–4 hold. Then:*

- (a) *If in addition H_0 holds, for each $p \in (0, 1)$, $P[\mathcal{J}_n > c(p, \hat{C}_n^0, \hat{L}_n^0, \hat{V}_n^0)] \rightarrow p$.*
- (b) *Suppose instead that H_0 is violated, that $R(\bar{\beta}, \cdot, 0) : \mathbb{R}^{k_1} \rightarrow \mathbb{R}$ has a unique maximizer on \mathbb{R}^{k_1} , and that $R(\bar{\beta}, 0, \cdot) : \mathbb{R}^k \rightarrow \mathbb{R}$ has a unique minimizer on \mathbb{R}^k . Then for each $p \in (0, 1)$, $P[\mathcal{J}_n > c(p, \hat{C}_n^0, \hat{L}_n^0, \hat{V}_n^0)] \rightarrow 1$.*

Because each quadratic form of normal random variables can be easily rewritten as a linear combination of χ^2 random variables using the eigenvalue decomposition, $c(p, \hat{C}_n, \hat{L}_n, \hat{V}_n)$ is the $(1 - \alpha)$ -quantile of a linear combination of χ^2 random variables. To compute $c(p, \hat{C}_n, \hat{L}_n, \hat{V}_n)$, we can numerically find the $(1 - \alpha)$ -quantile of the distribution of the linear combination, evaluating the distribution function by using Farebrother's (1984) algorithm.

6 Power under weak instruments

According to Theorem 5.3(b), our test proposed in the previous section is consistent in the regular asymptotic framework with strong instruments. In this section, we discuss the power properties of the test when the instruments are weak. For this purpose, we need a model describing how weak instruments arise in our problem. Before formalizing the notion of weak instruments in our problem setup, we first review the concept of weak instruments in the conventional IV regression. Staiger and Stock (1997) introduces weak instruments in a thought experiment in which the correlation between the endogenous regressors and the instruments becomes weaker as the sample size grows. More concretely, they relate the $k \times 1$ instrument vector Z_t to the $g \times 1$ endogenous regressor vector $Y_t^{(n)}$ through

$$Y^{(n)} = n^{-1/2} \Lambda Z_t + V_t, \quad t \in \{1, 2, \dots, n\}, n \in \mathbb{N}$$

where Λ is a $g \times k$ constant matrix, and V_t is a unobservable $g \times 1$ random vector such that Z_t is exogenous to V_t . The superscript “ (n) ” in $Y_t^{(n)}$ indicates dependence of $Y_t^{(n)}$ on n . The structural equation in the thought experiment is then

$$y_t^{(n)} = Y_t^{(n)'} \beta_0 + Z_{t,1}' \alpha_0 + U_t, \quad t \in \{1, 2, \dots, n\}, n \in \mathbb{N},$$

where $Z_{t,1}$ is a $k_1 \times 1$ subvector of Z_t , and the regression error U_t is orthogonal to Z_t . In this setup, Staiger and Stock investigates the asymptotic behavior of tests of the hypothesis that $H_0: \beta_0 = \bar{\beta}$, where $\bar{\beta}$ is a known constant in \mathbb{R}^g .

Define $W_t \equiv U_t - V_t'(\bar{\beta} - \beta_0)$ ($t \in \mathbb{N}$). Then it is straightforward to verify that the “null-restricted residual”, i.e., the residual evaluated with coefficients $(\bar{\beta}', \alpha_0)'$ is equal to

$$y_t^{(n)} - Y_t^{(n)'} \bar{\beta} - Z_{t,1}' \alpha_0 = W_t - Z_t' n^{-1/2} \Lambda' (\bar{\beta} - \beta_0). \quad (7)$$

Because

$$\mathbb{E}[Z_t W_t] = 0, \tag{8}$$

it follows that

$$\mathbb{E}[Z_t(y_t - Y_t' \bar{\beta} - Z_{t,1}' \alpha_0)] = -\mathbb{E}[Z_t Z_t'] n^{-1/2} \Lambda'(\bar{\beta} - \beta_0).$$

Thus, the null restricted residual violates the moment condition underlying the conventional IV estimator, but only in the order of $n^{-1/2}$. This is the essential feature of the setup that Staiger and Stock used to demonstrate that the behavior of the conventional tests of H_0 may be very different from what the conventional asymptotic analysis indicates, and why the AR test can be a better choice.

Note that, while the fact that the null-restricted residual violates the moment condition in the order of $n^{-1/2}$ hinges on (7) and (8), it does not matter for it what W_t is or where Λ comes from. Also, note that there is no natural universally agreeable reduced-form equation in our setup, unlike the conventional IV regression setup. In analyzing our test of H_0 with weak instruments, we therefore take as basis (7) and (8) suitably modified for constructing an environment with weak instruments in our setup, as found in the next assumption.

ASSUMPTION 5: *The triangle array $\{X_t^{(n)} \equiv (y_t^{(n)}, Y_t^{(n)}, Z_{t,1}, Z_{t,2}) : t \in \{1, 2, \dots, n\}, n \in \mathbb{N}\}$ consists of random vectors on a probability space (Ω, \mathcal{F}, P) , where $y_t^{(n)}$, $Y_t^{(n)}$, $Z_{t,1}$, and $Z_{t,2}$ are 1×1 , $g \times 1$, $k_1 \times 1$, and $k_2 \times 1$, respectively; $\bar{\beta}$ is a constant vector in \mathbb{B} that is a nonempty and compact subset of \mathbb{R}^g ; and τ is a known constant in $(0, 1)$. There exists $\beta_0 \in \mathbb{B}$, a $g \times k$ matrix Λ , $\bar{\alpha} \in \mathbb{R}^{k_1}$, and a sequence of random variables $\{W_t\}_{t \in \mathbb{N}}$ that satisfy that*

$$y_t^{(n)} - Y_t^{(n)'} \bar{\beta} - Z_{t,1}' \bar{\alpha} = W_t - Z_t' n^{-1/2} \Lambda'(\bar{\beta} - \beta_0), \quad t \in \{1, 2, \dots, n\}, n \in \mathbb{N}, \tag{9}$$

and that for each $t \in \mathbb{N}$, τ -quantile of W_t given Z_t is zero.

In Assumption 5, β_0 appears as some vector satisfying the required condition, rather than the true coefficient of $Y_t^{(n)}$, because our mathematical results do not depend on what β_0 is. Of course, our results are most useful when Assumption 5 holds with β_0 set equal to the true true coefficient of $Y_t^{(n)}$. If $\beta_0 = \bar{\beta}$, the conditional quantile restriction imposed upon $\{W_t\}_{t \in \mathbb{N}}$ is essentially the same as Assumption 2. The

equivalence of the two conditions can be achieved by setting $\bar{\alpha} = \alpha_0$ and $W_1 = U_1$, in particular when we require that $\{(Z'_t, W'_t)\}_{t \in \mathbb{N}}$ is i.i.d., as we will do below. When $\beta_0 \neq \bar{\beta}$, the assumption implies that the conditional τ -quantile of the null-restricted residual given Z_1 is local-to-zero. In general, the distribution of W_t depends on $\bar{\beta} - \beta_0$. Assumption 5 is clearly satisfied, if (U_t, V'_t) is independent from Z_1 in the setup of Staiger and Stock (1997) discussed above. The matrix Λ captures the strength of the instruments. For example, the instruments are irrelevant when $\Lambda = 0$.

In addition to Assumption 5, we impose the following conditions similar to Assumption 3:

ASSUMPTION 6: (a) $E\rho_\tau(W_1 - Z'_1\gamma)$ is uniquely minimized at $\gamma = 0_{k \times 1}$.

(b) A neighborhood $\Gamma_0 \subset \mathbb{R}^k$ of the origin satisfies the following conditions:

(i) The conditional distribution W_1 given Z_1 denoted by $F(\cdot | Z_1)$ has a pdf $f(\cdot | Z_1)$ at $Z'_1\gamma$ for each $\gamma \in \Gamma_0$ a.s.-P.

(ii) There exists a random variable $D : \Omega \rightarrow \mathbb{R}$ with a finite second moment such that for each $\gamma \in \Gamma_0$, $f(Z'_1\gamma | Z_1)|Z_1|^2 < D$ a.s.-P.

(c) $J_{\gamma\gamma} = \partial^2 R(\bar{\beta}, \bar{\alpha}, 0_{k \times 1}) / \partial\gamma\partial\gamma'$ is positive definite, and $J_{\gamma\alpha} = \partial^2 R(\bar{\beta}, \bar{\alpha}, 0_{k \times 1}) / \partial\gamma\partial\alpha'$ is of full column rank.

(d) $E[|Z_1|^2] < \infty$.

(e) $\{(W_t, Z'_t) : t = 1, \dots, n\}$ are independent and identically distributed.

The following theorem describes the asymptotic distribution of \mathcal{J}_n in the case of fixed alternatives ($\bar{\beta} - \beta_0$ is a fixed vector) and the weak IVs design assumed in Assumption 5.

THEOREM 6.1: Suppose that Assumptions 5 and 6 hold. Then

$$\mathcal{J}_n \overset{A}{\rightsquigarrow} (E\rho_\tau(W_1))^{-1} (\eta + J_{\gamma\gamma} \Lambda' (\bar{\beta} - \beta_0))' (J_{\gamma\gamma}^{-1} - C^0 (C^{0'} J_{\gamma\gamma} C^0)^{-1} C^{0'}) (\eta + J_{\gamma\gamma} \Lambda' (\bar{\beta} - \beta_0)),$$

where η is a $k \times 1$ random vector distributed with $N(0, \tau(1 - \tau)E[Z_1 Z'_1])$.

In the case of weak instruments and under fixed alternatives, the asymptotic distribution of \mathcal{J}_n is a non-central mixed- χ^2 random variable. The power of the test that rejects $H_0 : \bar{\beta} = \beta_0$ when $\mathcal{J}_n > c(\alpha, \hat{C}_n^0, \hat{L}_n^0, \hat{V}_n^0)$

depends on the magnitude of the the non-centrality parameter given by

$$(\bar{\beta} - \beta_0)' \Lambda (J_{\gamma\gamma} - J_{\gamma\gamma} C^0 (C^{0'} J_{\gamma\gamma} C^0)^{-1} C^{0'} J_{\gamma\gamma}) \Lambda' (\bar{\beta} - \beta_0),$$

where $J_{\gamma\gamma} - J_{\gamma\gamma} C^0 (C^{0'} J_{\gamma\gamma} C^0)^{-1} C^{0'} J_{\gamma\gamma}$ is a positive definite matrix by Assumption 6(c). Under H_0 , $\bar{\beta} - \beta_0 = 0$ and the test rejects asymptotically with probability α . Thus, the test has correct size regardless of the strength of the instruments. Under the fixed alternatives, the asymptotic rejection probability depends on the distance between $\bar{\beta}$ and β_0 and the strength of the instruments Λ . For example, the test has no power when the instruments are irrelevant and $\Lambda = 0$. The test also lacks power in certain directions if $\Lambda \neq 0$ however its rank is less than g .

Appendix A Mathematical Proofs

Given Assumption 1, write $\|\xi\|_{\rho_\tau} \equiv E[\rho_\tau(\xi)]$ for each $\xi \in \mathcal{L}_1(\Omega, \mathcal{F}, P)$. Then $\|\cdot\|_{\rho_\tau}$ is a pseudo norm on $\mathcal{L}_1(\Omega, \mathcal{F}, P)$. Using $\|\cdot\|_{\rho_\tau}$, R can be written as

$$R(\beta, \alpha, \gamma) = \|y_1 - Y_1' \beta - Z_{1,1}' \alpha - Z_1' \gamma\|_{\rho_\tau}, \quad (\beta, \alpha, \gamma) \in \mathbf{B} \times \mathbb{R}^{k_1} \times \mathbb{R}^k.$$

It follows that the minimization in the numerator of the ratio on the right-hand side of (3) is the $\|\cdot\|_{\rho_\tau}$ -metric projection of $y_1 - Y_1' \beta$ on Z_1 , while the minimization in the denominator is the $\|\cdot\|_{\rho_\tau}$ -metric projection of $y_1 - Y_1' \beta$ on $Z_{1,1}$. The norm $\|\cdot\|_{\rho_\tau}$ is closely related to the \mathcal{L}_1 norm $\|\cdot\|_1$. They actually generate the equivalent topologies, because

$$\|\xi\|_{\rho_\tau} \leq \|\xi\|_1 \leq \frac{1}{\min\{\tau, 1 - \tau\}} \|\xi\|_{\rho_\tau}$$

An important implication of the equivalence is that $\|\xi\|_{\rho_\tau} = 0$ if and only if $\|\xi\|_1 = 0$. Our analysis uses the equivalence of the two norms, mostly without mentioning it explicitly.

We show below that $\{\sup_{\alpha \in \mathbb{R}^{k_1}} \hat{Q}_n(\beta, \alpha, \cdot)\}_{n \in \mathbb{N}}$ converges to $\sup_{\alpha \in \mathbb{R}^{k_1}} Q(\beta, \alpha)$ uniformly in β on the compact set \mathbf{B} . We can then conclude that $\{\hat{\beta}_n\}_{n \in \mathbb{N}}$ is consistent for β_0 , because β_0 is the unique maximizer of $\beta \mapsto \sup_{\alpha \in \mathbb{R}^{k_1}} Q(\beta, \alpha)$ on \mathbf{B} . Once the consistency of $\hat{\beta}_n$ is established, we can also prove that $\{\hat{\alpha}_n\}_{n \in \mathbb{N}}$ converges a.s.- P to α_0 , at which $R(\beta_0, \cdot, 0) : \mathbb{R}^{k_1} \rightarrow \mathbb{R}$ is minimized, by utilizing the convexity of $\hat{R}_n(\hat{\beta}_n, \alpha, 0, \cdot)$ in α and the pointwise convergence of $\{\hat{R}_n(\hat{\beta}_n, \alpha, 0, \cdot)\}_{n \in \mathbb{N}}$ to $R(\beta_0, \alpha, 0)$ for each α .

We first establish a few lemmas. For later conveniences, some lemmas have more generality than we need for proving Theorem 4.1. The generality will be useful in our proof of 4.4.

LEMMA A.1: *Suppose that Assumptions 1 holds. Then for each $\beta \in \mathbf{B}$,*

$$\inf_{\gamma \in \mathbb{R}^k} \hat{R}_n(\beta, 0, \gamma, \cdot) - \inf_{\gamma \in \mathbb{R}^k} R(\beta, 0, \gamma) \rightarrow 0 \text{ a.s.-}P,$$

and

$$\inf_{\alpha \in \mathbb{R}^{k_1}} \hat{R}_n(\beta, \alpha, 0, \cdot) - \inf_{\alpha \in \mathbb{R}^{k_1}} R(\beta, \alpha, 0) \rightarrow 0 \text{ a.s.-}P.$$

Proof of LEMMA A.1: The two convergence results can be proved in similar manners. We only prove the first one. Let β be an arbitrary point in \mathbf{B} . Then the $\|\cdot\|_{\rho_\tau}$ -metric projection of $y_1 - Y_1'\beta$ on the linear subspace spanned by Z_1 exists and is in general a compact set. By the linear independence of Z_1 (Assumption 1), this further means that $\Gamma_1 \equiv \arg \min_{\gamma \in \mathbb{R}^k} R(\beta, 0, \gamma)$ is compact. It follows that there exists a closed ball Γ_2 containing Γ_1 in its interior. Now fix a point γ_1 in Γ_1 . By the Kolmogorov law of large numbers (Rao 1973, p. 115), $\{\hat{R}_n(\beta, 0, \gamma_1, \cdot)\}_{n \in \mathbb{N}}$ converges to $R(\beta, 0, \gamma_1)$ a.s.- P . Also, by Jennrich's uniform law of large numbers (Jennrich 1969, Theorem 2), $\{\hat{R}_n(\beta, 0, \gamma, \cdot) - R(\beta, 0, \gamma)\}_{n \in \mathbb{N}}$ converges to zero uniformly in γ on the boundary $\partial\Gamma_2$ of Γ_2 a.s.- P . Because $\hat{R}_n(\beta, 0, \gamma, \cdot)$ is convex as a function of γ , and

$$R(\beta, 0, \gamma_1) < \inf_{\gamma \in \partial\Gamma_2} R(\beta, 0, \gamma),$$

it follows from the above-mentioned facts that

$$\hat{R}_n(\beta, 0, \gamma_1, \cdot) < \inf_{\gamma \in \mathbb{R}^k \setminus \Gamma_2} \hat{R}_n(\beta, 0, \gamma, \cdot)$$

for almost all $n \in \mathbb{N}$ a.s.- P . On the other hand, by Jennrich's uniform law of large numbers, $\{\hat{R}_n(\beta, 0, \gamma, \cdot) - R(\beta, 0, \gamma)\}_{n \in \mathbb{N}}$ converges to zero uniformly in $\gamma \in \Gamma_2$ a.s.- P , so that

$$\inf_{\gamma \in \Gamma_2} \hat{R}_n(\beta, 0, \gamma, \cdot) \rightarrow R(\beta, 0, \gamma_1) \text{ a.s.-}P.$$

The desired result therefore follows. \square

LEMMA A.2: *Suppose that Assumptions 1 holds. Then*

$$\sup_{\beta \in \mathbf{B}} \left| \inf_{\gamma \in \mathbb{R}^k} \hat{R}_n(\beta, 0, \gamma, \cdot) - \inf_{\gamma \in \mathbb{R}^k} R(\beta, 0, \gamma) \right| \rightarrow 0 \text{ a.s.-}P,$$

and

$$\sup_{\beta \in \mathbf{B}} \left| \inf_{\alpha \in \mathbb{R}^{k_1}} \hat{R}_n(\beta, \alpha, 0, \cdot) - \inf_{\alpha \in \mathbb{R}^{k_1}} R(\beta, \alpha, 0) \right| \rightarrow 0 \text{ a.s.-}P.$$

Proof of LEMMA A.2: We only prove the first convergence result, as the second one can be shown in an analogous manner. Because Lemma A.1 has shown the corresponding pointwise a.s. convergence, and \mathbf{B} is compact, it suffices to show that the series in question is strongly stochastically equicontinuous (Andrews 1992, Theorem 2).

Let β_1 and β_2 be arbitrary points in \mathbf{B} . Also, let g_{nj} be Koenker and Bassett's (1978) estimator in τ -quantile regression of $y_t - Y_t' \beta_j$ on Z_t , i.e.,

$$\hat{R}_n(\beta_j, 0, g_{nj}, \cdot) = \inf_{\gamma \in \mathbb{R}^k} \hat{R}_n(\beta_j, 0, \gamma, \cdot),$$

for $j = 1, 2$. Then we have that for each $n \in \mathbb{N}$,

$$\begin{aligned} & \hat{R}_n(\beta_1, 0, g_{n1}, \cdot) - \hat{R}_n(\beta_2, 0, g_{n2}, \cdot) \\ &= (\hat{R}_n(\beta_1, 0, g_{n1}, \cdot) - \hat{R}_n(\beta_1, 0, g_{n2}, \cdot)) + (\hat{R}_n(\beta_1, 0, g_{n2}, \cdot) - \hat{R}_n(\beta_2, 0, g_{n2}, \cdot)) \\ &\leq \hat{R}_n(\beta_1, 0, g_{n2}, \cdot) - \hat{R}_n(\beta_2, 0, g_{n2}, \cdot), \end{aligned}$$

where the inequality holds, because $\hat{R}_n(\beta_1, 0, g_{n1}, \cdot) \leq \hat{R}_n(\beta_1, 0, g_{n2}, \cdot)$ for each $n \in \mathbb{N}$. We further have that

$$\begin{aligned} \hat{R}_n(\beta_1, 0, g_{n2}, \cdot) - \hat{R}_n(\beta_2, 0, g_{n2}, \cdot) &= n^{-1} \sum_{t=1}^n (\rho_\tau(y_t - Y_t' \beta_1 - Z_t' g_{n2}) - \rho_\tau(y_t - Y_t' \beta_2 - Z_t' g_{n2})) \\ &\leq n^{-1} \sum_{t=1}^n |Y_t' \beta_1 - Y_t' \beta_2| \leq |\beta_1 - \beta_2| n^{-1} \sum_{t=1}^n |Y_t|. \end{aligned}$$

It follows that for each $n \in \mathbb{N}$

$$\hat{R}_n(\beta_1, 0, g_{n1}, \cdot) - \hat{R}_n(\beta_2, 0, g_{n2}, \cdot) \leq |\beta_1 - \beta_2| n^{-1} \sum_{t=1}^n |Y_t|.$$

Analogously, we can also show that for each $n \in \mathbb{N}$

$$\hat{R}_n(\beta_2, 0, g_{n2}, \cdot) - \hat{R}_n(\beta_1, 0, g_{n1}, \cdot) \leq |\beta_1 - \beta_2| n^{-1} \sum_{t=1}^n |Y_t|.$$

Thus, it holds that for each $n \in \mathbb{N}$

$$\left| \inf_{\gamma \in \mathbb{R}^k} \hat{R}_n(\beta_2, 0, \gamma, \cdot) - \inf_{\gamma \in \mathbb{R}^k} \hat{R}_n(\beta_1, 0, \gamma, \cdot) \right| = |\hat{R}_n(\beta_2, 0, g_{n2}, \cdot) - \hat{R}_n(\beta_1, 0, g_{n1}, \cdot)| \leq |\beta_1 - \beta_2| n^{-1} \sum_{t=1}^n |Y_t|.$$

Because $\{n^{-1} \sum_{t=1}^n |Y_t|\}_{n \in \mathbb{N}}$ converges to $E[|Y_1|]$ a.s.- P by the Kolmogorov strong law of large numbers, the desired result follows by Andrews (1992, Lemma 2). \square

LEMMA A.3: *Suppose that Assumptions 1 and 2 hold. For each $\beta \in \mathbf{B}$, $\{\inf_{\alpha \in \mathbb{R}^k} \hat{Q}_n(\beta, \alpha, \cdot)\}$ converges to $\inf_{\alpha \in \mathbb{R}^k} Q(\beta, \alpha)$ uniformly in $\beta \in \mathbf{B}$ a.s.- P*

Proof of LEMMA A.3: Because the linear independence of the elements of $X_1 = (y_1, Y_1', Z_1')$ in Assumption 1 implies that for each $\beta \in \mathbf{B}$, the distance between $y_1 - Y_1'\beta$ and the $\|\cdot\|_{\rho_\tau}$ -metric projection of $y_1 - Y_1'\beta$ on the subspace spanned by Z_1 is positive, i.e., $\inf_{\alpha \in \mathbb{R}^{k_1}} R(\beta, \alpha, 0) > 0$. Because $\beta \mapsto \inf_{\alpha \in \mathbb{R}^{k_1}} R(\beta, \alpha, 0) : \mathbf{B} \rightarrow \mathbb{R}$ is continuous, it is bounded away from zero on \mathbf{B} . The desired results from this fact and Lemma A.2, because $(r_1, r_2) \mapsto r_1/r_2 : \mathbb{R} \times (a, \infty) \rightarrow \mathbb{R}$ is a Lipschitz function if $a > 0$. \square

LEMMA A.4: *Suppose that Assumptions 1 and 2 hold. Let $\{b_n\}_{n \in \mathbb{N}}$ be a sequence of \mathbf{B} -valued random vectors on (Ω, \mathcal{F}, P) converging to β_0 a.s.- P (in probability- P). Let $\{a_n\}_{n \in \mathbb{N}}$ be sequences of $k_1 \times 1$ vectors on (Ω, \mathcal{F}, P) satisfying that for each $n \in \mathbb{N}$, $\hat{R}_n(b_n, a_n, 0, \cdot) = \inf_{\alpha \in \mathbb{R}^{k_1}} \hat{R}_n(b_n, \alpha, 0, \cdot)$. Then:*

(a) *Then $a_n \rightarrow \alpha_0$ a.s.- P (in probability- P).*

(b) *Let $\{c_n\}_{n \in \mathbb{N}}$ be a sequence of $k \times 1$ random vectors on (Ω, \mathcal{F}, P) satisfying that for each $n \in \mathbb{N}$ $\hat{R}_n(b_n, a_n, c_n, \cdot) = \inf_{\gamma \in \mathbb{R}^k} \hat{R}_n(b_n, a_n, \gamma, \cdot)$. Then $c_n \rightarrow 0$ a.s.- P (in probability- P), provided that the minimizer of $R(\beta_0, \alpha_0, \cdot) : \mathbb{R}^k \rightarrow \mathbb{R}$ over \mathbb{R}^k is unique.*

Proof of LEMMA A.4: We only prove the result for $\{a_n\}_{n \in \mathbb{N}}$. The result for $\{c_n\}_{n \in \mathbb{N}}$ can be established in an analogous way.

Suppose that $b_n \rightarrow \beta_0$ a.s.- P . Then for each $\alpha \in \mathbb{R}^{k_1}$, $\{\hat{R}_n(b_n, \alpha, 0, \cdot)\}_{n \in \mathbb{N}}$ converges to $R(\beta_0, \alpha, 0)$ a.s.- P , because for each $\alpha \in \mathbb{R}^{k_1}$, $\{\hat{R}_n(\beta, \alpha, 0, \cdot)\}_{n \in \mathbb{N}}$ converges to $R(\beta, \alpha, 0)$ uniformly in $\beta \in \mathbf{B}$ by Jennrich's uniform law of large numbers (Jennrich 1969, Theorem 2). Further, we can apply Rockafellar (1970, Theorem 10.8) to show that the convergence is uniform in α over any compact subset of \mathbb{R}^{k_1} , because for each $n \in \mathbb{N}$, $\hat{R}_n(b_n, \alpha, 0, \cdot)$ is convex in α over \mathbb{R}^{k_1} .

Take an arbitrary compact subset \mathbf{A}_1 of \mathbb{R}^{k_1} that contain α_0 in its interior. Then $\{\hat{R}_n(b_n, \alpha_0, 0, \cdot)\}$ converges to $R(\beta_0, \alpha_0, 0)$ a.s.- P ; $\{\hat{R}_n(b_n, \alpha, 0, \cdot)\}$ converges to $R(\beta_0, \alpha, 0)$ uniformly on $\alpha \in \partial \mathbf{A}_1$ a.s.- P ; and $R(\beta_0, \alpha_0, 0) < \inf_{\alpha \in \partial \mathbf{A}_1} R(\beta_0, \alpha, 0)$, because α_0 is the unique minimizer of $R(\beta_0, \cdot, 0)$ on \mathbb{R}^{k_1} by Assumption 2.

Because $\hat{R}_n(b_n, \alpha, 0, \cdot)$ is convex in α , it follows that

$$\hat{R}_n(b_n, \alpha_0, 0, \cdot) < \inf_{\alpha \in \mathbb{R}^{k_1} \setminus \mathbf{A}_1} \hat{R}_n(b_n, \alpha, 0, \cdot)$$

for almost all $n \in \mathbb{N}$ a.s.- P . That is, $a_n \in A_1$ for almost all $n \in \mathbb{N}$ a.s.- P . Because A_1 is an arbitrary compact subset containing α_0 in its interior, this establishes the a.s.- P convergence of $\{a_n\}_{n \in \mathbb{N}}$ to α_0 . The convergence of $\{a_n\}_{n \in \mathbb{N}}$ in probability in the current lemma immediately follows from the result of the a.s. convergence of $\{a_n\}_{n \in \mathbb{N}}$ by using the subsequence theorem. \square

Proof of THEOREM 4.1: By Assumption 2, $\beta \mapsto \sup_{\alpha \in \mathbb{R}^k} Q(\beta, \alpha) : \mathbf{B} \rightarrow \mathbb{R}$ is uniquely maximized at β_0 . Because $\hat{\beta}_n$ maximizes $\sup_{\alpha \in \mathbb{R}^k} \hat{Q}_n(\beta, \alpha, \cdot)$ with respect to β over the compact subset \mathbf{B} , and $\{\sup_{\alpha \in \mathbb{R}^k} \hat{Q}_n(\beta, \alpha, \cdot)\}_{n \in \mathbb{N}}$ converges to $\sup_{\alpha \in \mathbb{R}^k} Q(\beta, \alpha)$ uniformly in $\beta \in \mathbf{B}$ a.s.- P , it follows by Pötscher and Prucha (1991, Lemma 4.2) that $\{\hat{\beta}_n\}_{n \in \mathbb{N}}$ converges to β_0 a.s.- P . Further, applying Lemma A.4(a) by setting $b_n = \hat{\beta}_n$ and $a_n = \hat{\alpha}_n$ establishes that the strong consistency of $\hat{\alpha}_n$ for α_0 . The result therefore follows. \square

In proving Lemmas 4.2, 4.3 and Theorem 4.4, we use the following lemma.

LEMMA A.5: *Suppose that Assumptions 1–3 hold, and let $\{\tilde{d}_{nj} \equiv (b'_{nj}, a_{nj}, g'_{nj})' : \Omega \rightarrow \mathbf{B} \times \mathbb{R}^{k_1} \times \mathbb{R}^k\}_{n \in \mathbb{N}}$ be a sequence of random vectors that converges in probability- P to $d_0 \equiv (\beta'_0, \alpha'_0, 0_{1 \times k})'$, $j = 1, 2$. Then*

$$\begin{aligned} & \hat{R}_n(b_{n2}, a_{n2}, g_{n2}, \cdot) - \hat{R}_n(b_{n1}, a_{n1}, g_{n1}, \cdot) \\ &= -n^{-1} \sum_{t=1}^n (\tau - 1(U_t < 0)) \tilde{X}'_t(\tilde{d}_{n2} - \tilde{d}_{n1}) + \frac{1}{2}(\tilde{d}_{n2} - d_0)' J(\tilde{d}_{n2} - d_0) - \frac{1}{2}(\tilde{d}_{n1} - d_0)' J(\tilde{d}_{n1} - d_0) \\ & \quad + o_P(n^{-1/2}(|d_{n2} - d_{n1}| + |d_{n1} - d_0|^2 + |d_{n2} - d_0|^2)), \end{aligned} \tag{10}$$

where $\tilde{X}_t \equiv (Y_t, Z'_{t,1}, Z'_t)'$, $t \in \mathbb{N}$.

Proof of LEMMA A.5: Define $r : \mathbb{R} \times \mathbb{R}^{g+k_1+k} \times \mathbb{R}^{g+k_1+k} \times \mathbb{R}^{g+k_1+k} \rightarrow \mathbb{R}$ by

$$\begin{aligned} r(y, \tilde{x}, d_1, d_2) &\equiv \frac{1}{|d_2 - d_1|} \left(\rho_\tau(y - \tilde{x}'d_2) - \rho_\tau(y - \tilde{x}'d_1) + (\tau - 1(y - \tilde{x}'d_0 < 0)) \tilde{x}'(d_2 - d_1) \right), \\ & (y, \tilde{x}, d_1, d_2) \in \mathbb{R} \times \mathbb{R}^{g+k_1+k} \times \mathbb{R}^{g+k_1+k} \times \mathbb{R}^{g+k_1+k}, \end{aligned}$$

with the rule that division by zero is zero. Also, following Pollard (1985), let ν_n denote the standardized sample average operator such that for each function $f : \mathbb{R} \times \mathbb{R}^{l+k} \rightarrow \mathbb{R}$ with $\mathbb{E}[|f(Y_1, \tilde{X}_1)|] < \infty$

$$\nu_n f(\cdot, \cdot) = n^{-1/2} \sum_{t=1}^n (f(Y_t, \tilde{X}_t) - \mathbb{E}[f(Y_1, \tilde{X}_1)]), \quad n \in \mathbb{N}.$$

By the definition of r , we obtain that

$$\begin{aligned} \hat{R}_n(d_2, \cdot) - \hat{R}_n(d_1, \cdot) &= R(d_2) - R(d_1) - \left(\ell + n^{-1} \sum_{t=1}^n (\tau - 1(U_t < 0)) \tilde{X}_t \right)' (d_2 - d_1) \\ &\quad + n^{-1/2} |d_2 - d_1| \nu_n r(\cdot, \cdot, d_1, d_2) \end{aligned}$$

for each $(d_1, d_2) \in \mathbb{R}^{l+k} \times \mathbb{R}^{l+k}$, where ℓ is the gradient of R at $(\beta_0, \alpha_0, 0_{1 \times k})$, which is equal to $-\mathbb{E}[(\tau - 1(U_1 < 0)) \tilde{X}_1]$. Taking the second-order Taylor expansion of $R(d_1)$ and $R(d_2)$ about d_0 on the right-hand side of this equality and replacing d_1 with \tilde{d}_{n1} and d_2 with \tilde{d}_{n2} in the resulting equality yields the desired result, if $\{\nu_n r(\cdot, \cdot, \tilde{d}_{n1}, \tilde{d}_{n2})\}_{n \in \mathbb{N}}$ converges to zero in probability- P . It thus suffices to show the convergence of $\{\nu_n r(\cdot, \cdot, \tilde{d}_{n1}, \tilde{d}_{n2})\}$ to zero in probability- P .

It is straightforward to verify that $r(y_1, \tilde{X}_1, \theta_1, \theta_2) \leq 2|\tilde{X}_1|$, from which it follows that

$$\mathbb{E} \left[\sup_{(d_1, d_2) \in \mathbb{R}^{g+k_1+k} \times \mathbb{R}^{g+k_1+k}} r(y_1, \tilde{X}_1, d_1, d_2)^2 \right] \leq 4\mathbb{E}[|\tilde{X}_1|^2] < \infty.$$

Also, $\{r(\cdot, \cdot, d_1, d_2) : (d_1, d_2) \in \mathbb{R}^{g+k_1+k} \times \mathbb{R}^{g+k_1+k}\}$ can be expressed as a sum of a fixed member of functions from a polynomial class. These facts imply that $\{\nu_n r(\cdot, \cdot, d_1, d_2)\}_{n \in \mathbb{N}}$ is stochastically equicontinuous at (d_0, d_0) (Pollard 1985, pp. 311–312). Further, $r(y_1, \tilde{X}_1, d_1, d_2)^2$ converges to zero as $(d_1, d_2) \rightarrow (d_0, d_0)$ a.s.- P , and $r(y_1, \tilde{X}_1, d_1, d_2)^2$ is dominated by $4|\tilde{X}_1|^2$ with a finite moment. It follows by the dominated convergence theorem that $\mathbb{E}[r(y_1, \tilde{X}_1, d_1, d_2)^2] \rightarrow 0$ as $(d_1, d_2) \rightarrow (d_0, d_0)$.

Now let $\{\mathbf{U}_n \subset \mathbb{R}^{g+k_1+k} \times \mathbb{R}^{g+k_1+k}\}_{n \in \mathbb{N}}$ be an arbitrary sequence of balls centered at $(d'_0, d'_0)'$ that shrinks down to $(d'_0, d'_0)'$. Then, as Pollard (1985, page. 309) explains, it follows from the above-mentioned facts that $\sup_{(d_1, d_2) \in \mathbf{U}_n} |\nu_n r(\cdot, \cdot, d_1, d_2)| \rightarrow 0$ in probability- P . Thus, $\{\nu_n r(\cdot, \cdot, \tilde{d}_{n1}, \tilde{d}_{n2})\}$ converges to zero in probability- P , given that $\{\tilde{d}_{nj}\}_{n \in \mathbb{N}}$ converges to d_0 in probability- P , $j = 1, 2$. \square

LEMMA A.6: *Let (Ω, \mathcal{F}, P) be a probability space. Suppose that a sequence of random vectors $\{\eta_n : \Omega \rightarrow \mathbb{R}^m\}_{n \in \mathbb{N}}$ and a sequence of random variables $\{\xi_n : \Omega \rightarrow \mathbb{R}\}_{n \in \mathbb{N}}$ satisfy that $-\eta_n' A \eta_n + \xi_n \geq 0$ for each $n \in \mathbb{N}$, where A is a positive definite $m \times m$ symmetric matrix. Also, let $\{\zeta_n : \Omega \rightarrow \mathbb{R}\}_{n \in \mathbb{N}}$ be a sequence of random variables. Suppose that $\xi_n = o_P(|\eta_n| + |\eta_n|^2 + |\zeta_n|)$ as $n \rightarrow \infty$. Then $|\eta_n| = o_P(|\zeta_n|^{1/2} + 1)$ as $n \rightarrow \infty$.*

We now prove Lemma 4.2.

Proof of LEMMA 4.2: The existence of $\{c_n\}$ follows immediately from the fact that the minimization of

$\hat{R}_n(b_n, a_n, \gamma, \cdot)$ in terms of γ is the $\|\cdot\|_{\rho_\tau}$ -metric projection of $(y_1 - Y_1' b_n - Z_{1,1}' a_n, y_2 - Y_2' b_n - Z_{1,2}' a_n, \dots, y_n - Y_n' b_n - Z_{1,n}' a_n)$ on the space spanned by the rows of (Z_1, Z_2, \dots, Z_n) .

To prove the second result, we first show that $\{c_n\}$ converges to 0 in probability- P , and then apply Lemmas A.5 and A.6. For each fixed $\gamma \in \mathbb{R}^k$, $\hat{R}_n(\beta, \alpha, \gamma, \cdot)$ is convex in β and α . By the Kolmogorov strong law of large numbers and Hjort and Pollard (1993, Lemma 1), $\{\hat{R}_n(\beta, \alpha, \gamma, \cdot)\}_{n \in \mathbb{N}}$ converges to $R(\beta, \alpha, \gamma)$ uniformly in (β', α') in each neighborhood of (β_0', α_0') in probability- P . Because $\{(b_n', a_n')'\}_{n \in \mathbb{N}}$ converges to (β_0', α_0') in probability- P by the assumption, it follows that $\{\hat{R}_n(b_n, a_n, \gamma, \cdot)\}_{n \in \mathbb{N}}$ converges to $R(\beta_0, \alpha_0, \gamma)$ for each $\gamma \in \mathbb{R}^k$. Under Assumptions 1-3(c), this fact implies by Hjort and Pollard (1993, Lemma 2) that $\{c_n\}$ converges to 0 in probability- P .

We now set b_n to both b_{n1} and b_{n2} , c_n to g_{n1} ,

$$g_n^\dagger \equiv C \begin{pmatrix} b_n - \beta_0 \\ a_n - \alpha_0 \end{pmatrix} + J_{\gamma\gamma}^{-1} n^{-1} \sum_{t=1}^n (\tau - 1(U_t < 0)) Z_t.$$

to g_{n2} in (10) and multiply the resulting equality by n to obtain that

$$\begin{aligned} 0 &\leq n(\hat{R}_n(b_n, a_n, g_n^\dagger, \cdot) - \hat{R}_n(b_n, a_n, c_n, \cdot)) \\ &= -\frac{1}{2} n^{1/2} (c_n - g_n^\dagger)' J_{\gamma\gamma} n^{1/2} (c_n - g_n^\dagger) \\ &\quad + o_P(n^{1/2} |c_n - g_n^\dagger| + n|b_n - \beta_0|^2 + n|a_n - \alpha_0|^2 + n|c_n|^2 + n|g_n^\dagger|^2) \\ &= -\frac{1}{2} n^{1/2} (c_n - g_n^\dagger)' J_{\gamma\gamma} n^{1/2} (c_n - g_n^\dagger) \\ &\quad + o_P(n^{1/2} |c_n - g_n^\dagger|) + n|c_n - g_n^\dagger|^2 + n|b_n - \beta_0|^2 + n|a_n - \alpha_0|^2 + 1, \end{aligned}$$

where the second equality holds because $|g_n^\dagger| = O_P(|b_n - \beta_0| + |a_n - \alpha_0| + 1)$ and $c_n = O_P(|c_n - g_n^\dagger| + |g_n^\dagger|)$.

The result follows from this inequality by Lemma A.6. \square

Proof of LEMMA 4.3: Let $\{c_n\}_{n \in \mathbb{N}}$ be as in Lemma 4.2. Note that the difference between $\{\hat{R}_n(b_n, a_n, c_n, \cdot)\}_{n \in \mathbb{N}}$ and $\{\hat{R}_n(b_n, a_n, 0, \cdot)\}_{n \in \mathbb{N}}$ converges to zero in probability- P . Applying the delta method with this fact, we

obtain that

$$\begin{aligned}
n \log \hat{Q}_n(b_n, a_n, \cdot) &= n(\log \hat{R}_n(b_n, a_n, c_n, \cdot) - \log \hat{R}_n(b_n, a_n, 0, \cdot)) \\
&= \frac{1}{R(\beta_0, \alpha_0, 0)} n(\hat{R}_n(b_n, a_n, c_n, \cdot) - \hat{R}_n(b_n, a_n, 0, \cdot)) \\
&\quad - \frac{1}{2R(\beta_0, \alpha_0, 0)^2} n(\hat{R}_n(b_n, a_n, c_n, \cdot) - R(\beta_0, \alpha_0, 0))^2 \\
&\quad + \frac{1}{2R(\beta_0, \alpha_0, 0)^2} n(\hat{R}_n(b_n, a_n, 0, \cdot) - R(\beta_0, \alpha_0, 0))^2 \\
&\quad + o_P(n(\hat{R}_n(b_n, a_n, c_n, \cdot) - R(\beta_0, \alpha_0, 0))^2 + n(\hat{R}_n(b_n, a_n, 0, \cdot) - R(\beta_0, \alpha_0, 0))^2).
\end{aligned} \tag{11}$$

We apply Lemma A.5 to each of the non-remainder terms on the right-hand side of this equality:

$$\begin{aligned}
&n(\hat{R}_n(b_n, a_n, c_n, \cdot) - \hat{R}_n(b_n, a_n, 0, \cdot)) \\
&= -\frac{1}{2}n^{1/2}c'_n J_{\gamma\gamma} n^{1/2}c_n + o_P(n^{1/2}|c_n| + n|b_n - \beta_0|^2 + n|a_n - \alpha_0|^2 + |c_n|^2) \\
&= -\frac{1}{2}n^{1/2}c'_n J_{\gamma\gamma} n^{1/2}c_n + o_P(n^{1/2}|b_n - \beta_0| + n^{1/2}|a_n - \alpha_0| + n|b_n - \beta_0|^2 + n|a_n - \alpha_0|^2 + 1),
\end{aligned}$$

$$\begin{aligned}
&n^{1/2}(\hat{R}_n(b_n, a_n, c_n, \cdot) - R(\beta_0, \alpha_0, 0)) \\
&= n^{1/2}(\hat{R}_n(b_n, a_n, c_n, \cdot) - \hat{R}_n(\beta_0, \alpha_0, 0, \cdot)) + n^{1/2}(\hat{R}_n(\beta_0, \alpha_0, 0, \cdot) - R(\beta_0, \alpha_0, 0)) \\
&= -(n^{-1} \sum_{t=1}^n (\tau - 1(U_t < 0)) (Y'_t, Z'_{t,1})')' n^{1/2}(\tilde{\theta}_n - \theta_0) + n^{1/2}(\hat{R}_n(\beta_0, \alpha_0, 0, \cdot) - R(\beta_0, \alpha_0, 0)) \\
&\quad + o_P(n^{1/2}|b_n - \beta_0| + n^{1/2}|a_n - \alpha_0| + n|b_n - \beta_0|^2 + n|a_n - \alpha_0|^2 + 1),
\end{aligned}$$

and

$$\begin{aligned}
&n^{1/2}(\hat{R}_n(b_n, a_n, 0, \cdot) - R(\beta_0, \alpha_0, 0)) \\
&= n^{1/2}(\hat{R}_n(b_n, a_n, 0, \cdot) - \hat{R}_n(\beta_0, \alpha_0, 0, \cdot)) + n^{1/2}(\hat{R}_n(\beta_0, \alpha_0, 0, \cdot) - R(\beta_0, \alpha_0, 0)) \\
&= -\left(n^{-1} \sum_{t=1}^n (\tau - 1(U_t < 0)) (Y'_t, Z'_{t,1})'\right)' n^{1/2}(b_n - \beta_0) + n^{1/2}(\hat{R}_n(\beta_0, \alpha_0, 0, \cdot) - R(\beta_0, \alpha_0, 0)) \\
&\quad + o_P(n^{1/2}|b_n - \beta_0| + n^{1/2}|a_n - \alpha_0| + n|b_n - \beta_0|^2 + n|a_n - \alpha_0|^2 + 1).
\end{aligned}$$

Substituting these into (11) and applying Lemma 4.2 yields the desired result. \square

Proof of THEOREM 4.4: Let

$$\tilde{\theta}_n^\dagger \equiv \theta_0 - \frac{1}{R(\beta_0, \alpha_0, 0)} K^{-1} C' n^{-1} \sum_{t=1}^n (\tau - 1(U_t < 0)) Z_t, \quad n \in \mathbb{N},$$

and let b_n^\dagger and a_n^\dagger denote the vectors containing the first g elements and the remaining elements of $\tilde{\theta}_n$, respectively. Then, by Lemma 4.3, we have that

$$\begin{aligned} 0 &\leq n \log \hat{Q}_n(\hat{\beta}_n, \hat{\alpha}_n, \cdot) - n \log \hat{Q}_n(b_n^\dagger, a_n^\dagger, \cdot) \\ &= \frac{1}{2} n^{1/2} (\hat{\theta}_n - \tilde{\theta}_n^\dagger)' K n^{1/2} (\hat{\theta}_n - \tilde{\theta}_n^\dagger) \\ &\quad + o_P(n^{1/2} |\hat{\beta}_n - \beta_0| + n^{1/2} |\hat{\alpha}_n - \alpha_0| + n |\hat{\beta}_n - \beta_0|^2 + n |\hat{\alpha}_n - \alpha_0|^2 + 1). \end{aligned}$$

The first result follows from this equality by Lemma A.6. For the second result, apply the central limit theorem (CLT) for i.i.d. random vectors (Rao 1973, p. 128) to show that $\{n^{-1/2} \sum_{t=1}^n (\tau - 1(U_t < 0)) Z_t\}_{n \in \mathbb{N}}$ is asymptotically distributed with $N(0, R(\beta_0, \alpha_0, 0)^2 V)$, and then apply the continuous mapping theorem. \square

Proof of THEOREM 4.5: To prove (a), let $\{\tilde{\theta}_n^\dagger\}_{n \in \mathbb{N}}$ be as in the proof of Theorem 4.4 and $\{\delta_n\}_{n \in \mathbb{N}}$ an arbitrary sequence of $(g + k_1) \times 1$ random vectors that converges to the origin in probability- P . Recall that the expression consisting of the second and third terms on the right-hand side of (5) is minimized when $(b'_n, a'_n)' = \tilde{\theta}_n^\dagger$, and that $\{n^{1/2}(\hat{\theta}_n - \tilde{\theta}_n^\dagger)\}_{n \in \mathbb{N}}$ converges to zero in probability- P by Theorem 4.4. Using these facts with Lemma 4.3, we can show that $n \log \hat{Q}_n(\hat{\theta}_n, \cdot) - n \log \hat{Q}_n(\tilde{\theta}_n^\dagger, \cdot) = o_P(1)$ and

$$\begin{aligned} n \log \hat{Q}_n(\hat{\theta}_n + \delta_n, \cdot) - n \log \hat{Q}_n(\tilde{\theta}_n^\dagger, \cdot) &= -\frac{1}{2} n^{1/2} (\hat{\theta}_n - \tilde{\theta}_n^\dagger + \delta_n)' K n^{1/2} (\hat{\theta}_n - \tilde{\theta}_n^\dagger + \delta_n) + o_P(n^{1/2} |\delta_n| + |\delta_n|^2 + 1) \\ &= -\frac{1}{2} n^{1/2} \delta_n' K n^{1/2} \delta_n + o_P(n^{1/2} |\delta_n| + |\delta_n|^2 + 1) \end{aligned} \tag{12}$$

By taking each of $\tau_n e_i + \tau_n e_j$, $-\tau_n e_i + \tau_n e_j$, $\tau_n e_i - \tau_n e_j$, and $-\tau_n e_i - \tau_n e_j$ for δ_n in this equality and using the resulting equalities in the definition of \hat{K}_{nij} ($i, j = 1, 2, \dots, l$), we obtain that

$$4n\tau_n^2 \hat{K}_{nij} = 4\tau_n^2 K_{ij} + o_P(n^{-1/2} \tau_n + \tau_n^2 + 1).$$

Dividing both sides of this equality by $4n\tau_n^2$ and applying Assumption 4 yields the desired result.

To prove (b), let δ be an arbitrary $(g + k_1) \times 1$ vector. By Lemma 4.2, we have that for each $i = 1, 2, \dots, k$ and each $j = 1, 2, \dots, l$, $\hat{\gamma}_{ni}(\hat{\theta}_n + \tau_n \delta) - \hat{\gamma}_{ni}(\hat{\theta}_n - \tau_n \delta) = 2\tau_n C \delta_n + o_P(\tau_n)$. It follows that

$$\frac{1}{2\tau_n} (\hat{\gamma}_n(\hat{\beta}_n + \tau_n \delta) - \hat{\gamma}_n(\hat{\beta}_n - \tau_n \delta)) = C \delta + o_P(1).$$

Taking e_j for δ for each $j = 1, 2, \dots, k$ in this equality completes the proof.

To show (c), notice that $\{\hat{R}_n(\beta, \alpha, 0, \cdot)\}_{n \in \mathbb{N}}$ converges to $R(\beta, \alpha, 0)$ uniformly in $(\beta', \alpha')'$ in each closed ball centered at $(\beta'_0, \alpha'_0)'$ in probability- P , because $\{\hat{R}_n(\beta, \alpha, 0, \cdot)\}$ converges to $R(\beta, \alpha, 0)$ in probability- P for each $(\beta, \alpha) \in \mathbf{B} \times \mathbb{R}^{k_1}$, and $\hat{R}_n(\beta, \alpha, 0, \cdot)$ is convex in β and α for each $n \in \mathbb{N}$ (Hjort and Pollard 1993, Lemma 1). It follows by Theorem 4.4 that $\{\hat{R}_n(\hat{\beta}_n, \hat{\alpha}_n, 0, \cdot)\}_{n \in \mathbb{N}}$ converges to $R(\beta_0, \alpha_0, 0)$. Also, $n^{-1} \sum_{t=1}^n Z_t Z_t'$ converges to $E[Z_1 Z_1']$ a.s.- P by the Kolmogorov strong law of large numbers and Hjort and Pollard (1993, Lemma 1). The desired result follows from these by the Slutsky theorem. Finally, the claim (d) immediately follows from a-c by the Slutsky theorem. \square

Proof of LEMMA 5.1: Using Lemma 4.3 and Theorem 4.4, the ρ_τ -IV regression of $V_t \equiv y_t - Y_t' \bar{\beta}$ on $Z_{t,1}$, taking Z_t for instruments, we obtain that

$$\begin{aligned} \mathcal{J}_n &= \frac{1}{R(\beta_0, \alpha_0, 0)} \left(n^{-1/2} \sum_{t=1}^n (\tau - 1(U_t < 0)) Z_t \right)' (L^{-1} - C^0 K^{0-1} C^{0'}) \\ &\quad \times \frac{1}{R(\beta_0, \alpha_0, 0)} \left(n^{-1/2} \sum_{t=1}^n (\tau - 1(U_t < 0)) Z_t \right) + o_P(1), \end{aligned}$$

where $K^0 \equiv R(\beta_0, \alpha_0, 0)^{-1} J_{\alpha\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\alpha}$. Because $R(\beta_0, \alpha_0, 0)^{-1} n^{-1/2} \sum_{t=1}^n (\tau - 1(U_t < 0)) Z_t \stackrel{A}{\sim} \eta$ by the central limit theorem (CLT) for i.i.d. random vectors (Rao 1973, p. 128), it follows by the continuous mapping theorem and the asymptotic equivalence lemma that

$$\mathcal{J}_n \stackrel{A}{\sim} \eta' (L^{-1} - C^0 K^{0-1} C^{0'}) \eta.$$

Applying the fact that $K^0 = C^{0'} L C^0$ to this result completes the proof. \square

Proof of LEMMA 5.2: Let $\{\delta_n\}_{n \in \mathbb{N}}$ be an arbitrary sequence of $k \times 1$ random vectors that converges to the origin in probability- P . By Lemma A.5 with Theorem 4.4 and Lemma 4.2, we have that

$$\hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0 + \delta_n, \cdot) - \hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0, \cdot) = \frac{1}{2} \delta_n' J_{\gamma\gamma} \delta_n + o_P \left(n^{-1/2} |\delta_n| + |\delta_n|^2 + n^{-1} \right).$$

By the delta method, it follows that

$$\begin{aligned} &\log \hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0 + \delta_n, \cdot) - \log \hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0, \cdot) \\ &= \frac{1}{R(\beta_0, \alpha_0, 0)} \left(\hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0 + \delta_n, \cdot) - \hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0, \cdot) \right) = \frac{1}{2} \delta_n' L \delta_n + o_P(1). \end{aligned}$$

The desired result follows from this equality in essentially the same way as Theorem 4.5(a) is proved using (12). \square

Proof of THEOREM 5.3: To prove (a), it suffices to show that for each $p \in (0, 1)$, $\{c(p, \hat{C}_n, \hat{L}_n, \hat{V}_n)\}_{n \in \mathbb{N}}$ is weakly consistent for $c(p, C, L, V)$. For each $k \times l$ matrix \tilde{C} , each $k \times k$ symmetric matrix \tilde{L} , and each $k \times k$ symmetric matrix \tilde{V} , let $F(\cdot, \tilde{C}, \tilde{L}, \tilde{V}) : \mathbb{R} \rightarrow \mathbb{R}$ denote the distribution function of $\tilde{\eta}'(\tilde{L}^+ - \tilde{C}(\tilde{C}'\tilde{L}\tilde{C})^+ \tilde{C}')\tilde{\eta}$ with $\tilde{\eta}$ distributed with $N(0, \tilde{V})$. Also, define $\{F_n : \mathbb{R} \times \Omega \rightarrow \mathbb{R}\}_{n \in \mathbb{N}}$ and $F^* : \mathbb{R} \rightarrow \mathbb{R}$ by $\hat{F}_n(x, \cdot) \equiv F(x, \hat{C}_n, \hat{V}_n, \hat{L}_n)$, $x \in \mathbb{R}$, $n \in \mathbb{N}$, and $F^*(x) \equiv F(x, C, L, V)$, $x \in \mathbb{R}$. Because F^* is positively sloped at $c(p, C, L, V)$ for each $p \in (0, 1)$, the convergence of $\{c(p, \hat{C}_n, \hat{L}_n, \hat{V}_n)\}_{n \in \mathbb{N}}$ in probability- P to $c(p, C, L, V)$ for each $p \in (0, 1)$ is implied by the convergence of $\{\hat{F}_n(x, \cdot)\}_{n \in \mathbb{N}}$ in probability- P to $F^*(x)$ for each $x \in \mathbb{R}$. We below prove this sufficient condition.

Notice that the characteristic function $\phi(\cdot, \tilde{C}, \tilde{L}, \tilde{V})$ of $F(\cdot, \tilde{C}, \tilde{L}, \tilde{V})$ is given by

$$\phi(s, \tilde{C}, \tilde{L}, \tilde{V}) \equiv \det(I - 2is\tilde{V}^{1/2}(\tilde{L}^+ - \tilde{C}(\tilde{C}'\tilde{L}\tilde{C})^+ \tilde{C}')\tilde{V}^{1/2}),$$

where i denotes the imaginary unit. Because $\hat{C}_n \rightarrow C$, $\hat{L}_n \rightarrow L$, and $\hat{V}_n \rightarrow V$ in probability- P , it follows from the subsequent theorem (Lukacs 1975, Theorem 2.4.4) that for each subsequence $\{n_i\}_{i \in \mathbb{N}}$ of $\{n\}_{n \in \mathbb{N}}$, there exists a further subsequence $\{n_{i_j}\}_{j \in \mathbb{N}}$ of it, along which $\hat{C}_{n_{i_j}} \rightarrow C$, $\hat{L}_{n_{i_j}} \rightarrow L$, and $\hat{V}_{n_{i_j}} \rightarrow V$ as $j \rightarrow \infty$ a.s.- P . Because V is positive definite, we have that

$$\phi(s, \hat{C}_{n_{i_j}}, \hat{L}_{n_{i_j}}, \hat{V}_{n_{i_j}}) = \det \hat{V}_{n_{i_j}} \det(\hat{V}_{n_{i_j}}^+ - 2is(\hat{L}_{n_{i_j}}^+ - \hat{C}_{n_{i_j}}(\hat{C}'_{n_{i_j}}\hat{L}_{n_{i_j}}\hat{C}_{n_{i_j}})^+ \hat{C}'_{n_{i_j}})), \quad s \in \mathbb{R}$$

for each $s \in \mathbb{R}$ and a.a. $n \in \mathbb{N}$ a.s.- P . By the continuity theorem (Billingsley 1968, Theorem 7.6), this implies that for each $x \in \mathbb{R}$, $\{F_{n_{i_j}}(x, \cdot)\}_{i \in \mathbb{N}}$ converges to $F^*(x)$ a.s.- P . By the subsequence theorem, it follows from this fact that for each $x \in \mathbb{R}$, $\{F_n(x, \cdot)\}_{n \in \mathbb{N}}$ converges to $F^*(x)$ in probability- P . The desired result therefore follows.

For (b), it suffices to show that $c(p, \hat{C}_n, \hat{L}_n, \hat{V}_n) = O_P(1)$, because for each $\bar{c} \in \mathbb{R}$, $P[\mathcal{J}_n > \bar{c}] \rightarrow 1$ if the function $Q(\bar{\beta}, \cdot) : \mathbb{R}^{k_1} \rightarrow \mathbb{R}$ attains one nowhere on \mathbb{R}^{k_1} . We only give a sketch of the proof here. It is possible to derive results similar to those of Lemmas 4.2, 4.3, and Theorem 4.4 without imposing H_0 in the ρ_τ -IV regression of $y_t - Y_t'\bar{\beta}$ on $Z_{t,1}$, taking Z_t for instruments, by employing essentially the same techniques. Although the constants in the resulting formulas are different from those found in Lemmas 4.2, 4.3, and Theorem 4.4, essentially the same logic delivers the convergence of $\{\hat{C}_n^0\}_{n \in \mathbb{N}}$, $\{\hat{L}_n^0\}_{n \in \mathbb{N}}$, and $\{\hat{V}_n^0\}_{n \in \mathbb{N}}$ in probability- P . The stochastic limit of $\{\hat{L}_n^0\}_{n \in \mathbb{N}}$ is the Hessian of $R(\bar{\beta}, \alpha^*, \cdot) : \mathbb{R}^k \rightarrow \mathbb{R}$ at γ^* , which is nonsingular by Assumption 3(c). Using the same argument as in the proof of (a), we can show that $\{c(p, \hat{C}_n, \hat{L}_n, \hat{V}_n)\}_{n \in \mathbb{N}}$ converges in probability- P , so that it is $O_P(1)$. \square

Proof of THEOREM 6.1: First, we show that $\hat{\alpha}_n^0 \equiv \arg \min_{\alpha \in \mathbb{R}^{k_1}} \hat{R}_n(\bar{\beta}, \alpha, 0, \cdot)$ converges to $\bar{\alpha}$ a.s.- P , where $\bar{\alpha}$ is defined in Assumption 5. It suffices to show that $\sup_{\alpha \in \mathbb{R}^{k_1}} |\hat{R}_n(\bar{\beta}, \alpha, 0, \cdot) - \mathbb{E}\rho_\tau(W_1 - Z'_{1,1}(\alpha - \bar{\alpha}))| \rightarrow 0$ a.s.- P . We have $|\hat{R}_n(\bar{\beta}, \alpha, 0, \cdot) - \mathbb{E}\rho_\tau(W_1 - Z'_{1,1}(\alpha - \bar{\alpha}))| \leq |\hat{R}_n(\bar{\beta}, \alpha, 0, \cdot) - R(\bar{\beta}, \alpha, 0)| + |\mathbb{E}\rho_\tau(W_1 - Z'_1\Lambda'(\bar{\beta} - \beta_0)n^{-1/2} - Z'_{1,1}(\alpha - \bar{\alpha})) - \mathbb{E}\rho_\tau(W_1 - Z'_{1,1}(\alpha - \bar{\alpha}))|$, where the first summand converges to zero a.s.- P and uniformly in α , as we argued in the proof of Lemma A.4, while the second summand can be bounded using Knight's identity by $n^{-1/2}\mathbb{E}|Z_1||\Lambda'(\bar{\beta} - \beta_0)|$ (Koenker 2005, equation (4.3), page 121).

Note that \mathcal{J}_n in (6) can also be written as

$$\mathcal{J}_n = -2n \log \frac{\hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0, \cdot)}{\hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, 0, \cdot)}, \quad \text{where } \hat{\gamma}_n^0 \equiv \arg \inf_{\gamma \in \mathbb{R}^k} \hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \gamma, \cdot).$$

In a similar manner as before, one can show that $\hat{\gamma}_n^0 \rightarrow 0$ a.s.- P .

Next, the result of Lemma A.5 can be also stated as follows. Let $\{(a'_{nj}, c'_{nj})' : \Omega \rightarrow \mathbb{R}^{k_1} \times \mathbb{R}^k\}_{n \in \mathbb{N}}$, $j = 1, 2$, be sequences of random vectors that converge to zero a.s.- P . Then,

$$\begin{aligned} & n^{-1} \sum_{t=1}^n \rho_\tau(W_t - Z'_{t,1}a_{n2} - Z'_t c_{n2}) - n^{-1} \sum_{t=1}^n \rho_\tau(W_t - Z'_{t,1}a_{n1} - Z'_t c_{n1}) \\ &= -n^{-1} \sum_{t=1}^n (\tau - 1(W_t < 0)) (Z'_{t,1}(a_{n2} - a_{n1}) + Z'_t(c_{n2} - c_{n1})) \\ &+ \frac{1}{2} \begin{pmatrix} a_{n2} \\ c_{n2} \end{pmatrix}' \begin{pmatrix} J_{\alpha\alpha} & J_{\alpha\gamma} \\ J_{\gamma\alpha} & J_{\gamma\gamma} \end{pmatrix} \begin{pmatrix} a_{n2} \\ c_{n2} \end{pmatrix} - \frac{1}{2} \begin{pmatrix} a_{n1} \\ c_{n1} \end{pmatrix}' \begin{pmatrix} J_{\alpha\alpha} & J_{\alpha\gamma} \\ J_{\gamma\alpha} & J_{\gamma\gamma} \end{pmatrix} \begin{pmatrix} a_{n1} \\ c_{n1} \end{pmatrix} \\ &+ o_P(n^{-1/2}|a_{n2} - a_{n1}| + n^{-1/2}|c_{n2} - c_{n1}| + |a_{n1}|^2 + |c_{n1}|^2 + |a_{n2}|^2 + |c_{n2}|^2). \end{aligned} \tag{13}$$

Define

$$g_n^\dagger \equiv -n^{-1/2}\Lambda'(\bar{\beta} - \beta_0) - J_{\gamma\gamma}^{-1}J_{\gamma\alpha}(\hat{\alpha}_n^0 - \bar{\alpha}) + J_{\gamma\gamma}^{-1}n^{-1} \sum_{t=1}^n (\tau - 1(W_t < 0))Z_t.$$

By (13) and as in the proof of Lemma 4.2,

$$\begin{aligned} & 0 \leq n(\hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, g_n^\dagger, \cdot) - \hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0, \cdot)) \\ &= -\frac{1}{2}n^{1/2}(\hat{\gamma}_n^0 - g_n^\dagger)' J_{\gamma\gamma} n^{1/2}(\hat{\gamma}_n^0 - g_n^\dagger) + o_P(n^{1/2}|\hat{\gamma}_n^0 - g_n^\dagger| + n|\hat{\gamma}_n^0 - g_n^\dagger|^2 + n|\hat{\alpha}_n^0 - \bar{\alpha}|^2 + 1), \end{aligned}$$

and, therefore, by Lemma A.6,

$$\begin{aligned} n^{1/2}\hat{\gamma}_n^0 &= -\Lambda'(\bar{\beta} - \beta_0) + C^0 n^{1/2}(\hat{\alpha}_n^0 - \bar{\alpha}) + J_{\gamma\gamma}^{-1}n^{-1/2} \sum_{t=1}^n (\tau - 1(W_t < 0))Z_t \\ &+ o_P(n^{1/2}|\hat{\alpha}_n^0 - \bar{\alpha}| + 1). \end{aligned} \tag{14}$$

Now, as in the proof of Lemma 4.3,

$$\begin{aligned}
n \log \hat{Q}_n(\bar{\beta}, \hat{\alpha}_n^0, \cdot) &= n(\log \hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0, \cdot) - \log \hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, 0, \cdot)) \\
&= \frac{1}{\mathbf{E}\rho_\tau(W_1)} n(\hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0, \cdot) - \hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, 0, \cdot)) \\
&\quad - \frac{1}{2(\mathbf{E}\rho_\tau(W_1))^2} n(\hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0, \cdot) - \mathbf{E}\rho_\tau(W_1))^2 \\
&\quad + \frac{1}{2(\mathbf{E}\rho_\tau(W_1))^2} n(\hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, 0, \cdot) - \mathbf{E}\rho_\tau(W_1))^2 \\
&\quad + o_P(n(\hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0, \cdot) - \mathbf{E}\rho_\tau(W_1))^2 + n(\hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, 0, \cdot) - \mathbf{E}\rho_\tau(W_1))^2).
\end{aligned} \tag{15}$$

For the difference term in (15) we have by (13),

$$\begin{aligned}
&n(\hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, \hat{\gamma}_n^0, \cdot) - \hat{R}_n(\bar{\beta}, \hat{\alpha}_n^0, 0, \cdot)) \\
&= \sum_{t=1}^n \rho_\tau(W_t - Z'_{t,1}(\hat{\alpha}_n^0 - \bar{\alpha}) - Z'_t(\hat{\gamma}_n^0 + \Lambda'(\bar{\beta} - \beta_0))) \\
&\quad - \sum_{t=1}^n \rho_\tau(W_t - Z'_{t,1}(\hat{\alpha}_n^0 - \bar{\alpha}) - Z'_t\Lambda'(\bar{\beta} - \beta_0)) \\
&= -n^{-1/2} \sum_{t=1}^n (\tau - 1(W_t < 0)) Z'_t n^{1/2} \hat{\gamma}_n^0 \\
&\quad + \frac{1}{2} n^{1/2} \hat{\gamma}_n^{0'} J_{\gamma\gamma} n^{1/2} \hat{\gamma}_n^0 + n^{1/2} (\bar{\beta} - \beta_0)' \Lambda J_{\gamma\gamma} n^{1/2} \hat{\gamma}_n^0 + n^{1/2} (\hat{\alpha}_n^0 - \bar{\alpha})' J_{\alpha\gamma} n^{1/2} \hat{\gamma}_n^0 \\
&\quad + o_P(n^{1/2} |\hat{\gamma}_n^0| + n |\hat{\alpha}_n^0 - \bar{\alpha}|^2 + n |\hat{\gamma}_n^0|^2 + 1) \\
&= -n^{1/2} \hat{\gamma}_n^{0'} J_{\gamma\gamma} n^{1/2} \hat{\gamma}_n^0 + o_P(n^{1/2} |\hat{\alpha}_n^0 - \bar{\alpha}| + n |\hat{\alpha}_n^0 - \bar{\alpha}|^2 + 1).
\end{aligned} \tag{16}$$

Define

$$\Psi_n \equiv n^{-1/2} \sum_{t=1}^n (\tau - 1(W_t < 0)) Z_t - J_{\gamma\gamma} \Lambda'(\bar{\beta} - \beta_0).$$

Using (14)–(16), the result analogous to that of Lemma 4.3 can be stated as

$$-2n \log \hat{Q}_n(\bar{\beta}, \hat{\alpha}_n^0, \cdot) \mathbf{E}\rho_\tau(W_1) \tag{17}$$

$$\begin{aligned}
&= \Psi_n' J_{\gamma\gamma}^{-1} \Psi_n - 2\Psi_n' J_{\gamma\gamma}^{-1} J_{\gamma\alpha} n^{1/2} (\hat{\alpha}_n^0 - \bar{\alpha}) + n^{1/2} (\hat{\alpha}_n^0 - \bar{\alpha})' J_{\alpha\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\alpha} n^{1/2} (\hat{\alpha}_n^0 - \bar{\alpha}) \\
&\quad + o_P(n^{1/2} |\hat{\alpha}_n^0 - \bar{\alpha}| + n |\hat{\alpha}_n^0 - \bar{\alpha}|^2 + 1) \\
&= |J_{\gamma\gamma}^{-1/2} (\Psi_n - J_{\gamma\alpha} n^{1/2} (\hat{\alpha}_n^0 - \bar{\alpha}))|^2 \\
&\quad + o_P(n^{1/2} |\hat{\alpha}_n^0 - \bar{\alpha}| + n |\hat{\alpha}_n^0 - \bar{\alpha}|^2 + 1).
\end{aligned} \tag{18}$$

Similarly to the result of Theorem 4.4, we have now:

$$n^{1/2}(\hat{\alpha}_n^0 - \bar{\alpha}) = (J_{\alpha\gamma}J_{\gamma\gamma}^{-1}J_{\gamma\alpha})^{-1}J_{\alpha\gamma}J_{\gamma\gamma}^{-1}\Psi_n + o_P(1), \quad (19)$$

and after substituting this into (18), we obtain:

$$\begin{aligned} & -2n \log \hat{Q}_n(\bar{\beta}, \hat{\alpha}_n^0, \cdot) \mathbb{E} \rho_\tau(W_1) \\ &= |(J_{\gamma\gamma}^{-1/2} - J_{\gamma\gamma}^{-1/2} J_{\gamma\alpha} (J_{\alpha\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\alpha})^{-1} J_{\alpha\gamma} J_{\gamma\gamma}^{-1}) \Psi_n|^2 + o_P(1). \end{aligned}$$

Next, note that $\Psi_n \rightarrow_d N(J_{\gamma\gamma} \Lambda'(\bar{\beta} - \beta_0), \tau(1 - \tau) \mathbb{E}[Z_1 Z_1'])$, $|J_{\gamma\gamma}^{-1/2} - J_{\gamma\gamma}^{-1/2} J_{\gamma\alpha} (J_{\alpha\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\alpha})^{-1} J_{\alpha\gamma} J_{\gamma\gamma}^{-1}|^2 = J_{\gamma\gamma}^{-1} - C^0(C^{0'} J_{\gamma\gamma} C^0)^{-1} C^{0'}$, and the result follows.

To show (19), define $n^{1/2}(\alpha_n^\dagger - \bar{\alpha}) \equiv (J_{\alpha\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\alpha})^{-1} J_{\alpha\gamma} J_{\gamma\gamma}^{-1} \Psi_n$. Now,

$$\begin{aligned} 0 &\leq -2n \log \hat{Q}_n(\bar{\beta}, \hat{\alpha}_n^0, \cdot) \mathbb{E} \rho_\tau(W_1) + 2n \log \hat{Q}_n(\bar{\beta}, \alpha_n^\dagger, \cdot) \mathbb{E} \rho_\tau(W_1) \\ &= -2\Psi_n' J_{\gamma\gamma}^{-1} J_{\gamma\alpha} n^{1/2} ((\hat{\alpha}_n^0 - \bar{\alpha}) - (\alpha_n^\dagger - \bar{\alpha})) + n^{1/2} (\hat{\alpha}_n^0 - \bar{\alpha})' J_{\alpha\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\alpha} n^{1/2} (\hat{\alpha}_n^0 - \bar{\alpha}) \\ &\quad - n^{1/2} (\alpha_n^\dagger - \bar{\alpha})' J_{\alpha\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\alpha} n^{1/2} (\alpha_n^\dagger - \bar{\alpha}) \\ &\quad + o_P(n^{1/2} |\hat{\alpha}_n^0 - \bar{\alpha}| + n |\hat{\alpha}_n^0 - \bar{\alpha}|^2 + 1) \\ &= -2\Psi_n' J_{\gamma\gamma}^{-1} J_{\gamma\alpha} (J_{\alpha\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\alpha})^{-1} J_{\alpha\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\alpha} n^{1/2} ((\hat{\alpha}_n^0 - \bar{\alpha}) - (\alpha_n^\dagger - \bar{\alpha})) \\ &\quad + n^{1/2} (\hat{\alpha}_n^0 - \bar{\alpha})' J_{\alpha\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\alpha} n^{1/2} (\hat{\alpha}_n^0 - \bar{\alpha}) \\ &\quad - n^{1/2} (\alpha_n^\dagger - \bar{\alpha})' J_{\alpha\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\alpha} n^{1/2} (\alpha_n^\dagger - \bar{\alpha}) \\ &\quad + o_P(n^{1/2} |\hat{\alpha}_n^0 - \bar{\alpha}| + n |\hat{\alpha}_n^0 - \bar{\alpha}|^2 + 1) \\ &= -2(\alpha_n^\dagger - \bar{\alpha})' J_{\alpha\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\alpha} n^{1/2} ((\hat{\alpha}_n^0 - \bar{\alpha}) - (\alpha_n^\dagger - \bar{\alpha})) \\ &\quad + n^{1/2} (\hat{\alpha}_n^0 - \bar{\alpha})' J_{\alpha\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\alpha} n^{1/2} (\hat{\alpha}_n^0 - \bar{\alpha}) \\ &\quad - n^{1/2} (\alpha_n^\dagger - \bar{\alpha})' J_{\alpha\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\alpha} n^{1/2} (\alpha_n^\dagger - \bar{\alpha}) \\ &\quad + o_P(n^{1/2} |\hat{\alpha}_n^0 - \bar{\alpha}| + n |\hat{\alpha}_n^0 - \bar{\alpha}|^2 + 1) \\ &= n^{1/2} (\hat{\alpha}_n^0 - \alpha_n^\dagger)' J_{\alpha\gamma} J_{\gamma\gamma}^{-1} J_{\gamma\alpha} n^{1/2} (\hat{\alpha}_n^0 - \alpha_n^\dagger) + o_P(n^{1/2} |\hat{\alpha}_n^0 - \bar{\alpha}| + n |\hat{\alpha}_n^0 - \bar{\alpha}|^2 + 1), \end{aligned}$$

and (19) follows by Lemma A.6. \square

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