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VARYING-COEFFICIENT PANEL DATA MODEL WITH INTERACTIVE FIXED EFFECTS

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Abstract: We propose a varying-coefficient panel-data model with unobservable multiple interactive fixed effects that are correlated with the regressors. We approximate each coefficient function using B-splines, and propose a robust nonlinear iteration scheme based on the least squares method to estimate the coefficient functions of interest. We also establish the asymptotic theory of the resulting estimators under certain regularity assumptions, including the consistency, convergence rate, and asymptotic distributions. To construct the pointwise confidence intervals for the coefficient functions, we propose a residual-based block bootstrap method that reduces the computational burden and avoids accumulative errors. We extend our proposed procedure to partially linear varying-coefficient panel-data models with unobservable multiple interactive fixed effects, and examine the problem of constant coefficients versus function coefficients. Simulation studies and a real-data analysis are used to assess the performance of the proposed methods.

Key words and phrases: Bootstrap, B-spline, hypothesis testing, interactive fixed effect, panel data, partially linear varying-coefficient model, varying-coefficient model.

1. Introduction

Panel-data models typically incorporate individual and time effects to control the heterogeneity in the cross-section and across periods. Panel-data analysis has attracted considerable attention in the literature. The methodology for a parametric panel-data analysis is relatively mature; see, for example, Arellano (2003), Hsiao (2003), Baltagi (2005), and the references therein. Individual and time effects may enter the model additively, or they can interact multiplicatively, leading to the so-called interactive effects or a factor structure. Panel-data models with interactive fixed effects are a useful modeling paradigm. In macroeconomics, incorporating interactive effects can account for the heterogenous effects of unobservable common shocks, while the regressors can be inputs, such as labor and

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capital. Panel-data models with interactive fixed effects are used to incorporate unmeasured skills or unobservable characteristics, or to study the individual wage rate (Su and Chen (2013)). In finance, a combination of unobserved factors and observed covariates can explain the excess returns of assets. Bai (2009) considered the following linear panel-data model with interactive fixed effects:

$$Y_{it} = X_{it}^{\tau}\boldsymbol{\beta} + \lambda_i^{\tau} F_t + \varepsilon_{it}, \quad i = 1, \ldots, N, \quad t = 1, \ldots, T, \tag{1.1}$$

where X_{it} is a $p \times 1$ vector of observable regressors, $\boldsymbol{\beta}$ is a $p \times 1$ vector of unknown coefficients, λ_i is an $r \times 1$ vector of factor loadings, F_t is an $r \times 1$ vector of common factors, such that $\lambda_i^{\tau} F_t = \lambda_{i1} F_{1t} + \cdots + \lambda_{ir} F_{rt}$, and ε_{it} are idiosyncratic errors. In this model, λ_i, F_t, and ε_{it} are unobserved, and the dimension r of the factor loadings does not depend on the cross-section size N or the time series length T.

A number of researchers have developed statistical methods to study panel-data models with interactive fixed effects. For example, Holtz-Eakin, Newey and Rosen (1988) estimated model (1.1) using quasi-differencing and lagged variables as instruments. Their approach, however, rules out time constant regressors. Coakley, Fuertes and Smith (2002) studied model (1.1) by augmenting the regression of Y on X with the principal components of the ordinary least squares residuals. However, Pesaran (2006) showed that this method is inconsistent unless X_{it} and λ_i tend to be uncorrelated or fully correlated as N tends to infinity. As an alternative, Pesaran (2006) developed a correlated common effects (CCE) estimator, in which model (1.1) is augmented with the cross-sectional averages of X_{it}. Although Pesaran's estimator is consistent, it does not allow for time-invariant individual regressors. Ahn, Lee and Schmidt (2001) developed a generalized method of moments (GMM) estimator for model (1.1). Their estimator is more efficient than the least squares estimator under a fixed T. However, being able to identify their estimator requires that X_{it} is correlated with λ_i, and it is impossible to test for the interactive random effects assumption. Bai (2009) studied the identification, consistency, and limiting distribution of the principal component analysis (PCA) estimators, showing that they are $\sqrt{NT}$-consistent. Bai and Li (2014) investigated the maximum likelihood estimation of model (1.1). Wu and Li (2014) conducted several tests for the existence of individual effects and time effects in model (1.1). Li, Qian and Su (2016) studied the estimation and inference of common structural breaks in panel-data models with interactive fixed effects using Lasso-type methods. More studies can be found in Moon and Weidner (2017), Lee, Moon and Weidner (2012), Su and Chen (2013), Moon and Weidner (2015), Lu and Su (2016), and many others.

Note that the aforementioned works focus on linear specifications of the regression relationships in panel-data models with interactive fixed effects. A natural extension of model (1.1) is to consider the following varying-coefficient panel-data model with interactive fixed effects:

$$Y_{it} = X_{it}^{\tau}\boldsymbol{\beta}(U_{it}) + \lambda_i^{\tau} F_t + \varepsilon_{it}, \quad i = 1, \ldots, N, \quad t = 1, \ldots, T, \tag{1.2}$$

where $\boldsymbol{\beta}(\cdot) = (\beta_1(\cdot), \ldots, \beta_p(\cdot))^{\tau}$ is a $p \times 1$ vector of unknown coefficient functions to be estimated. We allow for $\{X_{it}\}$ and/or $\{U_{it}\}$ to be correlated with $\{\lambda_i\}$ alone or with $\{F_t\}$ alone, or simultaneously correlated with $\{\lambda_i\}$ and $\{F_t\}$, or correlated with an unknown correlation structure. In fact, X_{it} can be a nonlinear function of λ_i and F_t. Hence, model (1.2) is a fixed-effects model, and assumes an interactive fixed-effects linear model for each fixed time t, but allows the coefficients to vary with the covariate U_{it}. This model is attractive because it has an intuitive interpretation, while retaining the unobservable multiple interactive fixed effects, general nonparametric characteristics, and explanatory power of the linear panel-data model.

Model (1.2) is fairly general, and encompasses various panel-data models as special cases. If $X_{it} \equiv 1$ and $p = 1$, model (1.2) reduces to the nonparametric panel-data model with interactive fixed effects, which has received much attention in recent years. Huang (2013) studied the local linear estimation of such models. Su and Jin (2012) extended the CCE method of Pesaran (2006) from a linear model to a nonparametric model using the method of sieves. Jin and Su (2013) constructed a nonparametric test for poolability in nonparametric regression models with interactive fixed effects. Su, Jin and Zhang (2015) proposed a consistent nonparametric test for linearity in a large-dimensional panel-data model with interactive fixed effects.

If $r = 1$ and $F_t \equiv 1$, model (1.2) reduces to the fixed individual effects panel-data varying-coefficient model:

$$Y_{it} = X_{it}^{\tau}\boldsymbol{\beta}(U_{it}) + \lambda_i + \varepsilon_{it}.$$

This model has also been widely studied in the literature. For example, Sun, Carroll and Li (2009) considered estimations using a local linear regression and kernel-based weights. Li, Chen and Gao (2011) considered a nonparametric time varying-coefficient model with fixed effects under the assumption of cross-sectional independence, and proposed methods for estimating the trend function and coefficient functions. Rodriguez-Poo and Soberon (2014) proposed a new technique to estimate the varying-coefficient functions based on the first-order differences and a local linear regression. Rodriguez-Poo and Soberon (2015) investigated the model using the mean transformation technique and a local linear regression. Li et al. (2015) considered variable selection for the model using the basis function approximations and the group nonconcave penalized functions. Malikov, Kumbhakar and Sun (2016) considered the problem of a varying-coefficient panel-data model in the presence of endogenous selectivity and fixed effects. In addition, if $\lambda_i \equiv 0$ or $F_t \equiv 0$, model (1.2) reduces to the varying-coefficient model with panel data. For the development of this model, refer to Chiang, Rice and Wu (2001), Huang, Wu and Zhou (2002), Huang, Wu and Zhou (2004), Xue and Zhu (2007), Cai (2007), Cai and Li (2008), Wang, Li and Huang (2008), Wang and Xia (2009), and Noh and Park (2010). Note, however, that most of these studies focus on a "large N small T" setting.

Despite the rich literature on panel data models with interactive fixed effects, to the best of our knowledge, there are few works on varying-coefficient panel-data models with interactive fixed effects. As such, the main goals of this study are to estimate the coefficient functions $\boldsymbol{\beta}(\cdot)$, and to establish the asymptotic theory for varying-coefficient panel-data models with interactive fixed effects when both N and T tend to infinity and there exist serial or cross-sectional correlations and heteroskedasticities of unknown form in ε_{it}. To achieve these goals, we first apply the B-spline expansion to estimate the smooth functions in model (1.2), owing to its simplicity. We then introduce a novel iterative least squares procedure to estimate the coefficient functions and the factor loadings, and derive some asymptotic properties for the proposed estimators. Nevertheless, the existence of the unobservable interactive fixed effects and the weak correlations and heteroskedasticities of unknown form in both dimensions make the estimation procedure and the asymptotic theory much more complicated than those in Huang, Wu and Zhou (2002). To apply the asymptotic normality to construct the pointwise confidence intervals for the coefficient functions, we need consistent estimators of the asymptotic biases and variances. To reduce the computational burden and to avoid accumulative errors, we propose a residual-based block bootstrap procedure to construct these confidence intervals.

Moreover, we extend the proposed estimation procedure to include partially linear varying coefficient models with interactive fixed effects, and show that the convergence rate for the estimation of the parametric components is of order $O_P((NT)^{-1/2})$. To determine whether a varying-coefficient model or partially linear varying-coefficient model is appropriate, we propose a test statistic to test between the two alternatives in practice. Numerical results confirm that our proposed estimation and testing procedures work well in a wide range of settings.

The remainder of the paper is organized as follows. In Section 2, we propose an estimation procedure for the coefficient functions and provide a robust iteration algorithm under the identification restrictions. In Section 3, we establish the asymptotic theory of the resulting estimators under some regularity assumptions as both N and T tend to infinity. In Section 4, we develop a residual-based block bootstrap procedure to construct the pointwise confidence intervals for the coefficient functions. In Section 5, we extend the estimation procedure to partially linear varying coefficient models and establish the asymptotic distribution of the estimator. In Section 6, a test statistic and the bootstrap procedure are developed. Finally, we conclude the paper in Section 7. Technical details are given in the online Supplementary Material, along with simulation studies and a real application to demonstrate the efficacy of our proposed methods.

2. Methodology

To estimate the coefficient functions $\beta_k(\cdot)$, for $1 \le k \le p$, we consider the widely used B-spline approximations. Let $B_k(u) = (B_{k1}(u), \ldots, B_{kL_k}(u))^\tau$ be the $(m+1)$th-order B-spline basis functions, where $L_k = l_k + m + 1$ is the number of basis functions in approximating $\beta_k(u)$, l_k is the number of interior knots for $\beta_k(\cdot)$, and m is the degree of the spline. The interior knots of the splines can be either equally spaced or placed on the sample number of observations between any two adjacent knots. With the above basis functions, the coefficient functions $\beta_k(u)$ can be approximated by

$$\beta_k(u) \approx \sum_{l=1}^{L_k} \gamma_{kl} B_{kl}(u), \quad k = 1, \ldots, p, \tag{2.1}$$

where γ_{kl} are the coefficients, and L_k represent the smoothing parameters, selected using "leave-one-subject-out" cross-validation.

Substituting (2.1) into model (1.2), we have the following approximation:

$$Y_{it} \approx \sum_{k=1}^{p} \sum_{l=1}^{L_k} \gamma_{kl} X_{it,k} B_{kl}(U_{it}) + \lambda_i^\tau F_t + \varepsilon_{it},\ i = 1, \ldots, N,\ t = 1, \ldots, T. \tag{2.2}$$

Model (2.2) is a standard linear regression model with interactive fixed effects.

Because each coefficient function $\beta_k(u)$ in model (1.2) is characterized by $\boldsymbol{\gamma}_k = (\gamma_{k1}, \ldots, \gamma_{kL_k})^\tau$, model (2.2) cannot be estimated directly, owing to the unobservable multiple interactive fixed effects. In what follows, we propose a robust nonlinear iteration scheme based on the least squares method to estimate the coefficient functions and deal with these fixed effects.

For the sake of convenience, we use vectors and matrices to present the model and perform the analysis. Let $\boldsymbol{Y}_i = (Y_{i1}, \ldots, Y_{iT})^\tau, \boldsymbol{F} = (F_1, \ldots, F_T)^\tau$, $\boldsymbol{\varepsilon}_i = (\varepsilon_{i1}, \ldots, \varepsilon_{iT})^\tau$, and $\Lambda = (\lambda_1, \ldots, \lambda_N)^\tau$ be an $N \times r$ matrix. Let

$$\boldsymbol{B}(u) = \begin{pmatrix} B_{11}(u) \cdots B_{1L_1}(u) & 0 \cdots 0 & 0 & \cdots & 0 \\ \vdots & \vdots & & \vdots & \\ 0 \quad \cdots \quad 0 & 0 \cdots 0 & B_{p1}(u) & \cdots & B_{pL_p}(u) \end{pmatrix},$$

$R_{it} = (X_{it}^\tau \boldsymbol{B}(U_{it}))^\tau$, and $\boldsymbol{R}_i = (R_{i1}, \ldots, R_{iT})^\tau$. Furthermore, let $\boldsymbol{\gamma} = (\boldsymbol{\gamma}_1^\tau, \ldots, \boldsymbol{\gamma}_p^\tau)^\tau$, where $\boldsymbol{\gamma}_k = (\gamma_{k1}, \ldots, \gamma_{kL_k})^\tau$. Then, model (2.2) can be rewritten as

$$\boldsymbol{Y}_i \approx \boldsymbol{R}_i \boldsymbol{\gamma} + \boldsymbol{F}\lambda_i + \boldsymbol{\varepsilon}_i, \quad i = 1, \ldots, N.$$

Owing to potential correlations between the unobservable effects and the regressors, we treat F_t and λ_i as the fixed-effects parameters to be estimated. To ensure the identifiability of the coefficient function $\boldsymbol{\beta}(\cdot) = (\beta_1(\cdot), \ldots, \beta_p(\cdot))^\tau$, we follow Bai (2009) and impose the following identification restrictions:

$$\frac{\boldsymbol{F}^\tau \boldsymbol{F}}{T} = I_r \quad \text{and} \quad \Lambda^\tau \Lambda = \text{diagonal}. \tag{2.3}$$

These two restrictions uniquely determine Λ and $\boldsymbol{F}$. We then define the objective function as

$$Q(\boldsymbol{\gamma}, \boldsymbol{F}, \Lambda) = \sum_{i=1}^N (\boldsymbol{Y}_i - \boldsymbol{R}_i\boldsymbol{\gamma} - \boldsymbol{F}\lambda_i)^\tau (\boldsymbol{Y}_i - \boldsymbol{R}_i\boldsymbol{\gamma} - \boldsymbol{F}\lambda_i), \tag{2.4}$$

subject to constraint (2.3). Taking partial derivatives of (2.4) with respect to λ_i and setting them equal to zero, we have

$$\widetilde{\lambda}_i = (\boldsymbol{F}^\tau \boldsymbol{F})^{-1} \boldsymbol{F}^\tau (\boldsymbol{Y}_i - \boldsymbol{R}_i\boldsymbol{\gamma}) = T^{-1} \boldsymbol{F}^\tau (\boldsymbol{Y}_i - \boldsymbol{R}_i\boldsymbol{\gamma}). \tag{2.5}$$

Replacing λ_i in (2.4) with (2.5), we have

$$\begin{aligned} Q(\boldsymbol{\gamma}, \boldsymbol{F}) &= \sum_{i=1}^{N} (\boldsymbol{Y}_i - \boldsymbol{R}_i\boldsymbol{\gamma} - \boldsymbol{F}\widetilde{\lambda}_i)^\tau (\boldsymbol{Y}_i - \boldsymbol{R}_i\boldsymbol{\gamma} - \boldsymbol{F}\widetilde{\lambda}_i) \\ &= \sum_{i=1}^{N} (\boldsymbol{Y}_i - \boldsymbol{R}_i\boldsymbol{\gamma})^\tau M_{\boldsymbol{F}} (\boldsymbol{Y}_i - \boldsymbol{R}_i\boldsymbol{\gamma}), \end{aligned}$$

where $M_{\boldsymbol{F}} = I_T - \boldsymbol{F}(\boldsymbol{F}^\tau\boldsymbol{F})^{-1}\boldsymbol{F}^\tau = I_T - \boldsymbol{F}\boldsymbol{F}^\tau/T$ is a projection matrix. For each given $\boldsymbol{F}$, if $\sum_{i=1}^N \boldsymbol{R}_i^\tau M_{\boldsymbol{F}} \boldsymbol{R}_i$ is invertible, the least squares estimator of $\boldsymbol{\gamma}$ can be uniquely obtained by minimizing $Q(\boldsymbol{\gamma}, \boldsymbol{F})$, as follows:

$$\hat{\boldsymbol{\gamma}}(\boldsymbol{F}) = \left(\sum_{i=1}^{N} \boldsymbol{R}_i^\tau M_{\boldsymbol{F}} \boldsymbol{R}_i \right)^{-1} \sum_{i=1}^{N} \boldsymbol{R}_i^\tau M_{\boldsymbol{F}} \boldsymbol{Y}_i. \tag{2.6}$$

Because the least squares estimator (2.6) of $\boldsymbol{\gamma}$ depends on the unknown common factors $\boldsymbol{F}$, the final solution of $\boldsymbol{\gamma}$ can be obtained by iteration between $\boldsymbol{\gamma}$ and $\boldsymbol{F}$ using the following nonlinear equations:

$$\hat{\boldsymbol{\gamma}} = \left(\sum_{i=1}^{N} \boldsymbol{R}_i^\tau M_{\hat{\boldsymbol{F}}} \boldsymbol{R}_i \right)^{-1} \sum_{i=1}^{N} \boldsymbol{R}_i^\tau M_{\hat{\boldsymbol{F}}} \boldsymbol{Y}_i, \tag{2.7}$$

$$\hat{\boldsymbol{F}} V_{NT} = \left[\frac{1}{NT} \sum_{i=1}^{N} (\boldsymbol{Y}_i - \boldsymbol{R}_i\hat{\boldsymbol{\gamma}})(\boldsymbol{Y}_i - \boldsymbol{R}_i\hat{\boldsymbol{\gamma}})^\tau \right] \hat{\boldsymbol{F}}, \tag{2.8}$$

where V_{NT} is a diagonal matrix consisting of the r largest eigenvalues of the matrix $(NT)^{-1}\sum_{i=1}^N (\boldsymbol{Y}_i - \boldsymbol{R}_i\hat{\boldsymbol{\gamma}})(\boldsymbol{Y}_i - \boldsymbol{R}_i\hat{\boldsymbol{\gamma}})^\tau$, arranged in decreasing order. As noted by Bai (2009), the iterated solution is somewhat sensitive to the initial values. Bai (2009) proposed starting with either the least squares estimator of $\boldsymbol{\gamma}$ or the principal components estimate of $\boldsymbol{F}$. From the numerical studies in the Supplementary Material, we find that the procedure is more robust when the principal components estimator of $\boldsymbol{F}$ is used for the initial values. In general, poor initial values result in an exceptionally large number of iterations. By (2.5), (2.7), and (2.8), we have

$$\hat{\Lambda} = (\hat{\lambda}_1, \ldots, \hat{\lambda}_N)^\tau = T^{-1}\Big(\hat{\boldsymbol{F}}^\tau (\boldsymbol{Y}_1 - \boldsymbol{R}_1\hat{\boldsymbol{\gamma}}), \ldots, \hat{\boldsymbol{F}}^\tau (\boldsymbol{Y}_N - \boldsymbol{R}_N\hat{\boldsymbol{\gamma}}) \Big)^\tau. \tag{2.9}$$

Once we obtain the estimator $\hat{\boldsymbol{\gamma}} = (\hat{\boldsymbol{\gamma}}_1^\tau, \ldots, \hat{\boldsymbol{\gamma}}_p^\tau)^\tau$ of $\boldsymbol{\gamma}$ with $\hat{\boldsymbol{\gamma}}_k = (\hat{\gamma}_{k1}, \ldots,$

$\hat{\gamma}_{kL_k})^\tau$, for $k = 1, \ldots, p$, we can estimate $\beta_k(u)$ as

$$\hat{\beta}_k(u) = \sum_{l=1}^{L_k} \hat{\gamma}_{kl} B_{kl}(u), \quad k = 1, \ldots, p.$$

In what follows, we present a robust iteration algorithm for estimating the parameters $(\boldsymbol{\gamma}, \boldsymbol{F}, \Lambda)$.

Step 1. Obtain an initial estimator $(\hat{\boldsymbol{F}}, \hat{\Lambda})$ of $(\boldsymbol{F}, \Lambda)$.

Step 2. Given $\hat{\boldsymbol{F}}$ and $\hat{\Lambda}$, compute $\hat{\boldsymbol{\gamma}}(\hat{\boldsymbol{F}}, \hat{\Lambda}) = \left(\sum_{i=1}^N \boldsymbol{R}_i^\tau \boldsymbol{R}_i \right)^{-1} \sum_{i=1}^N \boldsymbol{R}_i^\tau (\boldsymbol{Y}_i - \hat{\boldsymbol{F}} \hat{\lambda}_i)$.

Step 3. Given $\hat{\boldsymbol{\gamma}}$, compute $\hat{\boldsymbol{F}}$ according to (2.8) (multiplied by $\sqrt{T}$, owing to the restriction that $\boldsymbol{F}^\tau \boldsymbol{F}/T = I_r$), and calculate $\hat{\Lambda}$ using formula (2.9).

Step 4. Repeat Steps 2 and 3 until $(\hat{\boldsymbol{\gamma}}, \hat{\boldsymbol{F}}, \hat{\Lambda})$ satisfy the given convergence criterion.

3. Regularity Assumptions and Asymptotic Properties

To derive asymptotic properties for the proposed estimators, we let $\mathcal{F} \equiv \{\boldsymbol{F} : \boldsymbol{F}^\tau \boldsymbol{F}/T = I_r\}$ and

$$D(\boldsymbol{F}) = \frac{1}{NT} \sum_{i=1}^N \boldsymbol{R}_i^\tau M_{\boldsymbol{F}} \boldsymbol{R}_i - \frac{1}{T} \left[\frac{1}{N^2} \sum_{i=1}^N \sum_{j=1}^N \boldsymbol{R}_i^\tau M_{\boldsymbol{F}} \boldsymbol{R}_j a_{ij} \right],$$

where $a_{ij} = \lambda_i^\tau (\Lambda^\tau \Lambda/N)^{-1} \lambda_j$. To obtain a unique estimator of $\boldsymbol{\gamma}$ with probability tending to one, we require that the first term of $D(\boldsymbol{F})$ on the right-hand side is positive-definite when $\boldsymbol{F}$ is observable. The presence of the second term is because of the unobservable $\boldsymbol{F}$ and Λ. The reason for this particular form is the nonlinearity of the interactive effects (see Bai (2009)).

3.1. Regularity assumptions

In this section, we introduce a definition and present some regularity assumptions, which we use to establish the asymptotic theory of the resulting estimators.

Definition 1. Let $\mathcal{H}_d$ define the collection of all functions on the support $\mathcal{U}$ whose mth-order derivative satisfies the Hölder condition of order ν, with $d \equiv m + \nu$, where $0 < \nu \leq 1$. That is, for each $h \in \mathcal{H}_d$, there exists a constant $M_0 \in (0, \infty)$, such that $|h^{(m)}(u) - h^{(m)}(v)| \leq M_0 |u - v|^\nu$, for any $u, v \in \mathcal{U}$.

(A1) The random variable X_{it} is independent and identically distributed (i.i.d.) across the N individuals, and there exists a positive M, such that $|X_{it,k}| \leq M < \infty$, for all $k = 1, \ldots, p$. We further assume that $\{X_{it} : 1 \leq t \leq T\}$ is strictly stationary for each i. The eigenvalues $\rho_1(u) \leq \cdots \leq \rho_p(u)$ of $\Omega(u) = E(X_{it}X_{it}^{\tau}|U_{it} = u)$ are bounded away from zero and ∞ uniformly over $u \in \mathcal{U}$; that is, there exist positive constants ρ_0 and ρ^*, such that $0 < \rho_0 \leq \rho_1(u) \leq \cdots \leq \rho_p(u) \leq \rho^* < \infty$, for $u \in \mathcal{U}$.

(A2) The observation variables U_{it} are chosen independently according to a distribution F_U on the support $\mathcal{U}$. Moreover, the density function of U, $f_U(u)$, is uniformly bounded away from zero and ∞, and continuously differentiable uniformly over $u \in \mathcal{U}$.

(A3) $\beta_k(u) \in \mathcal{H}_d$, for all $k = 1, \ldots, p$.

(A4) Let $u_{k1}, \ldots, u_{kl_k}$ be the interior knots of the kth coefficient function over $u \in \mathcal{U} = [U_0, U_1]$, for $k = 1, \ldots, p$. Furthermore, let $u_{k0} = U_0$ and $u_{k(l_k+1)} = U_1$. There exists a positive constant C_0, such that

$$\frac{h_k}{\min_{1\leq i\leq l_k} h_{ki}} \leq C_0 \quad \text{and} \quad \frac{\max_{1\leq k\leq p} h_{ki}}{\min_{1\leq k\leq p} h_{ki}} \leq C_0,$$

where $h_{ki} = u_{ki} - u_{k(i-1)}$ and $h_k = \max_{1\leq i\leq l_k+1} h_{ki}$.

(A5) Suppose that $\inf\limits_{\boldsymbol{F}\in\mathcal{F}} D(\boldsymbol{F}) > 0$.

(A6) $E\|F_t\|^4 \leq M$ and $\sum_{t=1}^{T} F_tF_t^{\tau}/T \xrightarrow{P} \Sigma_F > 0$, for some $r \times r$ matrix Σ_F, as $T \to \infty$, where "$\xrightarrow{P}$" denotes convergence in probability.

(A7) $E\|\lambda_i\|^4 \leq M$ and $\Lambda^{\tau}\Lambda/N \xrightarrow{P} \Sigma_\Lambda > 0$, for some $r\times r$ matrix Σ_Λ, as $N \to \infty$.

(A8) (i) Suppose that ε_{it} are independent of X_{js}, U_{js}, λ_j, and F_s, for all i, t, j, and s with zero mean and $E(\varepsilon_{it})^8 \leq M$.

(ii) Let $\sigma_{ij,ts} = E(\varepsilon_{it}\varepsilon_{js})$. $|\sigma_{ij,ts}| \leq \rho_{ij}$ for all (t, s), and $|\sigma_{ij,ts}| \leq \varrho_{ts}$ for all (i, j), such that

$$\frac{1}{N}\sum_{i,j=1}^{N} \rho_{ij} \leq M, \quad \frac{1}{T}\sum_{t,s=1}^{T} \varrho_{ts} \leq M, \quad \frac{1}{NT}\sum_{i,j=1}^{N}\sum_{t,s=1}^{T} |\sigma_{ij,ts}| \leq M.$$

The smallest and largest eigenvalues of $\Omega_i = E(\boldsymbol{\varepsilon}_i\boldsymbol{\varepsilon}_i^{\tau})$ are bounded uniformly for all i and t, where $\boldsymbol{\varepsilon}_i = (\varepsilon_{i1}, \ldots, \varepsilon_{iT})^{\tau}$.

(iii) For every (t,s), $E\left|N^{-1/2}\sum_{i=1}^{N}[\varepsilon_{it}\varepsilon_{is}-E(\varepsilon_{it}\varepsilon_{is})]\right|^4 \le M$.

(iv) Moreover, we assume that $T^{-2}N^{-1}\sum_{t,s,u,v}\sum_{i,j}|\mathrm{cov}(\varepsilon_{it}\varepsilon_{is},\varepsilon_{ju}\varepsilon_{jv})| \le M$ and $T^{-1}N^{-2}\sum_{t,s}\sum_{i,j,m,l}|\mathrm{cov}(\varepsilon_{it}\varepsilon_{jt},\varepsilon_{ms}\varepsilon_{ls})| \le M$.

(A9) $\limsup_{N,T}(\max_k L_k/\min_k L_k) < \infty$.

Assumptions (A1)–(A4) are mild conditions that can be validated in many practical situations. These conditions have been widely assumed in studies on varying-coefficient models with repeated measurements, such as those of Huang, Wu and Zhou (2002), Huang, Wu and Zhou (2004), and Wang, Li and Huang (2008). Assumption (A5) is an identification condition for $\boldsymbol{\gamma}$, and $\boldsymbol{\gamma}$ can be uniquely determined by (2.7) if $D(\boldsymbol{F})$ is positive-definite. Assumptions (A6) and (A7) imply the existence of r factors. In this study, whether F_t or λ_i has a zero mean is not crucial, because they are treated as parameters to be estimated. Assumption (A8) allows for weak forms of both cross-sectional dependence and serial dependence in the error processes. Assumption (A9) can also be found in Noh and Park (2010), and is used for the system of general basis functions B_{kl}, which includes orthonormal bases, non-orthonormal bases, and B-splines.

Let $\|a\|_{L_2} = \{\int_{\mathcal{U}} a^2(u)\mathrm{d}u\}^{1/2}$ be the L_2 norm of any square integrable real-valued function $a(u)$ on $\mathcal{U}$, and let $\|A\|_{L_2} = \{\sum_{k=1}^{p}\|a\|_{L_2}^2\}^{1/2}$ be the L_2 norm of $A(u) = (a_1(u),\ldots,a_p(u))^\tau$, where $a_k(u)$ are real-valued functions on $\mathcal{U}$ (see Huang, Wu and Zhou (2002)). We define $\hat{\beta}_k(\cdot)$ as a consistent estimator of $\beta_k(\cdot)$ if $\lim_{N,T\to\infty}\|\hat{\beta}_k(\cdot)-\beta_k(\cdot)\|_{L_2} = 0$ holds in probability. Define $\delta_{NT} = \min[\sqrt{N},\sqrt{T}]$ and $L_N = \max_{1\le k\le p} L_k$, which tend to infinity as N or T tends to infinity. Let $\mathcal{D} = \{(X_{it}, U_{it}, \lambda_i, F_t), i = 1,\ldots,N, t = 1,\ldots,T\}$. We use $E_{\mathcal{D}}$ and $\mathrm{Var}_{\mathcal{D}}$ to denote the expectation and variance conditional on $\mathcal{D}$, respectively.

3.2. Asymptotic properties

Let $\boldsymbol{F}^0$ be the true value of $\boldsymbol{F}$. With an appropriate choice of L_k to balance the bias and variance, our proposed estimators have asymptotic properties including consistency, a convergence rate, and an asymptotic distribution.

Theorem 1. *Suppose assumptions* (A1)*–*(A9) *hold. If* $\delta_{NT}^{-2}L_N\log L_N \to 0$ *as* $N\to\infty$ *and* $T\to\infty$ *simultaneously, then*

(i) $\hat{\beta}_k(\cdot)$, *for* $k = 1,\ldots,p$, *are uniquely defined with probability tending to one.*

(ii) *The matrix* $\boldsymbol{F}^{0\tau}\hat{\boldsymbol{F}}/T$ *is invertible and* $\|P_{\hat{\boldsymbol{F}}} - P_{\boldsymbol{F}^0}\| \xrightarrow{P} 0$, *where* $P_A = A(A^\tau A)^{-1}A^\tau$ *for a given matrix* A.

Part (i) of Theorem 1 implies that, with probability tending to one, we can obtain unique estimators $\hat{\beta}_k(\cdot)$ for the unknown coefficient functions $\beta_k(\cdot)$ under some regularity assumptions, regardless of whether unobservable multiple interactive fixed effects exist in model (1.2). Part (ii) of Theorem 1 indicates that the spaces spanned by $\hat{\boldsymbol{F}}$ and $\boldsymbol{F}^0$ are asymptotically consistent. This is a key result that guarantees that the estimators $\hat{\beta}_k(\cdot)$ have good asymptotic properties, including the optimal convergence rate, consistency, and asymptotic normality.

Theorem 2. *Under the assumptions of Theorem* 1*, we further have*

$$\|\hat{\beta}_k(u) - \beta_k(u)\|_{L_2}^2 = O_P\left(\frac{L_N}{NT} + \frac{L_N}{T^2} + \frac{L_N}{N^2} + L_N^{-2d}\right), \quad k = 1, \ldots, p.$$

Theorem 2 gives the convergence rate of $\hat{\beta}_k(u)$, for all $k = 1, \ldots, p$, and, hence, establishes the consistency of our proposed estimators under the condition $\delta_{NT}^{-2} L_N \log L_N \to 0$ as $N \to \infty$ and $T \to \infty$ simultaneously. From the proof of Theorem 2, we note the following. The first term in the convergence rate is caused by the stochastic error. The second and third terms are caused by the estimation error of the fixed effects $\boldsymbol{F}^0$ and the presence of cross-sectional and serial correlation and heteroskedasticity, respectively. The last term is the error due to the basis approximation. If we take the appropriate relative rate $T/N \to c > 0$ as $N \to \infty$ and $T \to \infty$ simultaneously, then we have a more accurate convergence rate, as follows

$$\|\hat{\beta}_k(u) - \beta_k(u)\|_{L_2}^2 = O_P\left(\frac{L_N}{NT} + L_N^{-2d}\right), \quad k = 1, \ldots, p.$$

Furthermore, if we take $L_N = O((NT)^{1/(2d+1)})$, then

$$\|\hat{\beta}_k(u) - \beta_k(u)\|_{L_2}^2 = O_P\left((NT)^{-2d/(2d+1)}\right), \quad k = 1, \ldots, p.$$

This leads to the optimal convergence rate of order $O_P\left((NT)^{-2d/(2d+1)}\right)$, which holds for the i.i.d. data in Stone (1982).

Next, we establish the asymptotic distribution of $\hat{\boldsymbol{\beta}}(u)$. Let $\boldsymbol{Z}_i = M_{\boldsymbol{F}^0}\boldsymbol{R}_i - N^{-1}\sum_{j=1}^N a_{ij} M_{\boldsymbol{F}^0}\boldsymbol{R}_j$. The variance-covariance matrix of $\hat{\boldsymbol{\beta}}(u)$, conditioning on $\mathcal{D}$, is $\Sigma = \mathrm{Var}(\hat{\boldsymbol{\beta}}(u)|\mathcal{D}) = \boldsymbol{B}(u)\Phi\boldsymbol{B}(u)^\tau$, where Φ is the limit in probability of

$$\Phi^* = \left(\sum_{i=1}^N \boldsymbol{Z}_i^\tau \boldsymbol{Z}_i\right)^{-1} \Sigma_{NT1} \left(\sum_{i=1}^N \boldsymbol{Z}_i^\tau \boldsymbol{Z}_i\right)^{-1},$$

with $\Sigma_{NT1} = \sum_{i=1}^N \sum_{j=1}^N \sum_{t=1}^T \sum_{s=1}^T \sigma_{ij,ts} Z_{it} Z_{js}^\tau$. Let ϖ_k denote the unit vector in $\mathbb{R}^p$ with one in the kth coordinate, and zero in all other coordinates, for $k = 1, \ldots, p$. Then, the conditional variance of $\hat{\beta}_k(u)$ is

$$\Sigma_{kk} = \mathrm{Var}(\hat{\beta}_k(u)|\mathcal{D}) = \varpi_k^\tau \Sigma \varpi_k, \quad k = 1, \ldots, p.$$

To study the asymptotic distribution of $\hat{\boldsymbol{\beta}}(u)$, we add the following assumption.

(A10) Let Σ_1 be the limit in probability of $(1/NT)\Sigma_{NT1}$; then, $(1/\sqrt{NT}) \sum_{i=1}^N \boldsymbol{Z}_i^\tau \boldsymbol{\varepsilon}_i \xrightarrow{L} N(\mathbf{0}, \Sigma_1)$, where "$\xrightarrow{L}$" denotes convergence in distribution.

Denote $\tilde{\Sigma} = D_0^{-1} \Sigma_1 D_0^{-1}$, where $D_0 = \mathrm{plim}(L_N/NT) \sum_{i=1}^N \boldsymbol{Z}_i^\tau \boldsymbol{Z}_i$. The following theorem establishes the asymptotic distribution of $\hat{\boldsymbol{\beta}}(u)$.

Theorem 3. *Suppose that assumptions* (A1)*–*(A10) *hold. If* $\delta_{NT}^{-2} L_N \log L_N \to 0$*,* $L_N^{2d+1}/NT \to \infty$*, and* $T/N \to c$ *as* $N \to \infty$ *and* $T \to \infty$ *simultaneously, then*

$$\Sigma^{-1/2}(\hat{\boldsymbol{\beta}}(u) - \boldsymbol{\beta}(u)) \xrightarrow{L} N(\boldsymbol{b}(u), I_p),$$

where $\boldsymbol{b}(u) = \tilde{\Sigma}^{-1/2} c^{1/2} W_1^0 + \tilde{\Sigma}^{-1/2} c^{-1/2} W_2^0$*, and* W_1^0 *is the limit in probability of* W_1*, with*

$$W_1 = -\boldsymbol{B}(u) \left(L_N D(\boldsymbol{F}^0)\right)^{-1} \frac{1}{N} \sum_{i=1}^N \sum_{j=1}^N \frac{(\boldsymbol{R}_i - \boldsymbol{V}_i)^\tau \boldsymbol{F}^0}{T} \left(\frac{\boldsymbol{F}^{0\tau} \boldsymbol{F}^0}{T}\right)^{-1} \times \left(\frac{\Lambda^\tau \Lambda}{N}\right)^{-1} \lambda_j \left(\frac{1}{T} \sum_{t=1}^T \sigma_{ij,tt}\right),$$

and W_2^0 *is the limit in probability of* W_2*, with*

$$W_2 = -\boldsymbol{B}(u) \left(L_N D(\boldsymbol{F}^0)\right)^{-1} \frac{1}{NT} \sum_{i=1}^N \boldsymbol{R}_i^\tau M_{\boldsymbol{F}^0} \Omega \boldsymbol{F}^0 \left(\frac{\boldsymbol{F}^{0\tau} \boldsymbol{F}^0}{T}\right)^{-1} \left(\frac{\Lambda^\tau \Lambda}{N}\right)^{-1} \lambda_i,$$

where $\boldsymbol{V}_i = N^{-1} \sum_{j=1}^N a_{ij} \boldsymbol{R}_j$ *and* $\Omega = N^{-1} \sum_{i=1}^N \Omega_i$*.*

From the asymptotic normality in Theorem 3, we find that $\hat{\boldsymbol{\beta}}(u)$ has a bias term $\boldsymbol{b}(u)$, and $\boldsymbol{b}(u)$ has a complex structure. In order to improve the efficiency of a statistical inference, we propose a bias-corrected procedure to remove the bias term $\boldsymbol{b}(u)$. Noting that cross-sectional and serial dependence and heteroskedasticity are allowed in the error terms, we first estimate W_1 and W_2, as follows:

$$\hat{W}_1 = -\boldsymbol{B}(u)\hat{D}_0^{-1}\frac{1}{n}\sum_{i=1}^{n}\sum_{j=1}^{n}\frac{(\boldsymbol{R}_i - \hat{\boldsymbol{V}}_i)^\tau \hat{\boldsymbol{F}}}{T}\left(\frac{\hat{\Lambda}^\tau\hat{\Lambda}}{N}\right)^{-1}\hat{\lambda}_j\left(\frac{1}{T}\sum_{t=1}^{T}\hat{\varepsilon}_{it}\hat{\varepsilon}_{jt}\right),$$

$$\hat{W}_2 = -\boldsymbol{B}(u)\hat{D}_0^{-1}\frac{1}{NT}\sum_{i=1}^{N}\frac{1}{N}\sum_{k=1}^{N}\left(\boldsymbol{R}_i^\tau\hat{\Omega}_k\hat{\boldsymbol{F}} - T^{-1}\hat{\boldsymbol{F}}\hat{\boldsymbol{F}}^\tau\hat{\Omega}_k\hat{\boldsymbol{F}}\right)\left(\frac{\hat{\Lambda}^\tau\hat{\Lambda}}{N}\right)^{-1}\hat{\lambda}_i,$$

where n satisfies $n/N \to 0$, $n/T \to 0$, and $\hat{D}_0 = (L_N/NT)\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{Z}_{it}\hat{Z}_{it}^\tau$, with $\boldsymbol{F}^0$, λ_i, and Λ replaced with $\hat{\boldsymbol{F}}$, $\hat{\lambda}_i$, and $\hat{\Lambda}$ in $\hat{Z}_{it}$, respectively. Note that $\boldsymbol{R}_i^\tau\hat{\Omega}_k\hat{\boldsymbol{F}} = (I_{p_0}, \mathbf{0})(\boldsymbol{S}_i^\tau\hat{\Omega}_k\boldsymbol{S}_i)(\mathbf{0}^\tau, I_r)^\tau$ and $\hat{\boldsymbol{F}}^\tau\hat{\Omega}_k\hat{\boldsymbol{F}} = (\mathbf{0}, I_r)(\boldsymbol{S}_i^\tau\hat{\Omega}_k\boldsymbol{S}_i)(\mathbf{0}^\tau, I_r)^\tau$, where $p_0 = \sum_{k=1}^{p}L_k$ and $\boldsymbol{S}_i^\tau\hat{\Omega}_k\boldsymbol{S}_i = C_{0i} + \sum_{\nu=1}^{q}\left[1-\nu/(q+1)\right](C_{\nu i} + C_{\nu i}^\tau)$, $\boldsymbol{S}_i = (\boldsymbol{R}_i, \hat{\boldsymbol{F}})$, $C_{\nu i} = (1/T)\sum_{t=\nu+1}^{T}S_{it}\hat{\varepsilon}_{kt}\hat{\varepsilon}_{k,t-\nu}S_{i,t-\nu}$, and $q \to \infty$ and $q/T^{1/4} \to 0$ as $T \to \infty$. Thus, we define the bias-corrected estimator of $\boldsymbol{\beta}(u)$ as

$$\breve{\boldsymbol{\beta}}(u) = \hat{\boldsymbol{\beta}}(u) - \frac{L_N}{N}\hat{W}_1 - \frac{L_N}{T}\hat{W}_2.$$

The following theorem shows there is no bias term in the asymptotic distribution of the bias-corrected estimator $\breve{\boldsymbol{\beta}}(u)$.

Theorem 4. *Suppose that assumptions* (A1)*–*(A10) *hold. If* $\delta_{NT}^{-2}L_N\log L_N \to 0$, $L_N^{2d+1}/NT \to \infty$, *and* $T/N \to c$ *as* $N \to \infty$ *and* $T \to \infty$ *simultaneously, then*

$$\Sigma^{-1/2}(\breve{\boldsymbol{\beta}}(u) - \boldsymbol{\beta}(u)) \xrightarrow{L} N(0, I_p).$$

In particular, we have $\Sigma_{kk}^{-1/2}(\breve{\beta}_k(u) - \beta_k(u)) \xrightarrow{L} N(0,1)$, *for* $k = 1, \ldots, p$.

Next, we consider some special cases where the asymptotic bias can be simplified. (1) In the absence of serial correlation and heteroskedasticity, $E(\varepsilon_{it}\varepsilon_{jt}) = \sigma_{ij,tt} = \sigma_{ij}$, because it does not depend on t. It is easy to show that $W_2 = 0$. (2) In the absence of cross-sectional correlation and heteroskedasticity, $E(\varepsilon_{it}\varepsilon_{is}) = \sigma_{ii,ts} = \omega_{ts}$, because it does not depend on i, in which case, a simple calculation yields $W_1 = 0$. Let Π and Ξ be the probability limits, defined as, respectively,

$$\Pi = \mathrm{plim}\boldsymbol{B}(u)\left(\sum_{i=1}^{N}\boldsymbol{Z}_i^\tau\boldsymbol{Z}_i\right)^{-1}\Sigma_{NT2}\left(\sum_{i=1}^{N}\boldsymbol{Z}_i^\tau\boldsymbol{Z}_i\right)^{-1}\boldsymbol{B}(u)^\tau,$$

$$\Xi = \mathrm{plim}\boldsymbol{B}(u)\left(\sum_{i=1}^{N}\boldsymbol{Z}_i^\tau\boldsymbol{Z}_i\right)^{-1}\Sigma_{NT3}\left(\sum_{i=1}^{N}\boldsymbol{Z}_i^\tau\boldsymbol{Z}_i\right)^{-1}\boldsymbol{B}(u)^\tau,$$

where $\Sigma_{NT2}=\sum_{i=1}^{N}\sum_{j=1}^{N}\sigma_{ij}\sum_{t=1}^{T}Z_{it}Z_{jt}^{\tau}$ and $\Sigma_{NT3}=\sum_{t=1}^{T}\sum_{s=1}^{T}\omega_{ts}\sum_{i=1}^{N}Z_{it}Z_{is}^{\tau}$.

Corollary 1. *Suppose that assumptions* (A1)*–*(A10) *hold. If* $\delta_{NT}^{-2}L_N\log L_N\to 0$ *and* $L_N^{2d+1}/NT\to\infty$ *as* $N\to\infty$ *and* $T\to\infty$ *simultaneously, we have the following results:*

(i) *In the absence of serial correlation and heteroskedasticity and* $T/N\to 0$, $\Pi^{-1/2}(\hat{\boldsymbol{\beta}}(u)-\boldsymbol{\beta}(u))\xrightarrow{L}N(0,I_p)$.

(ii) *In the absence of cross-sectonal correlation and heteroskedasticity and* $N/T\to 0$, $\Xi^{-1/2}(\hat{\boldsymbol{\beta}}(u)-\boldsymbol{\beta}(u))\xrightarrow{L}N(0,I_p)$.

For model (1.2) with unobservable multiple interactive fixed effects, Theorem 4 establishes the asymptotic normality for the bias-corrected estimator $\breve{\beta}_k(\cdot)$ of $\beta_k(\cdot)$. Hence, if we can obtain a consistent estimator $\hat{\Sigma}_{kk}$ of Σ_{kk}, the asymptotic pointwise confidence intervals for $\beta_k(u)$ can be constructed as

$$\breve{\beta}_k(u)\pm z_{\alpha/2}\hat{\Sigma}_{kk}^{-1/2},\quad k=1,\ldots,p,$$

where $z_{\alpha/2}$ is the $(1-\alpha/2)$ quantile of the standard normal distribution.

4. A Residual-Based Block Bootstrap Procedure

In theory, we can construct the pointwise confidence intervals for the coefficient functions $\beta_k(\cdot)$ from Theorems 3 and 4. For Theorem 3, we first need to derive consistent estimators for the asymptotic biases and variances of the estimators $\hat{\beta}_k(\cdot)$, for $k=1,\ldots,p$. Nevertheless, because the asymptotic biases and variances involve the unknown fixed effects $\boldsymbol{F}$ and the covariance matrices Ω_i of $\boldsymbol{\varepsilon}_i$, it is difficult to obtain their consistent and efficient estimators, even if the plug-in method is used. For Theorem 4, it is difficult to show the consistency of the estimators $\hat{W}_1$ and $\hat{W}_2$, because cross-sectional and serial dependence and heteroskedasticity are allowed in the error terms.

Therefore, the standard nonparametric bootstrap procedure cannot be applied to construct the pointwise confidence intervals directly, because cross-sectional and serial correlations exist within the group in model (1.2). In addition to increasing the computational burden and causing accumulative errors, they make it more difficult to construct the pointwise confidence intervals. To overcome these limitations, we propose a residual-based block bootstrap bias-correction procedure to construct the pointwise confidence intervals for $\beta_k(\cdot)$. The algorithm follows.

Step 1. Fit model (1.2) using the methods proposed in Section 2, and estimate the residuals ε_{it} using

$$\hat{\varepsilon}_{it} = Y_{it} - \sum_{k=1}^{p}\sum_{l=1}^{L_k}\hat{\gamma}_{kl}X_{it,k}B_{kl}(U_{it}) + \hat{\lambda}_i^{\tau}\hat{F}_t,\ i=1,\ldots,N,\ t=1,\ldots,T.$$

Step 2. Generate the bootstrap residuals ε_{it}^* by $\hat{\varepsilon}_{it}$ using the block bootstrap method with a two-step procedure: (i) Choose the block lengths. In our block bootstrap procedure, similarly to Inoue and Shintani (2006), we choose block lengths of $l_1 = cT^{1/3}$ and $l_2 = cN^{1/3}$, respectively, for some $c > 0$. (ii) Resample the blocks and generate the bootstrap samples. The blocks can be overlapping or non-overlapping. According to Lahiri (1999), there is little difference in the performance for these two methods. We hence adopt the non-overlapping method, for simplicity. Then, we first divide the $N \times T$ residual matrix $\hat{\boldsymbol{\varepsilon}}$ into $m_1 = T/l_1$ blocks by column, and generate the bootstrap samples $N \times T$ matrix $\tilde{\boldsymbol{\varepsilon}}$ by resampling, with replacement, the m_1 blocks of columns of $\hat{\boldsymbol{\varepsilon}}$. Next, we divide $\tilde{\boldsymbol{\varepsilon}}$ into $m_2 = N/l_2$ blocks by row, and generate the bootstrap samples matrix $\boldsymbol{\varepsilon}^*$ by resampling, with replacement, the m_2 blocks of rows of $\tilde{\boldsymbol{\varepsilon}}$.

Step 3. We generate the bootstrap sample Y_{it}^* using the following model:

$$Y_{it}^* = \sum_{k=1}^{p}\sum_{l=1}^{L_k}\hat{\gamma}_{kl}X_{it,k}B_{kl}(U_{it}) + \hat{\lambda}_i^{\tau}\hat{F}_t + \varepsilon_{it}^*,\ i=1,\ldots,N,\ t=1,\ldots,T,$$

where $\hat{\gamma}_{kl}$, $\hat{F}_t$, and $\hat{\lambda}_i$ are the respective estimators of γ_{kl}, F_t, and λ_i, using the estimation procedure in Section 2. Based on the bootstrap sample $\{(Y_{it}^*, X_{it}, U_{it}), i = 1,\ldots,N, t = 1,\ldots,T\}$, we calculate the bootstrap estimator $\hat{\boldsymbol{\beta}}^{(b)}(\cdot)$, also using the estimation procedure in Section 2.

Step 4. Repeat Steps 2 and 3 B times to obtain a size B bootstrap estimator $\hat{\boldsymbol{\beta}}^{(b)}(u)$, for $b = 1,\ldots,B$. The bootstrap estimator $\mathrm{Var}^*(\hat{\boldsymbol{\beta}}(u)|\mathcal{D})$ of $\Sigma = \mathrm{Var}(\hat{\boldsymbol{\beta}}(u)|\mathcal{D})$ is taken as the sample variance of $\hat{\boldsymbol{\beta}}^{(b)}(u)$. Next, the bootstrap bias-corrected estimator of $\hat{\beta}_k(u)$ can be defined as

$$\breve{\beta}_k(u) = \hat{\beta}_k(u) - \left(\frac{1}{B}\sum_{b=1}^{B}\hat{\beta}_k^{(b)}(u) - \hat{\beta}_k(u)\right) = 2\hat{\beta}_k(u) - \frac{1}{B}\sum_{b=1}^{B}\hat{\beta}_k^{(b)}(u).$$

Intuitively, the bias of a bootstrap estimator is a good approximation to that of a true coefficient function estimator. Finally, we construct the asymptotic pointwise confidence intervals for $\beta_k(u)$ as

$$\breve{\beta}_k(u) \pm z_{\alpha/2}\{\mathrm{Var}^*(\hat{\beta}_k(u)|\mathcal{D})\}^{1/2}, \quad k = 1, \ldots, p,$$

where $z_{\alpha/2}$ is the $(1-\alpha/2)$ quantile of the standard normal distribution.

5. Partially Linear Varying-Coefficient Model

In this section, we consider a special case of model (1.2), where some components $\underline{X}_{it} = (X_{it,1}, \ldots, X_{it,q})^\tau$ of X_{it} are constant effects, and the rest $\overline{X}_{it} = (X_{it,q+1}, \ldots, X_{it,p})^\tau$ are varying effects, for $i = 1, \ldots, N$ and $t = 1, \ldots, T$. Then, model (1.2) becomes the following partially linear varying-coefficient model with interactive fixed effects:

$$Y_{it} = \underline{X}_{it}^\tau \boldsymbol{\beta}^{(1)}(U_{it}) + \overline{X}_{it}^\tau \boldsymbol{\theta} + \lambda_i^\tau F_t + \varepsilon_{it}, \tag{5.1}$$

where $\boldsymbol{\beta}^{(1)}(u) = (\beta_1(u)), \ldots, \beta_q(u))^\tau$ and $\boldsymbol{\theta} = (\beta_{q+1}, \ldots, \beta_p)^\tau$.

Similarly to the proposed estimation procedure in Section 2, we can define the following objective function:

$$Q(\boldsymbol{\gamma}^{(1)}, \boldsymbol{\theta}, \boldsymbol{F}) = \sum_{i=1}^{N} (\boldsymbol{Y}_i - \underline{\boldsymbol{R}}_i \boldsymbol{\gamma}^{(1)} - \overline{\boldsymbol{X}}_i \boldsymbol{\theta})^\tau M_{\boldsymbol{F}} (\boldsymbol{Y}_i - \underline{\boldsymbol{R}}_i \boldsymbol{\gamma}^{(1)} - \overline{\boldsymbol{X}}_i \boldsymbol{\theta}). \tag{5.2}$$

Thus, the estimators of $\boldsymbol{\gamma}^{(1)}$ and $\boldsymbol{\theta}$ can be obtained by iterating between $\boldsymbol{\gamma}^{(1)}$, $\boldsymbol{\theta}$, and $\boldsymbol{F}$ using the following nonlinear equations:

$$\begin{aligned}
\hat{\boldsymbol{\theta}} &= \left[\sum_{i=1}^{N} \overline{\boldsymbol{X}}_i^\tau M_{\hat{\boldsymbol{F}}} \left\{ I_T - \underline{\boldsymbol{R}}_i \left(\sum_{i=1}^{N} \underline{\boldsymbol{R}}_i^\tau M_{\hat{\boldsymbol{F}}} \underline{\boldsymbol{R}}_i \right)^{-1} \sum_{i=1}^{N} \underline{\boldsymbol{R}}_i^\tau M_{\hat{\boldsymbol{F}}} \right\} \overline{\boldsymbol{X}}_i \right]^{-1} \\
&\quad \times \sum_{i=1}^{N} \overline{\boldsymbol{X}}_i^\tau M_{\hat{\boldsymbol{F}}} \left\{ I_T - \underline{\boldsymbol{R}}_i \left(\sum_{i=1}^{N} \underline{\boldsymbol{R}}_i^\tau M_{\hat{\boldsymbol{F}}} \underline{\boldsymbol{R}}_i \right)^{-1} \sum_{i=1}^{N} \underline{\boldsymbol{R}}_i^\tau M_{\hat{\boldsymbol{F}}} \right\} \boldsymbol{Y}_i, \\
\hat{\boldsymbol{\gamma}}^{(1)} &= \left(\sum_{i=1}^{N} \underline{\boldsymbol{R}}_i^\tau M_{\hat{\boldsymbol{F}}} \underline{\boldsymbol{R}}_i \right)^{-1} \sum_{i=1}^{N} \underline{\boldsymbol{R}}_i^\tau M_{\hat{\boldsymbol{F}}} (\boldsymbol{Y}_i - \overline{\boldsymbol{X}}_i^\tau \hat{\boldsymbol{\theta}}), \\
\hat{\boldsymbol{F}} V_{NT} &= \left[\frac{1}{NT} \sum_{i=1}^{N} (\boldsymbol{Y}_i - \underline{\boldsymbol{R}}_i \hat{\boldsymbol{\gamma}}^{(1)} - \overline{\boldsymbol{X}}_i^\tau \hat{\boldsymbol{\theta}})(\boldsymbol{Y}_i - \underline{\boldsymbol{R}}_i \hat{\boldsymbol{\gamma}}^{(1)} - \overline{\boldsymbol{X}}_i^\tau \hat{\boldsymbol{\theta}})^\tau \right] \hat{\boldsymbol{F}}.
\end{aligned} \tag{5.3}$$

By the property of B-spline bases that $\sum_{l=1}^{L_k} B_{kl}(u) = 1$ if $\beta_k(u)$ is a constant

β_k, we set $\gamma_k = \beta_k \mathbf{1}_{L_k}$, where $\mathbf{1}_{L_k}$ is an $L_k \times 1$ vector with entries of one. With a slight abuse of notation, (5.2) can be rewritten as

$$Q(\boldsymbol{\gamma}^{(1)}, \boldsymbol{\theta}, \boldsymbol{F}) = Q(\boldsymbol{\gamma}, \boldsymbol{F}) = \sum_{i=1}^{N} (\boldsymbol{Y}_i - \boldsymbol{R}_i \boldsymbol{\gamma})^\tau M_{\boldsymbol{F}} (\boldsymbol{Y}_i - \boldsymbol{R}_i \boldsymbol{\gamma}), \tag{5.4}$$

where $\boldsymbol{\gamma} = (\gamma_1^\tau, \ldots, \gamma_q^\tau, \beta_{q+1}\mathbf{1}_{L_{q+1}}^\tau, \ldots, \beta_p \mathbf{1}_{L_p}^\tau)^\tau = (\boldsymbol{\gamma}^{(1)\tau}, \beta_{q+1}\mathbf{1}_{L_{q+1}}^\tau, \ldots, \beta_p \mathbf{1}_{L_p}^\tau)^\tau$. For each $k = q+1, \ldots, p$, we treat β_k as a function, and apply the estimation procedure in Section 2 to obtain the initial estimators of $\hat{\boldsymbol{\gamma}}^{(1)}$, $\hat{\boldsymbol{F}}$, and $\hat{\Lambda}$. Then, we propose the following robust iteration algorithm for estimating the parameters $(\boldsymbol{\gamma}^{(1)}, \boldsymbol{\theta}, \boldsymbol{F}, \Lambda)$.

Step 1. Start with an initial estimator $(\hat{\boldsymbol{\gamma}}^{(1)}, \hat{\boldsymbol{F}}, \hat{\Lambda})$.

Step 2. Given $\hat{\boldsymbol{\gamma}}^{(1)}$, $\hat{\boldsymbol{F}}$, and $\hat{\Lambda}$, compute

$$\hat{\boldsymbol{\theta}}(\hat{\boldsymbol{\gamma}}^{(1)}, \hat{\boldsymbol{F}}, \hat{\Lambda}) = \left(\sum_{i=1}^{N} \overline{\boldsymbol{X}}_i^\tau \overline{\boldsymbol{X}}_i \right)^{-1} \sum_{i=1}^{N} \overline{\boldsymbol{X}}_i^\tau (\boldsymbol{Y}_i - \underline{\boldsymbol{R}}_i \hat{\boldsymbol{\gamma}}^{(1)} - \hat{\boldsymbol{F}} \hat{\lambda}_i).$$

Step 3. Given $\hat{\boldsymbol{\theta}}$, $\hat{\boldsymbol{F}}$, and $\hat{\Lambda}$, compute

$$\hat{\boldsymbol{\gamma}}^{(1)}(\hat{\boldsymbol{\theta}}, \hat{\boldsymbol{F}}, \hat{\Lambda}) = \left(\sum_{i=1}^{N} \underline{\boldsymbol{R}}_i^\tau \underline{\boldsymbol{R}}_i \right)^{-1} \sum_{i=1}^{N} \underline{\boldsymbol{R}}_i^\tau (\boldsymbol{Y}_i - \overline{\boldsymbol{X}}_i \hat{\boldsymbol{\theta}} - \hat{\boldsymbol{F}} \hat{\lambda}_i).$$

Step 4. Given $\hat{\boldsymbol{\gamma}}^{(1)}$ and $\hat{\boldsymbol{\theta}}$, compute $\hat{\boldsymbol{F}}$ according to (5.3) (multiplied by $\sqrt{T}$, owing to the restriction that $\boldsymbol{F}^\tau \boldsymbol{F}/T = I_r$), and calculate $\hat{\Lambda}$ using formula (2.9), with $\hat{\boldsymbol{\gamma}} = (\hat{\boldsymbol{\gamma}}^{(1)\tau}, \hat{\beta}_{q+1}\mathbf{1}_{L_{q+1}}^\tau, \ldots, \hat{\beta}_p \mathbf{1}_{L_p}^\tau)^\tau$.

Step 5. Repeat Steps 2–4 until $(\hat{\boldsymbol{\gamma}}^{(1)}, \hat{\boldsymbol{\theta}}, \hat{\boldsymbol{F}}, \hat{\Lambda})$ satisfy the given convergence criterion.

In order to give the following asymptotic distribution, we first introduce some notation. Let

$$\overline{\boldsymbol{Z}}_i = M_{\boldsymbol{F}^0} \overline{\boldsymbol{X}}_i - \frac{1}{N} \sum_{j=1}^{N} M_{\boldsymbol{F}^0} \overline{\boldsymbol{X}}_j a_{ij}, \quad \underline{\boldsymbol{Z}}_i = M_{\boldsymbol{F}^0} \underline{\boldsymbol{R}}_i - \frac{1}{N} \sum_{j=1}^{N} M_{\boldsymbol{F}^0} \underline{\boldsymbol{R}}_j a_{ij},$$

$$\overline{\Phi} = \frac{1}{NT} \sum_{i=1}^{N} \overline{\boldsymbol{Z}}_i^\tau \overline{\boldsymbol{Z}}_i, \quad \underline{\Phi} = \frac{1}{NT} \sum_{i=1}^{N} \underline{\boldsymbol{Z}}_i^\tau \underline{\boldsymbol{Z}}_i, \quad \Psi = \frac{1}{NT} \sum_{i=1}^{N} \overline{\boldsymbol{Z}}_i^\tau \underline{\boldsymbol{Z}}_i,$$

and $\check{\boldsymbol{Z}}_i = \overline{\boldsymbol{Z}}_i - \underline{\boldsymbol{Z}}_i\underline{\Phi}^{-1}\Psi^\tau$. In addition, we define the following probability limits:

$$\Pi_1 = \text{plim}\frac{1}{NT}\sum_{i=1}^N \check{\boldsymbol{Z}}_i^\tau \check{\boldsymbol{Z}}_i = \text{plim}(\overline{\Phi} - \Psi\underline{\Phi}^{-1}\Psi^\tau),$$

$$\Pi_2 = \text{plim}\frac{1}{NT}\sum_{i=1}^N\sum_{j=1}^N\sum_{t=1}^T\sum_{s=1}^T \sigma_{ij,ts}\check{Z}_{it}\check{Z}_{js}^\tau.$$

The following theorem gives the asymptotic normality of the parametric components.

Theorem 5. *Suppose that assumptions* (A1)*–*(A10) *hold. If* $\delta_{NT}^{-2}L_N \log L_N \to 0$, $L_N^{2d+1}/NT \to \infty$, *and* $T/N \to c$ *as* $N \to \infty$ *and* $T \to \infty$ *simultaneously, then*

$$(NT)^{-1/2}(\hat{\boldsymbol{\theta}} - \boldsymbol{\theta}) \xrightarrow{L} N(\boldsymbol{b}, \Pi_1^{-1}\Pi_2\Pi_1^{-1}),$$

where $\boldsymbol{b} = c^{1/2}\check{S}_1^0 + c^{-1/2}\check{S}_2^0$, *and* $\check{S}_1^0$ *is the probability limit of* $\check{S}_1$, *with*

$$\check{S}_1 = -(\overline{\Phi} - \Psi\underline{\Phi}^{-1}\Psi^\tau)^{-1}\Bigg[\frac{1}{N}\sum_{i=1}^N\sum_{j=1}^N \frac{(\overline{\boldsymbol{X}}_i - \overline{\boldsymbol{V}}_i)^\tau \boldsymbol{F}^0}{T} G^0\lambda_j\left(\frac{1}{T}\sum_{t=1}^T \varepsilon_{it}\varepsilon_{jt}\right)$$
$$-\Psi\underline{\Phi}^{-1}\frac{1}{N}\sum_{i=1}^N\sum_{j=1}^N \frac{(\underline{\boldsymbol{R}}_i - \underline{\boldsymbol{V}}_i)^\tau \boldsymbol{F}^0}{T} G^0\lambda_j\left(\frac{1}{T}\sum_{t=1}^T \varepsilon_{it}\varepsilon_{jt}\right)\Bigg],$$

and $\check{S}_2^0$ *is the probability limit of* $\check{S}_2$, *with*

$$\check{S}_2 = -(\overline{\Phi} - \Psi\underline{\Phi}^{-1}\Psi^\tau)^{-1}\Bigg(\frac{1}{NT}\sum_{i=1}^N \overline{\boldsymbol{X}}_i^\tau M_{\boldsymbol{F}^0}\Omega\boldsymbol{F}^0 G^0\lambda_i$$
$$-\Psi\underline{\Phi}^{-1}\frac{1}{NT}\sum_{i=1}^N \underline{\boldsymbol{R}}_i^\tau M_{\boldsymbol{F}^0}\Omega\boldsymbol{F}^0 G^0\lambda_i\Bigg),$$

where $G^0 = (\boldsymbol{F}^{0\tau}\boldsymbol{F}^0/T)^{-1}(\Lambda^\tau\Lambda/N)^{-1}$ *and* $\overline{\boldsymbol{V}}_i = N^{-1}\sum_{j=1}^N a_{ij}\overline{\boldsymbol{X}}_j$.

It is easy to show that $\check{S}_1^0 = 0$ in the bias term $\boldsymbol{b}$ if the cross-sectional correlation and heteroskedasticity are absent. Similarly, $\check{S}_2^0 = 0$ if the serial correlation and heteroskedasticity are absent. We also show that both $\check{S}_1^0 = \check{S}_2^0 = 0$ if ε_{it} are i.i.d. over i and t. From Theorem 5, the convergence rate of $\hat{\boldsymbol{\theta}}$ is of order $O_P((NT)^{-1/2})$. Thus substituting $\hat{\boldsymbol{\theta}}$ for $\boldsymbol{\theta}$ in model (5.1) will have little effect on the estimation of $\beta_j(u)$, for $j = 1, \ldots, q$. This implies that the estimator $\hat{\beta}_j(u)$ will have similar asymptotic distributions in Theorems 3 and 4.

6. Hypothesis Testing

In practice, it is often of interest to test whether one or several coefficient functions are nonzero constants or are identically zero. We here propose a goodness-of-fit test that compares the residual sum of squares from least square fits under the null and alternative hypotheses.

We consider the null hypothesis that some of the coefficient functions are constants:

$$H_0: \quad \beta_{q+1}(u) = \beta_{q+1}, \ldots, \beta_p(u) = \beta_p,$$

for all $u \in \mathcal{U}$, where β_k $(k = q+1, \ldots, p)$ are unknown constants. Under H_0, model (1.2) reduces to the partially linear varying-coefficient panel-data model (5.1). Let $\hat{\boldsymbol{\gamma}}^{(1)*}$, $\hat{\boldsymbol{\theta}}$, $\hat{\boldsymbol{F}}^*$, and $\hat{\lambda}_i^*$ be the consistent estimators of $\boldsymbol{\gamma}^{(1)}$, $\boldsymbol{\theta}$, $\boldsymbol{F}$, and λ_i, respectively. Thus, the residual sum of squares under the null hypothesis H_0 is

$$\mathrm{RSS}_0 = \frac{1}{NT}\sum_{i=1}^{N}(\boldsymbol{Y}_i - \underline{\boldsymbol{R}_i}\hat{\boldsymbol{\gamma}}^{(1)*} - \overline{\boldsymbol{X}}_i\hat{\boldsymbol{\theta}} - \hat{\boldsymbol{F}}^*\hat{\lambda}_i^*)^\tau(\boldsymbol{Y}_i - \underline{\boldsymbol{R}_i}\hat{\boldsymbol{\gamma}}^{(1)*} - \overline{\boldsymbol{X}}_i\hat{\boldsymbol{\theta}} - \hat{\boldsymbol{F}}^*\hat{\lambda}_i^*).$$

Under the general alternative that all coefficient functions are allowed to vary with u, the residual sum of squares is defined by

$$\mathrm{RSS}_1 = \frac{1}{NT}\sum_{i=1}^{N}(\boldsymbol{Y}_i - \boldsymbol{R}_i\hat{\boldsymbol{\gamma}} - \hat{\boldsymbol{F}}\hat{\lambda}_i)^\tau(\boldsymbol{Y}_i - \boldsymbol{R}_i\hat{\boldsymbol{\gamma}} - \hat{\boldsymbol{F}}\hat{\lambda}_i). \tag{6.1}$$

We extend the generalized likelihood ratio in Fan, Zhang and Zhang (2001) to the current setting, and construct the test statistic under the null hypothesis H_0 as follows:

$$T_n = \frac{\mathrm{RSS}_0 - \mathrm{RSS}_1}{\mathrm{RSS}_1}, \tag{6.2}$$

where $\mathrm{RSS}_0 - \mathrm{RSS}_1$ indicates the difference of fit under the null and alternative hypotheses. If T_n is larger than an appropriate critical value, we reject the null hypothesis H_0. Let t_0 be the observed value of T_n. Then, the p-value of the test is defined as $p_0 = P_{H_0}(T_n > t_0)$, which denotes the probability of the event $\{T_n > t_0\}$. For a given significance level α_0, the null hypothesis H_0 is rejected if $p_0 \leq \alpha_0$.

Theorem 6. *Suppose that the conditions of Theorem* 3 *are satisfied. Under the null hypothesis H_0, $T_n \to 0$ in probability as $N \to \infty$ and $T \to \infty$. Otherwise, if* $\inf_{a\in\mathbb{R}} \|\beta_k(u) - a\|_{L_2} > 0$, *for some $k = q+1, \ldots, p$, then there exists a constant t_0, such that $T_n > t_0$ with probability approaching one as $N \to \infty$ and $T \to \infty$.*

Because it is difficult to develop the asymptotic null distribution of the statistic T_n, we use the following bootstrap procedure to evaluate the null distribution of T_n and compute the p-values of the test.

Step 1. We generate the bootstrap sample $\{(Y_{it}^*, X_{it}, U_{it}), i = 1, \ldots, N, t = 1, \ldots, T\}$, as described in Section 4, and calculate the bootstrap test statistic T_n^*.

Step 2. We repeat Step 1 many times to compute the bootstrap distribution of T_n^*.

Step 3. When the observed test statistic T_n is greater than or equal to the $\{100(1-\alpha_0)\}$th percentile of the empirical distribution T_n^*, we reject the null hypothesis H_0 at the significance level α_0. The p-value of the test is the empirical probability of the event $\{T_n^* \geq T_n\}$.

7. Conclusion

This study contributes to the literature by proposing an estimation procedure for a varying-coefficient panel-data model with interactive fixed effects. First, we use B-splines to approximate the coefficient functions for the model. With an appropriate choice of smoothing parameters, we propose a robust nonlinear iteration scheme based on the least squares method to estimate the coefficient functions. Then, we establish the asymptotic theory for the resulting estimators under some regularity assumptions, including their consistency, convergence rate, and asymptotic distribution. Second, to deal with the serial and cross-sectional correlation and heteroskedasticity within our model, which increases the computational burden and cause accumulative errors, we propose using a residual-based block bootstrap procedure to construct the pointwise confidence intervals for the coefficient functions. Third, we extend our proposed estimation procedure to include partially linear varying-coefficient models with interactive fixed effects, and study the asymptotic properties of the resulting estimator. In addition, we develop a test statistic for the constancy of the varying coefficient functions, and propose a bootstrap procedure to evaluate the null distribution of the test statistic. Finally, numerical studies demonstrate the satisfactory performance of our proposed methods in practice, and support our theoretical results.

Supplementary Material

The online Supplementary Material contains the numerical studies, proofs of Theorems 1–6 and Corollary 1, and Lemmas 1–7 and their proofs. In addition, we introduce the estimation procedure for a special model, namely, the varying-coefficient panel-data model with additive fixed effects.

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