Inference on Nonstationary Time Series with Moving Mean

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Abstract

A semiparametric model is proposed in which a parametric filtering of a non-stationary time series, incorporating fractionally differencing with short memory correction, removes correlation but leaves a nonparametric deterministic trend. Estimates of the memory parameter and other dependence parameters are proposed, and shown to be consistent and asymptotically normally distributed with parametric rate. Unit root tests with standard asymptotics are thereby justified. Estimation of the trend function is also considered. We include a Monte Carlo study of finite-sample performance.

Keywords: fractional time series; fixed design nonparametric regression; non-stationary time series; unit root tests.

JEL Classifications: C14, C22.

Proposed running head: Nonstationary Time Series

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1 INTRODUCTION

A long-established vehicle for smoothing a deterministically-trending time series $y_t$, $t = 1, ..., T$, is the fixed-design nonparametric regression model given by

$$y_t = g\left(\frac{t}{T}\right) + u_t, \ t = 1, ..., T,$$

(1)

where $g(x),\ x \in [0,1]$, is an unknown, smooth, nonparametric function, and $u_t$ is an unobservable sequence of random variables with zero mean. The dependence on sample size $T$ of $g(t/T)$ in (1) is to ensure sufficient accumulation of information to enable consistent estimation of $g(\tau)$ at any $\tau \in (0, 1)$. A more basic trend function is a polynomial in $t$ of given degree, as still frequently employed in various econometric models. A more general class of models than polynomials (and having analogy with the fractional stochastic trends we will employ in the current paper) involves fractional powers, i.e.

$$y_t = \beta_0 + \beta_1 t^{\gamma_1} + \ldots + \beta_p t^{\gamma_p} + u_t, \ t = 1, ..., T,$$

(2)

where all the $\beta_i$ and $\gamma_i$ are unknown and real-valued. Subject to identifiability and other restrictions, these parameters can be estimated consistently and asymptotically normally, e.g. by nonlinear least squares (Robinson (2012a)). Models such as (2) can be especially useful in modest sample sizes. However, and as with any parametric function of $t$, misspecification leads to inconsistent estimation, and a nonparametric treatment affords greater flexibility when $T$ is large (recognizing that nonparametric estimates converge more slowly than parametric ones).

With independent and identically distributed (iid) $u_t$, with finite variance, various kernel-type estimates of $g$ in (1) were developed by Gasser and Mueller (1979), Priestley and Chao (1972), with statistical properties established; in particular, under regularity conditions kernel estimates of $g(\tau)$ are consistent and asymptotically normally distributed as $T \to \infty$ (see e.g. Benedetti (1977)). A suitable choice of
kernel (and bandwidth) is an important ingredient in this theory, although kernel estimates are essentially an elaboration on simple moving window averages, which have a much longer history in empirical work. More recent empirical uses of (1) include Starica and Granger (2005) in modelling stock price series.

The iid assumption on $u_t$ is very restrictive, but similar asymptotic properties result when $u_t$ has weak dependence, for example is a covariance stationary process, generated by a linear process or satisfying suitable mixing conditions, and having finite and positive spectral density at degree zero (see e.g. Roussas, Tran and Ioannides (1992), Tran, Roussas, Yakowitz and Truong Van (1996)). The rate of convergence of kernel estimates is unaffected by this level of serial correlation, though the asymptotic variance differs from that in the iid case (unlike in the stochastic-design model in which the argument of $g$ in (1) is instead a weakly dependent stationary stochastic sequence).

Long-range dependence in $u_t$ has a greater impact on large-sample inference. If $u_t$ is a stationary and invertible fractional process, for example

$$(1 - L)^{\delta_0} u_t = \varepsilon_t, \quad |\delta_0| < 1/2;$$

$L$ being the lag operator and the $\varepsilon_t$ forming an iid sequence, or if $u_t$ has a "semiparametric" specification with spectral density $f(\lambda)$ having rate $\lambda^{-2\delta_0}$ as frequency $\lambda$ approaches zero from above, then the convergence rate of kernel estimates of $g(\tau)$ is slower when $\delta_0 > 0$ and faster when $\delta_0 < 0$. References dealing with (1) for such $u_t$ include Beran and Feng (2002), Csorgo and Mielniczuk (1995), Deo (1997), Guo and Koul (2007), Robinson (1997), Zhao, Zhang and Li (2013). The asymptotic variance of the kernel estimates depends on $\delta_0$ and any other time series parameters; for the "semiparametric" specification Robinson (1997) justified studentization using local Whittle estimates of $\delta_0$.

The restriction $\delta_0 < 1/2$ implies stationarity of $u_t$, so that $y_t$ given by (1) is non-
stationary only in the mean. Stochastic trends are also an important source of nonstationarity in many empirical time series. However, a nonstationary stochastic trend in $y_t$ generated by a nonstationary $u_t$, for example one having a unit root, would render $g(t/T)$ undetectable. An alternative, semiparametric, model which both incorporates a possibly nonstationary stochastic trend and enables estimation of a nonparametric deterministic trend is

$$
\Delta^\delta y_t = g \left( \frac{t}{T} \right) + u_t, \quad t = 1, ..., T, \quad (4)
$$

where $u_t$ is a sequence of uncorrelated, homoscedastic random variables and, for any real $\delta$, $\Delta^\delta_t$ is the truncated fractional differencing operator,

$$
\Delta^\delta_t = \sum_{j=0}^{t-1} \alpha_j(\delta) L^j, \quad t \geq 1, \quad (5)
$$

the $\alpha_j(\delta)$ being coefficients in the (possibly formal) expansion

$$
(1 - z)\delta = \sum_{j=0}^{\infty} \alpha_j(\delta) z^j,
$$

namely

$$
\alpha_j(\delta) = \frac{\Gamma(j - \delta)}{\Gamma(-\delta) \Gamma(j + 1)}. \quad (6)
$$

The truncation in (5) reflects non-observability of $y_t$ when $t \leq 0$, and avoids explosion of the moving average representation of (4) when $\delta_0 \geq 1/2$, the nonstationary region for $\delta_0$; it is this region with which we will be concerned.

One such $\delta_0$ has assumed wide empirical importance in connection with a variety of econometric models, the unit root case $\delta_0 = 1$, when (4) becomes

$$
(1 - L)y_t = g \left( \frac{t}{T} \right) + u_t, \quad t = 1, ..., T. \quad (7)
$$

The bulk of the econometric literature nests the unit root in autoregressive structures, which suggests treating (7) as a special case of

$$
(1 - \rho L)y_t = g \left( \frac{t}{T} \right) + u_t, \quad t = 1, ..., T, \quad (8)
$$
rather than (4). The autoregressive unit root literature suggests that estimates of 
\( \rho \) in (8) will have a nonstandard limit distribution under (7), but a normal one in 
the "stationary" region \( |\rho| < 1 \). By contrast we can anticipate, for example from 
literature concerning (4) with \( g(x) \) \emph{a priori} constant, that estimates of \( \delta_0 \) such as 
one\s optimizing an approximate pseudo-Gaussian likelihood, and Wald and other 
test statistics, will enjoy standard asymptotics, with the usual parametric conver-
gence rate, \( \sqrt{T} \), whatever the value of \( \delta_0 \), due essentially to smoothness properties of 
the fractional operator; tests are also expected to have the classical local efficiency 
properties. While (4) cannot, unlike (8), describe "explosive" behaviour (occurring 
when \( |\rho| > 1 \)), it describes a continuum of stochastic trends indexed by \( \delta_0 \). A conse-
quence of the \( T \)–dependence in \( g(t/T) \) is that the left side of (4) is also \( T \)-dependent, 
so the \( y_t = y_{t,T} \) in fact form a triangular array, but in common with the bulk of lit-
erature concerning versions of (1) we suppress reference to \( T \). The model (4) (which 
nests (1) with iid \( u_t \) on taking \( \delta_0 = 0 \)) supposes that the fractional filtering of \( y_t \) suc-
cessfully eliminates correlation, but possibly leaves a trend which we are not prepared 
to parameterize.

To provide greater generality than (4), the paper in fact considers the extended 
model

\[
\xi_t (L; \delta_0, \theta_0) y_t = g \left( \frac{t}{T} \right) + u_t, \quad t = 1, \ldots, T, \tag{9}
\]

where \( \theta_0 \) is an unknown \( p \)-dimensional column vector and

\[
\xi_t (z; \delta, \theta) = \sum_{j=0}^{t-1} \xi_j(\delta, \theta) z^j, \quad t \geq 1,
\]

where the \( \xi_j(\delta, \theta) \) are coefficients in the possibly formal expansion

\[
\xi (z; \delta, \theta) = \sum_{j=0}^{\infty} \xi_j(\delta, \theta) z^j,
\]

such that

\[
\xi (z; \delta, \theta) = (1 - z)^{\delta} \xi (z; \theta),
\]
where

$$\zeta (z; \theta) = \sum_{j=0}^{\infty} \zeta_j(\theta) z^j$$

is a known function of $z$ and $\theta$ that is at least continuous, and nonzero for $z$ on or inside the unit circle in the complex plane. When $\zeta (z; \theta) \equiv 1$, we have $\xi (z; \delta, \theta) = (1 - z)^{\delta}$. Leading examples of $\zeta (z; \theta)$ are stationary and invertible autoregressive moving average operators of known degree, for example the first order autoregressive operator $\zeta (z; \theta) = 1 - \theta z$, with $\theta$ here a scalar such that $|\theta| < 1$. In general $\zeta$ leaves the essential memory or degree of nonstationarity $\delta_0$ unchanged but allows otherwise richer dependence structure.

It would be possible to consider in effect a nonparametric $\zeta$, $\zeta (z)$, satisfying smoothness assumptions only near $z = 1$, and hence a "semiparametric" operator on $y_t$. This would lead to an estimate of $\delta_0$ with only a nonparametric convergence rate. However, establishing the parametric, $\sqrt{T}$, rate for estimating $\delta_0$ and $\theta_0$ seems actually more challenging and delicate, because of the presence of the nonparametric $g (t/T)$ in (9), estimates of which converge more slowly than $\sqrt{T}$. In particular, proving consistency requires establishing that certain (stochastic and deterministic) contributions to residuals, whose squares make up the objective function minimized by the parameter estimates, are negligible uniformly over the parameter space; these contributions are of larger order than would be the case with a parametric trend (and this fact also explains why we find ourselves unable to choose the parameter space for $\delta_0$ as large as is possible with a parametric trend). Then, corresponding contributions to scores evaluated at $\delta_0$, $\theta_0$ are also of larger order than in the parametric trend case and have to be shown to be negligible after being normalized by $\sqrt{T}$, rather than by a slower, nonparametric, rate, in order to prove asymptotic normality of the parameter estimates with $\sqrt{T}$ rate. Of course, the strong dependence in $y_t$ also impacts on the conditions, due to non-summability of certain weight sequences.
The following section proposes estimates of $\delta_0$ and $\theta_0$, and establishes their consistency and asymptotic normality, the proofs appearing in Appendices A and B. Section 3 develops unit root tests based on Wald, pseudo log likelihood ratio and Lagrange multiplier principles. Section 4 proposes estimates of $g(x)$ and establishes their asymptotic properties. A small Monte Carlo study of finite-sample performance is contained in Section 5. Section 6 concludes by describing further issues that might be considered.

2. ESTIMATION OF DEPENDENCE PARAMETERS

Were $g(x) \equiv 0$ in (9) a priori, a natural method of estimating $\delta_0$ and $\theta_0$ would be conditional-sum-of-squares, which approximates Gaussian pseudo-maximum-likelihood estimation. We modify this method by employing residuals, which requires preliminary estimation of $g(x)$. Note that under the conditions imposed below, $g(t/T) - g((t-1)/T) = O(T^{-1})$ \(2 \leq t \leq T\), so we could instead consider proceeding by first differencing in (9) as this would effectively eliminate the deterministic trend; however this also induces a moving average unit root on the right hand side.

Let $k(x), x \in \mathcal{R}$, be a user-chosen kernel function and $h$ a user-chosen positive bandwidth number. For any $\delta, \theta$ write $\xi_{t;\delta \theta}(z) = \xi_t(z; \delta, \theta)$ and introduce

$$
\hat{g}_{\delta \theta}(x) = \sum_{s=1}^{T} \xi_{s;\delta \theta}(L) y_s k\left(\frac{x - s/T}{h}\right) / \sum_{s=1}^{T} k\left(\frac{x - s/T}{h}\right),
$$

for any $x \in [0, 1]$. The corresponding estimate of Priestley and Chao (1972) type replaces the denominator by $Th$, but we prefer to use weights (of the $\xi_{s;\delta \theta}(L) y_s$) that exactly sum to 1 for all $x$. Define residuals

$$
\hat{u}_t(\delta, \theta) = \xi_{t;\delta \theta}(L) y_t - \hat{g}_{\delta \theta}(t/T) = \sum_{s=1}^{T} \frac{(\xi_{t;\delta \theta}(L) y_t - \xi_{s;\delta \theta}(L) y_s) k_{ts}}{\sum_{s=1}^{T} k_{ts}},
$$

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where $k_{ts} = k ((t - s) / Th)$. Denote by $\nabla_1$, $\nabla_2$ chosen real numbers such that $\nabla_1 < \nabla_2$, write $\nabla = [\nabla_1, \nabla_2]$, and let $\Theta$ be a suitably chosen compact subset of $\mathcal{R}^p$. We estimate $\delta_0$ and $\theta_0$ by

$$
\left( \hat{\delta}, \hat{\theta} \right) = \arg \min_{\delta \in \nabla, \theta \in \Theta} Q(\delta, \theta),
$$

(12)

where

$$
Q(\delta, \theta) = \sum_{t=1}^{T} \hat{u}_t^2(\delta, \theta).
$$

(13)

We first establish consistency of $\hat{\delta}$, $\hat{\theta}$, under the following regularity conditions.

**Assumption 1**

The $u_t$ are stationary and ergodic with finite fourth moment, and

$$
E \left( u_t | \mathcal{F}_{t-1} \right) = 0, \quad E \left( u_t^2 | \mathcal{F}_{t-1} \right) = \sigma^2
$$

almost surely, where $\mathcal{F}_t$ is the $\sigma$-field of events generated by $\varepsilon_s$, $s \leq t$, and conditional on $\mathcal{F}_{t-1}$ 3rd and 4th moments of $u_t$ equal corresponding unconditional moments.

**Assumption 2**

(i) $\theta_0 \in \Theta$;

(ii) $|\zeta(z; \theta)| \neq |\zeta(z; \theta_0)|$ for all $\theta \neq \theta_0$, $\theta \in \Theta$, on a $z$-set of positive Lebesgue measure;

(iii) for all $\theta \in \Theta$ and real $\lambda$, $\zeta(e^{i\lambda}; \theta)$ is differentiable in $\lambda$ with derivative in $\text{Lip}(\zeta)$, $\zeta > 1/2$;

(iv) for all $\lambda$, $\zeta(e^{i\lambda}; \theta)$ is continuous in $\theta$;

(v) for all $\theta \in \Theta$, $|\zeta(z; \theta)| \neq 0$, $|z| \leq 1$.

Assumption 1 is weaker than imposing independence and identity of distribution of $u_t$, and Assumption 2 is standard from the literature on parametric short memory models since Hannan (1973), ensuring identifiability of $\theta_0$ and easily covering stationary and invertible moving averages. These assumptions correspond to ones of Hualde and Robinson (2011), who established consistency of the same kind of estimates when
$g(x) \equiv 0$ in (9) \textit{a priori}. In that setting they were able to choose the set of admissible memory parameters (our $[\nabla_1, \nabla_2]$) arbitrarily large, to simultaneously cover stationary, nonstationary, invertible and non-invertible values. This is more difficult, and perhaps impossible, to achieve in the presence of the unknown, nonparametric $g$ in (9), which can only be estimated with a slow rate of convergence, and we impose:

\textbf{Assumption 3}

$$
\nabla_1 > 3/4, \nabla_2 < 5/4, \quad (14)
$$

$$
\delta_0 \in \nabla. \quad (15)
$$

As can be inferred from the proof of Theorem 1, strictly what is required instead of (14) is the weaker condition $\nabla_2 - \delta_0 < 1/2$, but since $\delta_0$ is known from (15) only to be no less than $\nabla_1$ the restriction $\nabla_2 - \nabla_1 < 1/2$ implied by (14) is appropriate. Inspection of our proofs indicates that they go through with $\nabla$ in Assumption 3 replaced by $[\varkappa, \varkappa + \omega]$, for any real $\varkappa$ and for $\omega \in (0, 1/2)$ (for example a subset of the stationary and invertible region $(-1/2, 1/2)$), but for the sake of clarity we fix on (14), which seems among the more empirically realistic possibilities, and covers the unit root case $\delta_0 = 1$.

We also need conditions on $g$, $k$ and $h$.

\textbf{Assumption 4}

\textit{The function $g(x)$ is twice boundedly differentiable on $[0, 1]$ and $g(0) = 0$.}

\textbf{Assumption 5}

\textit{The function $k(x)$ is even, differentiable at all but possibly finitely many $x$, with derivative $k'(x)$, and}

$$
\int_{\mathbb{R}} k(x)dx = 1,
$$

$$
k(x) = O((1 + x^{2+\eta})^{-1}), \quad k'(x) = O((1 + |x|^{1+\eta})^{-1}), \quad \text{some } \eta > 0.
$$

\textbf{Assumption 6}
As $T \to \infty$, the positive-valued sequence $h = h_T$ satisfies:

$$(Th)^{-1} + T^{2(\nabla_2 - \nabla_1)}h^3 \to 0.$$  \hfill (16)

Assumption 5 is virtually costless, covering many of the usual kernel choices. Assumption 6, however, represents a trade-off with Assumption 3: in the latter, $\nabla_2 - \nabla_1$ is desirably as close to $1/2$ as possible, but as it approaches $1/2$ from below the range of $h$ satisfying Assumption 6 reduces to $(Th)^{-1} + Th^3 \to 0$.

**Theorem 1**

Let (9) and Assumptions 1-6 hold. Then as $T \to \infty$,

$$\hat{\delta} \to_p \delta_0, \quad \hat{\theta} \to_p \theta_0.$$

The proof is in Appendix A. Asymptotic normality requires two further assumptions.

**Assumption 7**

(i) $\delta_0 \in (\nabla_1, \nabla_2)$; $\theta_0$ is an interior point of $\Theta$.

(ii) for all real $\lambda$, $\zeta(e^{i\lambda}; \theta)$ is twice continuously differentiable in $\theta$ on a closed neighbourhood of radius $< 1/2$ about $\theta_0$;

(iii) the matrix

$$\Omega = \begin{pmatrix}
\pi^2/6 & -\sum_{j=1}^{\infty} \psi_j' (\theta_0) / j \\
-\sum_{j=1}^{\infty} \psi_j (\theta_0) / j & \sum_{j=1}^{\infty} \psi_j (\theta_0) \psi_j' (\theta_0)
\end{pmatrix}$$

is non-singular, where

$$\psi_j (\theta) = \sum_{k=0}^{j-1} \phi_k (\theta) \frac{\partial \zeta_{j-k} (\theta)}{\partial \theta},$$

the $\phi_j (\theta)$ being coefficients in the expansion

$$\phi_j (z; \theta) = \zeta (z; \theta)^{-1} = \sum_{j=0}^{\infty} \phi_j (\theta) z^j.$$

This condition again is based on one of Hualde and Robinson (2011), but is similar to others in the literature, and practically unrestricted. However we have to strengthen the first component of Assumption 6 on $h$. 

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**Assumption 8**

As $T \to \infty$, $Th^2 / (\log T)^2 \to \infty$.

**Theorem 2**

*Let (9) and Assumptions 1-8 hold. Then as $T \to \infty$*

$$T^{1/2} \left( \begin{array}{c} \hat{\delta} - \delta_0 \\ \hat{\theta} - \theta_0 \end{array} \right) \to_d \mathcal{N}(0, \Omega^{-1}).$$

The proof is in Appendix B. Note that the same limit distributions results when $g$ is known or replaced by a parametric function. In the special case (4) of (9), we deduce that as $T \to \infty$

$$T^{1/2} \left( \hat{\delta} - \delta_0 \right) \to_d \mathcal{N}(0, 6/\pi^2).$$

**3. UNIT ROOT TESTING**

We first establish Wald tests for $\delta_0 = 1$ in (9), based on Theorem 2. Define

$$\Omega(\theta) = \begin{pmatrix} \pi^2/6 & -\sum_{j=1}^{\infty} \psi_j'(\theta)/j \\ -\sum_{j=1}^{\infty} \psi_j(\theta)/j & \sum_{j=1}^{\infty} \psi_j(\theta) \psi_j'(\theta) \end{pmatrix}$$

and denote by $\hat{\Omega}^{(1,1)}$ the element in the top left hand corner of $\Omega(\hat{\theta})^{-1}$. Put

$$W = T^{1/2} \left( \hat{\delta} - 1 \right) / \hat{\Omega}^{(1,1)1/2}.$$

**Theorem 3**

*Let $\delta_0 = 1$ in (9) and let Assumptions 1-8 hold. Then as $T \to \infty$,*

$$W \to_d \mathcal{N}(0, 1). \quad (17)$$

The theorem follows from Theorem 2 and $\Omega(\hat{\theta}) \to_p \Omega$, where the latter is implied by the proof of Theorem 2. We can reject the unit root null against more non-stationary alternatives when $W$ falls in the appropriate upper tail of the standard
normal density, and reject against less nonstationary alternatives when it falls in the appropriate lower tail.

Pseudo-log likelihood ratio tests can also be constructed. Define

$$\tilde{\theta} = \arg\min_{\theta \in \Theta} Q_T(1, \theta),$$

and

$$LR = \log \frac{Q_T(1, \tilde{\theta})}{Q_T(\delta, \tilde{\theta})}.$$  

**Theorem 4**

*Let $\delta_0 = 1$ in (9) and let Assumptions 1-8 hold. Then as $T \to \infty,$

$$LR \to_d \chi^2_1.$$*

The proof is standard, given Theorem 2 and a central limit theorem for $\tilde{\theta}$ (see e.g. Hannan (1973), or implied by Hualde and Robinson (2011)).

Though it of course does not use $\hat{\delta}, \hat{\theta},$ for completeness we also present a Lagrange multiplier-type test, as it and the Wald and pseudo-log likelihood tests are expected to have equal local power. Robinson (1994) developed Lagrange multiplier tests for unit root and other hypotheses against fractional alternatives for the disturbances in multiple linear regression models. The stress there was on frequency-domain tests, but starting from an initial time-domain statistic, and to avoid introducing considerable additional notation we stay in the time domain here.

Writing $\partial = \partial / \partial (\delta, \theta)^t,$ from (13)

$$\partial Q(\delta, \theta) = \frac{2}{T} \sum_{t=1}^T \hat{u}_t(\delta, \theta) \partial \hat{u}_t(\delta, \theta),$$

where

$$\partial \hat{u}_t(\delta, \theta) = \xi_{t\delta \theta}(L) y_t - \partial \hat{g}_{\delta \theta}(t/T) = \sum_{s=1}^T \frac{(\partial \xi_{t\delta \theta}(L) y_t - \partial \xi_{s\delta \theta}(L) y_s) k_{ts}}{\sum_{s=1}^T k_{ts}}.$$
in which

\[ \partial \xi_{t \delta \theta} (z) = \sum_{j=0}^{t-1} \partial \xi_j (\delta, \theta) z^j, \partial \xi_j (\delta, \theta) = \sum_{i=0}^{j} \left( \frac{\partial \alpha_i (\delta) / \partial \delta}{\partial \xi_j (\theta) / \partial \theta} \right) \left( \frac{\partial \alpha_i (\theta) / \partial \delta}{\partial \xi_j (\theta) / \partial \theta} \right) \right). \]  

(21)

In fact

\[ \partial \xi_{t \delta \theta} (1, \theta) = \left( \sum_{j=0}^{t-1} \sum_{l=1}^{j} \left( \frac{\partial \alpha_i (\theta) / \partial \delta}{\partial \xi_j (\theta) / \partial \theta} \right) \left( y_{t-j} - y_{t-j-1} \right) \right) \left( \sum_{j=0}^{t-1} \left( \frac{\partial \alpha_i (\theta) / \partial \delta}{\partial \xi_j (\theta) / \partial \theta} \right) \left( y_{t-j} - y_{t-j-1} \right) \right) \right). \]

Define

\[ LM = \frac{T}{4} \partial Q(1, \tilde{\theta})^\Omega \left( \tilde{\theta} \right)^{-1} \partial Q(1, \tilde{\theta}), \]

with \( \tilde{\theta} \) given by (18). The proof of the following theorem is straightforward given the sentence after Theorem 3.

**Theorem 5**

Let \( \delta_0 = 1 \) in (9) and let Assumptions 1-8 hold. Then as \( T \to \infty \),

\[ LM \to d \chi_1^2. \]

### 4. NONPARAMETRIC REGRESSION ESTIMATION

We can base estimation of \( g(x) \) on our estimates of \( \hat{\delta}, \hat{\theta} \) and (10), but in view of the stringent conditions on the bandwidth \( h \) in Theorems 1 and 2 we allow use of a possibly different bandwidth, \( b \), in

\[ \tilde{g}_{s \delta \theta} (x) = \sum_{s=1}^{T} \xi_{s \delta \theta} (L) y_s k \left( \frac{x - s/T}{b} \right) / \sum_{s=1}^{T} k \left( \frac{x - s/T}{b} \right), \]

(22)

We provide a multivariate central limit theorem for \( \tilde{g}_{s \delta \theta} (\tau_i), i = 1, 2, \ldots, q \), where \( \tau_i, i = 1, 2, \ldots, q \), are distinct fixed points, imposing also:

**Assumption 9**


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As $T \to \infty$, $(bT)^{-1} + b^5 T \to 0$.

The proof of the following theorem is omitted as univariate and multivariate central limit theorems for the $\tilde{g}_{\delta_0 \theta_0}(x_i)$ are already in the literature (see e.g. Benedetti (1977), Robinson (1997)) and from Theorem 2 it is readily shown that $\tilde{g}(x) - \tilde{g}_{\delta_0 \theta_0}(x) = O_p(T^{-1/2})$ for all $x$.

**Theorem 6**

Let (9) and Assumptions 1-9 hold. Then as $T \to \infty$, the $(bT)^{1/2} (\tilde{g}_{\delta_0 \theta_0}(\tau_i) - g(\tau_i))$, $i = 1, 2, ..., q$, converge in distribution to independent $N \left(0, \sigma^2 \int k(x)^2 dx \right)$ random variables, where $\sigma^2$ is consistently estimated by

$$\hat{\sigma}^2 = Q \left(\hat{\delta}, \hat{\theta} \right).$$

This is the same limit distribution as results if $\delta_0$ and $\theta_0$ are known, i.e. the same as in the model (1) with iid $u_t$.

**5. FINITE-SAMPLE PERFORMANCE**

A small Monte Carlo study was carried out to investigate the finite-sample behaviour of our parameter estimates, and of one of our unit root tests. To generate data, in (9) we took $g(x) = \sin(2\pi x), p = 1, \zeta(z; \theta) = 1 - \theta z$ (so $y_t$ was a FARIMA$(1, \delta_0, 0)$), for various values of $\delta_0$ and $\theta_0$, and $\varepsilon_t$ standard normally distributed. Throughout, parameter estimates were obtained taking $k$ to be the standard normal kernel. All results are based on 1000 replications.

Tables 1-3 contain Monte Carlo biases $b_{\delta}$ and $b_{\theta}$ of $\hat{\delta}$ and $\hat{\theta}$ and corresponding Monte Carlo standard deviations $s_{\delta}$ and $s_{\theta}$, for two $(\delta_0, \theta_0)$ choices each, the sets $\nabla$ and $\Theta$ varying with these. In Table 1 $(\delta_0, \theta_0) = \left(\frac{5}{8}, \frac{1}{2}\right)$ ($\nabla = [.4, 1.2], \Theta = [.1, .9]$) and $(\delta_0, \theta_0) = \left(\frac{7}{8}, 0\right)$ ($\nabla = [.4, 1.1], \Theta = [-.5,.5]$); in Table 2 $(\delta_0, \theta_0) = (1, \frac{1}{2})$ ($\nabla = [.5, 1.5], \Theta = [.1, .9]$) and $(\delta_0, \theta_0) = (1, 0)$ ($\nabla = [.5, 1.5], \Theta = [-.5,.5]$); in Table
3 \((\delta_0, \theta_0) = \left(\frac{13}{8}, \frac{1}{2}\right)\) \((\nabla = [.9, 2.1], \Theta = [.1, .9])\) and \((\delta_0, \theta_0) = \left(\frac{9}{8}, 0\right)\) \((\nabla = [.9, 1.5], \Theta = [-.5, .5])\). All but one of these \(\delta_0\), and none of these \(\nabla\), satisfy Assumption 3. Note that in the cases \(\theta_0 = 0\), one of which is included in each table, \(y_t\) reduces to a FARIMA\((0, \delta_0, 0)\), but we suppose that the practitioner does not know this. Three different \((T, Th)\) combinations were employed: \((250, 20), (600, 50)\) and \((1000, 80)\); these choices only partly represent an \(h\) sequence obeying our assumptions, because though \(Th\) increases with \(T\), \(h\) takes successive values 12.5, 12, 12.5. In the tables, the excessive biases for the smaller sample sizes whenever \(\theta_0 \neq 0\) may partly be due to the overly large \(\nabla\), as well as the deleterious effect of the nonparametric estimation, and the estimation procedure having some difficulty in distinguishing the long and short memory effects. The biases when \(\theta_0 = 0\) are much smaller, and generally both biases and standard deviations diminish with increasing \(T\), while there is a high stability across corresponding elements of the tables, especially Tables 1 and 2.

Table 1: Bias and standard deviation of \(\left(\hat{\delta}, \hat{\theta}\right), \left(\delta_0, \theta_0\right) = \left(\frac{5}{8}, \frac{1}{2}\right), \left(\frac{7}{8}, 0\right)\).

<table>
<thead>
<tr>
<th>((\delta_0, \theta_0))</th>
<th>((\frac{5}{8}, \frac{1}{2}))</th>
<th>((\frac{7}{8}, 0))</th>
</tr>
</thead>
<tbody>
<tr>
<td>(T)</td>
<td>(b_\delta)</td>
<td>(s_\delta)</td>
</tr>
<tr>
<td>250</td>
<td>0.3770</td>
<td>0.0537</td>
</tr>
<tr>
<td>600</td>
<td>0.1905</td>
<td>0.0436</td>
</tr>
<tr>
<td>1000</td>
<td>0.0669</td>
<td>0.0424</td>
</tr>
</tbody>
</table>

Table 2: Bias and standard deviation of \(\left(\hat{\delta}, \hat{\theta}\right), \left(\delta_0, \theta_0\right) = (1, \frac{1}{2}), (1, 0)\).

<table>
<thead>
<tr>
<th>((\delta_0, \theta_0))</th>
<th>( (1, \frac{1}{2}))</th>
<th>((1, 0))</th>
</tr>
</thead>
<tbody>
<tr>
<td>(T)</td>
<td>(b_\delta)</td>
<td>(s_\delta)</td>
</tr>
<tr>
<td>250</td>
<td>0.3713</td>
<td>0.0555</td>
</tr>
<tr>
<td>600</td>
<td>0.1925</td>
<td>0.0437</td>
</tr>
<tr>
<td>1000</td>
<td>0.0674</td>
<td>0.0443</td>
</tr>
</tbody>
</table>
Table 3: Bias and standard deviation of \( \left( \hat{\delta}, \hat{\theta} \right) \), \( (\delta_0, \theta_0) = \left( \frac{13}{8}, \frac{1}{2} \right), \left( \frac{9}{8}, 0 \right) \).

<table>
<thead>
<tr>
<th>( (\delta_0, \theta_0) )</th>
<th>( \left( \frac{13}{8}, \frac{1}{2} \right) )</th>
<th>( \left( \frac{9}{8}, 0 \right) )</th>
</tr>
</thead>
<tbody>
<tr>
<td>( T )</td>
<td>( b_\delta )</td>
<td>( s_\delta )</td>
</tr>
<tr>
<td>250</td>
<td>0.2666</td>
<td>0.0488</td>
</tr>
<tr>
<td>600</td>
<td>0.1853</td>
<td>0.0485</td>
</tr>
<tr>
<td>1000</td>
<td>0.0612</td>
<td>0.0457</td>
</tr>
</tbody>
</table>

Table 4 contains Monte Carlo sizes and powers for the LR unit root test described in Section 3, based on nominal 1% and 5% levels. Sizes were obtained using \( (\delta_0, \theta_0) = (1, \frac{1}{2}) \) and \( (T, T_h) = (250, 12), (600, 25), (1000, 52) \), powers using \( (\delta_0, \theta_0) = \left( \frac{15}{16}, \frac{1}{2} \right) \) and \( \left( \frac{17}{16}, \frac{1}{2} \right) \), with \( (T, T_h) = (250, 20), (600, 50), (1000, 80) \). Considering the serious biases found in Tables 1-3 when \( \theta_0 = \frac{1}{2} \), the sizes in Table 4 do not seem bad, and they do improve with increasing \( T \). Only one alternative in either direction from the unit root null is considered, but given that these are both close to 1 the differences between powers and corresponding sizes seem quite satisfactory, with slightly the greater sensitivity when \( \delta_0 = \frac{15}{16} \), and again there is improvement as \( T \) increases.

Table 4: Sizes and powers at nominal 1% and 5% levels

<table>
<thead>
<tr>
<th>( H_0 )</th>
<th>Size, ( (\delta_0, \theta_0) = (1, \frac{1}{2}) )</th>
<th>( \alpha = 1% )</th>
<th>( \alpha = 5% )</th>
</tr>
</thead>
<tbody>
<tr>
<td>( T )</td>
<td>( \alpha = 1% )</td>
<td>( \alpha = 5% )</td>
<td></td>
</tr>
<tr>
<td>250</td>
<td>0.019</td>
<td>0.068</td>
<td></td>
</tr>
<tr>
<td>600</td>
<td>0.014</td>
<td>0.061</td>
<td></td>
</tr>
<tr>
<td>1000</td>
<td>0.012</td>
<td>0.045</td>
<td></td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>( H_1 )</th>
<th>Power, ( (\delta_0, \theta_0) = \left( \frac{15}{16}, \frac{1}{2} \right) )</th>
<th>( \alpha = 1% )</th>
<th>( \alpha = 5% )</th>
</tr>
</thead>
<tbody>
<tr>
<td>( T )</td>
<td>( \alpha = 1% )</td>
<td>( \alpha = 5% )</td>
<td></td>
</tr>
<tr>
<td>250</td>
<td>0.133</td>
<td>0.284</td>
<td>0.102</td>
</tr>
<tr>
<td>600</td>
<td>0.303</td>
<td>0.562</td>
<td>0.211</td>
</tr>
<tr>
<td>1000</td>
<td>0.504</td>
<td>0.725</td>
<td>0.479</td>
</tr>
</tbody>
</table>
6. FINAL REMARKS

The paper has justified large sample inference on the fractional and short memory parameters and nonparametric regression function in a semiparametric model incorporating nonstationary stochastic and deterministic trends. For parametric inference, the restrictions on the admissible memory parameter interval and the range of bandwidths are relatively strong, due to the presence of the nonparametric function and the extent of the time series dependence; possibly one or both these restrictions could be relaxed by means of higher-order kernels, as have been used elsewhere in the semiparametric econometric literature. A variety of further avenues might be explored. As always when nonparametric estimation is involved, bandwidth choice is a practical issue, though possibly a less acute one than in the stochastic design setting in which the density of explanatory variables varies over their support. In our Monte Carlo study only one value of $h$ was used for each $T$, but sensitivity of estimates and tests to $h$, and $b$, can be gauged by carrying out the computations over a range of choices. With respect to automatic rules, in the model (1) a cross–validation choice of $b$ is known to minimize mean integrated squared error (MISE), and we can extend this property to our setting, using $\hat{\delta}, \hat{\theta}$, though as usual the minimum-MSE rate does not quite satisfy conditions (our Assumption 9) for asymptotic normality about $g$; for $h$, as is familiar in the semiparametric literature the minimum-MISE rate is clearly excluded by the conditions (our Assumption 8) for asymptotic normality of parameter estimates, and a more appropriate goal may be to make a selection that matches the orders of the next two terms after the normal distribution function in an Edgeworth expansion for distribution function of $\hat{\delta}, \hat{\theta}$, and thereby minimizes the departure from the normal limit and leads to better-sized tests and more accurate interval estimates; in some settings this problem has a neat solution, but we do not know whether this is the case in ours. Bootstrapping is also likely to improve finite-sample properties.
Inference issues that might be investigated include testing constancy or other parametric restrictions on $g(x)$. Possible model extensions that require non-trivial further work include adding a nonparametric function of explanatory variables to $g(t/T)$ in (9), and allowing for unconditional or conditional heteroscedasticity in $u_t$. Our work might also be extended to a panel setting, including individual effects and possible cross-sectional dependence.

ACKNOWLEDGMENTS

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APPENDIX A

Proof of Theorem 1

The model (9) refers to $y_t$, $u_t$, and $g(t/T)$ only for $t \geq 1$ so we can set $y_t = u_t = 0$, $t \leq 0$ and $g(x) = 0$, $x < 0$. Then for $t = 1, ..., T$

$$\xi_t (L; \delta, \theta) y_t = \xi (L; \delta, \theta) y_t$$

and

$$y_t = \xi (L; \delta_0, \theta_0)^{-1} \left( g \left( \frac{t}{T} \right) + u_t \right),$$

so

$$\xi_t (L; \delta, \theta) y_t = \xi (L; \delta, \theta) \xi (L; \delta_0, \theta_0)^{-1} y_t$$

$$= (1 - L)^{\delta - \delta_0} \tau (z; \theta) \left( g \left( \frac{t}{T} \right) + u_t \right),$$

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where
\[ \tau(z; \theta) = \zeta(z; \theta) \zeta(z; \theta_0)^{-1} = \sum_{j=0}^{\infty} \tau_j(\theta) z^j. \]

From Zygmund (1977, p. 46), Assumption 2 implies that the Fourier coefficients \( \tau_j(\theta) \) satisfy
\[ \sup_{\theta} |\tau_j(\theta)| = O(j^{-1-\epsilon}). \tag{A.1} \]

The Fourier coefficients \( \chi_j(\delta, \theta) \) of
\[ (1 - z)^{\delta - \delta_0} \tau(z; \theta) = \sum_{j=0}^{\infty} \chi_j(\delta, \theta) z^j \]
are given by
\[ \chi_j(\delta, \theta) = \sum_{l=0}^{j} \tau_l(\theta) \alpha_{j-l}(\delta - \delta_0). \]
(Note that \( \alpha_j(0) = \tau_j(\theta_0) = \chi_j(\delta_0, \theta_0) \equiv 0, j \geq 1 \). From (6), uniformly in \( \delta \in [\nabla_1, \nabla_2] \setminus \{\delta_0\} \)
\[ \alpha_j(\delta - \delta_0) = O(j^{\beta_0 - \delta - 1}), \text{ as } j \to \infty, \tag{A.2} \]
and so, using also (A.1)
\[
\begin{align*}
|\chi_j(\delta, \theta)| &\leq \left| \sum_{l=0}^{\lfloor j/2 \rfloor} \tau_l(\theta) \alpha_{j-l}(\delta - \delta_0) \right| + \left| \sum_{l=\lfloor j/2 \rfloor + 1}^{j} \tau_l(\theta) \alpha_{j-l}(\delta - \delta_0) \right| \\
&\leq K j^{\beta_0 - \delta - 1} \sum_{l=0}^{\infty} |\tau_l(\theta)| + K j^{1-\epsilon} \sum_{l=0}^{j} |\alpha_l(\delta - \delta_0)| \\
&\leq K j^{\beta_0 - \delta - 1} + K j^{\beta_0 - \delta - 1 - \epsilon} \leq K j^{\beta_0 - \delta - 1} \tag{A.3}
\end{align*}
\]
uniformly in \( \delta, \theta \), where \( K \) throughout denotes a generic finite, positive constant.
Also for future use note that from (6), uniformly in \( \delta \in [\nabla_1, \nabla_2] \setminus \{\delta_0\}, \theta \in \Theta, \)
\[ |\alpha_j(\delta - \delta_0) - \alpha_{j+1}(\delta - \delta_0)| = O(j^{\beta_0 - \delta - 2}), \text{ as } j \to \infty, \tag{A.4} \]
\[
\left| \chi_j(\delta, \theta) - \chi_{j+1}(\delta, \theta) \right| \leq \left| \sum_{l=0}^{j} \tau_l(\theta) (\alpha_{j-l}(\delta - \delta_0) - \alpha_{j+1-l}(\delta - \delta_0)) \right| + \left| \tau_{j+1}(\theta) \right|
\]
\[
\leq K_j^{\delta_0 - \delta - 2} \sum_{l=0}^{\infty} \left| \tau_l(\theta) \right| + K_j^{-1 - \zeta} \sum_{l=1}^{\infty} l^{\delta_0 - \delta - 2} + K_j^{-1 - \zeta}
\]
\[
\leq K_j^{\max(\delta_0 - \delta, 1 - \zeta) - 2}. \quad (A.5)
\]

With the abbreviations
\[
\chi_{t\delta\theta} = \sum_{j=0}^{t-1} \chi_j(\delta, \theta) L^j, \quad g_t = g \left( \frac{t}{T} \right), \quad k_t = \frac{1}{T h} \sum_{s=1}^{T} k_{ts}
\]
we have from (11)
\[
\hat{u}_t(\delta, \theta) = \frac{1}{T h} \sum_{s=1}^{T} (\chi_{t\delta\theta} g_t - \chi_{s\delta\theta} g_s) k_{ts}/k_t + \frac{1}{T h} \sum_{s=1}^{T} (\chi_{t\delta\theta} u_t - \chi_{s\delta\theta} u_s) k_{ts}/k_t
\]
\[
= \chi_{t\delta\theta} u_t + D_{t\delta\theta} - S_{t\delta\theta},
\]
where
\[
D_{t\delta\theta} = \frac{1}{T h} \sum_{s=1}^{T} (\chi_{t\delta\theta} g_t - \chi_{s\delta\theta} g_s) k_{ts}/k_t
\]
and
\[
S_{t\delta\theta} = \frac{1}{T h} \sum_{s=1}^{T} \chi_{s\delta\theta} u_s k_{ts}/k_t
\]
are respectively the deterministic and stochastic errors contributing to the residual, that are absent when \( g(t/T) \equiv 0 \) in (9). Thus
\[
Q(\delta, \theta) = \frac{1}{T} \sum_{t=1}^{T} (\chi_{t\delta\theta} u_t + D_{t\delta\theta} - S_{t\delta\theta})^2
\]
\[
= \frac{1}{T} \sum_{t=1}^{T} (\chi_{t\delta\theta} u_t)^2 + \frac{1}{T} \sum_{t=1}^{T} D_{t\delta\theta}^2 + \frac{1}{T} \sum_{t=1}^{T} S_{t\delta\theta}^2
\]
\[
+ \frac{2}{T} \sum_{t=1}^{T} (\chi_{t\delta\theta} u_t) D_{t\delta\theta} - \frac{2}{T} \sum_{t=1}^{T} (\chi_{t\delta\theta} u_t) S_{t\delta\theta}
\]
\[
- \frac{2}{T} \sum_{t=1}^{T} D_{t\delta\theta} S_{t\delta\theta}. \quad (A.6)
\]
Hualde and Robinson (2011) show that the estimates minimizing
\[ \frac{1}{T} \sum_{t=1}^{T} (\chi_{t\delta\theta}u_t)^2 \]  
are consistent for \( \delta_0, \theta_0 \). From their proof it suffices to show that as \( T \to \infty \),
\[ \sup \frac{1}{T} \sum_{t=1}^{T} D^2_{t\delta\theta} \to 0, \]  
\[ \sup \frac{1}{T} \sum_{t=1}^{T} S^2_{t\delta\theta} \to p 0, \]  
\[ \sup \left| \frac{1}{T} \sum_{t=1}^{T} (\chi_{t\delta\theta}u_t) D_{t\delta\theta} \right| \to p 0, \]  
\[ \sup \left| \frac{1}{T} \sum_{t=1}^{T} (\chi_{t\delta\theta}u_t) S_{t\delta\theta} \right| \to p 0, \]  
\[ \sup \left| \frac{1}{T} \sum_{t=1}^{T} D_{t\delta\theta} S_{t\delta\theta} \right| \to p 0, \]
where the suprema here and subsequently are over \( \delta \in [\nabla_1, \nabla_2], \theta \in \Theta \). Given
(A.8) and (A.9), and using the Cauchy inequality, (A.10)-(A.12) follow from the fact, implied by the proof of Hualde and Robinson (2011), that (A.7) is uniformly \( O_p(1) \).

To prove (A.8) note first that Lemma 3 of Robinson (2012) gives, for all sufficiently large \( T \),
\[ \inf_t |k_t| \geq \frac{1}{8}. \]  
Suppressing reference to \( \delta, \theta \) in \( \chi_j = \chi_j(\delta, \theta) \),
\[ \sum_{s=1}^{T} (\chi_{t\delta\theta}g_t - \chi_{s\delta\theta}g_s) k_{ts} = \sum_{s=1}^{T} \left( \sum_{j=0}^{t-1} \chi_j g_{t-j} - \sum_{j=0}^{s-1} \chi_j g_{s-j} \right) k_{ts} \]  
\[ = \sum_{j=0}^{T-1} \chi_j \sum_{s=1}^{T} (g_{t-j} - g_{s-j}) k_{ts}, \]
with the convention already adopted that \( g(x) = 0, x < 0 \), and \( g(0) = 0 \) from
Assumption 4. Then
\[
\sup_{\delta, \theta} \left| \sum_{s=1}^{T} (X_{t,s}g_t - X_{s,s} g_s) k_{ts} \right| \leq \sum_{j=0}^{T-1} \left| \sum_{s=1}^{T} (g_{t-j} - g_{s-j}) k_{ts} \right|.
\] (A.15)

From (A.3) and Assumption 3
\[
\sup |\chi_j| \leq K_j^{\nabla^2 \nabla^1 - 1}.
\] (A.16)

Applying Assumption 4 and with $g_t$ denoting the derivative of $g(x)$ at $x = t/T$,
\[
\sum_{s=1}^{T} (g_{t-j} - g_{s-j}) k_{ts} = \hat{g}_{t-j} \sum_{s=1}^{T} \left( \frac{t-s}{T} \right) k_{ts} + O \left( \sum_{s=1}^{T} \left( \frac{t-s}{T} \right)^2 |k_{ts}| \right),
\] (A.17)

where $\hat{g}_{t} = 0, t \leq 0$. By Lemma 2 of Robinson (2012)
\[
\sum_{s=1}^{T} \left( \frac{t-s}{T} \right) k_{ts} = Th^2 \left( \frac{1}{Th} \sum_{s=1}^{T} \left( \frac{t-s}{Th} \right) k_{ts} - \int u k(u) du \right) = O(h)
\] (A.18)

uniformly in $t \in (Th, T - Th)$. Uniformly in $t \leq Th, t \geq T - Th$,
\[
\sum_{s=1}^{T} \left( \frac{t-s}{T} \right) k_{ts} = O(Th^2)
\] (A.19)

from Lemma 1 of Robinson (2012). By the same lemma,
\[
\sum_{s=1}^{T} \left( \frac{t-s}{T} \right)^2 |k_{ts}| = O(Th^3)
\] (A.20)

uniformly in $t$. Thus from (A.17)-(A.20),
\[
\max_{j} \left| \sum_{s=1}^{T} (g_{t-j} - g_{s-j}) k_{ts} \right| = O \left( h + Th^3 \right), t \in (Th, T - Th)
\]
\[
= O \left( Th^2 \right), t \in (Th, T - Th),
\] (A.21)

uniformly. Using also (A.13), (A.15) and (A.16),
\[
\sup_{\delta, \theta} |D_{t,\delta}| \leq K(Th)^{-1} \left( h + Th^3 \right) \sum_{j=0}^{T-1} (1 + j)^{\delta_0 - \nabla^1 - 1}
\]
\[
\leq K(T^{\nabla^2 - \nabla^1} + T^{\nabla^2 - \nabla^1} h^2), t \in (Th, T - Th),
\]
and

$$\sup_{\delta, \theta} |D_{t, \delta\theta}| \leq K(Th)^{-1}Th^2 \sum_{j=0}^{T-1} (1 + j)^{\nabla_2 - \nabla_1}$$

$$\leq KT^{\nabla_2 - \nabla_1}h, \ t \leq Th, t \geq T - Th,$$

uniformly over the stated ranges of \(t\). Thus

$$\sup \frac{1}{T} \sum_{t=1}^{T} D_{t, \delta\theta}^2 \leq K(T^{2(\nabla_2 - \nabla_1)} + T^{2(\nabla_2 - \nabla_1)}h^4 + T^{2(\nabla_2 - \nabla_1)}h^3) \to 0$$

by Assumption 6, verifying (A.8).

To prove (A.9), we have

$$\sum_{s=1}^{T} \chi_{s, \delta\theta} u_s k_{t,s} = \sum_{j=0}^{T-1} \chi_j c_{tj} = \sum_{j=0}^{[Th]} \chi_j c_{tj} + \sum_{j=[Th]+1}^{T-1} \chi_j c_{tj}$$

where

$$c_{tj} = \sum_{r=1}^{T-j} u_r k_{t+r+j},$$

so, using (A.3),

$$\sup \left| \sum_{j=0}^{[Th]} \chi_j c_{tj} \right| \leq \sum_{j=0}^{[Th]} \left| \chi_j \right| |c_{tj}| \leq K \sum_{j=1}^{[Th]} j^{\nabla_2 - \nabla_1} |c_{tj}|$$

and thus

$$E \left( \sup \left| \sum_{j=0}^{[Th]} \chi_j c_{tj} \right| \right)^2 \leq K \sum_{j=1}^{[Th]} \sum_{l=1}^{[Th]} j^{\nabla_2 - \nabla_1} l^{\nabla_2 - \nabla_1} (E c_{ij}^2 E c_{il}^2)^{1/2}. \quad (A.22)$$

Now

$$Ec_{ij}^2 = \alpha^2 \sum_{r=1}^{T-j} k_{t+r+j}^2 = O(Th)$$

by Assumption 6, so (A.22) is \(O((Th)^2(\nabla_2 - \nabla_1)+1) = o((Th)^2)\) uniformly in \(t\), by Assumption 3.
By summation-by-parts

\[ \sum_{j=[Th]+1}^{T-1} \chi_j c_{\epsilon_j} = \sum_{j=[Th]+1}^{T-2} \left( \chi_j - \chi_{j+1} \right) d_{ij} + \chi_{T-1} d_{ji,T-1}. \]  \hspace{1cm} (A.23)

where

\[ d_{ij} = \sum_{l=0}^{j} c_{\epsilon l}. \]

Now (A.23) is bounded uniformly by

\[ \sum_{j=[Th]+1}^{T-2} \left( \sup_{\chi_j} |\chi_j - \chi_{j+1}| \right) |d_{ij}| + \left( \sup_{\chi_{T-1}} |\chi_{T-1}| \right) |d_{ji,T-1}| \]

\[ \leq K \sum_{j=[Th]+1}^{T-2} j^{\gamma-1} |d_{ij}| + K T^\gamma |d_{ji,T-1}| \]  \hspace{1cm} (A.24)

using (A.3) and (A.5) and writing \( \gamma = \max(\nabla_2 - \nabla_1, 1 - \varsigma) - 1 \). Rearranging,

\[ d_{ij} = \sum_{r=1}^{T} u_r \left( \sum_{s=r}^{\min(r+j,T)} k_{ts} \right), \]

so

\[ Ed_{ij}^2 = \sigma^2 \sum_{r=1}^{T} \left( \sum_{s=r}^{\min(r+j,T)} k_{ts} \right)^2 \leq K j \left( \sum_{s=1}^{T} |k_{ts}| \right)^2 \leq K j (Th)^2 \]

and (A.24) has second moment bounded by

\[ K (Th)^2 \sum_{j=[Th]+1}^{T-2} \sum_{l=[Th]+1}^{T-2} j^{\gamma-1/2} l^{\gamma-1/2} + K (Th)^2 T^{2\gamma+1} \]

\[ \leq K (Th)^{2\gamma+3} = o((Th)^2) \]

uniformly in \( t \), since \( \gamma < -1/2 \). We have established that \( E \sup S^2_{\epsilon \theta} = o(1) \) uniformly in \( t \), whence follows (A.9), to complete the proof of the theorem.

**APPENDIX B**

**Proof of Theorem 2**

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Writing $\partial = \partial / \partial (\delta, \theta')$, by the mean value theorem

$$0 = \partial Q(\delta, \theta)/2 = \partial Q(\delta_0, \theta_0)/2 + \Omega \left( \frac{\delta - \delta_0}{\theta - \theta_0} \right),$$

where $\partial Q(\delta, \theta)$ is given by (19)-(21) and $\Omega$ is obtained from the matrix $\partial^2 Q(\delta, \theta)/2 = \partial \partial' Q(\delta, \theta)/2$ by evaluating each row at a generally different $\delta, \theta$ such that $\left\| \delta - \delta_0, \theta' - \theta'_0 \right\| \leq \left\| \delta - \delta_0, \theta' - \theta'_0 \right\|$. The theorem follows if

$$\frac{T^{1/2}}{2} \partial Q(\delta_0, \theta_0) \rightarrow d \mathcal{N}(0, \Omega), \quad (B.1)$$

$$\Omega \rightarrow p \Omega. \quad (B.2)$$

From (11)

$$\dot{u}_t(\delta_0, \theta_0) = u_t + D_t - S_t$$

$$= u_t + \sum_{s=1}^{T} (g_t - g_s) k_{ts} - \sum_{s=1}^{T} u_s k_{ts}$$

where

$$D_t = \sum_{s=1}^{T} (g_t - g_s) k_{ts}/k_t, \quad S_t = \sum_{s=1}^{T} u_s k_{ts}/k_t.$$

From (20), (21)

$$\partial \dot{u}_t(\delta, \theta) = \partial \chi_{t\delta \theta} u_t + \partial D_t \partial \theta - \partial S_t \partial \theta.$$

Write

$$\pi_j = \partial \chi_j |_{\delta_0, \theta_0} = \sum_{l=0}^{j} \left( \tau_l(\theta_0) \partial \alpha_{j-l}(\delta_0)/\partial \delta \right) \left( \alpha_{j-l}(\delta_0) \partial \tau_l(\theta_0)/\partial \theta \right) = \left( j^{-1} \psi_j(\theta_0) \right), \quad j \geq 1$$

Then

$$\partial \dot{u}_t(\delta_0, \theta_0) = v_t + \partial D_t - \partial S_t.$$
where

\[ v_t = \partial \chi_{t \delta \theta} u_t \bigg|_{\delta_0, \theta_0} = \sum_{j=1}^{t-1} \pi_j u_{t-j}, \]

\[ \partial D_t = \partial D_{t \delta \theta} \bigg|_{\delta_0, \theta_0} = \sum_{s=1}^{T} \left( \sum_{j=1}^{t-1} g_{t-j} \pi_j - \sum_{j=1}^{s-1} g_{s-j} \pi_j \right) k_{ts} / k_t, \]

\[ \partial S_t = \partial S_{t \delta \theta} \bigg|_{\delta_0, \theta_0} = \sum_{s=1}^{T} \sum_{j=1}^{s-1} u_{s-j} \pi_j k_{ts} / k_t. \]

Thus from (19)

\[ \partial Q(\delta_0, \theta_0) = \frac{2}{T} \sum_{t=1}^{T} (u_t + D_t - S_t) (v_t + \partial D_t - \partial S_t). \]

Hualde and Robinson (2011) show that

\[ T^{-1/2} \sum_{t=1}^{T} u_t v_t \rightarrow_{d} \mathcal{N}(0, \Omega). \]

Thus (B.1) holds if all the following are \( o_p(1) \):

\[ T^{-1/2} \sum_{t=1}^{T} u_t \partial D_t, \ T^{-1/2} \sum_{t=1}^{T} u_t \partial S_t, \ T^{-1/2} \sum_{t=1}^{T} D_t v_t, \ T^{-1/2} \sum_{t=1}^{T} S_t v_t, \]

\[ T^{-1/2} \sum_{t=1}^{T} D_t \partial D_t, \ T^{-1/2} \sum_{t=1}^{T} D_t \partial S_t, \ T^{-1/2} \sum_{t=1}^{T} S_t \partial D_t, \ T^{-1/2} \sum_{t=1}^{T} S_t \partial S_t. \]  

(B.3)

Note first that, as in (A.14), and using (A.13), (A.21) and \( \| \pi_j \| = O(j^{-1}) \),

\[ \| \partial D_t \| \leq \frac{K}{Th} \left\| \sum_{j=1}^{T} \pi_j \sum_{s=1}^{T} (g_{t-j} - g_{s-j}) k_{ts} \right\| \leq \frac{K}{Th} \max_{t,j} \left| \sum_{s=1}^{T} (g_{t-j} - g_{s-j}) k_{ts} \right| \sum_{j=1}^{T} j^{-1} \]

\[ = O \left( (T^{-1} + h^2 \log T) \right), \ t \in (Th, T - Th) \]

\[ = O(\log T), \ t \leq Th, t \geq T - Th, \]  

(B.4)
uniformly over the stated ranges of $t$. Similarly but more easily, we derive, uniformly,

$$
D_t = O \left( (T^{-1} + h^2) \right), \quad t \in (Th, T - Th)
$$

$$
= O(h), \quad t \leq Th, t \geq T - Th.
$$

(B.5)

We check each claim of (B.3) in turn; for notational convenience, when $j \leq 0$ we take $\pi_j = 0$ and interpret $1/j$ to be 0.

First, using (B.4) and Assumption 6,

$$
E \left| T^{-1/2} \sum_{t=1}^{T} u_t \partial D_t \right|^2 \leq \frac{K}{T} \sum_{t=1}^{T} \left| \partial D_t \right|^2
$$

$$
\leq \frac{K}{T} \left( T \left( T^{-1} + h^2 \right)^2 \log^2 T + Th^3 \log^2 T \right)
$$

$$
= O \left( T^{-2} \log^2 T + h^4 \log^2 T + h^3 \log^2 T \right) = o(1).
$$

Next,

$$
\left| T^{-1/2} \sum_{t=1}^{T} u_t \partial S_t \right|^2 = \frac{1}{T} E \left\{ \sum_{t=1}^{T} u_t \sum_{r=1}^{T} u_r \sum_{s=1}^{T} u_{s-j} \pi_j' \frac{k_{ts} k_{tq} k_{ts}}{k_l} \sum_{q=1}^{T} u_{q-t} \pi_l' \frac{k_{ts}}{k_l} \right\}
$$

$$
= \frac{\sigma^4}{T} \sum_{t=1}^{T} \sum_{s=1}^{T} \sum_{j=1}^{T} \sum_{q=1}^{T} \pi_j' \pi_{q-s+j} \frac{k_{ts} k_{tq}}{k_l k_r}
$$

$$
+ \frac{\sigma^4}{T} \sum_{t=1}^{T} \sum_{r=1}^{T} \sum_{s=1}^{T} \sum_{q=1}^{T} \pi_r' \pi_{q-r} \frac{k_{ts} k_{tq}}{k_l k_r}
$$

$$
+ \frac{\sigma^4}{T} \sum_{t=1}^{T} \sum_{r=1}^{T} \sum_{s=1}^{T} \sum_{q=1}^{T} \pi_s' \pi_{q-t} \frac{k_{ts} k_{tq}}{k_l k_r}
$$

$$
+ \frac{\sigma^4}{T} \sum_{t=1}^{T} \sum_{r=1}^{T} \sum_{s=1}^{T} \sum_{q=1}^{T} \pi_s' \pi_{q-t} \frac{k_{ts} k_{tq}}{k_l k_r}.
$$

By boundedness of $k$, the final term in the last displayed expression is bounded by

$$
\frac{K}{T^3 h^2} \sum_{t=1}^{T} \sum_{s=1}^{T} \sum_{q=1}^{T} \frac{1}{s-t} \frac{1}{q-t} = O \left( \frac{\log^2 T}{T^2 h^2} \right),
$$

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while the other terms are bounded by

\[
\frac{K}{T^3h^2} \sum_{t=1}^{T} \sum_{s=1}^{T} \sum_{j=1}^{T} \sum_{q=1}^{T} \frac{|k_{ts}|}{j} |k_{tq}| + \frac{K}{T^3h^2} \sum_{r=1}^{T} \sum_{s=1}^{T} \sum_{j=1}^{T} \sum_{q=1}^{T} \frac{|k_{ts}|}{j} |k_{rj}| \leq \frac{KT^2h\log^2 T}{T^3h^2} = O\left(\frac{\log^2 T}{Th}\right) = o(1),
\]

by Assumption 8.

Next, noting that

\[
\sum_{t=1}^{T} D_t v_t = \sum_{t=1}^{T} u_t \sum_{s=t+1}^{T} D_s \pi_{s-t},
\]

from (B.5)

\[
E \left\| T^{-1/2} \sum_{t=1}^{T} D_t v_t \right\|^2 = \sigma^2 T^{-1} E \sum_{t=1}^{T} \left\| \sum_{s=t+1}^{T} D_s \pi_{s-t} \right\|^2 \leq KT^{-1} h^2 \sum_{t=1}^{T} \left( \sum_{s=t+1}^{T} (s-t)^{-1} \right)^2 \leq Kh^2 \log^2 T = o(1),
\]

by Assumption 6.

Next, using (B.4) and (B.5),

\[
T^{-1/2} \sum_{t=1}^{T} D_t \partial_t D_t = O \left(T^{-1/2}(T^{-1} + h^2)^2 \log T + T^{-1/2}Th^3 \log T \right) = O \left(T^{-3/2} \log T + T^{1/2}h^4 \log T + T^{1/2}h^3 \log T \right) = o(1),
\]

by Assumption 6.

Next,

\[
E \left\| T^{-1/2} \sum_{t=1}^{T} S_t v_t \right\|^2 = T^{-1} E \left\{ \sum_{t=1}^{T} \sum_{s=1}^{T} u_s^{t} k_{ts}^{t} \sum_{j=1}^{t-1} \pi_j^{t-j} u_{t-j} \sum_{q=1}^{T} \sum_{r=1}^{T} u_q^{r} k_{rq}^{r} \sum_{l=1}^{r-1} \pi_l^{r-l-1} \right\}
\]
which equals

\[
\sigma^4 T^{-1} \left( \sum_{t=1}^{T} \sum_{s=1}^{t-1} \frac{k_{ts}}{k_t} \pi'_{t-s} \right)^2 \\
+ \sigma^4 T^{-1} \sum_{t=1}^{T} \sum_{s=1}^{T} \sum_{r=1}^{T} \frac{k_{ts}}{k_t} \frac{k_{rs}}{k_r} \sum_{j=1}^{t-1} \pi'_j \pi_{r-t+j} \\
+ \sigma^4 T^{-1} \sum_{t=1}^{T} \sum_{s=1}^{T} \sum_{r=1}^{T} \sum_{q=1}^{T} \frac{k_{ts}}{k_t} \frac{k_{rs}}{k_r} \pi'_{t-q} \pi_{r-s} \\
+ \sigma^4 T^{-1} \sum_{t=1}^{T} \sum_{s=1}^{T} \sum_{r=1}^{T} \sum_{j=1}^{T} \frac{k_{ts}}{k_t} \frac{k_{rs}}{k_r} \pi'_{t-s} \pi_{r-s},
\]

and using boundedness of \(k\) this is bounded by \(KT^{-1} (Th)^{-2} T^2 \log^2 T + KT^{-1} (Th)^{-2} T \log^2 T \leq K \log^2 T / (Th^2) = o(1)\), by Assumption 8.

Next,

\[
E \left\| T^{-1/2} \sum_{t=1}^{T} D_t \partial S_t \right\|^2 = \sigma^2 T^{-1} \sum_{t=1}^{T} \sum_{r=1}^{T} D_t D_r \sum_{s=1}^{T} \sum_{q=1}^{T} \sum_{j=1}^{T} \frac{k_{ts}}{k_t} \frac{k_{rq}}{k_r} \\
\leq KT^{-1} (Th)^{-2} h^2 \sum_{t=1}^{T} \sum_{r=1}^{T} \sum_{s=1}^{T} \sum_{q=1}^{T} \sum_{j=1}^{T} \| \pi_{q-s+j} \| \| \pi_j \| \\
\leq KT^{-1} (Th)^{-2} h^2 T^2 (Th) \log^2 T \\
\leq Kh \log^2 T = o(1),
\]

by (B.5) and Assumption 6.

Next,

\[
E \left\| T^{-1/2} \sum_{t=1}^{T} S_t \partial D_t \right\|^2 = E \left\| T^{-1/2} \sum_{t=1}^{T} \partial D_t \sum_{s=1}^{T} u_s \frac{k_{ts}}{k_t} \right\|^2 \\
= \sigma^2 T^{-1} \sum_{t=1}^{T} \sum_{r=1}^{T} \partial D_t' \partial D_r \sum_{s=1}^{T} \frac{k_{ts}}{k_t} \frac{k_{rs}}{k_r} \\
\leq KT^{-1} (Th)^{-2} h^2 \log^2 T \sum_{t=1}^{T} \sum_{r=1}^{T} \sum_{s=1}^{T} \| k_{ts} k_{rs} \| \\
\leq KT^{-1} (Th)^{-2} h^2 (\log^2 T) T (Th)^2 \leq Kh^2 \log^2 T = o(1),
\]
by Assumption 6.

Finally,

$$E \left| T^{-1/2} \sum_{t=1}^{T} S_t \partial S_t \right|^2 = T^{-1} E \left( \sum_{t=1}^{T} \sum_{s=1}^{T} u_s \frac{k_{ts}}{k_t} \sum_{r=1}^{T} \sum_{j=1}^{r-1} u_{r-j} \frac{k_{tr}}{k_t} \sum_{q=1}^{T} \sum_{p=1}^{T} \frac{k_{qp}}{k_q} \sum_{n=1}^{T} \sum_{l=1}^{T} \frac{k_{tn}}{k_t} \frac{k_{qn}}{k_q} \right),$$

which equals

$$\sigma^4 T^{-1} \sum_{t=1}^{T} \sum_{s=1}^{T} \frac{k_{ts}}{k_t} \sum_{r=1}^{T} \sum_{j=1}^{r-1} \frac{k_{tr}}{k_t} \sum_{q=1}^{T} \sum_{p=1}^{T} \frac{k_{qp}}{k_q} \sum_{n=1}^{T} \sum_{l=1}^{T} \frac{k_{tn}}{k_t} \frac{k_{qn}}{k_q},$$

which is bounded by

$$KT^{-1} (Th)^{-4} \left( \sum_{t=1}^{T} \sum_{s=1}^{T} \frac{k_{ts}}{k_t} \sum_{r=1}^{T} \||\pi_{r-s}\|| \right)^2 + KT^{-1} (T h)^{-4} \sum_{t=1}^{T} \sum_{s=1}^{T} \left| k_{ts} \right| \sum_{r=1}^{T} \sum_{j=1}^{r-1} \||\pi_{j}\|| \sum_{q=1}^{T} \sum_{n=1}^{T} \||\pi_{n-r+j}\|| + KT^{-1} (T h)^{-4} \sum_{t=1}^{T} \sum_{s=1}^{T} \left| k_{ts} \right| \sum_{r=1}^{T} \sum_{p=1}^{T} \||\pi_{p-r}\|| \sum_{q=1}^{T} \sum_{n=1}^{T} \||\pi_{n-s}\|| + KT^{-1} (T h)^{-4} \sum_{t=1}^{T} \sum_{s=1}^{T} \left| k_{ts} \right| \sum_{r=1}^{T} \sum_{q=1}^{T} \||\pi_{r-s}\|| \sum_{n=1}^{T} \sum_{l=1}^{T} \||\pi_{n-l}\|| \leq KT^{-1} (Th)^{-4} \left( T^2 h \log T \right)^2 + KT^{-1} (Th)^{-4} T (Th \log T)^2 \leq K (Th^2)^{-1} \log^2 T = o(1),$$

by Assumption 8.
This completes the proof of (B.3), and thus of (B.1).

Finally (B.2) follows if

\[
\frac{1}{2} \partial^2 Q(\delta_0, \theta_0) \rightarrow_p \Omega \tag{B.6}
\]

and, given Theorem 1, for a neighbourhood \(\mathcal{N}\) of \(\delta_0, \theta_0\),

\[
\sup_{\mathcal{N}} \left\| \partial^2 Q(\delta, \theta) - \partial^2 Q(\delta_0, \theta_0) \right\| \rightarrow_p 0. \tag{B.7}
\]

The proof of (B.6) partly employs Theorem 2.2 of Hualde and Robinson (2011) and partly methods used above to deal with contributions from deterministic and stochastic errors, where these are less delicate than in Theorem 1’s proof because \(\nabla_2 - \nabla_1\) is replaced by an arbitrarily small positive number, and less delicate than the proof of Theorem 1 because \(T^{-1/2}\) norming is replaced by \(T^{-1}\) norming. The proof of (B.7) uses that of Theorem 2.2 of Hualde and Robinson (2011) and standard techniques. The full details of the proofs of (B.6) and (B.7) are very lengthy but straightforward relative to what has gone before and are thus omitted.

REFERENCES


