



# Saddlepoint techniques for the statistical analysis of time series

Davide La Vecchia

results from joint work with E. Ronchetti, A. Moor, C. Jiang & O. Scaillet

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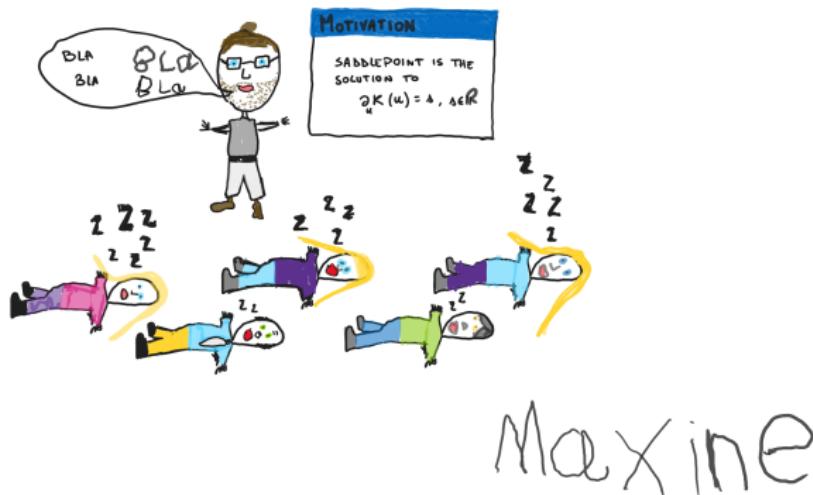
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- Illustrate that first-order asymptotic theory suffers from finite sample distortions
- Develop saddlepoint techniques (for CDF/pdf approximation,  $p$ -values, and testing) which perform well in small samples and feature higher-order accuracy
- Illustrate numerically the performance of some saddlepoint techniques (testing)

## Motivation from theoretical statistics

*Typical statistical problem:* For a given statistic/functional

$$T : \text{dom } T \rightarrow \mathbb{R}^p, p \geq 1$$

(it can be an estimator  $\hat{\theta}_n$ ), tail probabilities or quantiles at different levels are needed to carry out **statistical inference** (essentially, tests and confidence intervals).

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⇒ we have to rely on **approximations**.

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$$p(y|x, \Delta) = \frac{1}{2\pi i} \int_{\tau-i\infty}^{\tau+i\infty} \exp \left\{ K_{y|x}(\Delta, z; x) - zy \right\} dz$$

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Analytical and resampling techniques can achieve higher order refinements over the first order asymptotic theory

## First part: i.i.d. setting

# Analytical techniques: Edgeworth expansion

Edgeworth/Charlier series is obtained as follows:

1 Assume we are given:

Random Variable	$X$	$Y$
Distr.	$F_X$	$G$
Measures	$\mu$	$\nu$
Charact. fct.	$\chi(u)$	$\xi(u)$
Cumulants	$\beta_r$	$\gamma_r$

where the two Fourier transforms are

$$\chi(u) = \int e^{iux} dF_X(x), \quad \xi(u) = \int e^{iux} dG(x)$$

and the cumulants are (by definition)

$$\beta_r = (-i)^r \frac{d^r}{du^r} \ln \chi(u) \Big|_{u=0}, \quad \gamma_r = (-i)^r \frac{d^r}{du^r} \ln \xi(u) \Big|_{u=0}$$

# Analytical techniques: Edgeworth expansion

2 By **Taylor expansion of the cumulant generating function (c.g.f.) about  $u = 0$ :**

$$\ln \frac{\chi(u)}{\xi(u)} = \ln \chi(u) - \ln \xi(u) \stackrel{\text{Taylor at } u=0}{=} \sum_{r=1}^{\infty} (\beta_r - \gamma_r) \frac{(iu)^r}{r!},$$

thus,

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3 By Fourier inversion (under suitable assumptions on  $G$ , see **Wallace (Ann. Math. Stats, 1958)**):

$$F_X(x) = \exp \left\{ \sum_{r=1}^{\infty} (\beta_r - \gamma_r) \frac{(-D)^r}{r!} \right\} G(x),$$

$D$  denotes a differentiation operator (with respect to  $x$  and  $e^D = \sum_{j=0}^{\infty} D^j / j!$ ): it is such that the first term of the expansion of  $F_X$  is  $G$ .

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4 Differentiation yields an expansion for the density of  $X$ .

# Motivation from theoretical statistics

## Example (Sample mean)

Let  $X \sim \mu$  absolutely continuous w.r.t. the Lebesgue measure and having density  $f_X$ .

We are given a random sample  $\mathbf{X} = (X_1, \dots, X_n)$  of i.i.d. copies of  $X$ , whose mgf and cgf exist and:  $E_\mu[X] = 0$ ,  $V_\mu(X) = \sigma^2 < \infty$ .

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We make use of Step 1-3 to approximate the density  $f_n$  of the standardized mean via Edgeworth expansion...

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### On Some Connections Between Esscher's Tilting Saddlepoint Approximations, and Optimal Transportation: A Statistical Perspective

Davide La Vecchia, Elvezio Ronchetti, Andrej Ilijevski

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Statist. Sci. 38(1): 30-51 (February 2023). DOI: 10.1214/21-STSB47



## Motivation from theoretical statistics

### Example (cont'd)

... using as  $G(x)$  the standard normal, we obtain an expansion of  $f_n$  in powers of  $n^{-1/2}$ , where the leading term is the normal density and higher order terms correct for skewness, kurtosis:

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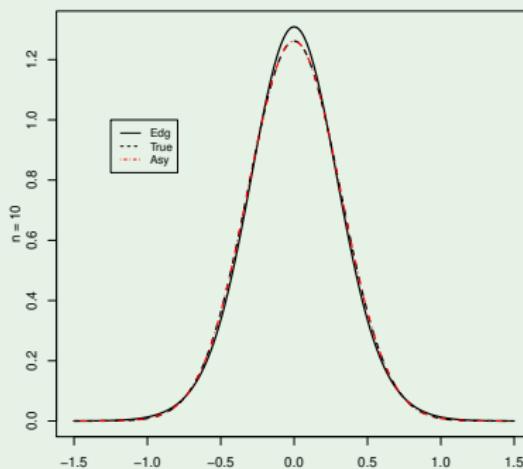
By construction, Edgeworth expansions provide in general a good approximation in the center of the density, BUT they can be inaccurate in the tails, where they can even become negative.

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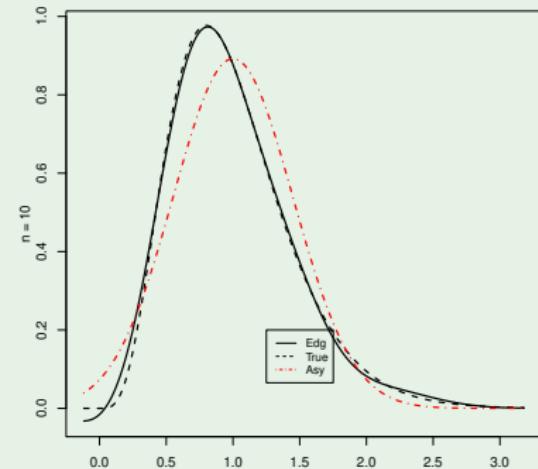
## Example (cont'd)

For Asy and Edg, consider  $\bar{X}_n$  for  $n = 10, 50, 250$ , for  $X_i \sim \mathcal{N}(0, 1)$  and  $X_i \sim \text{exp}(1)$

$\mathcal{N}(0, 1)(n = 10)$



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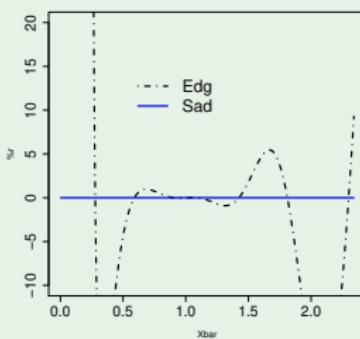


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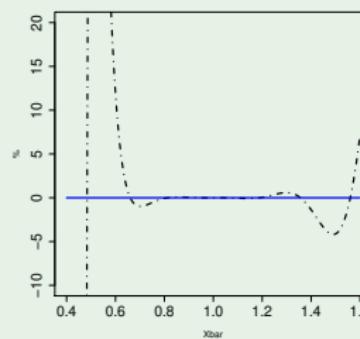
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for the **exponential case**, rel. err. =  $100 \cdot (\text{true} - \text{approx})/\text{true}$

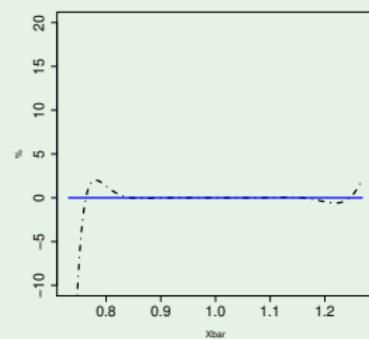
$n = 10$



$n = 50$



$n = 250$



Any other higher order technique to cope with these issues? saddlepoint approx...

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## Example (cont'd )

In this example about  $\bar{X}_n$ , we know the c.g.f. and the saddlepoint density approx  $g_n(s)$  is (Daniels (1954)):

$$g_n(s) = \left[ \frac{n}{2\pi \mathcal{K}''\{v(s)\}} \right]^{1/2} \exp \left( n \left[ \mathcal{K}\{v(s)\} - v(s)s \right] \right) \quad (1)$$

and  $v(s)$  (saddlepoint) is the solution to

$$\mathcal{K}'(v) - s = 0,$$

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namely, we look for  $v(s)$  such that  $X$  has expected value equal to  $s$ .

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## Example (cont'd)

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- for  $\sqrt{n}\bar{X}_n$ , (by Jacobian formula)  $g_n(s) = \left[\frac{1}{2\pi}\right]^{1/2} e^{-\frac{s^2}{2}}$

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- $g_n(s)$  is a “Gaussian-type” integral with both mean and variance that depends on  $s$ : it is a density-like object that cannot take on negative values ( $\neq$  Edg):  
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$$\mathcal{K}(v) = \frac{\nu^2}{2}, \quad \mathcal{K}'(v) = v \text{ and } \mathcal{K}''(v) = 1,$$

the saddlepoint is defined by  $\mathcal{K}'(v) - s = 0$ , thus  $v(s) = s$  and

- for  $\bar{X}_n$ ,  $g_n(s) = \left[\frac{n}{2\pi}\right]^{1/2} e^{-\frac{ns^2}{2}}$  pdf of  $\mathcal{N}(0, \frac{1}{n})$  it's exact!!
- for  $\sqrt{n}\bar{X}_n$ , (by Jacobian formula)  $g_n(s) = \left[\frac{1}{2\pi}\right]^{1/2} e^{-\frac{s^2}{2}}$  pdf of  $\mathcal{N}(0, 1)$  it's exact!!

# Motivation from theoretical statistics

## Example (cont'd)

- The saddlepoint density approximation  $g_n$  features relative error of order  $O(n^{-1})$  over the whole  $\mathbb{R}$

$$f_n(s) = g_n(s) \{1 + O(n^{-1})\} \quad (2)$$

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- The density  $g_n$  is obtained by approximating the Fourier inversion of  $M^n$ , which yields  $f_n$ :

$$\begin{aligned} f_n(s) &= \frac{n}{2\pi} \int_{-\infty}^{\infty} e^{-ivns} M^n(iv) dv \stackrel{(z=iv)}{=} \frac{n}{2\pi i} \int_{\mathcal{I}} e^{-nzs} M^n(z) dz \\ &= \frac{n}{2\pi i} \int_{\tau-i\infty}^{\tau+i\infty} e^{n(\mathcal{K}(z) - zs)} dz, \quad \tau \in \mathbb{R}, \end{aligned}$$

which may be obtained using a Taylor expansion of  $(\mathcal{K}(z) - zs)$  about  $v(s)$ .

## Motivation from theoretical statistics

The **sadd approx**  $g_n$  is obtained via the method of the steepest descent: this is a general technique in **complex analysis** applied to compute asymptotic expansions of integrals

$$\int_{\mathcal{P}} e^{\nu w(z)} \xi(z) dz,$$

with  $\nu \in \mathbb{R}^+$  is large,  $\xi$  and  $w$  being analytic functions of  $z \in \mathbb{C}$ .

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### Idea

*Deform the path of integration* (Cauchy's theorem) so that the new path of integration passes through the so-called **saddlepoint**, namely the zero of the derivative  $w'(z)$ . Then, we approximate the resulting integral using a series expansion (Watson's lemma). See *Daniels (AoMS, 1954)*.

Loosely speaking, one does a "Laplace-type approx" on  $\mathbb{C}$ .

[Jump to Laplace](#)

# Motivation from theoretical statistics

Alternative: derive  $g_n$  via convex analysis.

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- repeat this procedure for every  $s \in \mathbb{R}$

⇒ *saddlepoint density approximation* is a sequence of low-order local approximations; see *Easton & Ronchetti (1986), JASA, Wang (1992)* and *La Vecchia et al. (2023)*.

## Motivations related to dependent data analysis

Many macroeconomic time series display a persistent time trend and contain only a few observations recorded at annual frequency. Much controversy in macroeconomics has revolved around the suitability of ARIMA models; see the seminal paper of Nelson and Plosser (1982) and Gil-Alana and Robinson (1997) for a review of the literature.

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Within this setting, to model the slow decay of the autocorrelation function displayed by many macroeconomic time series, the use of (Gaussian) FARIMA models and first order Gaussian asymptotic theory (Wald-type test statistics) is routinely applied for confidence intervals and testing statistical hypotheses.

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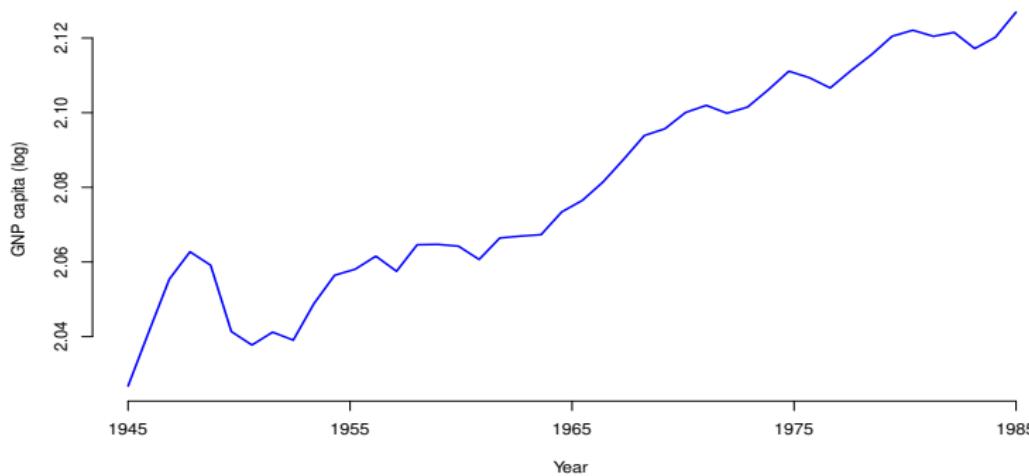
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Saddlepoint approximations for short and long memory time series: A frequency domain approach

Davide La Vecchia   , Elvezio Ronchetti

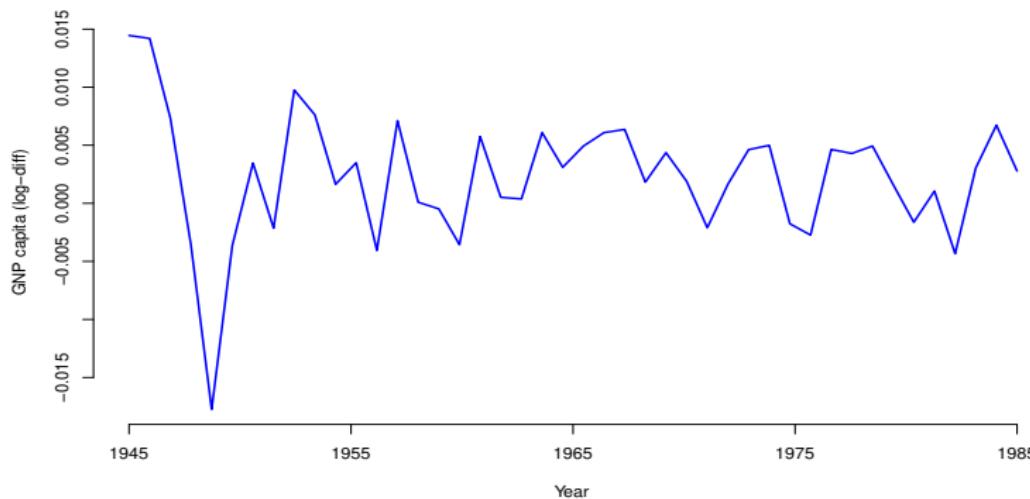
## Motivations related to dependent data analysis

Focus on the [extended Nelson and Plosser data set](#): plot log-GNP per capita (other time series available in the JoE paper)



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### Remark

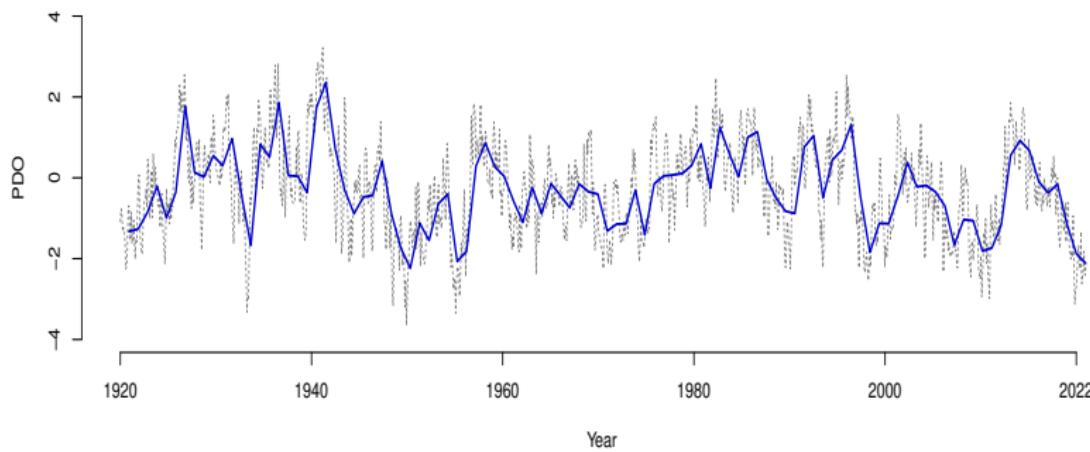
In the literature one is typically testing for the presence of long memory: ARFIMA models and

$$\mathcal{H}_0 : d = 0 \quad \text{vs} \quad \mathcal{H}_1 : d > 0$$

we resort on an M-estimator (Whittle), which is *asymptotically*  $\chi^2$  Wald-type test statistics ( $W_n$ ) are applied when  $n = 44$ . Is this a sensible procedure? Is the asymptotics suffering from size distortion due to the small sample size?

## Motivations related to dependent data analysis

The Pacific Decadal Oscillation (PDO) index measures the climatological situation of the Southern hemisphere: its extremes correspond to episodes of abnormal weather conditions.



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### Remark

*Whiting et al. (2003)* model the time series by an ARFIMA(0, d, 0). Data analysis and inference is conducted using **annual data**, from 1920 to 2022, so  $n = 122$ , relying on M-estimator (Whittle), which yields **Wald-type statistic  $W_n$  from first order asymptotic theory** to test

$$\mathcal{H}_0 : d = 0 \quad vs \quad \mathcal{H}_1 : d > 0.$$

## Motivations related to dependent data analysis

### Example (ARFIMA synthetic data)

Let  $\{Y_t, t \in \mathbb{Z}\}$  be an ARFIMA( $p, d, q$ ), having dynamics

$$\theta(L)(1 - L)^d Y_t = \phi(L)\epsilon_t, \quad (3)$$

where  $\forall t$ , the  $\{\epsilon_t\}$  are i.i.d. with zero mean and known  $\sigma_\epsilon^2 = 1$ .

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- We focus on a Gaussian ARFIMA of order  $p = 2$ ,  $d = 0$  and  $q = 0$  process, with AR coefficients  $\theta_1 = 0.1$  and  $\theta_2 = 0.2$ .

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- We consider increasing values of the sample size  $n = 250, 2500, 5000$ .
- We estimate  $\theta$  via the routinely applied Whittle's M-estimator, as implemented in the routine `WhittleEst` available in the R package `longmemo`.

# Motivations related to dependent data analysis

## Example (cont'd)

The goal of our inference is to test

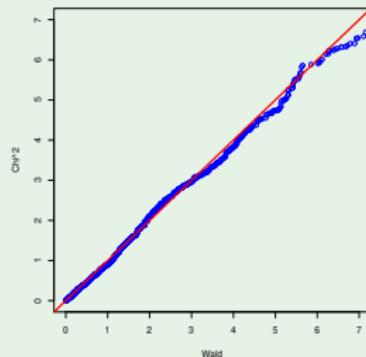
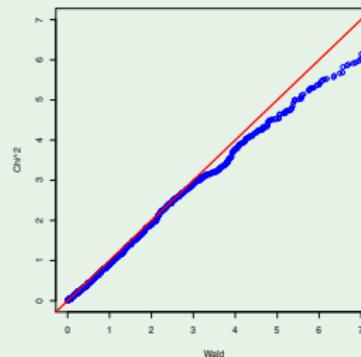
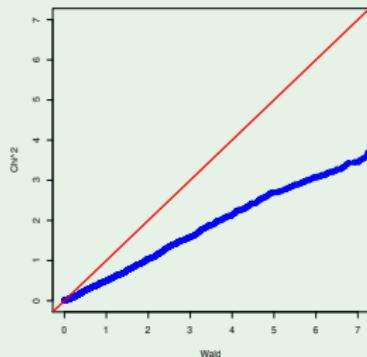
$$\mathcal{H}_0 : d = 0 \text{ vs. } \mathcal{H}_1 : d > 0,$$

and we resort on the **Wald test statistic**  $W_n$ , for Whittle's estimator, as available in the **statistical software**, comparing  $\chi^2$  quantiles to the true (as obtained by MC simulation).

$n = 250$

$n = 2500$

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# Motivations related to dependent data analysis

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$n = 250$

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## Remark

As conjectured, the first order asymptotic theory suffers from size distortion. Any **saddlepoint techniques?**



# Motivations related to dependent data analysis

Another example comes from the literature on Spatial Autoregressive processes (random fields):

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Theory and Methods

## Saddlepoint Approximations for Spatial Panel Data Models

Chaonan Jiang, Davide La Vecchia, Elvezio Ronchetti & Olivier Scaillet

Received 01 Jul 2020, Accepted 09 Sep 2021, Accepted author version posted online: 20 Sep 2021, Published online: 17 Nov 2021

Download citation https://doi.org/10.1080/01621459.2021.1981913 Check for updates

Full access

## **Second part:** time series setting

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- Literature: a bird's-eye view

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  - ▶ Monte Carlo results

## Literature: a bird's-eye view

- (i) Most of the results on **saddlepoint techniques** are available for the **iid setting**: see Field & Ronchetti (1990), Jensen (1995), Kolassa (2006), Butler (2007), or Brazzale et al. (2007) for book-length presentation.

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- (iii) **Higher order techniques in frequency domain (spectral analysis) for time series** are available: see Taniguchi (JMA, 1987, Edgeworth for Whittle under SRD), Franke & Härdle (Annals, 1992, FDB), Dahlhaus & Janas (Annals, 1996, FDB), Andrews & Lieberman (Econometric Theory, 2005, Edgeworth for Whittle under LRD).

## Some elements of spectral analysis

To start with, let me recall the autocovariance function

$$\gamma_Y(h) = \text{cov}(Y_{t+h}, Y_t) = E[(Y_{t+h} - \mu)(Y_t - \mu)]$$

for  $h \in \mathbb{Z}$  and with  $E(Y_t) = \mu, \forall t$ .

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Under suitable assumptions, we have (for  $i \in \mathbb{C}$ )

$$\gamma_Y(h) = \int_{-1/2}^{1/2} \exp\{2\pi i \lambda h\} f(\lambda) d\lambda, \quad h = 0, \pm 1, \pm 2 \dots$$

as the inverse Fourier transform of the **spectral density**  $f(\cdot)$ :

$$f(\lambda) = \sum_{h=-\infty}^{\infty} \gamma_Y(h) \exp\{-i2\pi\lambda h\}, \quad -1/2 \leq \lambda \leq 1/2.$$

# Some elements of spectral analysis

## Definition

Given time series data  $Y_1, \dots, Y_n$ , the discrete Fourier transform (DFT) is

$$d(\lambda_j) = n^{-1/2} \sum_{t=1}^n Y_t \exp\{-2\pi i \lambda_j t\},$$

for  $j = 0, 1, \dots, n - 1$ , where the frequencies  $\lambda_j = j/n$  are called Fourier or fundamental frequencies.

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The periodogram at  $\lambda_j$  is  $I(\lambda_j) = |d(\lambda_j)|^2$ , and we have that

$$I(\lambda_j) = \sum_{h=-(n-1)}^{n-1} \hat{\gamma}_Y(h) \exp\{-2i\pi \lambda_j h\},$$

where  $\hat{\gamma}_Y(h)$  is the empirical covariance.

## Some elements of spectral analysis

**Property.** The periodogram ordinates are such that

$$I(\lambda) \xrightarrow{\mathcal{D}} i.d. \xi f(\lambda), \quad \xi \sim \text{exp}^*(1) \quad (4)$$

### Remark

*The asymptotic iid-ness of the standardized periodogram ordinates allows to transform problems for dependent data into problems for iid data.*

## Some elements of spectral analysis

The Property allows to derive a frequency domain likelihood and parameter estimation is obtained maximizing this likelihood.

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This idea goes back to [Whittle \(1951\)](#): if there is a [parametric model](#) for  $f(\lambda, \theta)$ , then we may work on:

$$\mathcal{L}_W(\theta) = \frac{1}{2\pi} \left[ \int_{-\pi}^{\pi} \ln f(\lambda, \theta) d\lambda + \int_{-\pi}^{\pi} \frac{I(\lambda)}{f(\lambda, \theta)} d\lambda \right], \quad (5)$$

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which is obtained making use of Property 2 ( $\lambda$  is in radians, from now on).

The optimization of  $\mathcal{L}_W(\theta)$  (the Riemann-discretized version of  $\mathcal{L}_W$ ):

$$\hat{\theta}_n = \arg \max_{\theta} \mathcal{L}_W(\theta)$$

(or  $\nabla_{\theta} \mathcal{L}_W(\hat{\theta}_n) = 0$ ) defines an **M-estimator in the frequency domain, having estimating function  $\psi_j$** . Then,

$$\mathcal{V}_n = \sqrt{n}(\hat{\theta}_n - \theta^0)$$

and we want an approximation to its density  $f_{\hat{\theta}_n}$ .

## Incidentally....

At the other edge of the spectrum, the frequency domain approach can be suitably adapted and exploited for the statistical analysis of high-dimensional (large time, spatial, cross-sectional dimensions) random fields over a regular network:

The screenshot shows a Cornell University logo and the text "Cornell University". On the right, it says "We gratefully acknowledge" and has a "Search..." bar, "Help | About", and a "Log In" button. The main content area has a red header with "arXiv > stat > arXiv:2312.02591". Below it, the title "Statistics > Methodology" is shown, along with a submission date "Submitted on 5 Dec 2023". The full title of the paper is "General Spatio-Temporal Factor Models for High-Dimensional Random Fields on a Lattice". The authors listed are "Matteo Barigozzi, Davide La Vecchia, Hang Liu". The abstract begins with: "Motivated by the need for analysing large spatio-temporal panel data, we introduce a novel dimensionality reduction methodology for  $n$ -dimensional random fields observed across a number  $S$  spatial locations and  $T$  time periods. We call it General Spatio-Temporal Factor Model (GSTFM). First, we provide the probabilistic and mathematical underpinning needed for the representation of a random field as the sum of two components: the common component (driven by a small number  $q$  of latent factors) and the idiosyncratic component (mildly cross-correlated). We show that the two components are identified as  $n \rightarrow \infty$ . Second, we propose an estimator of the common component and derive its statistical guarantees (consistency and rate of convergence) as  $\min(n, S, T) \rightarrow \infty$ . Third, we propose an information criterion to determine the number of factors. Estimation makes use of Fourier analysis in the frequency domain and thus we fully exploit the information on the spatio-temporal covariance structure of the whole panel. Synthetic data examples illustrate the applicability of GSTFM and its advantages over the extant generalized dynamic factor model that ignores the spatial correlations."

## Setting: SRD and LRD

For the sake of exposition, suppose that  $\{Y_t\}$  admits a linear second order stationary process,  $\{Y_t, t \in \mathbb{Z}\}$  (see e.g. Hannan (1970), Ch. 2 or Beran et al. (2010), Ch. 2) with spectral density function

$$f(\lambda, \theta) = |\lambda|^{-2d} L(\lambda, \vartheta), \quad \lambda \in \Pi = (-\pi, \pi] \quad (6)$$

where  $d \in [0, 0.5]$ ,  $\vartheta \in \mathbb{R}^p$  with  $p \geq 1$  and  $\theta = (d, \vartheta)$ .

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### Definition

We classify the process  $\{Y_t\}$  as short-range dependent (SRD) or long-range dependent (LRD)

- when  $d = 0$  and the function  $L(\cdot, \vartheta)$  is bounded with  $L(0, \vartheta) \neq 0$ , then the process  $\{Y_t\}$  features SRD
- Otherwise, the process  $\{Y_t\}$  features LRD— $f$  has a pole at  $\lambda = 0$ .

# Saddlepoint approximation (exponential-based)

First order asymptotic theory implies

$$\mathcal{V}_n \xrightarrow{\mathcal{D}} \mathcal{N}(0, V).$$

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First order asymptotic theory implies

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To have a better density approximation to  $f_{\hat{\theta}_n}$ , we may derive the saddlepoint density approximation  $g_{\hat{\theta}_n}$  treating the periodogram ordinates as independently and exponentially distributed r.v.'s: we use it to approximate the c.g.f.  $\mathcal{K}_{\mathcal{V}_n}$  ...

## Saddlepoint approximation (exponential-based)

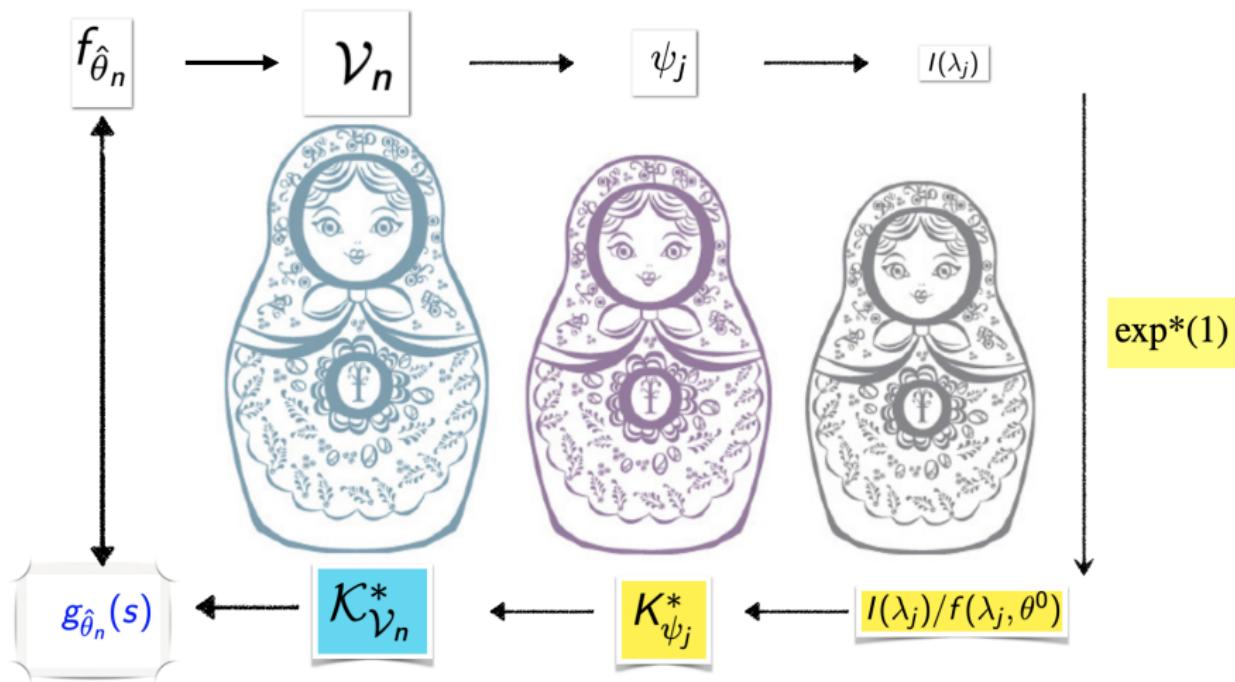
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$$\mathcal{V}_n \xrightarrow{\mathcal{D}} \mathcal{N}(0, V).$$

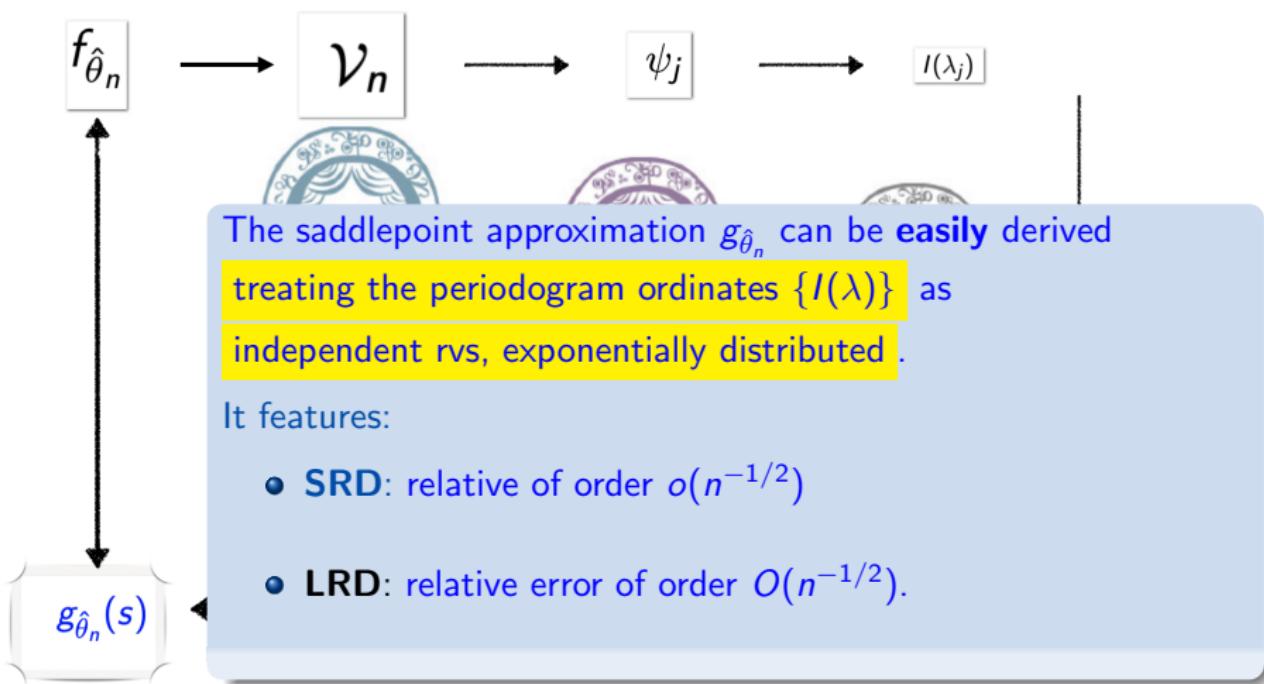
To have a better density approximation to  $f_{\hat{\theta}_n}$ , we may derive the saddlepoint density approximation  $g_{\hat{\theta}_n}$  treating the periodogram ordinates as independently and exponentially distributed r.v.'s: we use it to approximate the c.g.f.  $\mathcal{K}_{\mathcal{V}_n}$  ...

... let's visualize the idea ...

# Visualisation



# Visualisation



# Saddlepoint approximation (exponential-based)

Specifically:

- Whittle's estimating function is

$$\psi_j(I(\lambda_j), \theta) = \left( \frac{I(\lambda_j)}{f(\lambda_j, \theta)} - 1 \right) \nabla_{\theta} \ln f(\lambda_j, \theta),$$

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- define  $K_{\mathcal{V}_n}^*(v, s) = \sum_j K_{\psi_j}^*(v, s)$ , where

$$K_{\psi_j}^*(v, s) = \ln \left( E^* [\exp\{v\psi_j(I(\lambda_j), s)\}] \right),$$

with  $E^*$  computed treating  $I(\lambda_j)/f(\lambda_j, \theta^0) \sim \exp^*(1)$ .

## Saddlepoint approximation (exponential-based)

The saddlepoint density approximation is:

$$g_{\hat{\theta}_n}(s) = \left[ \frac{n}{2\pi \mathcal{K}_{\mathcal{V}_n}^{''}(v_0, s)} \right]^{1/2} e^{\mathcal{K}_{\mathcal{V}_n}^{*}(v_0, s)}, \quad (7)$$

and the saddlepoint  $v_0 = v_0(s)$  solves

$$\mathcal{K}_{\mathcal{V}_n}'(v, s) = 0.$$

### Remark

The advantage of using  $I(\lambda)/f(\lambda, \theta) \sim \exp(1)$  is that  $\mathcal{K}_{\mathcal{V}_n}^{*}$  is strictly convex, thus the saddlepoint equation admits a unique solution—which can be computed using standard methods, like the one based on the secant.

## Saddlepoint approximation (exponential-based)

### Example (ASY and frequency domain bootstrap (FDB))

Let us consider the AR(1):

$$Y_t = \theta_0 Y_{t-1} + \epsilon_t, \quad \epsilon_t \sim t_6 \quad \theta^0 = 0.4.$$

and the Whittle's estimator  $\hat{\theta}_n$ . Goal: approximate  $P_{\theta^0}(\hat{\theta}_n > t_0)$  (e.g. for  $p$ -value).

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12.5%

10%

5%

2.5%

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	12.5%	10%	5%	2.5%
$n = 36$				
SAD	12.2%	9.1%	4.4%	2.0%
ASY	15.0%	11.8%	6.4%	3.2%
FDB	—	—	—	—

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SAD	12.2%	9.1%	4.4%	2.0%
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$n = 150$				
SAD	12.7%	9.9%	4.9%	2.3%
ASY	12.1%	9.2%	4.4%	2.0%
FDB	13.5%	10.8%	5.6%	2.9%
$(q_1; q_3)$	$(10.5\%; 15.7\%)$	$(8.0\%; 12.7\%)$	$(4.0\%; 6.6\%)$	$(2.0\%; 3.5\%)$

## Saddlepoint approximation (exponential-based)

More generally, let  $\theta = (\theta^{(1)}, \theta^{(2)})$ , where  $\theta^{(2)} \in \mathbb{R}^{p_2}$ ,  $1 < p_2 < p$  and consider testing

$$\mathcal{H}_0 : \theta^{(2)} = 0 \text{ vs } \mathcal{H}_1 : \theta^{(2)} > 0$$

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- adapt the **univariate saddlepoint test statistic** of **Robinson et al (2003. AoS)**:

$$\tilde{\mathcal{K}}^\dagger(\hat{\theta}_n^{(2)}) = 2 \inf_{\theta^{(1)}} \left[ \sup_v \left\{ - \sum_j K_{\psi_j}^*(v; (\theta^{(1)}, \hat{\theta}_n^{(2)})) \right\} \right],$$

where  $v$  solves the saddlepoint equation. The distribution of  $\tilde{\mathcal{K}}^\dagger(\hat{\theta}_n^{(2)})$  under the null, can be approximated by a  $\chi^2_{p_2}$  and it

**is asymptotically first order equivalent to the Wald test.**

## Saddlepoint approximation (exponential-based)

### Example (Gaussian ARFIMA (0, d, 0))

Testing about the long-memory (simple hypothesis) for  $n = 100, 250$ :

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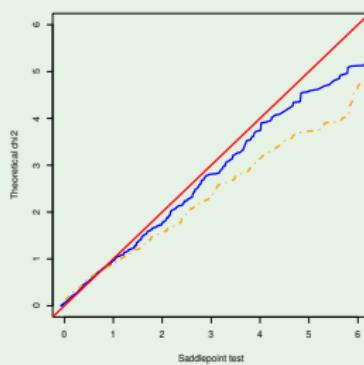
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Power



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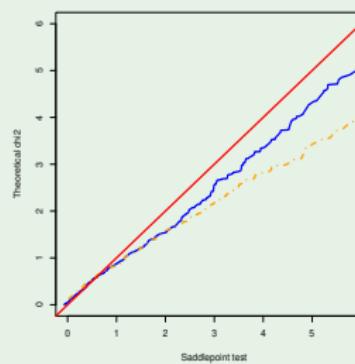
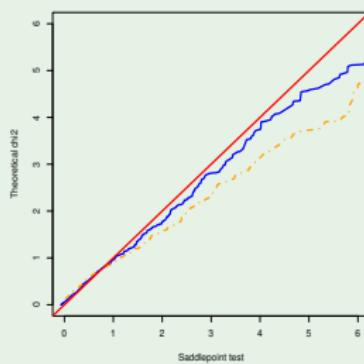
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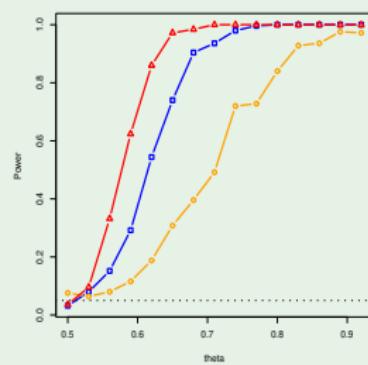
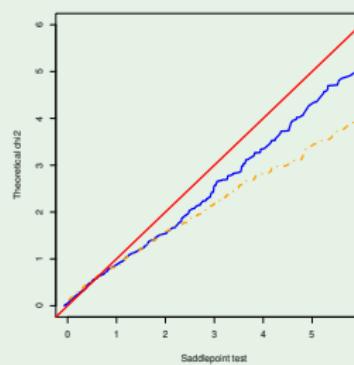
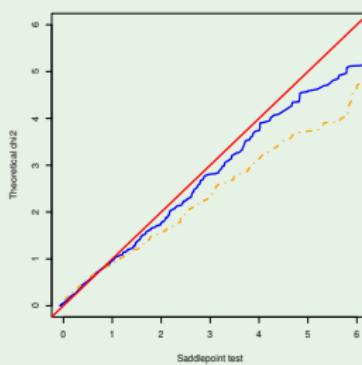
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- Dahlhaus & Janas (1996, AoS) (FDB)
- Monti (1997, Biom.) (FDEL) and Nordman & Lahiri (2006, AoS)
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*This suggests that the c.g.f. may be approximated using the empirical distribution of the periodogram ordinates and obtain novel saddlepoint techniques!*

# Saddlepoint approximation (empirical version)

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The screenshot shows a Cornell University logo and the text "Cornell University". Below it, the arXiv logo and the URL "arXiv > stat > arXiv:2403.12714" are visible. The title "Statistics > Methodology" is followed by the submission date "Submitted on 19 Mar 2024". The main title of the paper is "On the use of the cumulant generating function for inference on time series". The authors listed are "Alban Moor, Davide La Vecchia, Elvezio Ronchetti". The abstract begins with: "We introduce innovative inference procedures for analyzing time series data. Our methodology enables density approximation and composite hypothesis testing based on Whittle's estimator, a widely applied M-estimator in the frequency domain. Its core feature involves the (general Legendre transform of the) cumulant generating function of the Whittle likelihood score, as obtained using an approximated distribution of the periodogram ordinates. We present a testing algorithm that significantly expands the applicability of the state-of-the-art saddlepoint test, while maintaining the numerical accuracy of the saddlepoint approximation. Additionally, we demonstrate connections between our findings and three other prevalent frequency domain approaches: the bootstrap, empirical likelihood, and exponential tilting. Numerical examples using both simulated and real data illustrate the advantages and accuracy of our methodology." The footer includes "Subjects: Methodology (stat.ME); Computation (stat.CO)" and a small logo.

# Saddlepoint approximation (empirical version)

The empirical saddlepoint density approximation is

$$\hat{g}_{\hat{\theta}_n}(s) = \left(\frac{m}{2\pi}\right)^{p/2} \left| \det \hat{M}(s) \right| \left| \det \hat{\Sigma}(s) \right|^{-1/2} \exp\{m \hat{K}(s)\}, \quad (8)$$

where

$$\hat{K}(s) = \hat{K}(\hat{v}, s) = \ln \left[ \frac{1}{m} \sum_{j=1}^m \exp\{\hat{v}^T \psi_j(I_j, s)\} \right], \quad (9)$$

$$\hat{M}(s) = \frac{1}{m} \exp\{-\hat{K}(s)\} \sum_{j=1}^m \nabla_w \psi_j(I_j, w)|_{w=s} \exp\{\hat{v}^T \psi_j(I_j, s)\},$$

$$\hat{\Sigma}(s) = \frac{1}{m} \exp\{-\hat{K}(s)\} \sum_{j=1}^m \psi_j(I_j, s) \psi_j(I_j, s)^T \exp\{\hat{v}^T \psi_j(I_j, s)\}$$

and the empirical saddlepoint  $\hat{v}$  satisfies:

$$\sum_{j=1}^m \psi_j(I_j, s) \exp\{\hat{v}^T \psi_j(I_j, s)\} = 0. \quad (10)$$

## Saddlepoint approximation (empirical version)

The empirical saddlepoint is based on the c.g.f.  $\hat{K}$  as an approximation to the true c.g.f.: it is the key tool needed to compute  $\hat{g}_{\hat{\theta}_n}$  and it unveils important connection with the FDEL.

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Indeed, FDEL solves the system of (tilted) estimating equations

$$\sum_{j=1}^m \psi_j(I_j, s)[1 + \hat{\xi}^T \psi_j(I_j; s)]^{-1} = 0, \quad (11)$$

where we use the shorthand notation  $\hat{\xi} = \hat{\xi}(s)$ . Then, Monti defines a FD version of Owen's statistics as

$$\hat{W}(s) = 2 \sum_{j=1}^m \ln\{1 + \hat{\xi}^T \psi_j(I_j; s)\},$$

and  $\hat{W}$  is first-order equivalent to  $W_n$ .

## Saddlepoint approximation (empirical version)

Now notice that

- the saddlepoint satisfies (Taylor expansion of the exp) the equation

$$\sum_{j=1}^m \psi_j(I_j; s)[1 + \hat{v}^T \psi_j(I_j; s)] = O_P(n^{-1}),$$

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### Remark

The empirical saddlepoint and the empirical likelihood solve at the order  $O_P(n^{-1})$  the same equation.

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Building on this remark, we prove that:

$$-2n \underbrace{\hat{K}(s)}_{\text{Emp Sadd Test}} = 2 \underbrace{\hat{W}(s)}_{\text{Owen stat}} - \frac{2m^{-1/2}}{3} \sum_{j=1}^m \left\{ u^T \hat{M}^T \hat{\Sigma}^{-1} \psi_j(I_j; \hat{\theta}_n) \right\}^3 + R_n$$

where, under some conditions,  $R_n = O_P(n^{-1})$ ,  $\hat{\Sigma} = \hat{\Sigma}(\hat{\theta}_n)$  and  $\hat{M} = \hat{M}(\hat{\theta}_n)$ .

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- (ii) it illustrates that the difference between  $\hat{K}$  and  $\hat{W}$  depends on the third moment of the Whittle's score: both correct  $W_n$  for the skewness but in a different way
- (iii) it yields a nonparametric approximation to the density of Whittle's estimator based on the FDEL

## Saddlepoint approximation (empirical version)

On the practical side: use the empirical saddlepoint under  $\mathcal{H}_0$  to approximate the distribution of Wald-type (or EL, ET) test statistics, where

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To this end,

- We define the Wald-type statistic, with  $\hat{V} = \hat{M}^{-1}\hat{\Sigma}\hat{M}^{-1}$  (estimate of asym var of Whittle estim.),

$$W_n(\theta) = n(\hat{\theta}_n - \theta)^T \hat{V}^{-1}(\hat{\theta}_n - \theta).$$

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- In contrast, we make use of  $\hat{g}_{\hat{\theta}_n}$  to obtain

$$P[W_n(\theta^0) > \tilde{w}(\theta^0) | \mathcal{H}_0] \approx 1 - \int_{\mathcal{B}} \hat{g}_{\hat{\theta}_n}(\theta) d\theta, \quad (12)$$

where  $\tilde{w}(\theta^0)$  is the observed value of the test statistic and

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- To compute the integral in (12), we suggest to use an importance sampling scheme based on an instrumental Gaussian distribution.

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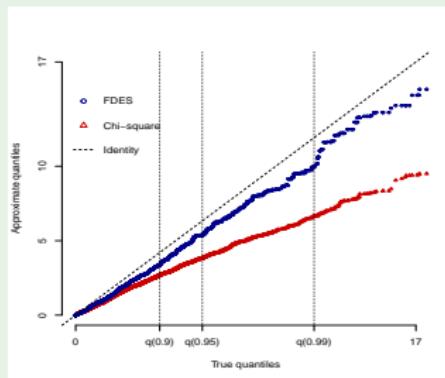
## Example

We consider **ARFIMA**(0,  $d$ , 0) model test for

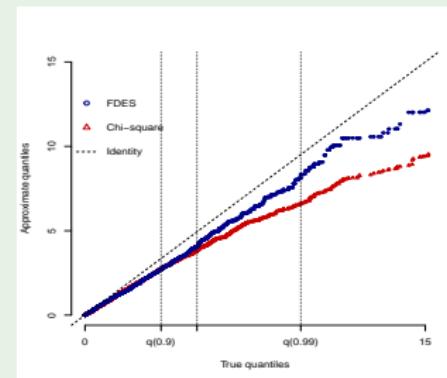
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using the empirical saddlepoint. We compare the approx quantiles to true quantiles (as obtained by MC simulations), for the **saddlepoint technique** and **first-order asymptotic theory ( $\chi_1^2$ )**.

$n = 30$



$n = 250$



# Saddlepoint approximation (empirical version)

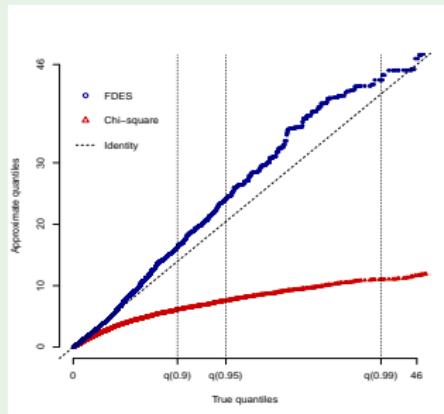
## Example

We consider an **ARFIMA(1,d,1)** with  $\theta^0 = (0.5, 0.25, 0.5)$  and test

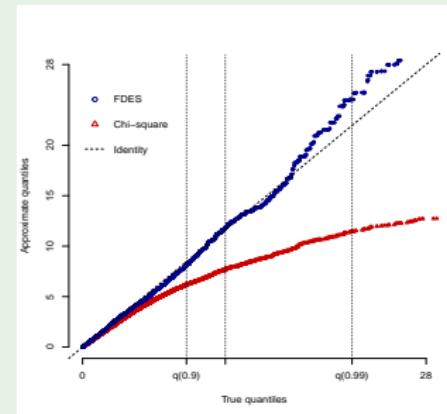
$$\mathcal{H}_0 : \theta = \theta^0 \text{ vs. } \mathcal{H}_1 : \theta \neq \theta^0$$

using the empirical saddlepoint. We compare the approx quantiles to true quantiles (as obtained by MC simulations), for the **saddlepoint technique** and **first-order asymptotic theory ( $\chi_3^2$ )**.

$n = 100$



$n = 500$



# Saddlepoint approximation (empirical version)

## Remark

- **Testing in the presence of nuisance.** There are cases where only certain components of  $\theta$  have to be tested. Namely, taking the partition  $(\theta_{(1)} \ \theta_{(2)})$ , we test

$$\mathcal{H}_0 : \theta_{(1)} = \theta_{(1)}^0$$

w.l.o.g. To perform this type of test, we can simply modify our procedure, redefining the integration set as

$$\mathcal{B} \leftarrow \left\{ \theta \in \Theta \mid \tilde{w} \left( \theta_{(1)}, \hat{\theta}_{(2),n} \right) > \tilde{w} \left( \theta_{(1)}^0, \hat{\theta}_{(2),n} \right) \right\},$$

instead of

$$\mathcal{B} \leftarrow \left\{ \theta \in \Theta \mid \tilde{w}(\theta) > \tilde{w}(\theta^0) \right\}.$$

⇒ our numerical integration via importance sampling avoids to take the infimum w.r.t. to the nuisance parameters.

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- **Numerical accuracy.** Also using the empirical distribution of the periodogram ordinates, the saddlepoint technique yields an improvement on the first order asymptotic theory.

Thank you

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# Laplace in brief

The Laplace method is typically applied to approximate integrals of type:

$$\int_a^b e^{v k(x)} dx,$$

where  $k(\cdot)$  has unique maximum at  $x_0 \in (a, b) \subset \mathbb{R}$  and  $v \in \mathbb{R}^+$  is large.

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