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DISCRETE CHOICES WITH PANEL DATA

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This paper reviews the existing approaches to deal with panel data binary choice models with individual effects. Their relative strengths and weaknesses are discussed. Much theoretical and empirical research is needed in this area, and the paper points to several aspects that deserve further investigation. In particular, I illustrate the usefulness of asymptotic arguments in providing both approximately unbiased moment conditions, and approximations to sampling distributions for panels of different sample sizes.

Keywords: Binary choice, panel data, fixed effects, modified likelihood, asymptotic corrections.

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1. Introduction

The use of fixed effects is a simple and well understood way of dealing with endogenous explanatory variables in linear panel data models. In such a context, least squares or instrumental variable methods for errors in differences provide consistent estimates that control for unobserved heterogeneity in short panels of large cross-sections (small T, large N). However, the situation is fundamentally different in models with nonlinear errors; for example, when one intends to use fixed effects to deal with an endogenous explanatory variable in a probit model. In those cases, estimates of the parameters of interest, jointly estimated with the effects, are typically inconsistent if T is fixed (incidental parameter problem). Moreover, fixed effects estimates in a spirit similar to differencing in the linear case are not available for many models of practical importance.

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There are also random effect methods that achieve fixed T consistency subject to a particular specification of the form of the dependence between the explanatory variables and the effects, but they rely on strong and untestable auxiliary assumptions, and even these methods are often out of reach. Without auxiliary assumptions, the common parameters of certain nonlinear fixed effects models are simply unidentifiable in a fixed T setting, so that fixed-T consistent estimation is not possible at all. In other cases, although identifiable, fixed-T consistent estimation at the standard root-N rate is impossible.

An alternative reaction to the fact that micro panels are short is to ask for estimators with small biases as opposed to no bias at all; specifically, estimators with biases of order $1/T^2$ instead of the standard magnitude of $1/T$. This alternative approach has the potential of overcoming some of the fixed-T identification difficulties and the advantage of generality.

The purpose of this lecture is twofold. First I review the incidental parameter problem (Sections 2 and 3), fixed-T solutions (Section 4), and identification problems (Section 5), all in the context of the static binary choice model with explanatory variables that are correlated with an individual effect. Second, I discuss the modified concentrated likelihood of Cox and Reid (1987), its role in achieving consistency up to a certain order of magnitude in T (Section 6), and a double asymptotic formulation which provides an effective discrimination between estimators with and without bias reduction (Section 7).

I focus on the static binary case for simplicity and because many results are only available for this case. Thus, dynamic models, multinomial choice, and models with random effects that are uncorrelated with the explanatory variables are all left out (see Arellano and Honoré (2001) for a fuller survey of the fixed-T panel data discrete choice literature). My intention is to exhibit the strengths and weaknesses of fixed-T approaches, and to illustrate the usefulness of double asymptotic arguments in providing both approximately unbiased moment conditions, and approximations to sampling distributions even for fairly short panels, which is the main theme of the paper.

2. Model and parameters of interest

I begin by considering the following static binary choice model

$$y_{it} = 1\left\{x'_{it}\beta_0 + \eta_i + v_{it} \geq 0\right\} \quad (t = 1, ..., T; i = 1, ..., N) \qquad [1]$$

where the errors v_{it} are independently distributed with *cdf* F conditional on η_i and $x_i = (x'_{i1}, ..., x'_{iT})'$, so that

$$\Pr(y_{it} = 1 \mid x_i, \eta_i) = F\left(x'_{it}\beta_0 + \eta_i\right). \qquad [2]$$

The linear model as a Benchmark. In a linear model of the form

$$E(y_{it} \mid x_i, \eta_i) = x'_{it}\beta_0 + \eta_i, \qquad [3]$$

β_0 is identifiable from the regression in first differences or deviations from means in a cross-sectional population for fixed T, regardless of the form of the distribution of $\eta_i \mid x_i$. That is, we have

$$\operatorname*{plim}_{N\to\infty} \frac{1}{TN} \sum_{i=1}^{N} \sum_{t=1}^{T} (x_{it} - \overline{x}_i) \left[(y_{it} - \overline{y}_i) - (x_{it} - \overline{x}_i)' \beta_0\right] = 0, \qquad [4]$$

which is uniquely satisfied by the true value β_0 provided

$$\operatorname*{plim}_{N\to\infty} \frac{1}{TN} \sum_{i=1}^{N} \sum_{t=1}^{T} (x_{it} - \overline{x}_i)(x_{it} - \overline{x}_i)' \qquad [5]$$

is non-singular. So, the value $\widehat{\beta}$ that solves

$$\frac{1}{TN} \sum_{i=1}^{N} \sum_{t=1}^{T} (x_{it} - \overline{x}_i) \left[(y_{it} - \overline{y}_i) - (x_{it} - \overline{x}_i)' \widehat{\beta}\right] = 0 \qquad [6]$$

(the "within-group" estimator) is a consistent estimator of β_0 for large N, no matter how small is T as long as $T \geq 2$ (see, for example, Arellano, 2003, or Hsiao, 2003).

This is of economic interest if one hopes that by conditioning on η_i, β_0 measures a more relevant (causal or structural) effect of x on y. The consistency result matters because one wants to make sure that gets the right answer when calculating $\widehat{\beta}$ from a large cross-sectional panel with a small time series dimension, which is a typical situation in microeconometrics.

The motivation and aim in a binary choice fixed effects model is to get similar results as in the linear case when the form of the model is given by [1]. In our context, the term "fixed effects" has nothing to do with the nature of sampling. It just refers to a model for the effect of x on y given x and η, in which we observe y and x but not η, and

the distribution of $\eta \mid x$ is left unrestricted. Following the usage in the econometric literature, the term "random effects" will be reserved for models in which some knowledge about the form of the distribution of $\eta \mid x$ is assumed.

Parameters of interest. The micropanel data literature has emphasized the large-N-short-T identification of β_0 with an unspecified distribution of $\eta_i \mid x_i$. However, a natural parameter of interest is the mean effect on the probability of $y_{it} = 1$ of changing x_{1it} from z_a to z_b, say. A consistent estimator of this is:

$$\frac{1}{N}\sum_{i=1}^{N}\int \left[F\left(z_a\beta_{01} + x'_{2it}\beta_{02} + \eta\right) - F\left(z_b\beta_{01} + x'_{2it}\beta_{02} + \eta\right)\right] dG(\eta \mid x_{2it}) \qquad [7]$$

where $G(. \mid x_{2it})$ is the *cdf* of η_i conditional on x_{2it}, and x_{1it} denotes the first component of x_{it}. Thus, measuring this effect would require us to specify G, which is not in the nature of the fixed effects approach.[1]

The direct information we can get from the β coefficients only concerns the relative impacts of explanatory variables on the probabilities. If x_{1it} and x_{2it} are continuous variables we have:

$$\frac{\beta_{02}}{\beta_{01}} = \frac{\partial \Pr ob(y_{it} = 1 \mid x_i, \eta_i)}{\partial x_{2it}} \Big/ \frac{\partial \Pr ob(y_{it} = 1 \mid x_i, \eta_i)}{\partial x_{1it}}. \qquad [8]$$

3. The problem

The log-likelihood function from [1] assuming that the y_{it} are independent conditional on x_i and η_i is given by

$$\sum_{i=1}^{N} \ell_i\left(\beta, \eta_i\right) \qquad [9]$$

where

$$\ell_i\left(\beta, \eta_i\right) = \sum_{t=1}^{T} \left\{y_{it} \log F_{it} + (1 - y_{it}) \log\left(1 - F_{it}\right)\right\} \qquad [10]$$

[1] An alternative is to obtain the difference in probabilities for specific values of η and x_{2t} (e.g. their means), but this may only be relevant for a small part of the population (see Chamberlain, 1984).

and $F_{it} = F(x'_{it}\beta + \eta_i)$. Moreover, the scores are

$$d_{\eta i}(\beta, \eta_i) \equiv \frac{\partial \ell_i(\beta, \eta_i)}{\partial \eta_i} = \sum_{t=1}^{T} \frac{f_{it}}{F_{it}(1 - F_{it})}(y_{it} - F_{it}) \quad [11]$$

$$d_{\beta i}(\beta, \eta_i) \equiv \frac{\partial \ell_i(\beta, \eta_i)}{\partial \beta} = \sum_{t=1}^{T} \frac{f_{it}}{F_{it}(1 - F_{it})} x_{it}(y_{it} - F_{it}) \quad [12]$$

where f_{it} denotes the *pdf* corresponding to F_{it}.

For the logit model F is the logistic *cdf* $\Lambda(r) = e^r/(1 + e^r)$ and we have

$$\frac{f_{it}}{F_{it}(1 - F_{it})} = 1$$

so that in this case the scores are simply $d_{\eta i}(\beta, \eta_i) = \sum_{t=1}^{T}(y_{it} - F_{it})$ and $d_{\beta i}(\beta, \eta_i) = \sum_{t=1}^{T} x_{it}(y_{it} - F_{it})$.

Let the MLE of η_i for given β be

$$\widehat{\eta}_i(\beta) = \arg\max_{\eta} \ell_i(\beta, \eta_i) \quad [13]$$

so that $\widehat{\eta}_i(\beta)$ solves

$$d_{\eta i}(\beta, \widehat{\eta}_i(\beta)) = 0. \quad [14]$$

Therefore, the MLE of β is given by the maximizer of the concentrated (or profile) log-likelihood

$$\widehat{\beta} = \arg\max_{\beta} \sum_{i=1}^{N} \ell_i(\beta, \widehat{\eta}_i(\beta)) \quad [15]$$

which solves the first order conditions

$$\begin{aligned} b_{TN}(\beta) &= \frac{1}{TN} \sum_{i=1}^{N} \left\{ d_{\beta i}(\beta, \widehat{\eta}_i(\beta)) + d_{\eta i}(\beta, \widehat{\eta}_i(\beta)) \frac{\partial \widehat{\eta}_i(\beta)}{\partial \beta} \right\} \\ &= \frac{1}{TN} \sum_{i=1}^{N} d_{\beta i}(\beta, \widehat{\eta}_i(\beta)) \end{aligned} \quad [16]$$

The problem is that $b_{TN}(\beta)$ evaluated at $\beta = \beta_0$ does not converge to zero in probability when $N \to \infty$ for T fixed (although it does converge to zero when $T \to \infty$). This situation is known as the *incidental parameters problem* since Neyman and Scott (1948). A discussion of this problem for discrete choice models is in Heckman (1981).

An example. As a classic illustration let us consider a model in which $T = 2$, f is symmetric, β is scalar, and x_{it} is a time dummy such that $x_{i1} = 0$ and $x_{i2} = 1$ (Andersen, 1973; Heckman, 1981). For observations with $(y_{i1}, y_{i2}) = (0,0)$ we have $\widehat{\eta}_i(\beta) \to -\infty$ and $\ell_i(\beta, \widehat{\eta}_i(\beta)) = \log F(-\widehat{\eta}_i(\beta)) + \log F(-\widehat{\eta}_i(\beta) - \beta) \to 0$. For observations with $(y_{i1}, y_{i2}) = (1,1)$ we have $\widehat{\eta}_i(\beta) \to \infty$ and $\ell_i(\beta, \widehat{\eta}_i(\beta)) = \log F(\widehat{\eta}_i(\beta)) + \log F(\widehat{\eta}_i(\beta) + \beta) \to 0$. Finally, for $(0,1)$ or $(1,0)$ observations we have, respectively,

$$\ell_i(\beta, \eta) = \log F(-\eta) + \log F(\eta + \beta)$$

or

$$\ell_i(\beta, \eta) = \log F(\eta) + \log F(-\eta - \beta),$$

which in both cases are maximized at

$$\widehat{\eta}_i(\beta) = -\frac{\beta}{2}. \qquad [17]$$

The implication is that the contributions of observations $(0,0)$ and $(1,1)$ to the concentrated log-likelihood are equal to zero, a $(0,1)$ observation contributes a term of the form $2\log F(\beta/2)$, and a $(1,0)$ observation contributes with $2\log[1 - F(\beta/2)]$. So the concentrated log-likelihood is given by

$$2\sum_{i=1}^{N} \{d_{10i} \log[1 - F(\beta/2)] + d_{01i} \log F(\beta/2)\} \qquad [18]$$

where $d_{10i} = 1(y_{i1} = 1, y_{i2} = 0)$ and $d_{01i} = 1(y_{i1} = 0, y_{i2} = 1)$.

Moreover, the MLE of $p = F(\beta/2)$ is

$$\widehat{p} = \frac{\sum_{i=1}^{N} d_{01i}}{\sum_{i=1}^{N} 1(y_{i1} + y_{i2} = 1)}, \qquad [19]$$

so that

$$\widehat{\beta} = 2F^{-1}(\widehat{p}). \qquad [20]$$

Note that $\widehat{p}$ is the sample counterpart of $p_0 = \Pr(y_{i1} = 0, y_{i2} = 1 \mid y_{i1} + y_{i2} = 1)$. Thus the MLE $\widehat{\beta}$ satisfies

$$\operatorname*{plim}_{N\to\infty} \widehat{\beta} = 2F^{-1}(p_0). \qquad [21]$$

Moreover, we have

$$p_0 = \int \Pr(y_{i1} = 0, y_{i2} = 1 \mid y_{i1} + y_{i2} = 1, \eta)\, dG(\eta \mid y_{i1} + y_{i2} = 1).$$

For the logit model $\Pr(y_{i1} = 0, y_{i2} = 1 \mid y_{i1} + y_{i2} = 1, \eta)$ does not depend on η and it turns out that $p_0 = \Lambda(\beta_0)$ where β_0 is the true value. Therefore, in such a case $\operatorname{plim}_{N\to\infty} \widehat{\beta} = 2\Lambda^{-1}[\Lambda(\beta_0)] = 2\beta_0$, so that ML would be estimating a relative log odds ratio that is twice as large as its true value.

More generally, using Bayes formula

$$p_0 = \int \frac{\Pr(y_{i1} = 0, y_{i2} = 1 \mid \eta)}{\Pr(y_{i1} + y_{i2} = 1 \mid \eta)} \frac{\Pr(y_{i1} + y_{i2} = 1 \mid \eta)}{\Pr(y_{i1} + y_{i2} = 1)} dG(\eta)$$

or

$$p_0 = \frac{E_\eta[\Pr(y_{i1} = 0, y_{i2} = 1 \mid \eta)]}{E_\eta[\Pr(y_{i1} + y_{i2} = 1 \mid \eta)]}, \qquad [22]$$

so that

$$\operatorname*{plim}_{N\to\infty} \widehat{\beta} = 2F^{-1}\left\{\frac{E_\eta[F(-\eta)F(\beta+\eta)]}{E_\eta[F(-\eta)F(\beta+\eta)] + E_\eta[F(\eta)F(-\beta-\eta)]}\right\}. \qquad [23]$$

Thus, in general the form of the asymptotic bias of $\widehat{\beta}$ depends on the distribution of the individual effects.

For probit, under $\eta \sim \mathcal{N}(0, \sigma_\eta^2)$ an explicit expression is available. Letting $\beta^* = \beta/(1+\sigma_\eta^2)^{1/2}$ and $\rho = \sigma_\eta^2/(1+\sigma_\eta^2)$ we have $E_\eta[\Phi(\eta)] = \Phi(0) = 0{,}5$, $E_\eta[\Phi(\beta+\eta)] = \Phi(\beta^*)$, and $E_\eta[\Phi(\eta)\Phi(\beta+\eta)] = \Phi_2(0, \beta^*; \rho)$, so that

$$\operatorname*{plim}_{N\to\infty} \widehat{\beta} = 2\Phi^{-1}\left[\frac{\Phi(\beta^*) - \Phi_2(0, \beta^*; \rho)}{\Phi(\beta^*) + 0{,}5 - 2\Phi_2(0, \beta^*; \rho)}\right] \qquad [24]$$

where $\Phi_2(.,.;\rho)$is the *cdf* of the standardized bivariate normal density with correlation coefficient ρ. As noted by Heckman (1981), we get $p_0 \to F(\beta)$ as $\sigma_\eta \to 0$, in which case we get a similar bias result as for the logistic in a limiting situation. Numerical calculations of [24] are reported below in Section 6.

4. Fixed T solutions

4.1. Conditional MLE

A sufficient statistic for η_i, S_i say, is a function of the data such that the distribution of the data given S_i does not depend on η_i. The idea is to use the likelihood conditioned on S_i to make inference about β_0 (Andersen, 1970). This works as long as β_0 is identified from the conditional likelihood of the data, which obviously requires that the conditional likelihood depends on β_0. Unfortunately, this is not the case except for the logit model.

In the logit model $\sum_{t=1}^{T} y_{it}$ is a sufficient statistic for η_i. Indeed, we have

$$\Pr\left(y_{i1},...,y_{iT} \mid \sum_{t=1}^{T} y_{it}, x_i\right) = \frac{\exp\left(\sum_{t=1}^{T} y_{it} x_{it}'\beta_0\right)}{\sum_{(d_1,..,d_T)\in B_i} \exp\left(\sum_{t=1}^{T} d_t x_{it}'\beta_0\right)} \qquad [25]$$

where B_i is the set of all $0-1$ sequences such that $\sum_{t=1}^{T} d_t = \sum_{t=1}^{T} y_{it}$. This result was first obtained by Rasch (1960, 1961) (for surveys see Chamberlain, 1984, or Arellano and Honoré, 2001). For example, with $T = 2$ we have

$$\Pr\left(y_{i1}, y_{i2} \mid y_{i1} + y_{i2}, x_i\right) = \begin{cases} 1 \quad \text{if } (y_{i1}, y_{i2}) & = (0,0) \text{ or } (1,1) \\ 1 - \Lambda\left(\Delta x_{i2}'\beta_0\right) & \text{if } (y_{i1}, y_{i2}) = (1,0) \\ \Lambda\left(\Delta x_{i2}'\beta_0\right) & \text{if } (y_{i1}, y_{i2}) = (0,1). \end{cases} \qquad [26]$$

Therefore, the log-likelihood conditioned on $y_{i1} + y_{i2}$ is given by[2]

$$L_c(\beta) = \sum_{i=1}^{N} \left\{ d_{10i} \log\left[1 - \Lambda\left(\Delta x_{i2}'\beta\right)\right] + d_{01i} \log \Lambda\left(\Delta x_{i2}'\beta\right) \right\} \qquad [27]$$

and the score takes the form

$$\frac{\partial L_c(\beta)}{\partial \beta} = \sum_{i=1}^{N} \Delta x_{i2} \left\{ d_{01i} - \Lambda\left(\Delta x_{i2}'\beta\right) 1(y_{i1} + y_{i2} = 1) \right\}. \qquad [28]$$

4.2. Maximum score estimation

The previous technique crucially relied on the logit assumption. Manski (1987) considered a more general model of the form [1] in which the

[2]The contributions of $(0,0)$ or $(1,1)$ observations is zero.

cdf of $-v_{it} \mid x_i, \eta_i$ was non-parametric and could depend on x_i and η_i in a time-invariant way. Namely, for all t and s

$$\Pr(-v_{it} \leq r \mid x_i, \eta_i) = \Pr(-v_{is} \leq r \mid x_i, \eta_i) = F(r \mid x_i, \eta_i), \qquad [29]$$

so that $F(r \mid x_i, \eta_i)$ does not change with t but is otherwise unrestricted.

This assumption imposes stationarity and strict exogeneity, but allows for serial dependence in the errors v_{it}. It also allows for a certain kind of conditional heteroskedasticity, though not a very plausible one, since $Var(v_{it} \mid x_i, \eta_i)$ may depend on x_i but v_{it} is not allowed to be more sensitive to x_{it} than to other x's. Similarly if the expectations $E(v_{it} \mid x_i, \eta_i)$ exist, they may depend on x_i but not their first-differences $E(\Delta v_{it} \mid x_i, \eta_i) = 0$.

The time-invariance of F implies that for $T = 2$:[3]

$$med(y_{i2} - y_{i1} \mid x_i, y_{i1} + y_{i2} = 1) = \operatorname{sgn}(\Delta x'_{i2}\beta_0). \qquad [30]$$

To see this note that, given $y_{i1} + y_{i2} = 1$, the difference $y_{i2} - y_{i1}$ can only equal 1 or -1. So the median will be one or the other depending on whether $\Pr(y_{i2} = 1, y_{i1} = 0 \mid x_i) \lesseqgtr \Pr(y_{i2} = 0, y_{i1} = 1 \mid x_i)$. Thus[4]

$$\begin{aligned} & med(y_{i2} - y_{i1} \mid x_i, y_{i1} + y_{i2} = 1) \\ = \; & \operatorname{sgn}[\Pr(y_{i2} = 1, y_{i1} = 0 \mid x_i) - \Pr(y_{i2} = 0, y_{i1} = 1 \mid x_i)] \\ = \; & \operatorname{sgn}[\Pr(y_{i2} = 1 \mid x_i) - \Pr(y_{i1} = 1 \mid x_i)]. \end{aligned}$$

Moreover, from the model's specification, i.e.

$$\begin{aligned} \Pr(y_{i1} = 1 \mid x_i, \eta_i) &= F(x'_{i1}\beta_0 + \eta_i \mid x_i, \eta_i) \\ \Pr(y_{i2} = 1 \mid x_i, \eta_i) &= F(x'_{i2}\beta_0 + \eta_i \mid x_i, \eta_i), \end{aligned}$$

[3]The sign function is defined as

$$\operatorname{sgn}(u) = 1(u > 0) - 1(u < 0),$$

i.e. $\operatorname{sgn}(u) = -1$ if $u < 0$, $\operatorname{sgn}(u) = 0$ if $u = 0$ and $\operatorname{sgn}(u) = 1$ if $u > 0$.

[4]The second equality follows from

$$\begin{aligned} \Pr(y_{i2} = 1 \mid x_i) &= \Pr(y_{i2} = 1, y_{i1} = 0 \mid x_i) + \Pr(y_{i2} = 1, y_{i1} = 1 \mid x_i) \\ \Pr(y_{i1} = 1 \mid x_i) &= \Pr(y_{i2} = 0, y_{i1} = 1 \mid x_i) + \Pr(y_{i2} = 1, y_{i1} = 1 \mid x_i). \end{aligned}$$

and the monotonicity of F, we have that for any η_i (the constancy of F over time becomes crucial at this point):

$$\Pr(y_{i2} = 1 \mid x_i, \eta_i) \gtreqless \Pr(y_{i1} = 1 \mid x_i, \eta_i) \Leftrightarrow x'_{i2}\beta_0 \gtreqless x'_{i1}\beta_0.$$

Therefore, the implication also holds unconditionally relative to η_i:

$$\Pr(y_{i2} = 1 \mid x_i) \gtreqless \Pr(y_{i1} = 1 \mid x_i) \Leftrightarrow x'_{i2}\beta_0 \gtreqless x'_{i1}\beta_0.$$

or

$$\operatorname{sgn}[\Pr(y_{i2} = 1 \mid x_i) - \Pr(y_{i1} = 1 \mid x_i)] = \operatorname{sgn}\left(\Delta x'_{i2}\beta_0\right).$$

Manski showed that the true value of β_0 uniquely maximizes (up to scale) the expected agreement between the sign of $\Delta x'_{i2}\beta$ and that of Δy_{i2} conditioned on $y_{i1}+y_{i2} = 1$. This identification result required an unbounded support for at least one of the explanatory variables with a non-zero coefficient. That is, letting $x'_{it} = (z_{it}, w'_{it})$ and $\beta'_0 = (\gamma_0, \alpha'_0)$, the minimal requirement for identification is that z_{it} has unbounded support and $\gamma_0 \neq 0$. Identification fails at $\gamma_0 = 0$, so that $\gamma_0 = 0$ is not a testable hypothesis. Manski's identification result implies that we can learn about the relative effects of the variables w_{it} under the maintained assumption that $\gamma_0 \neq 0$.

Manski then proposed to estimate β_0 by selecting the value that matches the sign of $\Delta x'_{i2}\beta$ with that of Δy_{i2} for as many observations as possible in the subsample with $y_{i1} + y_{i2} = 1$. The suggested estimator is

$$\widehat{\beta} = \arg\max_{\beta} \sum_{i=1}^{N} \operatorname{sgn}\left(\Delta x'_{i2}\beta\right)(y_{i2} - y_{i1}) \qquad [31]$$

subject to the normalization $\| \beta \| = 1$.[5] This is the maximum score estimator applied to the observations with $y_{i1} + y_{i2} = 1$ (notice that the estimation criterion is unaffected by removing observations having $y_{i1} = y_{i2}$). It is consistent under the assumption that there is at least one unbounded continuous regressor, but it is not root-N consistent, and not asymptotically normal.

An alternative form of the score objective function is

$$S_N(\beta) = \sum_{i=1}^{N} \left\{ d_{10i} 1\left(\Delta x'_{i2}\beta < 0\right) + d_{01i} 1\left(\Delta x'_{i2}\beta \geq 0\right) \right\}. \qquad [32]$$

[5] In the logit case the scale normalization is imposed through the variance of the logistic distribution. More generally, if F is a known distribution a priori, the scale normalization is determined by the form of F. Comparisons can be made by considering ratios of coefficients.

The score $S_N(\beta)$ gives the number of correct predictions we would make if we predicted (y_{i1}, y_{i2}) to be $(0,1)$ whenever $\Delta x'_{i2}\beta \geq 0$. In contrast, $\sum_{i=1}^{N} \operatorname{sgn}(\Delta x'_{i2}\beta)\,\Delta y_{i2}$ gives the number of successes minus the number of failures. Yet another form of the estimator suggested by the median regression interpretation is as the minimizer of the number of failures, which is given by

$$\frac{1}{2}\sum_{i=1}^{N} 1(y_{i1} \neq y_{i2})\left|\Delta y_{i2} - \operatorname{sgn}(\Delta x'_{i2}\beta)\right|. \qquad [33]$$

Smoothed maximum score. It is possible to consider a smoothed version of the maximum score estimator along the lines of Horowitz (1992), which does have an asymptotic normal distribution, although the rate of convergence remains slower than root-N (Charlier, Melenberg and van Soest, 1995, and Kyriazidou, 1997).[6] The idea is to replace $S_N(\beta)$ with a smooth function $S_N^*(\beta)$ whose limit *a.s.* as $N \to \infty$ is the same as $S_N(\beta)$. This is of the form

$$S_N^*(\beta) = \sum_{i=1}^{N} \left\{ d_{10i}\left[1 - K\left(\Delta x'_{i2}\beta/\gamma_N\right)\right] + d_{01i}K\left(\Delta x'_{i2}\beta/\gamma_N\right)\right\} \qquad [34]$$

where $K(.)$ is analogous to a *cdf* and γ_N is a sequence of positive numbers such that $\lim_{N\to\infty}\gamma_N = 0$.

4.3. Random effects

In general

$$\Pr(y_{i1}, ..., y_{iT} \mid x_i) = \int \Pr(y_{i1}, ..., y_{iT} \mid x_i, \eta_i)\, dG(\eta_i \mid x_i) \qquad [35]$$

where $G(\eta_i \mid x_i)$ is the *cdf* of $\eta_i \mid x_i$. The substantive model specifies $\Pr(y_{i1}, ..., y_{iT} \mid x_i, \eta_i)$, but only $\Pr(y_{i1}, ..., y_{iT} \mid x_i)$ has an empirical counterpart. For example, we may have specified

$$\Pr(y_{i1}, ..., y_{iT} \mid x_i, \eta_i) = \prod_{t=1}^{T} \Pr(y_{it} \mid x_i, \eta_i) = \prod_{t=1}^{T} F_{it}^{y_{it}}(1 - F_{it})^{(1-y_{it})}.$$

In a fixed effects model we seek to make inferences about parameters in $\Pr(y_{i1}, ..., y_{iT} \mid x_i, \eta_i)$ without restricting the form of G. In a random

[6] Chamberlain (1986) showed that there is no root-N consistent estimator of β under the assumptions of Manski for his maximum score method.

effects model G is typically parametric or semiparametric, and the parameters of interest may or may not be identified with G unrestricted. Thus a fixed effects model can be regarded as a random effects model that leaves the distribution of the effects unrestricted.

The choice between fixed and random effects models often involves a trade-off between robustness in the specification of $\Pr(y_{i1}, ..., y_{iT} \mid x_i, \eta_i)$ and robustness in G, in the sense that achieving fixed-T identification with unrestricted G usually requires a more restrictive specification of $\Pr(y_{i1}, ..., y_{iT} \mid x_i, \eta_i)$.

Chamberlain (1980, 1984) considered a random effects model in which the effects are of the form

$$\eta_i = \mu(x_i) + \varepsilon_i \qquad [36]$$

and ε_i is independent of x_i. He also made the normality assumptions

$$v_{it} \mid x_i, \eta_i \sim \mathcal{N}(0, \omega_{tt}) \qquad [37]$$

$$\varepsilon_i \mid x_i \sim \mathcal{N}(0, \sigma_\eta^2), \qquad [38]$$

which imply that

$$\Pr(y_{it} = 1 \mid x_i) = \Phi\left[\sigma_t^{-1}\left(x_{it}'\beta_0 + \mu(x_i)\right)\right]. \qquad [39]$$

where $\sigma_t^2 = \sigma_\eta^2 + \omega_{tt}$ and $\Phi(.)$ is the standard normal *cdf*. In this model the v_{it} may be serially dependent and heteroskedastic over time.

Chamberlain assumed a linear specification $\mu(x_i) = \lambda_0 + x_i'\lambda$, and Newey (1994) generalized the model to a non-parametric $\mu(x_i)$. In the linear case, β_0, λ_0, λ, and the σ_t^2 can be estimated subject to the normalization $\sigma_1^2 = 1$ by combining the period-by-period probit likelihood functions (see Bover and Arellano, 1997, for a discussion of alternative estimators). In the semi-parametric case, Newey used the fact that

$$\sigma_t \Phi^{-1}\left[\Pr(y_{it} = 1 \mid x_i)\right] - \sigma_{t-1}\Phi^{-1}\left[\Pr\left(y_{i(t-1)} = 1 \mid x_i\right)\right] = \Delta x_{it}'\beta_0 \qquad [40]$$

together with non-parametric estimates of the probabilities $\Pr(y_{it} = 1 \mid x_i)$ to obtain an estimator of β_0 and the relative scales. A further generalization of the model is to drop the normality assumptions and allow the distribution of the errors $\varepsilon_i + v_{it} \mid x_i$ to be unknown. This case has been considered by Chen (1998).

Another semi-parametric approach has been followed by Lee (1999). Under certain assumptions on the joint distribution of x_i and η_i, Lee proposed a maximum rank correlation-type estimator which is $\sqrt{N}$-consistent and asymptotically normal.

5. Identification problems with fixed T

It would be useful to know which models for $\Pr(y_{i1}, ..., y_{iT} \mid x_i, \eta_i)$ are identified without placing restrictions in the form of $G(\eta_i \mid x_i)$ (i.e. *fixed-effects identification with fixed* T) and which are not.

A model is given by a $2^T \times 1$ vector $p(x_i, \eta_i, \beta_0)$ with elements that specify the probabilities

$$\Pr((y_{i1}, ..., y_{iT}) = d_j \mid x_i, \eta_i) \quad (j = 1, ..., 2^T) \qquad [41]$$

where d_j is a $0-1$ sequence of order T. Let the true *cdf* of $\eta_i \mid x_i$ be $G_0(\eta \mid x)$. Identification will fail at β_0 if for all x in the support of x_i there is a *cdf* $G^*(\eta \mid x)$ and $\beta^* \neq \beta_0$ in the parameter space, such that

$$\int p(x, \eta, \beta_0)\, dG_0(\eta \mid x) = \int p(x, \eta, \beta^*)\, dG^*(\eta \mid x). \qquad [42]$$

If this is so, (β_0, G_0) and (β^*, G^*) give the same conditional distribution for $(y_{i1}, ..., y_{iT})$ given x_i. Therefore, they are observationally equivalent relative to such distribution.

Chamberlain (1992) studied the identification of a fixed effects binary choice model with $T = 2$. He considered the model

$$y_{it} = 1(x'_{it}\beta_0 + \eta_i + v_{it} \geq 0) \quad (t = 1, 2)$$

together with the assumption that the $-v_{it}$ are independent of x_i, η_i and are i.i.d. over time with a known *cdf* F. The distribution F is strictly increasing on the whole line, with a bounded, continuous derivative. Moreover, we have the partitions $x'_{it} = (d_t, z'_{it})$ and $\beta'_0 = (\alpha_0, \gamma'_0)$, where d_t is a time dummy such that $d_1 = 0$ and $d_2 = 1$, and z_i is a continuous random vector with bounded support.

With these assumptions Chamberlain showed that if F is not logistic, then there is a value of α such that identification fails for all β_0 in a neighborhood of $(\alpha, 0)$. This seems puzzling since Manski (1987) proved identification under less restrictive assumptions. He required, however, the presence of an explanatory variable with unbounded support. Indeed, the difference between the identification result of Manski

and the underidentification result of Chamberlain is due to the bounded support for the explanatory variables.

The line between identification and underidentification in this context is very subtle. Under Manski's assumptions identification will fail at $\beta_0' = (\alpha_0, 0)$ even if z_{it} has unbounded support, but there will be identification as long as a component of γ_0 is different from zero. Chamberlain shows that if z_{it} is bounded β_0 is underidentified not only when $\beta_0' = (\alpha_0, 0)$, but also for all β_0 in a neighborhood of $(\alpha_0, 0)$ for a certain value of α_0. So it seems to be a case of local underidentification at zero versus local underidentification in a neighborhood around zero.

The lesson from these findings is the fragility of fixed-T identification results and the special role of the logistic assumption. Chamberlain (1992) also showed that when the support of z_{it} is unbounded (so that identification holds to the exclusion of $\gamma_0 = 0$ from the parameter space) the information bound for β_0 is zero unless F is logistic. Thus, root-N consistent estimation is possible only for the logit model.

Chamberlain's proof can be sketched as follows. In his case $p(x, \eta, \beta_0)$ is

$$p(x, \eta, \beta_0) = \begin{pmatrix} (1 - F_1)(1 - F_2) \\ (1 - F_1) F_2 \\ F_1 (1 - F_2) \\ F_1 F_2 \end{pmatrix}$$

where $F_1 = F(z_1'\gamma_0 + \eta)$ and $F_2 = F(\alpha_0 + z_2'\gamma_0 + \eta)$.

Let $\beta^* = (\alpha, 0)$ and define the 4×4 matrix

$$H(x, \eta_1, ..., \eta_4, \beta^*) = [p(x, \eta_1, \beta^*), ..., p(x, \eta_4, \beta^*)].$$

which does not vary with x when evaluated at β^*.

The proof proceeds by showing that unless $H(x, \eta_1, ..., \eta_4, \beta^*)$ is singular for every α and $\eta_1, ..., \eta_4$, there will be lack of identification for all β_0 in a neighborhood of some β^*. Next it is shown that $H(x, \eta_1, ..., \eta_4, \beta^*)$ can only be singular if F is logistic.

Let us choose a *pmf* $\pi^* = (\pi_1^*, ..., \pi_4^*)'$, $\pi_j^* > 0$, $\sum_{j=1}^4 \pi_j^* = 1$. If for some other *pmf* $\pi_0(x)$ we have

$$H(x, \eta_1, ..., \eta_4, \beta_0) \pi_0(x) = H(x, \eta_1, ..., \eta_4, \beta^*) \pi^*,$$

then the models characterized by $(\beta_0, \pi_0(x))$ and (β^*, π^*) give the same unconditional choice probabilities, hence creating an identification

problem. To rule this out we have to rule out that H is invertible. To see this, suppose that $H(x, \eta_1, ..., \eta_4, \beta^*)$ is nonsingular for some α and $\eta_1, ..., \eta_4$. Since x is bounded, for $\beta_0 \neq \beta^*$ in a neighborhood of β^*, $H(x, \eta_1, ..., \eta_4, \beta_0)$ will also be nonsingular for *all* admissible values of x. We can now define

$$\pi_0(x) = H(x, \eta_1, ..., \eta_4, \beta_0)^{-1} H(x, \eta_1, ..., \eta_4, \beta^*) \pi^*,$$

such that $\pi_{0j}(x) > 0$ for all admissible x. Moreover, since $\iota' H = \iota'$ where ι is a 4×1 vector of ones, we also have $\iota' H^{-1} = \iota'$ and $\iota' \pi_0(x) = 1$. Therefore,

$$\sum_{j=1}^{4} p(x, \eta_j, \beta_0) \pi_{0j}(x) = \sum_{j=1}^{4} p(x, \eta_j, \beta^*) \pi_j^*$$

which implies that β_0 cannot be distinguished from β^*.

The singularity of $H(x, \eta_1, ..., \eta_4, \beta^*)$ requires that

$$\psi_1 [1 - F(\eta)] [1 - F(\alpha + \eta)] + \psi_2 [1 - F(\eta)] F(\alpha + \eta)$$

$$+\psi_3 F(\eta) [1 - F(\alpha + \eta)] + \psi_4 F(\eta) F(\alpha + \eta) = 0$$

for all η and some scalars $\psi_1, ..., \psi_4$ that are not all zero. Taking limits as η tends to $\pm\infty$ gives $\psi_1 = \psi_4 = 0$. Thus we are left with

$$\psi_2 Q(\alpha + \eta) + \psi_3 Q(\eta) = 0$$

where $Q \equiv F/(1 - F)$. For $\eta = 0$ we obtain $\psi_3/\psi_2 = -Q(\alpha)/Q(0)$. Therefore the singularity of H requires that for all α and η we have

$$q(\alpha + \eta) = q(\alpha) + q(\eta) - q(0).$$

This can only happen if the log odd ratios $q \equiv \log Q$ are linear or equivalently if F is logistic.

6. Adjusting the concentrated likelihood

Cox and Reid (1987) considered the general problem of doing inference for a parameter of interest in the absence of knowledge about nuisance parameters. They proposed a first-order adjustment to the concentrated likelihood to take account of the estimation of the nuisance parameters (the *modified profile likelihood*). Their formulation

required information orthogonality between the two types of parameters. That is, that the expected information matrix be block diagonal between the parameters of interest and the nuisance parameters; something that may be achieved by transformation of the latter (Cox and Reid explained how to construct orthogonal parameters). A discussion of orthogonality in the context of panel data models and a Bayesian perspective have been given by Lancaster (2000, 2002). The nature of the adjustment in a fixed effects model and some examples are also discussed in Cox and Reid (1992).

6.1. Orthogonalization

Let $\ell_i(\beta, \eta_i)$ be the log-likelihood for unit i (conditional on x_i and η_i). A strong form of orthogonality arises when for some parameterization of η_i we have

$$\ell_i(\beta, \eta_i) = \ell_{1i}(\beta) + \ell_{2i}(\eta_i), \qquad [43]$$

for in this case the MLE of η_i for given β does not depend on $\beta, \widehat{\eta}_i(\beta) = \widehat{\eta}_i$. The implication is that the MLE of β is unaffected by lack of knowledge of η_i. In this case $\partial^2 \ell_i(\beta, \eta_i) / \partial\beta\partial\eta_i = 0$ for all i. Unfortunately, such factorization does not hold for binary choice models. In contrast, information orthogonality just requires the cross derivatives to be zero on average.

Suppose that a reparameterization is made from (β, η_i) to (β, λ_i) chosen so that β and λ_i are information orthogonal. Thus $\eta_i = \eta(\beta, \lambda_i)$ is chosen such that the reparameterized log likelihood

$$\ell_i^*(\beta, \lambda_i) = \ell_i(\beta, \eta(\beta, \lambda_i)) \qquad [44]$$

satisfies (at true values):

$$E\left(\frac{\partial^2 \ell_i^*(\beta_0, \lambda_i)}{\partial\beta\partial\lambda_i} \mid x_i, \eta_i\right) = 0. \qquad [45]$$

Since we have

$$\frac{\partial \ell_i^*}{\partial \beta} = \frac{\partial \ell_i}{\partial \beta} + \frac{\partial \eta_i}{\partial \beta}\frac{\partial \ell_i}{\partial \eta_i} \qquad [46]$$

and[7]

$$\begin{aligned} & E\left(\frac{\partial^2 \ell_i^*}{\partial\beta\partial\lambda_i} \mid x_i, \eta_i\right) \\ = \; & \frac{\partial \eta_i}{\partial \lambda_i} E\left(\frac{\partial^2 \ell_i}{\partial\beta\partial\eta_i} \mid x_i, \eta_i\right) + \frac{\partial \eta_i}{\partial \lambda_i}\frac{\partial \eta_i}{\partial \beta} E\left(\frac{\partial^2 \ell_i}{\partial \eta_i^2} \mid x_i, \eta_i\right), \end{aligned} \qquad [47]$$

[7]Note that there is a term that vanishes: $(\partial^2 \eta_i / \partial\beta\partial\lambda_i) E(\partial \ell_i / \partial \eta_i \mid x_i, \eta_i) = 0$.

following Cox and Reid (1987) and Lancaster (2002), the function $\eta(\beta, \lambda_i)$ must satisfy the partial differential equations

$$\frac{\partial \eta_i}{\partial \beta} = -E\left(\frac{\partial^2 \ell_i}{\partial \beta \partial \eta_i} \mid x_i, \eta_i\right) / E\left(\frac{\partial^2 \ell_i}{\partial \eta_i^2} \mid x_i, \eta_i\right). \qquad [48]$$

Orthogonal effects in binary choice. Let us now consider the form of information orthogonal fixed effects for model [1]-[2]. These have been obtained by Lancaster (1998, 2000). For binary choice we have

$$E\left(\frac{\partial^2 \ell_i(\beta_0, \eta_i)}{\partial \beta \partial \eta_i} \mid x_i, \eta_i\right) = -\sum_{t=1}^{T} h\left(x_{it}'\beta_0 + \eta_i\right) x_{it} \qquad [49]$$

$$E\left(\frac{\partial^2 \ell_i(\beta_0, \eta_i)}{\partial \eta_i^2} \mid x_i, \eta_i\right) = -\sum_{t=1}^{T} h\left(x_{it}'\beta_0 + \eta_i\right) \qquad [50]$$

where

$$h(r) = \frac{f(r)^2}{F(r)\left[1 - F(r)\right]}. \qquad [51]$$

Since in general [49] is different from zero, β and η_i are not information orthogonal. In view of [48], an orthogonal transformation of the effects will satisfy

$$\frac{\partial \eta_i}{\partial \beta} = -\frac{1}{\sum_{t=1}^{T} h_{it}} \sum_{t=1}^{T} h_{it} x_{it} \qquad [52]$$

where $h_{it} = h\left(x_{it}'\beta + \eta_i\right)$.

Moreover, letting $\phi(r) = h'(r)$ and $\phi_{it} = \phi\left(x_{it}'\beta + \eta_i\right)$, since

$$\frac{\partial^2 \eta_i}{\partial \beta \partial \lambda_i} = -\frac{\partial \eta_i}{\partial \lambda_i}\left[\frac{1}{\sum_{t=1}^{T} h_{it}} \sum\nolimits_{t=1}^{T} \phi_{it}\left(x_{it} + \frac{\partial \eta_i}{\partial \beta}\right)\right]$$

and

$$\frac{\partial^2 \eta_i}{\partial \beta \partial \lambda_i} \Big/ \frac{\partial \eta_i}{\partial \lambda_i} = \frac{\partial}{\partial \beta} \log \left|\frac{\partial \eta_i}{\partial \lambda_i}\right|, \qquad [53]$$

it turns out that

$$\frac{\partial \eta_i}{\partial \lambda_i} = \frac{1}{\sum_{t=1}^{T} h_{it}}. \qquad [54]$$

Hence, Lancaster's orthogonal reparameterization is

$$\lambda_i = \sum_{t=1}^{T} \int_{-\infty}^{x'_{it}\beta+\eta_i} h(r)\, dr. \qquad [55]$$

When $F(r)$ is the logistic distribution $h(r)$ coincides with the logistic density, so that an orthogonal effect for the logit model is

$$\lambda_i = \sum_{t=1}^{T} \Lambda\left(x'_{it}\beta + \eta_i\right). \qquad [56]$$

6.2. Modified profile likelihood

The modified profile log likelihood function of Cox and Reid (1987) can be written as

$$L_M(\beta) = \sum_i \ell_{Mi}(\beta)$$

and

$$\ell_{Mi}(\beta) = \ell_i^*\left(\beta, \widehat{\lambda}_i(\beta)\right) - \frac{1}{2}\log\left[-d^*_{\lambda\lambda i}\left(\beta, \widehat{\lambda}_i(\beta)\right)\right], \qquad [57]$$

where $\widehat{\lambda}_i(\beta)$ is the MLE of λ_i for given β, and $d^*_{\lambda\lambda i}(\beta, \lambda_i) = \partial^2 \ell_i^* / \partial \lambda_i^2$. Intuitively, the role of the second term is to penalize values of β for which the information about the effects is relatively large.

An individual's modified score is of the form

$$d_{Mi}(\beta) = d_{\mathrm{Ci}}(\beta) - \frac{d^*_{\lambda\lambda\beta i}\left(\beta, \widehat{\lambda}_i(\beta)\right) + d^*_{\lambda\lambda\lambda i}\left(\beta, \widehat{\lambda}_i(\beta)\right)\left[\partial\widehat{\lambda}_i(\beta)/\partial\beta\right]}{2d^*_{\lambda\lambda i}\left(\beta, \widehat{\lambda}_i(\beta)\right)} \qquad [58]$$

where $d_{\mathrm{Ci}}(\beta)$ is the standard score from the concentrated likelihood function, $d^*_{\lambda\lambda\beta i}(\beta, \lambda_i) = \partial^3 \ell_i^* / \partial \lambda_i^2 \partial \beta$ and $d^*_{\lambda\lambda\lambda i}(\beta, \lambda_i) = \partial^3 \ell_i^* / \partial \lambda_i^3$.

The function [57] was derived by Cox and Reid as an approximation to the conditional likelihood given $\widehat{\lambda}_i(\beta)$. Their approach was motivated by the fact that in an exponential family model, it is optimal to condition on sufficient statistics for the nuisance parameters, and these can be regarded as the MLE of nuisance parameters chosen in a form to be orthogonal to the parameters of interest. For more general problems the idea was to derive a concentrated likelihood for β conditioned on the MLE $\widehat{\lambda}_i(\beta)$, having ensured via orthogonality that $\widehat{\lambda}_i(\beta)$ changes slowly with β.

Another motivation for using [57] is that the corresponding expected score has a bias of a smaller order of magnitude than the standard ML score (cf. Liang, 1987, McCullagh and Tibshirani, 1990, and Ferguson, Reid, and Cox, 1991). Seen in this way, the objective of the adjustment is to center the concentrated score function to achieve consistency up to a certain order of magnitude in T. Specifically, while the difference between the score with known λ_i and the concentrated score is in general of order $O_p(1)$, the corresponding difference with the modified concentrated score is of order $O_p\left(T^{-1/2}\right)$ (see Appendix). This leads to a bias of order $O\left(T^{-1}\right)$ in the expected modified score, as opposed to $O(1)$ in the concentrated score without modification.

The adjustment in terms of the original parameterization. Cox and Reid's motivation for modifying the concentrated likelihood relied on the orthogonality between common and nuisance parameters. Nevertheless, the *mpl* function [57] can be expressed in terms of the original parameterization. Firstly, note that because of the invariance of MLE $\widehat{\eta}_i(\beta) = \eta(\beta, \widehat{\lambda}_i(\beta))$ and

$$\ell_i^*\left(\beta, \widehat{\lambda}_i(\beta)\right) = \ell_i\left(\beta, \widehat{\eta}_i(\beta)\right). \qquad [59]$$

Next, the term $d^*_{\lambda\lambda i}\left(\beta, \widehat{\lambda}_i(\beta)\right)$ can be calculated as the product of the Fisher information in the (β, η_i) parameterization and the square of the Jacobian of the transformation from (β, η_i) to (β, λ_i) (Cox and Reid, 1987, p. 10). That is, since the second derivatives of ℓ_i^* and ℓ_i are related by the expression

$$\frac{\partial^2 \ell_i^*}{\partial \lambda_i^2} = \frac{\partial^2 \ell_i}{\partial \eta_i^2}\left(\frac{\partial \eta_i}{\partial \lambda_i}\right)^2 + \frac{\partial \ell_i}{\partial \eta_i}\left(\frac{\partial^2 \eta_i}{\partial \lambda_i^2}\right),$$

and $\partial \ell_i / \partial \eta_i$ vanishes at $\widehat{\eta}_i(\beta)$, letting $d_{\eta\eta i}(\beta, \eta_i) = \partial^2 \ell_i / \partial \eta_i^2$ we have

$$d^*_{\lambda\lambda i}\left(\beta, \widehat{\lambda}_i(\beta)\right) = d_{\eta\eta i}\left(\beta, \widehat{\eta}_i(\beta)\right)\left(\frac{\partial \eta_i}{\partial \lambda_i}\,|_{\lambda_i = \widehat{\lambda}_i(\beta)}\right)^2. \qquad [60]$$

Thus, the *mpl* can be written as

$$\ell_{Mi}(\beta) = \ell_i\left(\beta, \widehat{\eta}_i(\beta)\right) - \frac{1}{2}\log\left[-d_{\eta\eta i}\left(\beta, \widehat{\eta}_i(\beta)\right)\right] + \log\left(\frac{\partial \lambda_i}{\partial \eta_i}\,|_{\eta_i = \widehat{\eta}_i(\beta)}\right). \qquad [61]$$

Finally, in view of [48] and [53], the derivative with respect to β of the Jacobian term (the required term for the modified score) can be expressed as

$$\frac{\partial}{\partial\beta}\log\left|\frac{\partial\lambda_i}{\partial\eta_i}\right| = -\frac{\partial}{\partial\eta_i}q_i\left(\beta,\eta_i\right), \qquad [62]$$

where $q_i\left(\beta,\eta_i\right) = -\kappa_{\beta\eta i}\left(\beta,\eta_i\right)/\kappa_{\eta\eta i}\left(\beta,\eta_i\right)$ and

$$\kappa_{\beta\eta i}\left(\beta_0,\eta_i\right) = E\left[\frac{1}{T}d_{\beta\eta i}\left(\beta_0,\eta_i\right) \mid x_i,\eta_i\right] \qquad [63]$$

$$\kappa_{\eta\eta i}\left(\beta_0,\eta_i\right) = E\left[\frac{1}{T}d_{\eta\eta i}\left(\beta_0,\eta_i\right) \mid x_i,\eta_i\right]. \qquad [64]$$

Modified profile likelihood for binary choice. Replacing [54] in [61] we have

$$\ell_{Mi}\left(\beta\right) = \ell_i\left(\beta,\widehat{\eta}_i\left(\beta\right)\right) - \frac{1}{2}\log\left[-d_{\eta\eta i}\left(\beta,\widehat{\eta}_i\left(\beta\right)\right)\right] + \log\left(\sum_{t=1}^{T}\widehat{h}_{it}\left(\beta\right)\right) \qquad [65]$$

where $\widehat{h}_{it}\left(\beta\right) = h\left(x_{it}'\beta + \widehat{\eta}_i\left(\beta\right)\right)$,

$$\ell_i\left(\beta,\eta_i\right) = \sum_{t=1}^{T}\left\{y_{it}\log F_{it} + \left(1-y_{it}\right)\log\left(1-F_{it}\right)\right\}$$

and

$$d_{\eta\eta i}\left(\beta,\eta_i\right) = -\sum_{t=1}^{T}\left[h_{it} - \rho_{it}\left(y_{it}-F_{it}\right)\right] \qquad [66]$$

where $\rho_{it} = \rho\left(x_{it}'\beta+\eta_i\right)$ and

$$\rho\left(r\right) = \frac{f'\left(r\right) - h\left(r\right)\left[1-2F\left(r\right)\right]}{F\left(r\right)\left[1-F\left(r\right)\right]}. \qquad [67]$$

For logit, the MLE $\widehat{\lambda}_i\left(\beta\right)$ for given β solves

$$\widehat{\lambda}_i\left(\beta\right) = \sum_{t=1}^{T}\Lambda\left(x_{it}'\beta + \widehat{\eta}_i\left(\beta\right)\right) = \sum_{t=1}^{T}y_{it} \qquad [68]$$

so that it does not vary with β. Therefore, the likelihood conditioned on $\widehat{\lambda}_i\left(\beta\right)$ coincides with the conditional logit likelihood given a sufficient statistic for the fixed effect discussed in Section 4.1.

For the logistic distribution $\rho(r) = 0$. The modified profile likelihood (*mpl*) for logit is therefore

$$\ell_{Mi}(\beta) = \ell_i(\beta, \widehat{\eta}_i(\beta)) + \frac{1}{2}\log\left(\sum_{t=1}^{T} f_\Lambda\left(x_{it}'\beta + \widehat{\eta}_i(\beta)\right)\right) \qquad [69]$$

where $f_\Lambda(r) = \Lambda(r)[1 - \Lambda(r)]$ is the logistic density and $\ell_{Mi}(\beta)$ is defined for observations such that $\sum_{t=1}^{T} y_{it}$ is not zero or T.[8]

6.3. *Numerical comparisons for Logit and Probit*

Comparisons for the two-period logit model. The *mpl* for logit [69] differs from Andersen's conditional likelihood, and the estimator $\widehat{\beta}_{MML}$ that maximizes the *mpl* is inconsistent for fixed T. Pursuing the example in Section 3, we compare the large-N biases of ML and MML for $T = 2$ and $\Delta x_{i2} = 1$. Thus we are assessing the value of the large-T adjustment in [69] when $T = 2$.

When $T = 2$, for individuals who change state $\widehat{\eta}_i(\beta) = -\beta/2$ so that the second term in [69] becomes

$$\frac{1}{2}\log\left[f_\Lambda(-\beta/2) + f_\Lambda(\beta/2)\right]. \qquad [70]$$

Collecting terms and ignoring constants, the modified profile log-likelihood takes the form

$$\begin{aligned}
\frac{1}{N}\sum_{i=1}^{N}\ell_{Mi}(\beta) &= \frac{1}{N}\sum_{i=1}^{N}\{2d_{10i}\log[1-\Lambda(\beta/2)] + 2d_{01i}\log\Lambda(\beta/2) \\
&\quad + (d_{10i} + d_{01i})\frac{1}{2}(\log\Lambda(\beta/2) + \log[1-\Lambda(\beta/2)])\} \\
&\propto \frac{1}{N}\sum_{i=1}^{N}\{(5d_{10i} + d_{01i})\log[1-\Lambda(\beta/2)] \\
&\quad + (5d_{01i} + d_{10i})\log\Lambda(\beta/2)\} \\
&\propto (5 - 4\widehat{p})\log[1-\Lambda(\beta/2)] + (4\widehat{p} + 1)\log\Lambda(\beta/2) \qquad [71]
\end{aligned}$$

[8] If $\widehat{\eta}_i(\beta) \to \pm\infty$, then $\log\left(\sum_{t=1}^{T} f_\Lambda\left(x_{it}'\beta + \widehat{\eta}_i(\beta)\right)\right)$ tends to $-\infty$ for any β. So observations for individuals that never change state are uninformative about β.

where $d_{10i} = 1(y_{i1} = 1, y_{i2} = 0)$, $d_{01i} = 1(y_{i1} = 0, y_{i2} = 1)$ and $\widehat{p}$ is as defined in [19]. This is maximized at

$$\widehat{\beta}_{MML} = 2\Lambda^{-1}\left(\frac{4\widehat{p}+1}{6}\right) = 2\log\left(\frac{4\widehat{p}+1}{5-4\widehat{p}}\right). \qquad [72]$$

Therefore,

$$\operatorname*{plim}_{N\to\infty} \widehat{\beta}_{MML} = 2\log\left(\frac{4p_0+1}{5-4p_0}\right) = 2\log\left(\frac{4\Lambda(\beta_0)+1}{5-4\Lambda(\beta_0)}\right). \qquad [73]$$

FIGURE 1
Probability limits for a logit model with $T = 2$

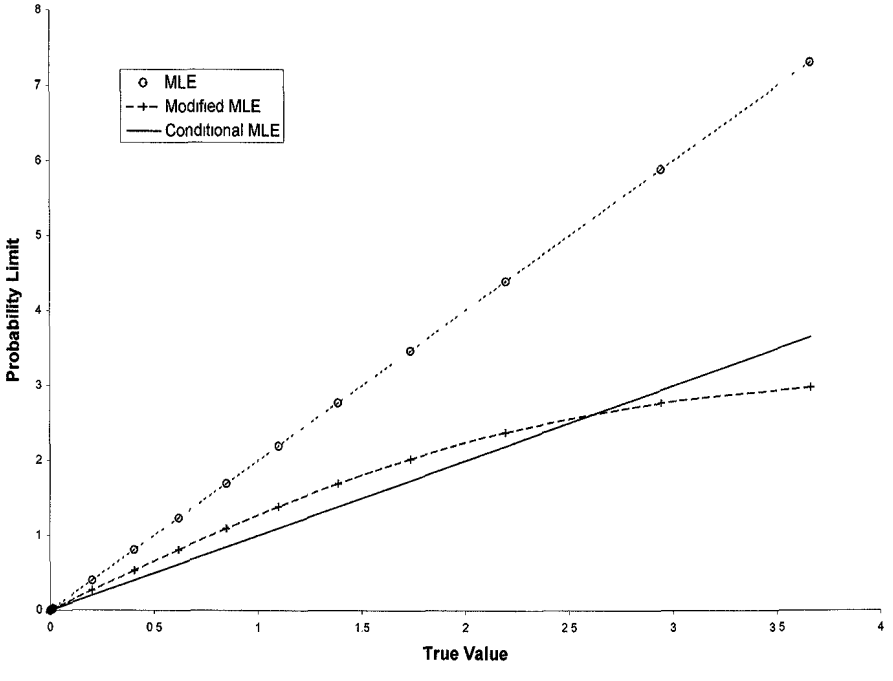

Figure 1 shows the probability limits of MML for positive values of β_0, together with those of ML (the $2\beta_0$ line) and conditional ML (the 45° line) for comparisons.[9] In this example the adjustment produces a surprisingly good improvement given that we are relying on a large T argument with $T = 2$. For example, for $p_0 = 0{,}65$, we have $\beta_0 = 0{,}62$, $\beta_{ML} = 1{,}24$ and $\beta_{MML} = 0{,}81$. Since the MML biases are of order $O\left(1/T^2\right)$, the result suggests that, although the biases are not negligible for $T = 2$, they may be so for values of T as small as 5 or 6.

[9] See McCullagh and Tibshirani (1990, pp. 337-8) for a similar exercise using different adjusted likelihood functions.

Comparisons for the two-period probit model. If $f(r)$ is the standard normal *pdf* we have $h(-r) = h(r)$ and $\rho(-r) = -\rho(r)$. Thus, in the two-period case, $\widehat{h}_{i1} = h[\widehat{\eta}_i(\beta)] = h(-\beta/2) = h(\beta/2)$ and $\widehat{h}_{i2} = h[\beta + \widehat{\eta}_i(\beta)] = h(\beta/2)$. Also $\widehat{F}_{i1} = F(-\beta/2)$ and $\widehat{F}_{i2} = F(\beta/2)$. Finally, $\widehat{\rho}_{i1} = \rho(-\beta/2) = -\rho(\beta/2)$ and $\widehat{\rho}_{i2} = \rho(\beta/2)$.

Therefore, for observations with $y_{i1} + y_{i2} = 1$ we have:

$$\begin{aligned} \ell_{Mi}(\beta) &= \ell_i(\beta, \widehat{\eta}_i(\beta)) \\ &\quad -\frac{1}{2}\log\left[\widehat{h}_{i1} - \widehat{\rho}_{i1}\left(y_{i1} - \widehat{F}_{i1}\right) + \widehat{h}_{i2} - \widehat{\rho}_{i2}\left(y_{i2} - \widehat{F}_{i2}\right)\right] \\ &\quad +\log\left(\widehat{h}_{i1}(\beta) + \widehat{h}_{i2}(\beta)\right) \end{aligned} \tag{74}$$

and

$$\begin{aligned} \ell_{Mi}(\beta) \propto\ & 2\left[d_{10i}\log F(-\beta/2) + d_{01i}\log F(\beta/2)\right] \\ & -\frac{1}{2}d_{10i}\log\left[2h(\beta/2) + 2\rho(\beta/2)F(\beta/2)\right] \\ & -\frac{1}{2}d_{01i}\log\left[2h(\beta/2) + 2\rho(-\beta/2)F(-\beta/2)\right] + \log h(\beta/2). \end{aligned} \tag{75}$$

Next, collecting terms, ignoring constants, averaging over observations with $y_{i1} + y_{i2} = 1$, and using the notation

$$q(r) = 1 + \frac{\rho(r)F(r)}{h(r)} = \frac{\Phi(r)}{1-\Phi(r)} - \frac{r\Phi(r)}{\phi(r)},$$

we have

$$\begin{aligned} & \frac{1}{N_1}\sum_{i=1}^{N_1}\ell_{Mi}(\beta) \\ =\ & (1-\widehat{p})\left[2\log F(-\beta/2) + \frac{1}{2}\log h(\beta/2) - \frac{1}{2}\log q(\beta/2)\right] \\ & +\widehat{p}\left[2\log F(\beta/2) + \frac{1}{2}\log h(\beta/2) - \frac{1}{2}\log q(-\beta/2)\right]. \end{aligned} \tag{76}$$

Thus, the probability limit of $\widehat{\beta}_{MML}$ for probit maximizes the limiting modified log likelihood as follows

$$\begin{aligned} & \operatorname*{plim}_{N\to\infty} \widehat{\beta}_{MML} \\ = \; & \arg\max_{\beta} \left\{ p_0 \left[2\log F(\beta/2) + \frac{1}{2}\log h(\beta/2) - \frac{1}{2}\log q(-\beta/2) \right] \right. \\ & \left. + (1-p_0) \left[2\log F(-\beta/2) + \frac{1}{2}\log h(\beta/2) - \frac{1}{2}\log q(\beta/2) \right] \right\}. \end{aligned} \tag{77}$$

FIGURE 2
Probability limits for a probit model with $T = 2$

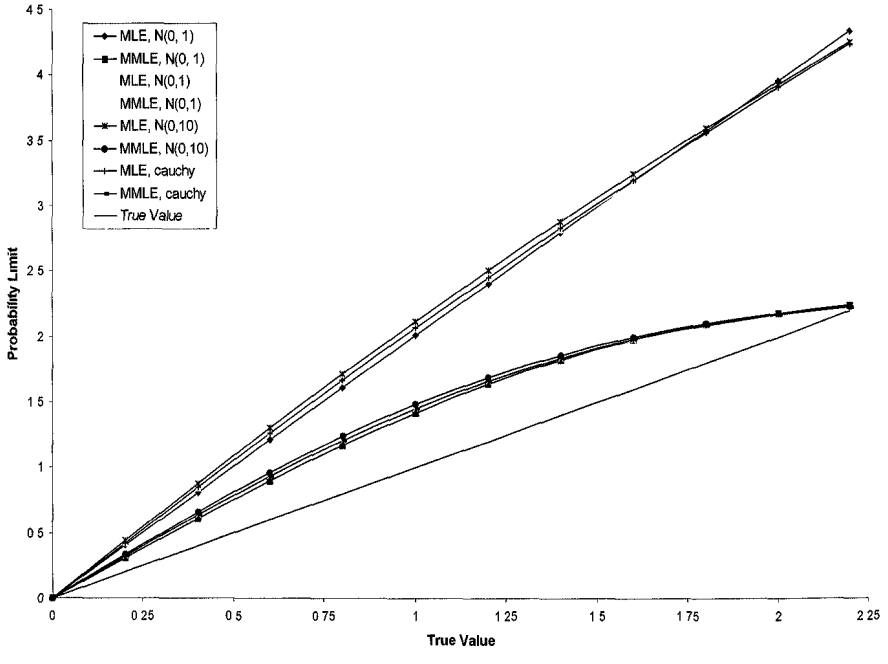

Figure 2 shows the probability limits of probit ML and MML for normally distributed individual effects with variances 0,1, 1, and 10, as well as for Cauchy distributed effects. The range of values of β has been chosen for comparability with Figure 1, in the sense that both figures cover similar intervals of p_0 values. The impact of changing the distribution of the effects is noticeably small for both ML and MML. The adjustment for probit also produces a good improvement given that T is only two, although less so than in the logit case. For exam-

ple, for $\beta_0 = 0{,}60$, the relevant ranges of values are $[1{,}21, 1{,}30]$ for β_{ML}, $[0{,}90, 0{,}96]$ for β_{MML}, and $[0{,}73, 0{,}74]$ for p_0.

7. N and T asymptotics

The panel data literature has probably overemphasized the quest for fixed-T large-N consistent estimation of non-linear models with fixed effects. We have already seen the difficulties that arise in trying to obtain a root-N consistent estimator for a simple static fixed effects probit model. Not surprisingly, the difficulties become even more serious for dynamic binary choice models. In a sense, insisting on fixed T consistency has similarities with (and may be as restrictive as) requiring exactly unbiased estimation in non-linear models. Panels with $T = 2$ are more common in theoretical discussions than in econometric practice. For a micro panel with 7 or 8 time series observations, whether estimation biases are of order $O(1/T)$ or $O\left(1/T^2\right)$ may make all the difference. So it seems useful to consider a wider class of estimation methods than those providing fixed-T consistency, and assess their merits with regard to alternative N and T asymptotic plans. There are multiple possible asymptotic formulations, and it is a matter of judgement to decide which one provides the best approximation for the sample sizes involved in a given application.

Here we consider the asymptotic properties of the estimators that maximize the concentrated likelihood (ML) and the modified concentrated likelihood (MML) when T/N tends to a constant (related results for autoregressive models are in Alvarez and Arellano, 2003, and Hahn and Kuersteiner, 2002).[10]

Consistency. The ML estimator of β can be shown to be consistent as $T \to \infty$ regardless of N using the arguments and the consistency theorem in Amemiya (1985, pp. 270-72). The consistency of MML follows from noting that the concentrated likelihood and the *mpl* converge to the same objective function uniformly in probability as $T \to \infty$.

[10] Since this lecture was first written I have become aware of recent work on double asymptotic formulations for nonlinear fixed effect models by Woutersen (2001) and Li, Lindsay, and Waterman (2002). Moreover, a modified ML estimator for dynamic binary choice models has been developed in Carro (2003) and its properties investigated in simulations and empirical calculations.

Letting $\widehat{\rho}_{it}(\beta) = \rho(x'_{it}\beta + \widehat{\eta}_i(\beta))$ and $\widehat{F}_{it}(\beta) = F(x'_{it}\beta + \widehat{\eta}_i(\beta))$, from (65) we have

$$\begin{aligned} & p\lim_{T\to\infty}\frac{1}{T}\ell_{Mi}(\beta) \qquad [78] \\ = \; & p\lim_{T\to\infty}\frac{1}{T}\ell_i(\beta,\widehat{\eta}_i(\beta)) + p\lim_{T\to\infty}\frac{1}{T}\log\left(\frac{1}{T}\sum_{t=1}^{T}\widehat{h}_{it}(\beta)\right) \\ & -p\lim_{T\to\infty}\frac{1}{T}\log\left(\frac{1}{T}\sum_{t=1}^{T}\left[\widehat{h}_{it}(\beta) - \widehat{\rho}_{it}(\beta)\left(y_{it} - \widehat{F}_{it}(\beta)\right)\right]\right)^{1/2}, \end{aligned}$$

where the convergence is uniform in β in a neighborhood of β_0, and the last two terms vanish.

Asymptotic normality. When $T/N \to c$, $0 < c < \infty$, both ML and MML are asymptotically normal but, unlike MML, the ML estimator has a bias in the asymptotic distribution. An informal calculation of the terms arising in the asymptotic distributions is given in the Appendix. The results are as follows:

$$\left(H'_{NT}V_{NT}^{-1}H_{NT}\right)^{1/2}\sqrt{NT}\left(\widehat{\beta}_{ML} - \beta_0 + \frac{1}{T}H_{NT}^{-1}b_N\right) \xrightarrow{d} \mathcal{N}(0,I) \quad [79]$$

$$\left(H_{NT}^{\dagger\prime}V_{NT}^{-1}H_{NT}^{\dagger}\right)^{1/2}\sqrt{NT}\left(\widehat{\beta}_{MML} - \beta_0\right) \xrightarrow{d} \mathcal{N}(0,I). \quad [80]$$

where

$$\kappa^*_{\beta\lambda\lambda i} = E\left[T^{-1}d^*_{\beta\lambda\lambda i}(\beta_0,\lambda_{i0}) \mid x_i,\lambda_i\right],$$

$$\kappa^*_{\lambda\lambda i} = E\left[T^{-1}d^*_{\lambda\lambda i}(\beta_0,\lambda_{i0}) \mid x_i,\lambda_i\right],$$

$$b_N = \frac{1}{N}\sum_{i=1}^{N}\left(\frac{\kappa^*_{\beta\lambda\lambda i}}{2\kappa^*_{\lambda\lambda i}}\right), \quad [81]$$

$$V_{NT} = \frac{1}{NT}\sum_{i=1}^{N} d^*_{\beta i}(\beta_0,\lambda_{i0})\, d^*_{\beta i}(\beta_0,\lambda_{i0})', \quad [82]$$

$$H_{NT} = \frac{1}{NT}\sum_{i=1}^{N}\frac{\partial}{\partial\beta'}d^*_{\beta i}\left(\beta_0,\widehat{\lambda}_i(\beta_0)\right), \quad [83]$$

and

$$H_{NT}^{\dagger} = \frac{1}{NT}\sum_{i=1}^{N}\frac{\partial}{\partial\beta'}d_{Mi}(\beta_0). \quad [84]$$

Thus, the asymptotic distribution of the ML estimator will contain a bias term unless $\kappa^*_{\beta\lambda\lambda t} = 0$.

8. Concluding remarks

In this paper we have considered ML and modified ML estimators, but the estimation problem can be put more generally in terms of moment conditions in a GMM framework. Fixed-T consistent estimators rely on exactly unbiased moment conditions. When T/N tends to a constant, a GMM estimator from moment conditions with a $O(1/T)$ bias will typically exhibit a bias in the asymptotic distribution, but not if the estimator is based on moment conditions with a $O(1/T^2)$ bias. Thus, in the context of binary choice and other non-linear microeconometric models, a search for optimal orthogonality conditions that are unbiased to order $O(1/T^2)$ or greater seems a useful research agenda.

But do these biases really matter? Heckman (1981) reported a Monte Carlo experiment for ML estimation of a probit model with strictly exogenous variables and fixed effects, $T = 8$ and $N = 100$. Using a random effects estimator as a benchmark, he concluded that the MLE of the common parameters (jointly estimated with the effects) performed well. According to this, it would seem that even for fairly small panels there is not much to be gained from the use of fixed-T unbiased or approximately unbiased orthogonality conditions. For models with only strictly exogenous explanatory variables this may well be the case. But these are models that are found to be too restrictive in many applications.

When modelling panel data, state dependence, predetermined regressors, and serial correlation often matter. Heckman (1981) found that when a lagged dependent variable was included the ML probit estimator performed badly. This is not surprising since similar problems occur with linear autoregressive models. The difference is that while standard tools are available in the literature that ensure fixed T consistency for linear dynamic models, very little is known for dynamic binary choice.[11] This is therefore a promising area of application of asymptotic arguments to both the construction of estimating equations and useful approximations to sampling distributions.

[11] See Keane (1994), Hyslop (1999), Honoré and Kyriazidou (2000), Magnac (2000), Honoré and Lewbel (2002), Arellano and Carrasco (2003), and Arellano and Honoré (2001) for a survey and more references.

Appendix A1

Expansion for the score of the concentrated likelihood. Let us consider a second order expansion of the score of the concentrated likelihood around the true value of the orthogonal effect.

The log likelihood is $\ell_i^*(\beta, \lambda_i)$; its vector of partial derivatives with respect to β is $d_{\beta i}^*(\beta, \lambda_i) = \partial \ell_i^*(\beta, \lambda_i) / \partial \beta$; the concentrated likelihood is $\ell_i^*\left(\beta, \widehat{\lambda}_i(\beta)\right)$ and its score is given by $d_{\beta i}^*\left(\beta, \widehat{\lambda}_i(\beta)\right)$. An approximation at β_0 around the true value λ_{i0} is

$$\begin{aligned} d_{\beta i}^*\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right) &= d_{\beta i}^*(\beta_0, \lambda_{i0}) + d_{\beta\lambda i}^*(\beta_0, \lambda_{i0})\left(\widehat{\lambda}_i(\beta_0) - \lambda_{i0}\right) \quad [A1.1] \\ &\quad + \frac{1}{2} d_{\beta\lambda\lambda i}^*(\beta_0, \lambda_{i0})\left(\widehat{\lambda}_i(\beta_0) - \lambda_{i0}\right)^2 + O_p\left(T^{-1/2}\right) \end{aligned}$$

where $d_{\beta\lambda i}^*(\beta, \lambda_i) = \partial^2 \ell_i^*(\beta, \lambda_i) / \partial\beta\partial\lambda_i$ and $d_{\beta\lambda\lambda i}^*(\beta_0, \lambda_{i0}) = \partial^3 \ell_i^*(\beta, \lambda_i) / \partial\beta\partial\lambda_i^2$. In general, the first three terms are $O_p\left(T^{1/2}\right)$, $O_p\left(T^{1/2}\right)$, and $O_p(1)$, but because of orthogonality $d_{\beta\lambda i}^*(\beta_0, \lambda_{i0})$ is $O_p\left(\sqrt{T}\right)$ as opposed to $O_p(T)$.[12]

Expansion for $\widehat{\lambda}_i(\beta_0) - \lambda_{i0}$. Letting $d_{\lambda i}^*(\beta, \lambda_i) = \partial \ell_i^*(\beta, \lambda_i) / \partial \lambda_i$, the estimator $\widehat{\lambda}_i(\beta_0)$ solves $d_{\lambda i}^*\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right) = 0$. Let us also introduce notation for the terms:

$$\begin{aligned} \kappa_{\lambda\lambda i}^* &\equiv \kappa_{\lambda\lambda}^*(\beta_0, \lambda_{i0}) = E\left[\frac{1}{T} d_{\lambda\lambda i}^*(\beta_0, \lambda_{i0}) \mid x_i, \lambda_i\right] \\ \kappa_{\beta\lambda\lambda i}^* &\equiv \kappa_{\beta\lambda\lambda}^*(\beta_0, \lambda_{i0}) = E\left[\frac{1}{T} d_{\beta\lambda\lambda i}^*(\beta_0, \lambda_{i0}) \mid x_i, \lambda_i\right] \end{aligned}$$

Note that $\kappa_{\lambda\lambda i}^*$ and $\kappa_{\beta\lambda\lambda i}^*$ are individual specific because they depend on λ_{i0}, but they do not depend on the y's.[13] Moreover, from the information matrix identity

$$E\left[\frac{1}{T} d_{\lambda i}^*(\beta_0, \lambda_{i0}) d_{\lambda i}^*(\beta_0, \lambda_{i0}) \mid x_i, \lambda_i\right] = -\kappa_{\lambda\lambda i}^*.$$

[12] Since $\sqrt{T}\left[\frac{1}{T} d_{\beta\lambda i}^*(\beta_0, \lambda_{i0}) - 0\right] = O_p(1)$, we have $d_{\beta\lambda i}^*(\beta_0, \lambda_{i0}) = O_p\left(\sqrt{T}\right)$.

[13] Also $\frac{1}{T} d_{\lambda\lambda i}^*(\beta_0, \lambda_{i0}) = \kappa_{\lambda\lambda}^*(\beta_0, \lambda_{i0}) + O_p\left(\frac{1}{\sqrt{T}}\right)$, which holds as $\sqrt{T}\left(\frac{1}{T} d_{\lambda\lambda i}^*(\beta_0, \lambda_{i0}) - \kappa_{\lambda\lambda}^*(\beta_0, \lambda_{i0})\right) = O_p(1)$.

Expanding $T^{-1/2}d^*_{\lambda i}\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right)$ in the usual way we obtain

$$\begin{aligned} 0 &= \frac{1}{\sqrt{T}} d^*_{\lambda i}\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right) \\ &= \frac{1}{\sqrt{T}} d^*_{\lambda i}(\beta_0, \lambda_{i0}) \\ &\quad + \frac{1}{T} d^*_{\lambda\lambda i}(\beta_0, \lambda_{i0}) \sqrt{T}\left(\widehat{\lambda}_i(\beta_0) - \lambda_{i0}\right) + O_p\left(\frac{1}{\sqrt{T}}\right) \end{aligned}$$

or

$$0 = \frac{1}{\sqrt{T}} d^*_{\lambda i}(\beta_0, \lambda_{i0}) + \kappa^*_{\lambda\lambda i}\sqrt{T}\left(\widehat{\lambda}_i(\beta_0) - \lambda_{i0}\right) + O_p\left(\frac{1}{\sqrt{T}}\right),$$

Hence, also

$$\sqrt{T}\left(\widehat{\lambda}_i(\beta_0) - \lambda_{i0}\right) = -\frac{1}{\kappa^*_{\lambda\lambda i}} \frac{1}{\sqrt{T}} d^*_{\lambda i}(\beta_0, \lambda_{i0}) + O_p\left(\frac{1}{\sqrt{T}}\right), \quad [A1.2]$$

and

$$\begin{aligned} T\left(\widehat{\lambda}_i(\beta_0) - \lambda_{i0}\right)^2 &= \frac{1}{\left(\kappa^*_{\lambda\lambda i}\right)^2} \frac{1}{T} \left[d^*_{\lambda i}(\beta_0, \lambda_{i0})\right]^2 + O_p\left(\frac{1}{\sqrt{T}}\right) \\ &= -\frac{1}{\kappa^*_{\lambda\lambda i}} + O_p\left(\frac{1}{\sqrt{T}}\right). \end{aligned} \quad [A1.3]$$

Combining [A1.1], [A1.2] and [A1.3]:

$$\begin{aligned} d^*_{\beta i}\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right) &= d^*_{\beta i}(\beta_0, \lambda_{i0}) \\ &\quad - \frac{1}{\kappa^*_{\lambda\lambda i}} d^*_{\beta\lambda i}(\beta_0, \lambda_{i0}) \left[\frac{1}{T} d^*_{\lambda i}(\beta_0, \lambda_{i0}) + O_p\left(\frac{1}{T}\right)\right] \\ &\quad + \frac{1}{2} d^*_{\beta\lambda\lambda i}(\beta_0, \lambda_{i0}) \frac{1}{T}\left[-\frac{1}{\kappa^*_{\lambda\lambda i}} + O_p\left(\frac{1}{\sqrt{T}}\right)\right] \\ &\quad + O_p\left(\frac{1}{\sqrt{T}}\right) \\ &= d^*_{\beta i}(\beta_0, \lambda_{i0}) - \frac{1}{\kappa^*_{\lambda\lambda i}} \frac{1}{T} d^*_{\beta\lambda i}(\beta_0, \lambda_{i0}) d^*_{\lambda i}(\beta_0, \lambda_{i0}) \\ &\quad - \frac{\kappa^*_{\beta\lambda\lambda i}}{2\kappa^*_{\lambda\lambda i}} + O_p\left(\frac{1}{\sqrt{T}}\right) \\ &= d^*_{\beta i}(\beta_0, \lambda_{i0}) + \frac{\kappa^*_{\beta\lambda\lambda i}}{2\kappa^*_{\lambda\lambda i}} + O_p\left(\frac{1}{\sqrt{T}}\right) \end{aligned} \quad [A1.4]$$

where we have made use of the facts that due to the orthogonality between λ_i and β we have $d^*_{\beta\lambda i}(\beta_0, \lambda_{i0}) = O_p\left(\sqrt{T}\right)$ and[14]

$$E\left[\frac{1}{T} d^*_{\beta\lambda i}(\beta_0, \lambda_{i0})\, d^*_{\lambda i}(\beta_0, \lambda_{i0})\right] = -\kappa^*_{\beta\lambda\lambda i}.$$

Finally, given the zero-mean property of the score

$$E\left[d^*_{\beta i}(\beta_0, \lambda_{i0}) \mid x_i, \lambda_i\right] = 0$$

the bias of the concentrated score is $O(1)$ and can be written as

$$E\left[d^*_{\beta i}\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right) \mid x_i, \lambda_i\right] = \frac{\kappa^*_{\beta\lambda\lambda i}}{2\kappa^*_{\lambda\lambda i}} + O\left(\frac{1}{T}\right).$$

The remainder is $O\left(T^{-1}\right)$ since the $O_p\left(T^{-1/2}\right)$ terms in the concentrated score have zero mean (cf. Ferguson et al., 1991, p. 290).

Expansion for the score of the modified concentrated likelihood. The *mpf* is given by

$$\ell_{Mi}(\beta) = \ell^*_i\left(\beta, \widehat{\lambda}_i(\beta)\right) - \frac{1}{2}\log\left[-d^*_{\lambda\lambda i}\left(\beta, \widehat{\lambda}_i(\beta)\right)\right]$$

and the *mpf* score

$$d_{Mi}(\beta) = d^*_{\beta i}\left(\beta, \widehat{\lambda}_i(\beta)\right) - \frac{1}{2}\frac{d}{d\beta}\log\left[-d^*_{\lambda\lambda i}\left(\beta, \widehat{\lambda}_i(\beta)\right)\right].$$

[14]Let $f = f(x; \beta, \lambda)$ and write information orthogonality as

$$\int \frac{\partial^2 \log f}{\partial\beta\partial\lambda} f dx = 0.$$

Taking derivatives with respect to λ we obtain:

$$\int \frac{\partial^3 \log f}{\partial\beta\partial\lambda^2} f dx + \int \frac{\partial^2 \log f}{\partial\beta\partial\lambda}\frac{\partial \log f}{\partial\lambda} f dx = 0.$$

Thus,

$$E\left(\frac{\partial^2 \log f}{\partial\beta\partial\lambda}\frac{\partial \log f}{\partial\lambda}\right) = -E\left(\frac{\partial^3 \log f}{\partial\beta\partial\lambda^2}\right).$$

Let us consider the form of the difference between the modified and ordinary concentrated scores at β_0:

$$\begin{aligned} & d_{Mi}(\beta_0) - d^*_{\beta i}\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right) \\ = & \frac{-1}{2\frac{1}{T}d^*_{\lambda\lambda i}\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right)} \\ & \left(\frac{1}{T}d^*_{\lambda\lambda\beta i}\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right) + \frac{1}{T}d^*_{\lambda\lambda\lambda i}\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right)\frac{\partial\widehat{\lambda}_i(\beta_0)}{\partial\beta}\right). \end{aligned}$$

Since $\widehat{\lambda}_i(\beta_0) = \lambda_{i0} + O_p\left(T^{-1/2}\right)$ we have

$$\begin{aligned} & d_{Mi}(\beta_0) - d^*_{\beta i}\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right) \\ = & -\frac{1}{2\kappa^*_{\lambda\lambda i}}\left(\kappa^*_{\beta\lambda\lambda i} + \kappa^*_{\lambda\lambda\lambda i}\frac{\partial\widehat{\lambda}_i(\beta_0)}{\partial\beta}\right) + O_p\left(\frac{1}{\sqrt{T}}\right) \end{aligned}$$

where $\kappa^*_{\lambda\lambda\lambda i} = E\left[T^{-1}d^*_{\lambda\lambda\lambda i}(\beta_0, \lambda_{i0}) \mid x_i, \lambda_i\right]$.

Now, differentiating $d^*_{\lambda i}\left(\beta, \widehat{\lambda}_i(\beta)\right) = 0$ we obtain

$$d^*_{\beta\lambda i}\left(\beta, \widehat{\lambda}_i(\beta)\right) + d^*_{\lambda\lambda i}\left(\beta, \widehat{\lambda}_i(\beta)\right)\frac{\partial\widehat{\lambda}_i(\beta)}{\partial\beta} = 0$$

or

$$\frac{\partial\widehat{\lambda}_i(\beta)}{\partial\beta} = -\frac{d^*_{\beta\lambda i}\left(\beta, \widehat{\lambda}_i(\beta)\right)}{d^*_{\lambda\lambda i}\left(\beta, \widehat{\lambda}_i(\beta)\right)}.$$

Therefore,

$$\frac{\partial\widehat{\lambda}_i(\beta_0)}{\partial\beta} = -\frac{\kappa^*_{\beta\lambda i}}{\kappa^*_{\lambda\lambda i}} + O_p\left(\frac{1}{\sqrt{T}}\right),$$

but because of orthogonality $\kappa^*_{\beta\lambda i} = E\left[T^{-1}d^*_{\beta\lambda i}(\beta_0, \lambda_{i0}) \mid x_i, \lambda_i\right] = 0$, so that $\partial\widehat{\lambda}_i(\beta_0)/\partial\beta$ is $O_p\left(T^{-1/2}\right)$ and

$$d_{Mi}(\beta_0) - d^*_{\beta i}\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right) = -\frac{\kappa^*_{\beta\lambda\lambda i}}{2\kappa^*_{\lambda\lambda i}} + O_p\left(\frac{1}{\sqrt{T}}\right).$$

Finally, combining this result with [A1.4] we obtain

$$d_{Mi}(\beta_0) = d^*_{\beta i}(\beta_0, \lambda_{i0}) + O_p\left(\frac{1}{\sqrt{T}}\right). \qquad [A1.5]$$

Thus, the difference between the concentrated likelihood and the modified concentrated likelihood depends primarily on the value of $\kappa^*_{\beta\lambda\lambda i}$. If $\kappa^*_{\beta\lambda\lambda i} = 0$ the scores from both functions will have biases of the same order of magnitude (Cox and Reid, 1992).

Asymptotic normality of the ML estimator. Let us begin by assuming that, as $T/N \to c$, $0 < c < \infty$, a standard central limit theorem applies to the true score $d^*_{\beta i}(\beta, \lambda_i) = \partial \ell^*_i(\beta, \lambda_i)/\partial\beta$, so that we have

$$V_{NT}^{-1/2} \frac{1}{\sqrt{NT}} \sum_{i=1}^{N} d^*_{\beta i}(\beta_0, \lambda_{i0}) \xrightarrow{d} \mathcal{N}(0, I) \qquad \text{[A1.6]}$$

where $V_{NT} = (NT)^{-1} \sum_{i=1}^{N} d^*_{\beta i}(\beta_0, \lambda_{i0}) d^*_{\beta i}(\beta_0, \lambda_{i0})'$.

Using [A1.4] we can write

$$\frac{1}{\sqrt{NT}} \sum_{i=1}^{N} d^*_{\beta i}\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right) = \frac{1}{\sqrt{NT}} \sum_{i=1}^{N} d^*_{\beta i}(\beta_0, \lambda_{i0}) + \sqrt{\frac{N}{T}} b_N + \sqrt{\frac{N}{T^2}} a_N$$

where $b_N = N^{-1} \sum_{i=1}^{N} \left[\kappa^*_{\beta\lambda\lambda i}/(2\kappa^*_{\lambda\lambda i})\right]$, $a_N = N^{-1} \sum_{i=1}^{N} a_i$, and a_i is an $O_p(1)$ term. Therefore,

$$V_{NT}^{-1/2} \{ \frac{1}{\sqrt{NT}} \sum_{i=1}^{N} d^*_{\beta i}\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right) - \sqrt{\frac{N}{T}} b_N \} \xrightarrow{d} \mathcal{N}(0, I). \qquad \text{[A1.7]}$$

Next, from a first order expansion of the concentrated score around the true value, we obtain

$$H_{NT} \sqrt{NT} \left(\widehat{\beta} - \beta_0\right) = -\frac{1}{\sqrt{NT}} \sum_{i=1}^{N} d^*_{\beta i}\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right) + O_p\left(\frac{1}{\sqrt{NT}}\right) \qquad \text{[A1.8]}$$

where

$$H_{NT} = \frac{1}{NT} \sum_{i=1}^{N} \frac{\partial}{\partial \beta'} d^*_{\beta i}\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right).$$

Combining [A1.7] and [A1.8] we can write

$$V_{NT}^{-1/2} H_{NT} \sqrt{NT} \left(\widehat{\beta} - \beta_0 + \frac{1}{T} H_{NT}^{-1} b_N\right) =$$

$$-V_{NT}^{-1/2} \{ \frac{1}{\sqrt{NT}} \sum_{i=1}^{N} d^*_{\beta i}\left(\beta_0, \widehat{\lambda}_i(\beta_0)\right) - \sqrt{\frac{N}{T}} b_N \} + O_p\left(\frac{1}{\sqrt{NT}}\right).$$

and finally,

$$\left(H'_{NT}V_{NT}^{-1}H_{NT}\right)^{1/2}\sqrt{NT}\left(\widehat{\beta}-\beta_0+\frac{1}{T}H_{NT}^{-1}b_N\right)\xrightarrow{d}\mathcal{N}(0,I).$$

Asymptotic normality of the MML estimator. We now turn to consider the asymptotic distribution of the modified ML estimator as $T/N\to c$, $0<c<\infty$. In view of [A1.5], given [A1.6] we have

$$V_{NT}^{-1/2}\frac{1}{\sqrt{NT}}\sum_{i=1}^{N}d_{Mi}(\beta_0)\xrightarrow{d}\mathcal{N}(0,I). \qquad \text{[A1.9]}$$

Next, from a first order expansion of the modified score around the true value, we obtain

$$H_{NT}^{\dagger}\sqrt{NT}\left(\widehat{\beta}_{MML}-\beta_0\right)=-\frac{1}{\sqrt{NT}}\sum_{i=1}^{N}d_{Mi}(\beta_0)+O_p\left(\frac{1}{\sqrt{NT}}\right) \qquad \text{[A1.10]}$$

where

$$H_{NT}^{\dagger}=\frac{1}{NT}\sum_{i=1}^{N}\frac{\partial d_{Mi}(\beta_0)}{\partial\beta'}.$$

Finally, combining [A1.9] and [A1.10] we can write

$$\begin{aligned}&V_{NT}^{-1/2}H_{NT}^{\dagger}\sqrt{NT}\left(\widehat{\beta}_{MML}-\beta_0\right)\\ =\;&-V_{NT}^{-1/2}\frac{1}{\sqrt{NT}}\sum_{i=1}^{N}d_{Mi}(\beta_0)+O_p\left(\frac{1}{\sqrt{NT}}\right)\end{aligned}$$

and

$$\left(H_{NT}^{\dagger\prime}V_{NT}^{-1}H_{NT}^{\dagger}\right)^{1/2}\sqrt{NT}\left(\widehat{\beta}_{MML}-\beta_0\right)\xrightarrow{d}\mathcal{N}(0,I).$$

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Resumen

Este trabajo examina los métodos existentes para el análisis con datos de panel de modelos de elección binaria con efectos individuales, valorando sus ventajas e inconvenientes. En esta área se necesita más investigación tanto teórica como empírica y el artículo señala varios aspectos que requieren atención. En especial, se ilustra la utilidad de los argumentos asintóticos para obtener condiciones de momentos aproximadamente insesgadas y aproximaciones a las distribuciones muestrales de los estimadores para paneles de distintos tamaños.

Palabras clave: elección binaria, datos de panel, efectos fijos, verosimilitud modificada, correcciones asintóticas.

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