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Abstract An accurate estimation of mortality rates is essential to make decisions. For example, the ensures companies, banks, among others, plans its operations according to estimations based on these rates. However, the vast number of variables and its intricate relation implies a challenge for the estimation of these rates.

We assume the following hypothesis: the mortality rate has a strong relationship with its owns past. In this line, we propose a stochastic model with long-term memory that describes mortality. Then, using data from Italy, we provide statistical evidence via Hurst parameter estimation that does not reject our hypothesis. Further, we extract a subset of data to evaluate its forecasting performance, and we observe a good estimation.

To the best of our knowledge, our contribution is the first attempt that includes long term memory in the formulation of a model to describe mortality rates. Our results suggest that the hypothesis of an imperfect correlation intensity across generations would be more realistic.

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1 Introduction

Future planning in the demographic, economic, and actuarial areas is crucial. For instance, proper planning in social programs, government budgets, cost of insurance, and others depends on the forecast. However, constant changes in technology, lifestyle, migration, to name a few, make predicting a demanding task. Mortality impacts directly in cash and therefore need a reliable future projection.

Previous work has only focused on deterministic or stochastic models without memory effects. Pitacco et al. (2009) review the first mortality tables and models. Milevsky and Promislow (2001) report a linear SDE driven by Brownian motion that describes the mortality hazard rate. In Giacometti et al. (2011), the authors extend the Mikevesky model to an SDE with time-dependent diffusion and study a type of autoregressive model for the logarithm of the hazard rate. Jevtić et al. (2013) formulate a cohort-based model with imperfect correlation across generations and use data from UK to estimate parameters.

This paper aims to show statistical evidence that mortality rates follow a stochastic process with long-range dependence (LRD), that is, a stochastic process with memory. According to Prakasa Rao (2010), a stochastic process or time series is LRD, if it has persistence behavior—below we give a formal definition.

We base our stochastic formulation in the fractional Brownian Motion (fBM) (Mandelbrot and Van Ness, 1968, see). Since fBM is a generalization of the standard Brownian Motion (BM) that still satisfies self-similarity and is LRD, fBM results to be a natural noise model to describe LRD behavior.

The main idea is to extend the stochastic model reported by Milevesky and Promislow to a model with LRD and verify its performance to fitting and forecasting with real data.

We obtain statistical evidence—via Hurst parameter estimation—that Italy mortality rate data is LRD. Our model captures women's mortality rate dynamics. But we observe that the Italian man has a higher mortality rate and variance. These results suggest that stratification by gender would be a direction for future formulations.

Our work exhorts to explore the long time memory effects of dynamics under uncertainty. Also these ideas could improve mathematical models, which considers mortality in its formulation, for example, models for infectious disease spread or population growth.

After this brief introduction, we formulate our fractional model in Section 2. Section 3 reviews the fBM and presents the fractional Ornstein-Uhlenbeck (fOU) process. Section 4 outlines the applied method for parameter estimation. In Section 5, we implement our formulation to fitting data of mortality from Italy. Further, in this section, we also run forecasting to a subset of the data and evaluate its efficiency. Finally, we conclude in Section 7.

2 Model formulation

Our formulation relies on the model of Milevsky and Promislow (2001), bellow we review the main ideas.

Let Y_t the solution process to the stochastic differential equation driven by standard Brownian motion.

$$\begin{aligned} dY_t =& - \lambda Y_t dt + \sigma dB_t, \\ & \lambda, \sigma > 0,\ Y_0 = 0,\ t \in [t_0, T]. \end{aligned} \tag{1}$$

According to the Milevsky-Promislow model, h_x denotes the stochastic force of mortality at time t of an individual aged x. Letting $h_x(t) = h(t)$, and conforming with SDE (1), this ratio is described by

$$\begin{aligned} h(t) =& h_0 \exp(\alpha_0 t + \alpha_1 Y_t) \\ & h_0, \alpha_0, \alpha > 0. \end{aligned} \tag{2}$$

Next, the survival probability of an individual aged x in the period $[t, T]$ as

$$S(t,T) := \mathbb{E}\Big[\exp\Big(- \int_t^T h_x(u) du \Big) \Big| \mathcal{F}_t \Big], \tag{3}$$

where $\{\mathcal{F}_t\}_{t \geq 0}$ is a filtration which represent the information until time t.

Our formulation extends the Milevsky model eqs. (1) to (3) by utilizing fractional Brownian motion. That is, instead of the Ornstein–Uhlenbeck process (OU) defined by SDE (1), we employ its fractional version (fOU)

$$\begin{aligned} dY_t^H =& -\lambda Y_t^H dt + \sigma dB_t^H, \\ & \lambda, \sigma > 0,\ Y_0 = 0,\ H \in [1/2, 1). \end{aligned} \tag{4}$$

In fact we implement the integral form of the above fractional SDE

$$Y_t^H = -\lambda \int_0^t Y_s^H ds + \sigma B_t^H. \tag{5}$$

fOU process is long time-dependent when $H \in [1/2, 1)$ (for details see Cheridito et al., 2003; Anh et al., 2002; Hu and Nualart, 2005; Kleptsyna, 2002). Then, we use the fOU to explore the long time dependence of mortality hazard-rates.

3 Fractional Gaussian processes

3.1 Fractional Brownian motion

We consider the Gaussian process$\{B_t^H, t \geq 0\}$, with $H \in (0, 1)$, zero-mean and covariance function given by

Introductory paragraph

$$R_H(t,s) := \mathbb{E}(B_s^H B_t^H) = \tfrac{1}{2}\big(t^{2H} + s^{2H} - |t-s|^{2H}\big). \tag{6}$$

This stochastic process is the so called *fractional Brownian motion* (fBM). Kolmogoroff introduced this fractional process in 1940 and Mandelbrot and Van Ness gives an in-depth study in 1968. The parameter H is called Hurst index in honor to the statistical analysis developed by the climatologist Harold Edwin Hurst.

The fBm is a generalization of Brownian motion without independent increments, in fact, it is a continuous-time Gaussian process wich has the properties of self-similarity and stationary increments. Its sample-paths are almost nowhere differentiable and almost-all Hölder continuous for any order strictly less than H. That is, for each trajectory, there exists a finite constant C such that for every$\epsilon > 0$

$$\mathbb{E}\big(|B_t^H - B_s^H|\big) \leq C|t-s|^{H-\epsilon}.$$

For $H = \frac{1}{2}$, the covariance of fBM follows $R_{1/2}(t,s) = \min(s,t)$, thus process $B_t^{1/2}$, is equivalent to the standard Brownian motion. However, if $H \neq \frac{1}{2}$, then this increments are dependent. Let $X_n = B_n^H - B_{n-1}^H$, $n \geq 1$. Then $\{X_n, n \geq 1\}$ is a Gaussian stationary sequence with unit variance and covariance function

$$\begin{aligned}\rho_H(n) &= \frac{1}{2}\Big((n+1)^{2H} + (n-1)^{2H} - (2n)^{2H}\Big)\\ &\approx H(2H-1)n^{2H-2} \to 0, \quad n \to \infty.\end{aligned}$$

Therefore,

- if $H > \frac{1}{2}$, then $\rho_H(n) > 0$ for n large enough and $\sum_{n=1}^{\infty} \rho_H(n) = \infty$. In this case, we say that process X_n is persistent with positive correlation and that X_n has the *long-range dependence* property
- if $H < \frac{1}{2}$, then $\rho_H(n) < 0$ for n large enough and $\sum_{n=1}^{\infty} \rho_H(n) < \infty$. Thus we say that X_n is an anti-persistent process with negative correlation.

For further information on fBM see Prakasa Rao (2010); Nualart (2006); Mishura (2008).

3.2 The Fractional Ornstein-Uhlenbeck process

fOU can be defined as the unique solution of a Langevin SDE driven by an fBm.Cheridito et al. introduces this fractional process and shows that

$$Y_t^H = \sigma \int_0^t e^{-\lambda(t-u)} dB_u^H, \tag{7}$$

is the unique continuous solution of the fractional Langevin SDE (5). For $H \in (1/2, 1)$, the fOU process is neither Markovian nor a semimartingale but Gaussian and ergodic. Moreover, when $H \in (1/2, 1)$, Y_t longe-range dependent. For a more in-depth discussion of this fractional process, we refer the reader to Cheridito et al. (2003); Prakasa Rao (2010); Dudley and Norvaiša (2011).

According to Zeng et al. (2012), the variance of the fOU process Y_t is given by the following expression:

$$\operatorname{Var}(Y_t) = \sigma^2 2He^{-2\lambda t}\int_0^t s^{2H-1}e^{2\lambda s}ds. \tag{8}$$

Notice that when $H = 1/2$ we get

$$\operatorname{Var}(Y_t) = \frac{\sigma^2}{2\lambda}\big(1 - e^{-2\lambda t}\big), \tag{9}$$

which is the variance of the standard Ornstein-Uhlenbeck process (see Mikosch, 1998, p. 143).

Let $\alpha_1 = T^{-H}$, thus from eq. (8), we deduce

$$\begin{aligned}\operatorname{Var}(\alpha_1 Y_t) &= \alpha_1^2 \operatorname{Var}(Y_t) = \alpha_1^2\sigma^2 2H\int_0^t s^{2H-1}e^{-2\lambda(t-s)}ds \\ &\leq \alpha_1^2\sigma^2 2H\int_0^t s^{2H-1}ds = \alpha_1^2\sigma^2 2H\frac{s^{2H}}{2H}\bigg|_{s=0}^t \\ &= \alpha_1^2\sigma^2 t^{2H} = \sigma^2(t/T)^{2H} \\ &\leq \sigma^2.\end{aligned} \tag{10}$$

That is, a constant independent of time bounds the variance of $\alpha_1 Y_t$. Thus we can use $\alpha_1 = T^{-H}$ and σ to modulate the variance of fOU.

4 Estimation of the parameters

In this section, we describe the applied methodology to estimate parameters. According to (2), we need to estimate α_0, α_1 and for fractional SDE model (1) σ, λ, and Hurst index H involved in the fBm. To estimate this index, we appeal to empirical evidence that H in equation (4) is invariant, that is, the value of the Hurst index in the equation (4), for the fBm B_t^H and the Hurst parameter for fOU Y_t^H are equivalent. We also assume that $\alpha_1 = T^{-H}$.

4.1 Estimation of the parameter α_0.

To estimate α_0 we assume $\alpha_1 = 1$. Moreover, h_0 is the initial value of the observed rates $h(t)$. Thus, from (2) we obtain

$$\ln h(t) = \ln h_0 + \alpha_0 t + Y_t. \tag{11}$$

We minimize the sum of the square errors to estimate parameter α_0. Let

$$S := \sum_{t_{initial}}^{t_{final}} \Big(\ln h(t) - \ln h_0 - \widehat{\alpha_0}t\Big)^2.$$

Taking derivative of S with respect to $\widehat{\alpha_0}$ we get

$$\frac{\partial S}{\partial \widehat{\alpha_0}} = -2 \sum_{t_{initial}}^{t_{final}} \Big(\ln h(t) - \ln h_0 - \widehat{\alpha_0} t \Big) t = 0,$$

and from this equation we deduce that $\widehat{\alpha_0}$ satisfies:

$$\widehat{\alpha_0} = \frac{\sum_t t \ln h(t) - \ln h(0) \sum_t t}{\sum_t t^2}. \tag{12}$$

Once we have estimated α_0 we proceed to estimate the Hurst parameter, σ and λ.

4.2 The relation between the Hurst parameter and the H-index in the FOU

In this subsection, we outline a method to estimate the parameter H. We use the following empirical fact. Suppose that an fOU process is driving with an fBM with a given Hurst parameter H_0. Yerlikaya-Özkurt et al. (2014) provide a relationship between the Hurst parameter H of the fractional Brownian motion and the Hurst parameter of the fractional Gaussian noise given by an SDE. They provide statistical evidence that the fOU should have the same value H0 that the fBM. Thus, under this empirical evidence, we assume the same value of the parameter H for both processes. A formal result of this fact, up to our knowledge, still is an open question.

The following sections present several methods to estimate the Hurst parameter for the fBM.

4.3 Estimation of the self-similarity index H for the fBM

The last subsection allows us to estimate the parameter H as follows. According to equation (11), the residuals are

$$\hat{Y}_t = \ln h(t) - \ln h_0 - \widehat{\alpha}_0 t.$$

Note taht $\hat{Y}_t$ is a fractional Gaussian process that we can use to estimate H. Afterwards we use estimation $\hat{H}$ to approximate the Hurst parameters of the fractional Brownian motion B_t^H. To this end, we review some methods to estimate the parameter H, our main refernces are Weron (2002), .

complete references

4.3.1 R / S Analysis

Following Weron (2002), we coniser a finite time serie of lenght L indexed by i Z_i. Thus, let $\{Z_{i,m}\}$, $m = 1, \ldots, d$ be a finite sequence of d series of lenght n, such that $L = n \times d$. Next, for each subseries $\{Z_{i,m}\}$, $m = 1, \ldots, d$:

1. Find the mean E_m and standard deviation S_m.

2. Normalize the data $Z_{i,m}$ by subtracting the sample mean $X_{i,m} = Z_{i,m} - E_m$ for $i = 1, \ldots, n$.
3. Create a cumulative time series $Y_{i,m} = \sum_{j=1}^{i} X_{j,m}$. for $i = 1, \ldots, n$
4. Find the range $R_m = \max\{Y_{1,m}, \ldots, Y_{n,m}\} - \min\{Y_{1,m}, \ldots, Y_{n,m}\}$;
5. Rescale the range R_m/S_m .
6. Calculate the mean value of the rescaled range for all subseries of length n

$$(R/S)_n = \frac{1}{d} \sum_{m=1}^{d} R_m/S_m.$$

Accoridng to Mandelbrot (1974/75), the R/S statistic asymptotically follows the relation:

$$(R/S)_n \sim cn^H,$$

where c is a constant. Thus, we can estimate H by linear regression over a sample of increasing time horizons

$$\log(R/S)_n = \log c + H \log n.$$

Equivalently, we can plot the $(R/S)_n$ statistics against n on a double-logarithmic scale. If the returns process is white noise, then the plot is roughly a straight line with slope 0.5. If the process is persistent, then the slope H is greater than 0.5. If the processes is anti-persistent, then the slope H is less than 0.5. The "significance" level of the estimated parameter H usually is chosen to be inverse proportional to the square root of sample length, that is, the standard deviation of a Gaussian white noise.

A major drawback of the R/S analysis is that no asymptotic distribution theory has been derived for the Hurst parameter H. The only results known are for the rescaled (but not by standard deviation) range R_m itself, see Lo (1991).

4.3.2 Method of rescaled range analysis R/S

Here we follow Prakasa Rao (2010, Chap. 9). This method was suggested by Hurst (1951). The series $\{X_j, 1 \leq j \leq N-2\}$ is divided into K nonoverlapping blocks such that each block contains M elements where M is the integer part of N/K. Let $t_i = M(i-1)$, where $t_i = M(i-1)$ is the starting point of the ith block for $i = 1, \ldots, K$. Define

$$R(t_i, r) = \max[W(t_1, 1), \ldots, W(t_i, r)] - \min[W(t_1, 1), \ldots, W(t_i, r)],$$

where r takes values in natural number whenever r satisfy the inequality $t_i + r \leq N$. Moreover, $W(t_i, k)$ is set as

$$W(t_i, k) = \sum_{j=0}^{k-1} X_{t_i+j} - k \left(\frac{1}{r} \sum_{j=0}^{r-1} X_{t_i+j} \right), \quad k = 1, \ldots, r.$$

Note that $R(t_i, r) \geq 0$ since $W(t_i, r) = 0$ and the quantity $R(t_i, r)$ can be computed only when $t_i + r \leq N$. Define

$$S^2(t_i, r) = \frac{1}{r} \sum_{j=0}^{r-1} X_{t_i+j}^2 - \left(\frac{1}{r} \sum_{j=0}^{r-1} X_{t_i+j} \right)^2 .$$

The ratio $R(t_i, r)/S(t_i, r)$ is called the rescaled adjusted range. It is computed for a number of values of r that makes sense according to the definition. Observe that, for each value of r, we obtain a number of R/S samples. The number of samples decrease as r increases. However, the resulting samples are not independent. It is believed that the R/S-statistic is proportional to r^H as $r \to \infty$ for the fractional Gaussian noise. Assuming this property, it is possible to regress $log(R/S)$ against $log(r)$ to obtain an estimator for H.

4.3.3 FDWhittle Estimator

Following Park et al. (2011). The Local Whittle Estimator (LWE) is a semi-parametric Hurst parameter estimator based on the periodogram. LWE assumes that the spectral density $f(\omega)$ of the process can be approximated by the function

$$f_{c,H}(\omega) = c\omega^{1-2H}, \tag{13}$$

for frequencies ω in a neighborhood of the origin, c is a constant. The periodogram of a time series $\{X_t, 1 \geq t \geq N\}$ is defined by

$$I_N(\omega) = \frac{1}{2\pi N} \left| \sum_{t=1}^{N} X_t e^{i\omega t} \right|^2 ,$$

where $i = \sqrt{-1}$. Usually, it is evaluated at the Fourier Frequencies $\omega_{j,N} = \frac{2\pi j}{N}$, $0 \leq j \leq [N/2]$. Note that the periodogram is the norm of the Discrete Fourier transform of the time series (see, for example, Priestley, 1981, sect. 6.1.2).

The LWE of the Hurst parameter, $\hat{H}_{LWE}(m)$ is implicitly the result of minimizing

$$\sum_{j=1}^{m} \log f_{c,H}(\omega j, N) + \frac{I_N(\omega j, N)}{f_{c,H}(\omega j, N)},$$

with respect to c and H, with $f_{c,H}$ defined in (13).

4.4 Estimation of σ and λ

There are several methods to estimate parameters σ and λ. For instance see Prakasa Rao (2010) or the references in Neuenkirch and Tindel (2014) or in Kubilius and Mishura (2012). In the following section we will do a brief review of some of these methods.

4.4.1 Estimation σ.

Brouste and Iacus (2013) proposed some consistent and asymptotically Gaussian estimators for the parameters σ, λ and H of the discretely observed fractional Ornstein-Uhlenbeck process solution expressed in the stochastic differential equation. There is a restriction on the estimation of the parameter λ —the results are valid only when $1/2 < H < 3/4$.

The key point of this estimation method is that the Hurst exponent H and the diffusion coefficient σ can be estimated without estimating λ. We use this method to estimate parameters σ and λ.

Let $\boldsymbol{a} = (a_0, \ldots, a_K)$ be a discrete filter of order $L \geq 1$ and length $K+1$, $K \in \mathbb{N}$ and we require $L \leq K$, i.e.

$$\sum_{k=0}^{K} a_k k^j = 0 \quad \text{for } 0 \leq j \leq L-1 \quad \text{and} \quad \sum_{k=0}^{K} a_k k^L \neq 0.$$

Let it be normalized

$$\sum_{k=0}^{K} (-1)^{1-k} a_k = 1.$$

We will also consider a *dilated* filter $\boldsymbol{a}^{(2)}$ associated to $\boldsymbol{a}$. For $0 \leq k \leq K$ we define

$$a_k^{(2)} = \begin{cases} a_{k'}, & \text{if } k = 2k'0, \\ & \text{otherwise.} \end{cases}$$

Since $\sum_{k=0}^{2K} a_k^2 k^j = 2^j \sum_{k=0}^{K} a_k k^j$ then the filter $\boldsymbol{a}^{(2)}$ has the same order than $\boldsymbol{a}$.

Let $Y^T = (Y_t : 0)$ be the sample path of the solution of (7). Consider the discretization of Y^T

$$(X_n := Y_{n\Delta_N}, n = 0, \ldots, N), \qquad N \in \mathbb{N},$$

where $\Delta_N = T/N$ and N is the number of observations of Y_t. We denote by

$$V_{N,\boldsymbol{a}} := \sum_{i=0}^{N-K} \left(\sum_{k=0}^{K} a_k X_{i+k} \right)^2,$$

the *generalized quadratic variation* associated to the filter $\boldsymbol{a}$ Istas and Lang (see, for example 1997). Then, define the following estimators for H and σ.

$$\hat{H}_N := \tfrac{1}{2} \log_2 \left(\frac{V_{N,\boldsymbol{a}^2}}{V_{N,\boldsymbol{a}}} \right), \tag{14}$$

$$\hat{\sigma}_N := \left(-2 \frac{V_{N,\boldsymbol{a}}}{\sum_{k,l} a_k a_l |k-l|^{2\hat{H}_N} \Delta_N^{2\hat{H}_N}} \right)^{1/2}. \tag{15}$$

Brouste and Iacus (2013) proved the following result.

Theorem 1 *Let $\boldsymbol{a}$ be a filter of order $L \geq 2$. Then, both estimators $\hat{H}_N$ and $\hat{\sigma}_N$ are strongly consistent, that is,*

$$(\hat{H}_N, \hat{\sigma}_N) \to (H, \sigma) \quad a.s, \; as \quad N \to \infty.$$

Moreover, for all $H \in (0,1)$, if $N \to \infty$, then

$$\sqrt{N}(\hat{H}_N - H) \xrightarrow{\mathcal{L}} N(0, \Gamma_1(\boldsymbol{a}, \sigma, H)),$$
$$\frac{\sqrt{N}}{\log N}(\hat{\sigma}_N - \sigma) \xrightarrow{\mathcal{L}} N(0, \Gamma_2(\boldsymbol{a}, \sigma, H))$$

where Γ_1 and Γ_2 are symmetric positive definite matrices depending on σ, H and the filter $\boldsymbol{a}$.

With this result we obtain an estimator for parameterσ. Consier the following filters.

- Classical filter. Let $K > 0$ and define
$$a_k := \frac{(-1)^{1-k}}{2^k}\binom{K}{k} = \frac{(-1)^{1-k}}{2^k}\frac{K!}{k!(K-k)!} \qquad \text{for } 0 \leq k \leq K.$$
- Daubechies filters (see Daubechies, 1992, for original definition). Let $K = 4$ and define the Daubechies filter by
$$\frac{1}{\sqrt{2}}\begin{bmatrix} 0.482\,962\,913\,144\,53 \\ -0.836\,516\,303\,737\,8 \\ 0.224\,143\,868\,042\,01 \\ 0.129\,409\,522\,551\,26 \end{bmatrix}.$$

4.4.2 Estimation of the drift parameter λ H and σ are known

Hu and Nualart (2010) proved that

$$\lim_{t\to\infty} Var(Y_t) = \lim_{t\to\infty} \frac{1}{t}\int_0^t Y_t^2 dt = \frac{\sigma^2 \Gamma(2H+1)}{2\lambda^{2H}} := \mu_2.$$

This equation gives a λ estimator, namely

$$\hat{\lambda}_N = \left(\frac{2\hat{\mu}_{2,N}}{\hat{\sigma}_N^2 \Gamma(2\hat{H}_N + 1)}\right)^{-\frac{1}{2\hat{H}_N}} \tag{16}$$

where $\hat{\mu}_{2,N}$ is the empirical moment of order 2, i.e

$$\hat{\mu}_{2,N} = \frac{1}{N}\sum_{n=1}^{N} X_N^2.$$

Set $T_N = N\Delta_N$. We have the next result.

Theorem 2 *Let $H \in \left(\frac{1}{2}, \frac{3}{4}\right)$ and a mesh satisfying the condition $N\Delta_N^p \to 0$, $p > 1$, and $\Delta_N(logN)^2 \to 0$ as $N \to +\infty$. Then, as $N \to +\infty$,*

$$\hat{\lambda}_N \to \lambda, \ \ a.s.$$

and

$$\sqrt{T_N}(\hat{\lambda}_N - \lambda) \xrightarrow{\mathcal{L}} N(0, \Gamma_3(\sigma, H)),$$

where $\Gamma_3(\sigma, H) = \lambda\left(\frac{\sigma_H}{2H}\right)^2$ and

$$\sigma_H^2 = (4H+1)\left(1 + \frac{\Gamma(1-4H)\Gamma(4H-1)}{\Gamma(2-2H)\Gamma(2H)}\right).$$

For the proof see (Brouste and Iacus, 2013, Thm. 2).

5 Results

In this section we present our estimated mortality rates according to obtain from Italy—the website of Human Mortality Database for the Italian population between 1950 to 2004.

We present the estimation of the H parameter. In second place, we present the results on simulated mortality rates using equations (15)-(16) to estimate the parameters σ, λ. The parameter α_0 has been fixed with the use of equation (12).

We sample 10 000 paths of the solution process to SDE (5) to estimate mortality rates of woman and man between 0 and 90 year old. To this end, we simulate paths of fractional Brownian motion $B_t^{\hat{H}}$ and using equation (**??**) we have estimated the mortality rate. To run the fBm simulations we have used the function **fbm** which is includes in the R library **somebm**. We also include a 95.5% confidence interval.

We present the results for women and men in sections 5.2 and 5.3, respectively.

Estimations and predictions were performed using R Ver. 3.2.3 (R Core Team, 2015), and specialized packages Fractal (Time Series Modeling and Analysis Version 2.0-1, 2016), Pracma (Practical Numerical Math Functions 2.0.7, 2017) and somebm (some Brownian motions simulation function Version 0.1, 2016).

5.1 Hurst parameter estimation

For the Hurst parameter estimate we have used three R routines: FDWhittle, RoverS and hurstexp. The two first routines are from the fractal library, while the latter is fromthe pracma library.

The former routine estimate the Hurst parameter by Whittle's method as was described in the subsection 4.3. RoverS routine estimate $\hat{H}$ by rescaled range (R/S) method. The hurstexp routine estimate $\hat{H}$ using R/S analysis.

Finally, figures 1 and 5.1 show the estimated Hurst parameter for women and men separately.

Fig. 1 Estimated Hurst parameter using R-routines.

With the rescaled range `R/S` and `hurst_exp` methods we obtain a consistent estimator for the Hurst parameter in the sense that they do not present dramatic changes through the time. Moreover, the H estimated with these two methods take values in the interval (0.57,0.80) approximately. This tells us that the data has the long memory property as was mentioned in section 3. Same results are obtained for the men and women. Notice that the estimated parameters for H using Whittle method have high variation through the time, in opposition to those obtained with the other two methods so that the estimated Hurst parameters using this method do not perform well in the simulations.

The high variation on the Hurst estimated values could be explained because Whittle method uses the periodogram to estimate H while the other two methods uses the raw data. The two approaches using Whittle likelihood and raw data are very different and hence they give very different estimates of H. Since rescaled range R/S and R/S methods have estimated very similar H, we decide to use the Hurst coefficients obtained with the method of R/S to perform the mortality rates simulations.

Fig. 2 Estimated Hurst parameter using R - routines.

5.2 Results for women

We present the results for 10 000 path simulations of the mortality rates for ages: $0, 5, 25, 50, 60, 70, 80, 90$. We graph the historical mortality rate, the mean of all simulations and the 95% confidence interval. See figures 3 and **??**.

In general, for all ages, the model is well fitted, in particular, after the 80's. Nevertheless, there are some time periods where the model is not so good as we want to. For instance, the model underestimates the mortality rate for women age 0 during the period of 50s to 80s and for women age 50, 60, 80 and 90 during 60s to 80s approximately. Moreover, it also overestimates the mortality rate for women age 25 during the period of 50s to 80s.

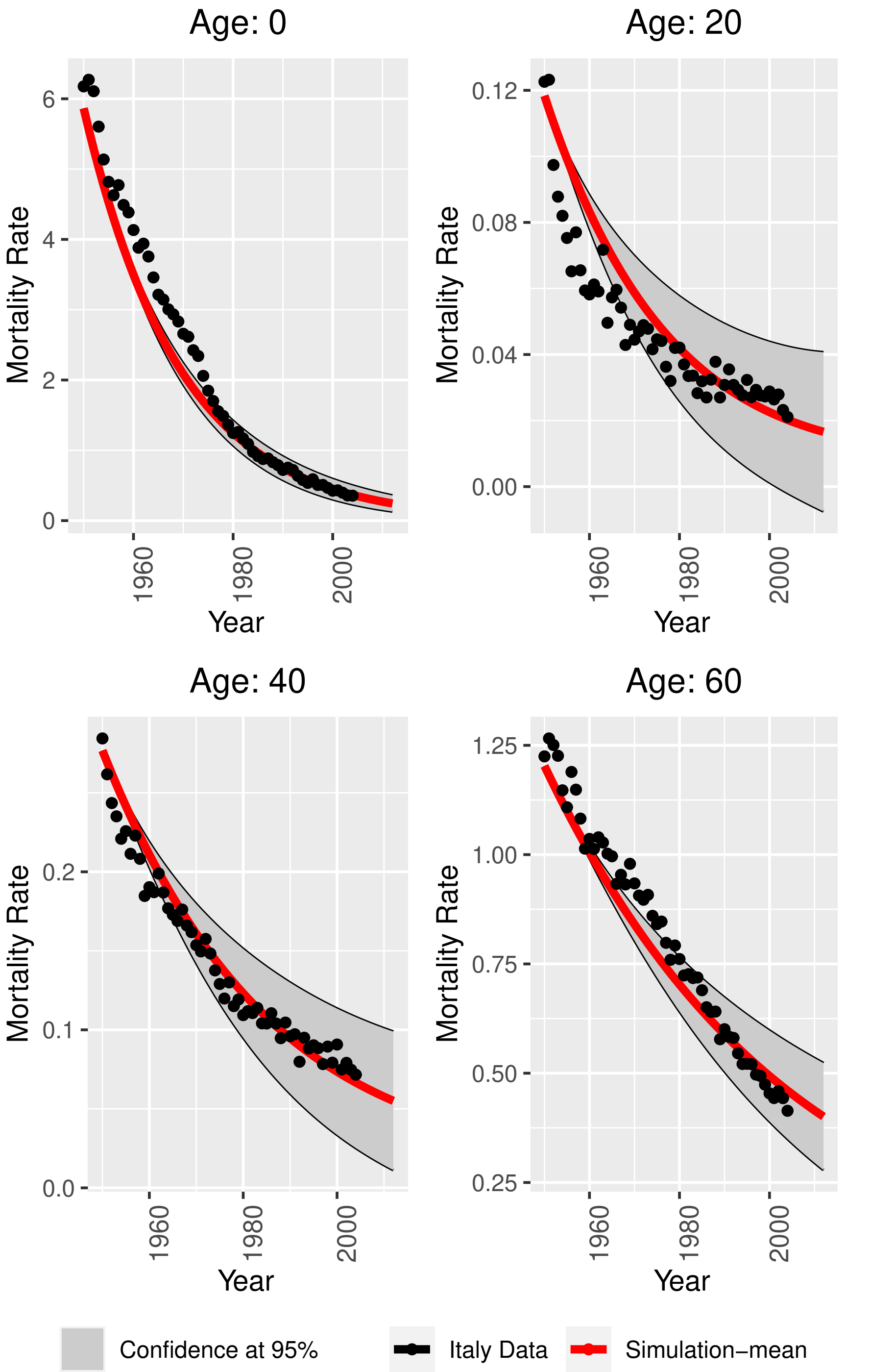

Fig. 3 Simulations for the rate mortality with the fOU model: ages 0, num20, 40, 60 and $N = 10\,000$ simulation paths.

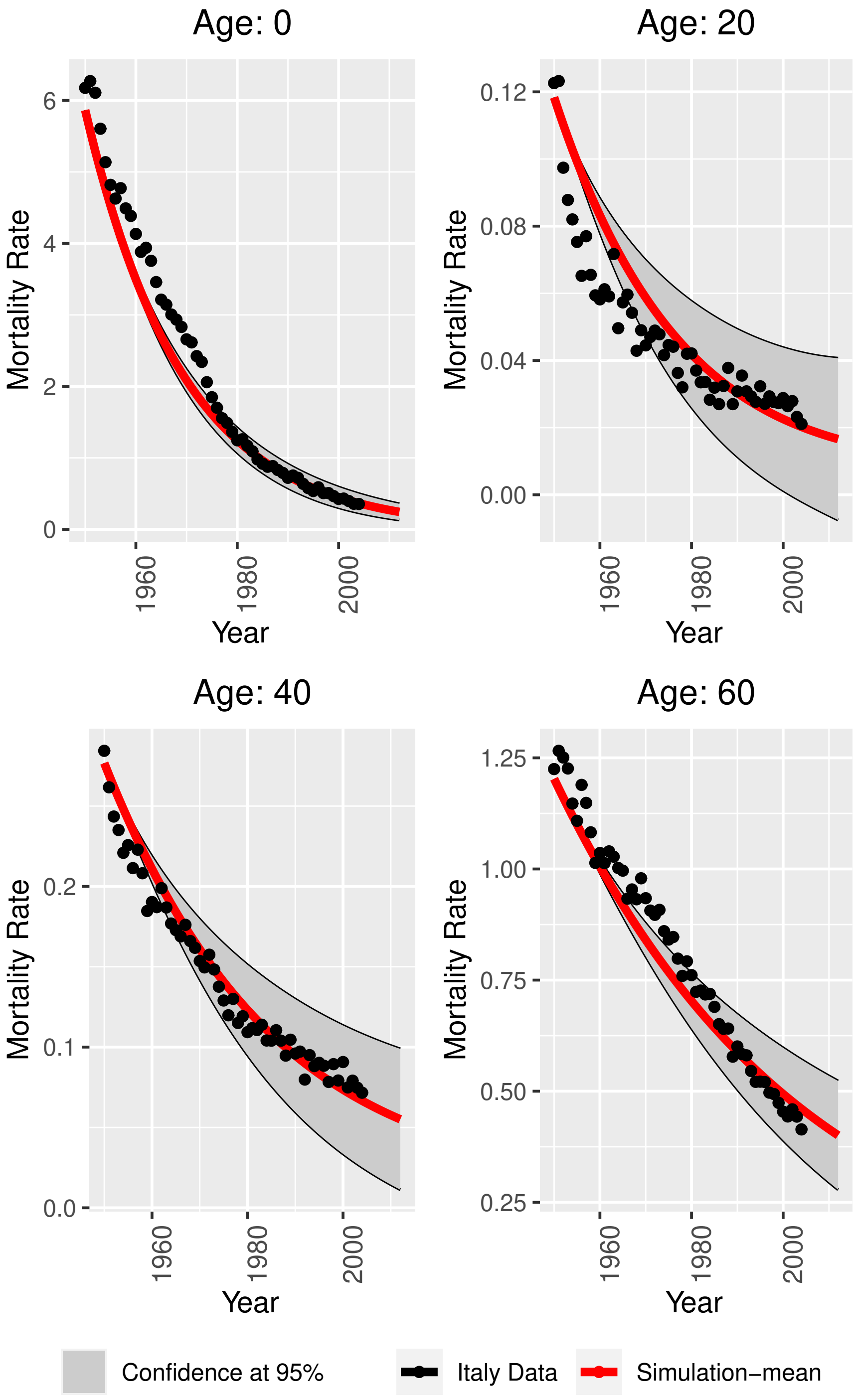

Fig. 4 Simulations for the rate mortality with the fOU model: ages 0, num20, 40, 60 and $N = 10\,000$ simulation paths.

For older ages (see figure **??**) we observe that for ages 60 and 70 the estimation is well fitted trough the years. We notice that predicted rates are not so far away and that the historical rates are inside the confidence interval. For the very oldest ages the estimation is not so good as for earlier ages. The main difference is in 50's when the absolute number of living persons arriving to those ages were small so that the variability of the estimates is larger.

All these suggest that a better model could include a short and a long-term memory process, so that the model could help us to control the short-term variations in a better way.

5.3 Results for men

As in the case for women, we present results for 10000 simulations of the mortality rates for ages: $0, 5, 25, 50, 60, 70, 80, 90$. We graph the historical rate mortality, the mean of all simulations and the 95.5% confidence interval. See figures **??** and **??**.

As in the case for women, the proposed model for men is well fitted. We observe an increase in the rates mortality at ages between 25 to 35, this caused a overestimation in the first 35 years and latter a underestimation of the mortality rates. As was mentioned before, if we include in the model a short-term process, we believe the model could be better fitted. The main example of a short-term process to try is a $\mathrm{AR}(p)$ with $p \leq 2$ or 3.

For older ages, we observed that the data is irregular, so it is necessary to use a more complex model to fit this data.

6 Forecast

When we use our model to forecast and compare with the real data between the years 2005 to 2014. In general, from our results we observe that the forecast for women are good for almost all ages we tested. For men the variability of the results is strong and the behavior of the forecast is in general not good as those for women; for instance for ages smaller than 10 the results are quite similar than for women, even for ages between 10 and 45 it is possible to consider the results just good. However, for ages greater than 50, the results are not good, in fact, for older ages the results are bad: the model overstimate the mortality rates.

We present the results in the figures **??** for women at ages $0, 25, 50, 90$. For men we present more ages to ilustrate that the results are bad as the ages increases. see figures **??** and **??** at ages $0, 25, 50, 90$.

7 Conclusions

We have applied our proposed model to the Italian mortality rates with a geometric-type fractional Ornstein-Uhlenbeck process. Our main hypothesis was that, for a fixed age, the mortality rates changes through the time slowly, so that a stochastic differential equations that captures the long-range dependence could be a good model. We have used a stochastic differential equation with a fractional Brownian motion as a driven noise with $H \in (0.5, 1)$ in order to satisfy the long-range dependence property. With the data we have fixed the Hurst coefficient and we have confirmed our hypothesis since we have found that the estimated Hurst, for all ages, is in $(0.58, 0.8)$.

Notice that we have consider a more general model that the one used in Giacometti et al. (2011). This is because we have included the possibility that the Hurst parameter could be equal to 1/2, which is the case when the fractional Brownian motion becomes a standard Brownian motion. Therefore, when $H = 1/2$ we recover the Giacometti, Ortobelli and Bertocchi model.

The model is, specially for women, well behaved. For men at some ages we found some shortcomings that suggest the use of more terms in order to improve the model. From this results, and in opposition of the European normative on insurance that does not discriminate by gender, we conclude from our model, applied to the Italian case, that modelling the mortality by gender could improve the risk management of the insurance companies.

The long-range dependence model proposed in this paper is good enough to reproduce the mortality rates. If we add some extra terms to make it more flexible to reproduce the cases where the mortality rates have more variations then it will generate a more accurate model. We are starting to work on this extension of the model. Moreover, a multiplicative noise model will be the subject of a future research.

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