

Syracuse University

## SURFACE at Syracuse University

---

Center for Policy Research

Maxwell School of Citizenship and Public  
Affairs

---

2001

### Asymptotic Inference in Censored Regression Models Revisited

Chihwa Kao

*Syracuse University. Center for Policy Research, cdkao@maxwell.syr.edu*

Follow this and additional works at: <https://surface.syr.edu/cpr>



Part of the [Mathematics Commons](#)

---

#### Recommended Citation

Kao, Chihwa, "Asymptotic Inference in Censored Regression Models Revisited" (2001). *Center for Policy Research*. 119.

<https://surface.syr.edu/cpr/119>

This Working Paper is brought to you for free and open access by the Maxwell School of Citizenship and Public Affairs at SURFACE at Syracuse University. It has been accepted for inclusion in Center for Policy Research by an authorized administrator of SURFACE at Syracuse University. For more information, please contact [surface@syr.edu](mailto:surface@syr.edu).

**Center for Policy Research  
Working Paper No. 36**

**ASYMPTOTIC INFERENCE IN CENSORED  
REGRESSION MODELS REVISITED**

**Chihwa Kao**

**Center for Policy Research  
Maxwell School of Citizenship and Public Affairs  
Syracuse University  
426 Eggers Hall  
Syracuse, New York 13244-1020  
(315) 443-3114 | Fax (315) 443-1081  
e-mail: [ctrpol@syr.edu](mailto:ctrpol@syr.edu)**

**February 2001**

**\$5.00**

Up-to-date information about CPR's research projects and other activities is available from our World Wide Web site at [www-cpr.maxwell.syr.edu](http://www-cpr.maxwell.syr.edu). All recent working papers and Policy Briefs can be read and/or printed from there as well.

# CENTER FOR POLICY RESEARCH – Spring 2001

**Timothy M. Smeeding, Director**  
**Professor of Economics & Public Administration**

---

## Associate Directors

Margaret M. Austin  
Associate Director,  
Budget and Administration

Douglas Holtz-Eakin  
Chair, Professor of Economics  
Associate Director, Center for Policy Research

Douglas Wolf  
Professor of Public Administration  
Associate Director, Aging Studies Program

John Yinger  
Professor of Economics and Public Administration  
Associate Director, Metropolitan Studies Program

## SENIOR RESEARCH ASSOCIATES

Scott Allard .....	Public Administration	Eric Kingson .....	Social Work
Dan Black .....	Economics	Thomas Kniesner .....	Economics
Stacy Dickert-Conlin.....	Economics	Jeff Kubik.....	Economics
William Duncombe .....	Public Administration	Jerry Miner .....	Economics
Gary Engelhardt .....	Economics	John Moran.....	Economics
Deborah Freund.....	Public Administration	Jan Ondrich .....	Economics
Vernon Greene .....	Public Administration	John Palmer .....	Public Administration
Leah Gutierrez .....	Public Administration	Lori Ploutz-Snyder.....	Health and Physical Education
Madonna Harrington Meyer .....	Sociology	Grant Reeher .....	Political Science
Christine Himes .....	Sociology	Stuart Rosenthal .....	Economics
Jacqueline Johnson .....	Sociology	Jodi Sandfort .....	Public Administration
Bernard Jump.....	Public Administration	Michael Wasylenko .....	Economics
Duke Kao .....	Economics	Assata Zerai.....	Sociology

## GRADUATE ASSOCIATES

Reagan Baughman .....	Economics	Kwangho Jung.....	Public Administration
Robert Bifulco.....	Public Administration	James Laditka.....	Public Administration
Caroline Bourdeaux .....	Public Administration	Xiaoli Liang.....	Economics
Christine Caffrey .....	Sociology	Donald Marples .....	Economics
Christopher Cunningham .....	Economics	Neddy Matshalaga.....	Sociology
Tae Ho Eom .....	Public Administration	Suzanne Plourde .....	Economics
Seth Giertz .....	Economics	Nora Ranney .....	Public Administration
Andrzej Grodner.....	Economics	Catherine Richards .....	Sociology
Rain Henderson .....	Public Administration	Adriana Sandu .....	Public Administration
Pam Herd.....	Sociology	Mehmet Serkan Tosun.....	Economics
Lisa Hotchkiss .....	Public Administration	Mark Trembley.....	Public Administration
Peter Howe .....	Economics	James Williamson .....	Economics
Benjamin Johns .....	Public Administration	Bo Zhao .....	Economics
Anil Kumar .....	Economics		

## STAFF

JoAnna Berger .....	Receptionist	Mary Santy .....	Administrative Secretary
Martha Bonney .....	Publications/Events Coordinator	Amy Storfer-Isser.....	Computer Support, .....Microsim Project
Karen Cimilluca.....	Librarian/Office Coordinator	Debbie Tafel .....	Secretary to the Director
Kati Foley .....	Administrative Assistant, LIS	Ann Wicks .....	Administrative Secretary
Esther Gray.....	Administrative Secretary	Lobrenzo Wingo.....	Computer Consultant
Kitty Nasto .....	Administrative Secretary		
Denise Paul.....	Editorial Assistant, <i>NTJ</i>		

## Abstract

This paper establishes that regressors in the models with censored dependent variables need not be bounded for the standard asymptotic results to apply. Thus regressors which grow monotonically with the observation index may be acceptable. It also purports to provide an upper bound on the rate at which regressors may grow.

We show that if  $\|x_t\| \leq c$  for all  $t$ , then  $\lambda_{\min} \sum_{t=1}^T x_t x_t^T \rightarrow \infty$  is a sufficient condition for the

consistency and asymptotic normality of the MLE in censored regression models, which are different from those used by Amemiya (1973). For the case of growing regressors, we show that the sufficient conditions for the consistency and asymptotic normality of the MLE are  $\|x_t\|^2 = o(\log t)$  and

$\lambda_{\min} \sum_{t=1}^T x_t x_t^T \geq cT^\alpha$ , for some  $\alpha > 0$  and  $c > 0$ , but only for one-half of the parameter space. The

admissible growth rate given above implies that the Fisher Information matrix diverges, which seems to be an indispensable requirement for asymptotic inference for the censored regression models. More importantly, it represents a critical upper bound in the Fisher information matrix if it is exceeded monotonically. It also implies that  $x_t = (1, t)^T$  in Judge et al. (1985:791) for the Tobit is not admissible.

# 1. Introduction

Consider the linear regression model

$$y_t = x_t' \beta + \varepsilon_t, \quad (t = 1, \dots, T) \quad (1)$$

where the  $\varepsilon_t$  are i.i.d. normal random variables having mean zero and finite variance  $\sigma^2$ ,  $x_t$  is a  $k \times 1$  vector of regressors and  $\beta$  is a  $k \times 1$  vector of parameter of interest. However, for each  $t$ , instead of observing  $(y_t, x_t)$ , one observes  $(z_t, \delta_t, x_t)$ , where  $z_t = \max\{y_t, 0\}$ , and  $\delta_t$  is an indicator variable taking the value 1 if  $y_t > 0$  and 0 if  $y_t \leq 0$ .

The consistency and asymptotic normality of the maximum likelihood estimator (MLE) of  $(\beta^T, \sigma^2)$  have been studied by Amemiya (1973) under assumptions of  $\{x_t\}$ , e.g.,  $x_t$  is bounded, i.e.,  $\|x_t\| \leq c$  for all  $t$ , and  $\lim_{T \rightarrow \infty} \frac{1}{T} \sum_{t=1}^T x_t x_t' = V$ , a positive definite matrix. However, the conditions given by Amemiya (1973) above may be too strong to study the cases where growing regressors are of interest (e.g., time trend model, i.e.,  $x_t = (1, t)'$ , see Judge et al. 1985:791).

The likelihood function of this model is

$$L = \prod_{t=1}^T [1 - \Phi(x_t' \beta, \sigma^2)]^{1-\delta_t} \left[ \frac{1}{\sqrt{2\pi\sigma^2}} e^{-\frac{1}{2\sigma^2}(y_t - x_t' \beta)^2} \right]^{\delta_t} \quad (2)$$

with

$$\Phi(x_t' \beta, \sigma^2) = \int_{-\infty}^{x_t' \beta} \frac{1}{\sqrt{2\pi\sigma^2}} e^{-\frac{1}{2\sigma^2}(\lambda/\sigma)^2} d\lambda \quad (3)$$

In section 2, we confirmed Fahrmeir's (1987:103) result that if  $\|x_t\| \leq c$  for all  $t$ , then

$$\lambda_{\min} \sum_{t=1}^T x_t x_t' \rightarrow \infty \quad (4)$$



is a sufficient condition for the consistency and asymptotic normality of the MLE for the censored model but only for one-half of the parameter space, where  $\lambda_{\min} A$  is the smallest eigenvalues of a symmetric matrix  $A$ . In Section 3 the case of growing regressors is discussed. We show that the sufficient conditions for the consistency and asymptotic normality of the MLE are

$$\|x_t\|^2 = o(\log t) \tag{5}$$

$$\lambda_{\min} \sum_{t=1}^T x_t x_t' \geq cT^\alpha, \text{ for some } \alpha > 0 \text{ and } c > 0, \tag{6}$$

for one-half of the parameter space. The summary is given in the Section 4.

**Remark 1:** In the classical linear regression model with i.i.d errors and the sequence  $\{x_t\}$  of regressors,  $\lambda_{\min} \sum_{t=1}^T x_t x_t' \rightarrow \infty$  is necessary and sufficient for weak (Drygas 1976) and strong (Lai, Robbins, and Wei 1979) consistency (also see Amemiya 1985:95).

**Remark 2:** Note that  $\Phi(x_t' \beta, \sigma^2)$  will tend to one or zero if some regressors are growing monotonically to  $+\infty$  or  $-\infty$ . Thus, for large  $T$  nearly all response  $y_t$  will fall into one category and there will be too less information to draw inference about the relevant parameters. The admissible growth rate in (5) and (6) assure that enough information is available and the asymptotic theory works.

**Remark 3:** The admissible growth rate,  $\|x_t\|^2 = o(\log t)$ , given in (5) implies that the Fisher Information matrix diverges, which seems to be an indispensable requirement for asymptotic inference for the censored regression models. More importantly, it represents a critical upper bound in the Fisher information matrix if it is exceeded monotonically

**Remark 4:** Fahrmeir and Kaufmann (1986) and Gourieroux and Monfort (1981) have discussed the sharp upper bounds on the admissible growth of regressors for logit, probit, cumulative logit, and loglinear, and linear Poisson model.

## 2. Statistical Inference

Olsen (1978) proved the global concavity of  $\log L$  in the Tobit model in terms of the transformed parameter  $\alpha = \beta/\sigma$  and  $h = 1/\sigma$ . Without loss of generality (see Remark 7), we assume  $\sigma = 1$ . The  $\log L$  in terms of the new parameters can be written as (see Olsen 1978)

$$\log L = \sum_0 \log[1 - \Phi(x_i'\beta)] - \frac{1}{2} \sum_1 (y_t - x_t'\beta)^2 \quad (7)$$

where  $\sum_0$  is the summation over all observations where  $y_t = 0$ ,  $\sum_1$  is the summation over all observations where  $y_t > 0$ , and  $T_1$  is the number of observations where  $y_t > 0$ . The score function  $s_t(\beta)$  and the information matrix  $F_t(\beta)$  are

$$\begin{aligned} s_t(\beta) &= -\sum_0 x_t \frac{\phi(x_t'\beta)}{1 - \Phi(x_t'\beta)} + \sum_1 x_t (y_t - x_t'\beta) \\ &= \sum_t x_t \left[ \delta_t (y_t - x_t'\beta) - (1 - \delta_t) \left( \frac{\phi(x_t'\beta)}{1 - \Phi(x_t'\beta)} \right) \right] \\ &= \sum_t x_t \left\{ \left[ \delta_t y_t + (1 - \delta_t) \left( x_t'\beta - \frac{\phi(x_t'\beta)}{1 - \Phi(x_t'\beta)} \right) \right] - x_t'\beta \right\} \\ &= \sum_t x_t (\hat{y}_t - x_t'\beta), \text{ say,} \end{aligned} \quad (8)$$



$$\begin{aligned}
F_t(\beta) &= \text{cov}_{\beta} s_t(\beta) = E_{\beta} H_t(\beta) = E_{\beta} \left[ \sum_0 \frac{\phi(x_t' \beta)}{1 - \Phi(x_t' \beta)} \left( \frac{\phi(x_t' \beta)}{1 - \Phi(x_t' \beta)} - x_t' \beta \right) x_t x_t' + \sum_1 x_t x_t' \right] \\
&= E_{\beta} \left[ \sum_t \left\{ (1 - \delta_t) \left[ \frac{\phi(x_t' \beta)}{1 - \Phi(x_t' \beta)} \left( \frac{\phi(x_t' \beta)}{1 - \Phi(x_t' \beta)} - x_t' \beta \right) \right] + \delta_t \right\} x_t x_t' \right] \\
&= \sum_t \left( \left\{ \frac{\phi(x_t' \beta)}{1 - \Phi(x_t' \beta)} - x_t' \beta \right\} \phi(x_t' \beta) + \Phi(x_t' \beta) \right) x_t x_t' \tag{9}
\end{aligned}$$

where  $H_t(\beta) = -\frac{\partial^2 \log L}{\partial \beta \partial \beta'}$  and  $\frac{\phi(x_t' \beta)}{1 - \Phi(x_t' \beta)} - x_t' \beta > 0$ . Therefore  $H_t(\beta)$  and  $F_t(\beta)$  are positive definite.

**Remark 5:**  $E \hat{y}_t = E \left[ \delta_t y_t + (1 - \delta_t) \left( x_t' \beta - \frac{\phi(x_t' \beta)}{1 - \Phi(x_t' \beta)} \right) \right] = x_t' \beta$  and

$\text{Var}(\hat{y}_t) = \sigma_t^2 = \sigma^2 - \text{Var}(y_t | y_t < -x_t' \beta)$ . In particular,  $\sup_{t \geq 1} \sigma_t^2 \leq \sigma^2 < \infty$  (see James and Smith 1984).

Let  $A^{1/2}$  be a unique symmetric positive definite matrix associated with a symmetric positive definite matrix  $A$  such that  $(A^{1/2})^2 = A$ . We approach the problem of the distribution of  $F_t^{-1/2} s_t$  first, and move to the large sample distribution of  $\hat{\beta}$ , MLE of  $\beta$ .

**Assumption 1:**  $\|x_t\| \leq c$  for all  $t$ ,

**Assumption 2:**  $\lambda_{\min} \sum_{t=1}^T x_t x_t' \rightarrow \infty$ .

Although Fahrmeir (1987:103) did not fully present a separate and detailed proof regarding the sufficient condition for the Tobit model with bounded regressors, he indicated that the stated sufficient conditions in Assumption 1 and Assumption 2 can be directly proved by his general results presented in his paper. Here, we merely confirm his conjecture.

**Lemma 1:** Under Assumption 1 and Assumption 2, the normed score function is asymptotically

normal:  $F_t^{-1/2} s_t \xrightarrow{d} N(0, I)$ .

Proof: We use the Lindeberg-Feller Theorem for triangular arrays. Fix  $\lambda$  with  $\lambda^T \lambda = 1$ .

$$z_{t,i} = \lambda^T F_t^{-1/2} x_i (\hat{y}_i - x_i^T \beta)$$

we have  $E z_{t,i} = 0$ ,  $\text{var} \sum_{i=1}^t z_{t,i} = \text{var} \lambda^T F_t^{-1/2} s_t = 1$ , i.e.,  $z_{t,i}$  are independent and  $\sum_{i=1}^t z_{t,i}$  has mean 0 and

variance 1. In order that  $z_{t,i}$  obey the central limit theorem, it is sufficient that the Lindeberg condition

(see Billingsley 1986:369) is satisfied, i.e., for any  $\varepsilon > 0$ ,

$$\lim_{t \rightarrow \infty} \sum_{i=1}^t \int_{\{z_{t,i}^2 > \varepsilon^2\}} z_{t,i}^2 dP = 0. \quad (11)$$

where  $P$  is the distribution of  $z_{t,i}$ . Let  $\alpha'_{t,i} = \lambda^T F_t^{-1/2} x_i$ . By the Cauchy-Schwarz inequality, we have

$$z_{t,i}^2 \leq \alpha'_{t,i} \alpha_{t,i} (\hat{y}_i - x_i^T \beta)^2 = \alpha'_{t,i} \alpha_{t,i} e_i^2, \text{ where } e_i = \hat{y}_i - x_i^T \beta. \text{ This gives}$$

$$\sum_{i=1}^t \int_{\{z_{t,i}^2 > \varepsilon^2\}} z_{t,i}^2 dP \leq \sum_{i=1}^t \alpha_{t,i}^T \alpha_{t,i} \int_{B(t,i)} e^2 dG_x \quad (12)$$

where  $G_x$  is the distribution of  $e$  for a given  $x$ , and  $B(t,i)$  is the set  $\left\{ e^2 > \frac{\varepsilon^2}{\alpha'_{t,i} \alpha_{t,i}} \right\}$ . From Assumption

1 and 2, we have  $\lambda_{\min} F_t \rightarrow \infty$ . Define

$$h_c(x) = \sup \int_{\{e^2 > e_c^2\}} e^2 dG_x \quad (13)$$

Under Assumption 1,  $\sum_{i=1}^t \alpha'_{t,i} \alpha_{t,i} \leq K < \infty$  with a constant  $K$ . Note that  $\sum_{i=1}^t \alpha'_{t,i} \alpha_{t,i} \int_{B(t,i)} e^2 dG_x$  is

bounded above by  $K h_c(x)$ .

$$\sum_{i=1}^t \int_{\{z_{t,i}^2 > \varepsilon^2\}} z_{t,i}^2 dP \leq K h_c(x) \rightarrow 0$$

because  $h_c(x) \rightarrow 0$  as  $c \rightarrow \infty$ .

**Theorem 1:** Under Assumption 1 and Assumption 2

$$(i) \hat{\beta} \xrightarrow{p} \beta$$

$$(ii) F_t^{1/2}(\hat{\beta} - \beta) \xrightarrow{d} N(0, I)$$

Proof: The proof is similar to the Theorem 4 in Fahrmeir and Kaufmann (1985:364).

**Remark 6:** A main step is the verification of

$$F_t^{-1/2}(\beta) H_t(\beta) F_t^{-1/2}(\beta) \xrightarrow{p} I. \quad (14)$$

This assertion is equivalent to  $\frac{\lambda' H_t(\beta) \lambda}{\lambda' F_t(\beta) \lambda} \rightarrow 1$ , uniformly all  $\lambda \neq 0$ , i.e., Condition (13) is a continuity

condition on  $H_t(\beta)$  as well as a convergence condition on the asymptotic relation between  $F_t(\beta)$  and

$H_t(\beta)$ , requiring that the ratio between observed information  $H_t(\beta)$  and expected information  $F_t(\beta)$

converges to one.

**Remark 7:** The above results hold even if  $\sigma^2$  is unknown. Note that

$$\lambda_{\min} F_t \geq \lambda_{\min} \begin{bmatrix} \sum_0 \left( \left\{ \frac{\phi(x_t' \alpha)}{1 - \Phi(x_t' \alpha)} - x_t' \alpha \right\} \phi(x_t' \alpha) + \Phi(x_t' \alpha) \right) x_t x_t' & 0 \\ 0 & \frac{T_1}{h^2} \end{bmatrix}.$$

Therefore  $\lambda_{\min} F_t \rightarrow \infty$  if and only if  $\lambda_{\min} \sum_{t=1}^T x_t x_t' \rightarrow \infty$  because

$$\lambda_{\min} \sum_{t=1}^T x_t x_t' \rightarrow \infty \text{ iff } \lambda_{\min} \sum_0 x_t x_t' \rightarrow \infty.$$

### 3. Growing Regressors

However, there are situations where growing regressors are of interest, e.g., time trend models (e.g., Judge et al. 1985:791). Fahrmeir and Kaufmann (1986:187) have given a sharp upper bound for admissible growth of regressors for the Probit. Note that Tobit like-likelihood in (2) is expressible as the sum of the probit log-likelihood and the truncated likelihood and the Probit MLE is asymptotically normal and  $\sqrt{T}$ —consistent for the original Tobit parameter vector divided by the standard deviation of the Tobit error term, one may think that the sufficient conditions on the growing regressors in the Tobit model and in the Probit model coincide.

It is true that  $\Phi(x_i'\beta, \sigma^2)$  in (2) will tend to one or zero if some regressors are growing monotonically to  $+\infty$  or  $-\infty$ . However, the statement by Fahrmeir and Kaufmann (1986:189) about the Probit: “Thus for large  $T$  nearly all response  $y_t$  will fall into one category and there will be too less information to draw inference about the relevant parameters” is only one-half right for the Tobit. Specifically, if the variable  $x_{it}$  is growing monotonically with  $t$ , and if its associated coefficient  $\beta_1$  is negative, then for large  $T$  nearly all responses  $y_t$  will be zero, so that additional observations will be indeed add too little information for asymptotic theory to work. That is the case that looks like Probit. But if the coefficient  $\beta_1$  is positive, then nearly all  $y_t > 0$  for large  $T$ , meaning new observations will add as much information as they would in the classical linear model. Thus, the upper bound on the growth rates for regressors in the censored normal model, is true as stated in (5) and (6), but only for half of the parameter space to assure that enough information is available and the asymptotic theory works.

The following theorem gives a sharp upper bound for admissible growth of regressors for half of the parameter space.

**Theorem 2:** If  $\|x_t\|^2 = o(\log t)$  and  $\lambda_{\min} \sum_{t=1}^T x_t x_t' \geq cT^\alpha$ , for some  $\alpha > 0$  and  $c > 0$ . Then

$$\lambda_{\min} F_t \rightarrow \infty.$$

Proof: The information matrix  $F_t(\beta)$  can also be written as  $F_t(\beta) = \sum_{t=1}^T x_t x_t' a_t^2$ , with

$$a_t^2 = \left( \left\{ \frac{\phi(x_t' \beta)}{(1 - \Phi(x_t' \beta))} - x_t' \beta \right\} \phi(x_t' \beta) + \Phi(x_t' \beta) \right) \quad (15)$$

From Magnus and Neudecker (1988:204), we have

$$\lambda_{\min} F_t(\beta) = \lambda_{\min} \sum_{t=1}^T x_t x_t' a_t^2 \geq \lambda_{\min} \sum_{t=1}^T \left[ \Phi(x_t' \beta) \right] x_t x_t'. \quad (16)$$

We also note that

$$\Phi(x_t' \beta) \geq c \exp(-\{x_t' \beta\}^2) \quad (17)$$

for some  $c > 0$ . The admissible growth rate of  $x_t$ ,  $\|x_t\|^2 = o(\log t)$ , in Theorem 2 is equivalent to

$$\exp\left(-\|x_t\|^2 \|\beta\|^2\right) \geq t^{-\delta} \text{ for all } \beta, \delta > 0, t \geq t_1(\beta, \delta). \quad (18)$$

In combination with (16) and (17),

$$\lambda_{\min} F_t(\beta) \geq \lambda_{\min} \sum_{t=1}^T x_t x_t' t^{-\delta}.$$

Since  $t^{-\delta} \geq T^{-\delta}$  for  $t \leq T$ ,  $\delta > 0$ . Under  $\lambda_{\min} \sum_{t=1}^T x_t x_t' \geq cT^\alpha$ , therefore, we have

$$\lambda_{\min} F_t(\beta) \geq cT^{\alpha-\delta}, \quad T \geq T_1$$

With  $\delta = \alpha/2$ , this implies that  $\lambda_{\min} F_t \rightarrow \infty$ .

**Remark 8:** If  $|x_t^T \beta|^2 \geq c \log t$ ,  $c > 1$ , then  $\lambda_{\min} F_t$  converges.

**Example:** Let  $x_t = (1, w_t)'$  with  $w_t^2 = (\log t)^\alpha$  for the Tobit Model in (1).  $\|x_t\|^2 = o(\log t)$  holds if  $0 < \alpha < 1$ . If  $\alpha \geq 1$ , then  $\lambda_{\min} F_t$  converges for some  $\beta$ .

**Remark 9:** It implies that  $x_t = (1, t)'$  in Judge et al. (1985:791) for the Tobit is not admissible.

**Remark 10:** It also implies that when  $x_t = (1, w_t)$  has a unit root with the drift, i.e.,  $w_t = \mu + w_{t-1} + \omega_t$

for the Tobit is not admissible. Since  $w_t$  can be written as  $w_t = w_0 + ut + \sum_{i=1}^t \omega_i$ .

## 4. Summary

Regression models for censored data have found numerous applications. Statistical analysis of these models relies heavily on large sample theory, i.e., asymptotic properties of the MLE. However, previously published conditions assuring these properties may be too strong. Consistency and Asymptotic normality of the MLE are shown under weak and easily verifiable requirements. This paper gives a sharp upper bound on the admissible growth of regressors.

## References

- Amemiya, T. 1973. "Regression Analysis when the Dependent Variable is Truncated Normal," *Econometrica*, 41: 997-1016.
- Amemiya, T. 1985. *Advanced Econometrics*. Cambridge, MA: Harvard University Press.
- Billingsley, P. 1986. *Probability and Measure*. New York: John Wiley.
- Drygas, R.A. 1976. "Weak and Strong Consistency of the Least Squares Estimators in Regression Models," *Zeitschrift Fur Wahrscheinlichkeitstheorie und Verwandte Gebiete*, 34: 119-127.
- Fahrmeir, L. and H. Kaufmann. 1985. "Consistency and Asymptotic Normality of the Maximum Likelihood Estimator in Generalized Linear Model," *Annals of Statistics*, 13: 342-368; Correction in *Annals of Statistics*, 14: 1643.
- Fahrmeir, L. and H. Kaufmann. 1986. "Asymptotic Inference in Discrete Response Models," *Statistische Hefte*, 27: 179-205.
- Fahrmeir, L. 1987. "Asymptotic Likelihood Inference for Nonhomogeneous Observations," *Statistische Hefte*, 28: 81-116.
- Gourieroux, C. and A. Monfort. 1981. "Asymptotic Properties of the Maximum Likelihood Estimator in Dichotomous Logit Models," *Journal of Econometrics*, 17: 83-97.
- James, I.R. and P.J. Smith. 1984. "Consistency Results for Linear Regression with Censored Data," *Annals of Statistics*, 12: 590-600.
- Judge, G., W.E. Griffiths, R.C. Hill, H. Lutkepohl, and T.-C. Lee. 1985. *The Theory and Practice and Econometrics*. New York: John Wiley.
- Lai, T.L., H. Robbins, and C.Z. Wei. 1979. "Strong Consistency of Least Squares Estimates in Multiple Regression II," *Journal of Multivariate Analysis*, 9: 343-361.
- Magnus, J.R. and H. Neudecker. 1988. *Matrix Differential Calculus with Applications in Statistics and Econometrics*. New York: John Wiley.
- Olsen, R.J. 1978. "Note on the Uniqueness of the Maximum Likelihood Estimator for the Tobit Model," *Econometrica*, 46: 1211-1215.