

Artificial data through calibration and empirical copulas

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Abstract: The aim of this study is to investigate artificial data generation possibilities offered by calibration estimators and empirical copulas, for continuous or mixed support variables with high skewness and kurtosis. After a brief discussion of the method, results of exercises on Survey Household Income data released by the Bank of Italy are shown and a global protection indicator is proposed.

1 Introduction

Simulation of multivariate economic data is a challenging task, from a methodological and computational point of view. Concerning the former, models lacking in appropriate constraints generally suffer of inadequate ability to fit features as high skewness and kurtosis. For the latter, moment based models (like maximum entropy ones) request a computational effort growing more than linearly with the number of variables. Difficulties due to the curse of dimensionality problem are obvious, but also the discretization of a wide univariate support can be troublesome: the simplification of circumventing the resort to integration is paid with the necessity of a fine grid resolution. Moreover for variables having probability masses concentrated to discrete points of the support many computational resources could be saved focusing only on support values relevant to efficiently describe examined phenomena. Disregarding “true” Generating Data process (GDP) oriented investigations, a proposal to cope with these problems retaining main observed data features is presented in the section two; simulation studies performed on Survey Household Income data released by the Bank of Italy are described in section three and a global protection measure is proposed; section four contains a summary and outlook on future research.

2 Artificial data generation

The merit of the simulation method would be the possibility of avoiding parametric assumptions on the GDP in a multivariate framework, maintaining some dependence relationships between variables and (eventually for time series or panel data) between observations. Subsections 2.1 and 2.2 discuss this matter.

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2.1 Empirical likelihoods and copulas

The intention of simulating multivariate economic data efficiently, from a computational and statistical point of view, can be pursued using calibration weights and empirical copulas. The first are instrumental to the reduction of the univariate discretized support dimension for each density to approximate, while the second is necessary to recovery a joint probability law from the univariate ones. An interesting set of calibration weights is constituted by those which maximize the empirical likelihood (Owen, 2001). The matter is treated in several works (i.e. C. Wu 2004). Although such estimates are first order equivalent to those obtained by linear regression (that is minimizing euclidean squared distances between original and constrained weights), avoiding their explicit dependence from the cross-product data matrix positively affects the robustness against influential observations. Given constraints which can be put in the form of orthogonality conditions (let $\boldsymbol{\psi}_i$ be the row-vector of conditions for the i^{th} observation, $i=1, \dots, n$), weights w_i are achieved minimizing the Kullback-Leibler divergence w.r.t those assigned to each statistical unit in the sample. Within the simple random sampling scheme, the Empirical Maximum Likelihood \boldsymbol{w} minimizes $D_{KL} = \sum (1/n) \ln[(1/n)/w_i]$. The Lagrangean is

$$\mathcal{L} = -\sum \frac{1}{n} \ln(nw_i) - \gamma(\sum w_i - 1) - \sum w_i \boldsymbol{\lambda}' \boldsymbol{\psi}_i$$

first order conditions equations are:

$$\partial \mathcal{L} / \partial w_i \equiv -\frac{1}{nw_i} - \gamma - \boldsymbol{\lambda}' \boldsymbol{\psi}_i = 0; \quad \partial \mathcal{L} / \partial \gamma \equiv (1 - \sum w_i) = 0; \quad \partial \mathcal{L} / \partial \boldsymbol{\lambda} \equiv \sum w_i \boldsymbol{\psi}_i' = 0$$

From the first, it results:

$$w_i = \frac{1}{n - \gamma - \boldsymbol{\lambda}' \boldsymbol{\psi}_i}$$

Since $\sum [-1/nw_i - \gamma - \boldsymbol{\lambda}' \boldsymbol{\psi}_i] w_i = 0$, due to constraints, $-\gamma=1$ follows. As we see later, in the proposed framework EML weights are used to calibrate a subset of the original sample to fit several moment conditions. The study of copula functions has received growing attention in recent years and a wide review is presented in Nelsen (2006). The Sklar's theorem (Rueschendorf, 2009), establishes that given an m -dimensional distribution function F , with marginals F_1, F_2, \dots, F_m , there exists an m -dimensional distribution function C (the copula function) on $[0,1]^m$ with uniform margins such that

$$F(x_1, x_2, \dots, x_m) = C[F_1(x_1), \dots, F_m(x_m)]$$

For a continuous random variable X , $F(X) \equiv p(X < x)$ and posing $U \equiv U(0,1)$ the distributional transform can be defined as $X = F^{-1}(U)$; then

$$p[\cap(X_j \leq x_j)] = p\{\cap[F_j^{-1}(U_j) \leq x_j]\} = p\{\cap[U_j \leq F_j(x_j)]\} \equiv C[F_1(x_1), \dots, F_m(x_m)]$$

If marginals F_j are not known, their sample counterparts, empirical *CDFs*, can be used and the inverse distributional transform delivers vectors proportional to observed ranks:

$$n\hat{U}_j = n\hat{F}_j(X_j) = R_j \quad (j = 1, \dots, m)$$

Hence, the Empirical Copula Function is

$$\hat{C}(\mathbf{t}) = \frac{1}{n} \sum I[\cap(\hat{U}_j \leq t_j)] = \frac{1}{n} \sum I[\cap(R_j \leq nt_j)] \quad \mathbf{t} \in [0, 1]^m$$

This circumstance clarifies the key role of the observed ranks matrix in conveying dependence relationships featuring observed data. If $\{x_{(1)j}, x_{(2)j}, \dots, x_{(n)j}\}$ and $\{r_{1j}, r_{2j}, \dots, r_{nj}\}$ are respectively the order statistics and the ranks for the variable j^{th} , dependence structure preserving is (approximately) accomplished posing

$$x_{ij} \equiv x_{(r_{ij})j} \quad (i = 1, \dots, n \quad j = 1, \dots, m)$$

Hence, no rank swapping is performed: ECF does not play any role of data perturbation.

2.2 A synthesis of the procedure

A sequence of steps summarizes the proposed data generation method:

- a) for each variable, by means of $k-1$ empirical percentiles, k strata of the observed support are delimited;
- b) the observed range is enlarged by values obtained adding to the original maximum (minimum) a positive (negative) term related to the data sparsity; that step is needed to ensure the existence of a solution for the constrained optimization problem used to achieve calibration weights;
- c) for each stratum, h values are drawn from a uniform law; hence a discretized version of each variable support is defined. kh is the number of points of the discretized support; that step allows, together with step (a), to control the trade-off between accuracy w.r.t. distributional features and data protection; calibration is needed to make the distribution a satisfactory approximation of the density which generates data;
- d) moments up to order p are used to calibrate weights for the artificial discrete support of length kh generated in (c); so doing the information loss due to the support compression of steps (a) and (c) is partially accommodated; it is necessary to stress that calibration weights are only used to draw samples from the compressed support satisfying moment constraints;

- e) according to calibration weights, samples of length n (with repetition if $n > kh$) are drawn;
- f) each sample from (e) is ordered according to observed ranks; that step is performed one variable at a time, maintains univariate dependence relationships between statistical units and implicitly recovers multivariate links between variables included in the data set; no rank swapping is performed;

This procedure is somehow similar to the maximum entropy bootstrap (Vinod, 2004), implicitly embedding a clustering task, points (a) and (c), and explicitly imposing moment constraints, point (d). Moment based estimations of probability laws are discussed in several papers (i.e. X. Wu 2003). In our framework, aims other than observed features replication are not considered, omitting any inferential purpose about “true” probability laws. At least first four moments are included, introducing further constraints (that is higher moment conditions) until the null hypothesis of equal distributions is not refused.

3 Simulated data for the survey of household income

3.1 Original data and simulation design

Data gathered by the 2006 Survey on Household Income and Wealth, freely available by the Bank of Italy Economic and Financial Statistics Department (see <http://www.bancaditalia.it/statistiche/indcamp/bilfait/dismicro>), seem interesting to assess the fitting accuracy for extremely sparse variables. The absence of socio-economic interpretative intentions leads to consider a subset of variables, resulting by aggregations based on income origins; grossing up weights are ignored. Data, in euro currency units, contains 7768 records and five variables: Y_{cf} (income from financial assets), Y_l (payroll income), Y_t (pensions and net transfers), Y_m (net self-employment income), Y_{ca} (income from real-estate). Extreme degrees of data compression are considered for investigating the trade-off between accuracy and protection. Relevant parameters are the number of support slices and sample fractions. Couples of settings (3 strata with 20% of sampled units and 100 strata with 50% of sampled units) are separately considered in 2500 replications, repeating steps (e)-(f) 50 times for each outcome of steps (a)-(d).

	Y_{cf}	Y_l	Y_t	Y_m	Y_{ca}
<i>Min</i>	-25432.36	0.00	0.00	-20000.00	0.00
<i>p25</i>	3.42	0.00	0.00	0.00	2400.00
<i>p50</i>	67.25	6105.00	6518.00	0.00	6000.00
<i>p75</i>	298.20	20000.00	14690.00	0.00	8400.00
<i>Max</i>	99789.76	251000.00	429770.00	800000.00	152000.00

Table 3.1 Some order statistics for survey data.

	<i>Ycf</i>	<i>Yl</i>	<i>Yt</i>	<i>Ym</i>	<i>Yca</i>
<i>Mean</i>	280.54	12054.03	8781.26	4327.05	6564.37
<i>Std. Dev.</i>	2742.24	15355.02	11194.32	19483.57	7457.06
<i>Skewness</i>	11.44	2.11	7.88	18.21	5.66
<i>Kurtosis</i>	322.74	17.07	261.62	549.23	72.47

Table 3.2 First four cumulants for survey data

	<i>Ycf</i>	<i>Yl</i>	<i>Yt</i>	<i>Ym</i>	<i>Yca</i>
<i>Ycf</i>	1.00	-0.02	0.15	0.12	0.25
<i>Yl</i>	-0.02	1.00	-0.37	-0.07	0.11
<i>Yt</i>	0.15	-0.37	1.00	-0.09	0.16
<i>Ym</i>	0.12	-0.07	-0.09	1.00	0.24
<i>Yca</i>	0.25	0.11	0.16	0.24	1.00

Table 3.3 Correlation matrix for survey data

3.1.1 Accuracy statistics

Focusing on the univariate field, survey data peculiarities appear to be satisfactory taken by statistical units simulated with the second setting (100 strata, 50% of sampled units). Since the calibration is only performed on moment conditions because of the simple implementation of constraints, as a by-product of the cumulants reproduction ability, similarity between observed and simulated position indices is achieved. By comparison of tables 3.1 and 3.4, the first setting delivers less accurate results than the second and simulated ranges are generally reduced; for *Ym* in the first setting a strong shrinking of the maximum can be noticed; the overshooting for the minimum of *Ym* is instrumental to the approximation (from lower values) of skewness and kurtosis; for *Yl*, the median and 3rd quartile are oppositely distorted. In the second setting maximum of *Ym* is better reproduced and the overshooting for the minimum is softened. In the second setting position indeces are generally closer to those observed; confidence intervals (*CI*s) referred to *Yca* quartiles do not comprise those observed and the same happens, with lower intensity, for the minimum of *Ym*; the *CI* for the 3rd quartile of *Yl* weakly exceeds the sample estimation as well values simulated for 2nd and 3rd quartile of *Ycf*. *Ym* is the extreme paradigm of the protection-accuracy conflict: for a variable whose tails contain all relevant information for describing the distribution, all intensities are fully involved by the distortion due to the protection. A better behaviour can be noticed about cumulants (tables 3.2 and 3.5). In the first setting, *Ym* skewness and kurtosis are downwardly approximated; adjusting skewness and kurtosis leads to underestimate the mean value of *Yt*; *Ycf* markedly underestimates 3rd and 4th cumulants. In the second setting simulated quantities are closer to those observed; a weak underestimation for the *Yca* mean value is noticed but *CI* are shorter. Better results could be achieved in the first setting tuning the number of moment constraints for the variables whose density approximation should be improved; such exercise has not been performed the focus

being on methods suited for to a large number of variables, as this is more relevant for statistical institutes.

		3 strata, 20% sampled units					100 strata, 50% sampled units				
		<i>Min</i>	<i>p25</i>	<i>p50</i>	<i>p75</i>	<i>Max</i>	<i>Min</i>	<i>p25</i>	<i>p50</i>	<i>p75</i>	<i>Max</i>
<i>Ycf</i>	<i>p2.5</i>	-20238.3	0.0	48.1	108.4	21686.5	-29517.5	0.0	69.2	313.5	43879.0
	<i>Mean</i>	-16460.8	0.0	50.9	109.7	25547.8	-25814.2	2.4	74.2	333.5	89611.9
	<i>p97.5</i>	-13423.6	0.0	53.6	111.8	32111.3	-19539.3	6.1	77.9	360.0	112583.3
<i>Yl</i>	<i>p2.5</i>	0.0	0.0	9092.2	14245.9	79349.2	0.0	0.0	4514.6	20425.8	119573.4
	<i>Mean</i>	0.0	0.0	9479.2	14463.7	84551.9	0.0	0.0	7082.0	21146.1	219428.3
	<i>p97.5</i>	0.0	0.0	9608.0	14722.1	97549.2	0.0	0.0	8305.3	21802.5	278502.7
<i>Yt</i>	<i>p2.5</i>	0.0	0.0	3568.8	9419.1	89644.0	0.0	0.0	6503.1	14633.9	107957.3
	<i>Mean</i>	0.0	0.0	3674.6	9431.8	136548.4	0.0	0.0	6653.3	15130.7	314007.7
	<i>p97.5</i>	0.0	0.0	3979.0	9528.1	643438.6	0.0	0.0	7082.6	15526.5	503211.3
<i>Ym</i>	<i>p2.5</i>	-29171.0	0.0	0.0	0.0	394210.3	-23823.2	0.0	0.0	0.0	388214.7
	<i>Mean</i>	-29170.5	0.0	0.0	0.0	591466.1	-23823.2	0.0	0.0	0.0	679318.0
	<i>p97.5</i>	-29171.0	0.0	0.0	0.0	622144.9	-23823.2	0.0	0.0	0.0	923022.9
<i>Yca</i>	<i>p2.5</i>	308.3	2389.1	5250.4	6538.0	79316.5	0.0	0.0	5387.2	8580.8	115848.5
	<i>Mean</i>	629.4	2684.9	5272.3	6551.8	142172.2	0.0	0.0	5596.2	8756.4	146122.7
	<i>p97.5</i>	777.8	2901.2	5277.7	6601.6	160687.3	0.0	0.0	5865.9	9122.0	162847.4

Table 3.4 Order statistics summary

		3 strata, 20% sampled units				100 strata, 50% sampled units			
		<i>Mean</i>	<i>Std. Dev.</i>	<i>Skewness</i>	<i>Kurtosis</i>	<i>Mean</i>	<i>Std. Dev.</i>	<i>Skewness</i>	<i>Kurtosis</i>
<i>Ycf</i>	<i>p2.5</i>	112.7	2409.5	1.8	16.6	252.2	2041.7	5.8	126.7
	<i>Mean</i>	167.3	2537.5	2.3	19.4	309.7	2614.1	13.3	389.9
	<i>p97.5</i>	223.3	2663.5	2.7	22.8	369.5	3265.0	19.9	691.7
<i>Yl</i>	<i>p2.5</i>	11303.2	14757.8	1.4	1.1	11836.9	14641.9	1.5	4.3
	<i>Mean</i>	11562.4	15070.7	1.4	1.4	12150.7	15269.1	2.1	13.8
	<i>p97.5</i>	11860.2	15424.8	1.5	1.7	12479.2	16006.4	3.1	30.6
<i>Yt</i>	<i>p2.5</i>	7177.2	11097.6	2.4	6.3	8574.9	9830.9	1.6	8.3
	<i>Mean</i>	7431.8	11635.8	3.3	45.8	8813.1	11078.5	6.6	187.8
	<i>p97.5</i>	7694.0	13634.5	14.9	633.6	9065.2	13349.1	15.4	531.6
<i>Ym</i>	<i>p2.5</i>	3882.5	14953.7	6.7	96.3	4190.8	14490.9	11.4	237.4
	<i>Mean</i>	4287.1	18892.0	13.7	327.7	4583.6	18848.6	17.1	480.4
	<i>p97.5</i>	4729.9	23116.8	17.9	498.6	5018.3	24351.6	23.4	909.1
<i>Yca</i>	<i>p2.5</i>	6097.0	6806.6	2.8	9.3	5855.3	7147.9	3.5	30.9
	<i>Mean</i>	6372.9	7377.6	4.5	47.6	6040.8	7805.7	5.1	59.1
	<i>p97.5</i>	6585.1	8011.4	6.1	79.5	6215.4	8551.7	6.5	86.3

Table 3.5 Cumulants summary

Comparing tables 3.3 and 3.6 pairwise correlations are satisfactory close, an important circumstance for a multivariate validation of the simulation framework. In both simulation settings, correlation signs are always maintained; in the first one, magnitudes are closer to those measured on data, but correlation often exceed the

simulated *CI* boundaries. However, in both settings *CI* lengths look short and, with the exception of that related to (Yl, Yt) in the second setting (length of 0.11), all others are not greater than 0.06.

	3 strata, 20% units sampled			100 strata, 50% units sampled		
	<i>p</i> 2.5	<i>Mean</i>	<i>p</i> 97.5	<i>p</i> 2.5	<i>Mean</i>	<i>p</i> 97.5
<i>Ycf, Yl</i>	-0.06	-0.05	-0.04	-0.03	-0.01	0.00
<i>Ycf, Yt</i>	0.18	0.21	0.22	0.11	0.14	0.17
<i>Ycf, Ym</i>	0.07	0.10	0.12	0.09	0.12	0.15
<i>Ycf, Yca</i>	0.19	0.21	0.23	0.18	0.21	0.24
<i>Yl, Yt</i>	-0.31	-0.30	-0.26	-0.42	-0.38	-0.31
<i>Yl, Ym</i>	-0.09	-0.08	-0.07	-0.09	-0.07	-0.06
<i>Yl, Yca</i>	0.11	0.12	0.13	0.11	0.12	0.13
<i>Yt, Ym</i>	-0.09	-0.07	-0.06	-0.13	-0.10	-0.07
<i>Yt, Yca</i>	0.19	0.21	0.22	0.14	0.15	0.17
<i>Ym, Yca</i>	0.19	0.22	0.24	0.20	0.23	0.26

Table 3.6 Correlation coefficients summary

Hints about the retention ability for more complex dependence relationships can be gathered from hypothesis tests as:

$$H_0 : \frac{\hat{m}_{a,b,c,d,e}}{m_{a,b,c,d,e}} - 1 = 0, \quad m_{a,b,c,d,e} \equiv \left(\frac{1}{n} \sum Ycf_i^a \cdot Yl_i^b \cdot Yt_i^c \cdot Ym_i^d \cdot Yca_i^e \right)^{\frac{1}{a+b+c+d+e}}$$

In details, 500 mixed moments have been selected raising each variable to a power given by a random integer in $[0, 4]$. The hypothesis test is intended to assess if simulated mixed moments are significantly different from original ones; the test is based on *CI*s estimated by means of 2,500 replications of data simulation. The Mean Absolute Percentage Error (MAPE) is calculated for each mixed moment.

Order	3	4	5	6	7	8	9	10	11	12	13	14	15	16	17	19	
Frequency	3	9	16	34	47	48	58	59	61	61	40	31	18	8	6	1	
3 strata, 20% s.u.	<i>H₀</i> not false	0.33	0.56	0.56	0.50	0.57	0.58	0.59	0.64	0.69	0.57	0.55	0.45	0.67	0.38	0.67	0.00
	<i>MAPE</i>	0.10	0.20	0.19	0.15	0.17	0.15	0.14	0.13	0.10	0.11	0.07	0.07	0.05	0.04	0.04	0.09
100 strata, 50% s.u.	<i>H₀</i> not false	1.00	1.00	1.00	1.00	1.00	1.00	1.00	0.98	0.98	0.98	1.00	1.00	1.00	1.00	1.00	1.00
	<i>MAPE</i>	0.04	0.10	0.10	0.12	0.12	0.13	0.14	0.13	0.11	0.10	0.10	0.08	0.09	0.07	0.08	0.08

Table 3.7 Proportions of null not refused and MAPEs w.r.t. moment orders

Table 3.7 shows percentage of instances in which the null is not refused at the level $\alpha=0.05$ and MAPEs, both grouped w.r.t. moment orders. Frequencies about the hypothesis test related to a certain order are calculated as ratios between the number of instances in which 0 falls within *CI* boundaries and the number of mixed moments of the given order. Analogously, MAPEs referred to a certain order are averaged. In

the first setting, the dispersion around test statistics is only partially accounted for, while averaged MAPEs are greater than those calculated in the second setting up to the order 8th. Beyond the order 8, lower support compressions do not allow better point estimations: in both settings supports discretization leads to a partially unsatisfactory description of mixed moment domains. Hence, simulated data can preserve linear relationships between couples of variables as well fairly high mixed moments (in the examined exercise at least those up to the 8th order). Figure 3.1 depicts simulated *CIs* for a sample of mixed moments randomly drawn from those tested in the second setting (in x-axis exponents are shown, respectively referred to *Ycf, Yl, Yt, Ym, Yca*).

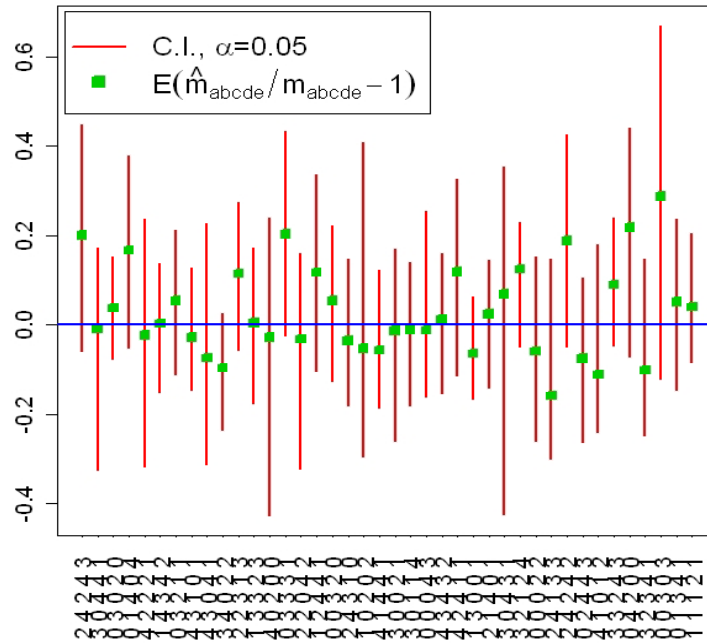


Fig 3.1 Some mixed moment tests (100 strata, 50% sampled units)

Concluding accuracy analyses, a test which provides a general measure of equality for multivariate distributions is suitable; that proposed by Székely and Rizzo[†] (2004) has been used for each simulated sample. Under the null, a random permutation of pooled original and simulated observations is equal in distribution to a random sample drawn from the mixture of sources. Due to data size, in each simulation, 100 tests on pooled random subsamples (100 observations from each source) have been performed; overall averaged p-values are 0.328 for the first setting and 0.999 for the second.

[†] For two samples \mathbf{x} and \mathbf{y} having respectively size n_1 and n_2 the test statistic is:

$$\mathcal{E}_{n_1, n_2} \equiv \left(\frac{n_1 n_2}{n_1 + n_2} \right) n_1^{-2} n_2^{-2} \sum_i \sum_j \sum_l \sum_h \|x_i - y_l\| + \|x_j - y_m\| - \|x_i - x_j\| - \|y_l - y_m\|$$

3.1.2 A global protection indicator for simulated data

A proposal for quantifying the global protection implied by simulated data can be scheduled as follows:

- a) the original dataset (with observations x_{ij} , $i=1, \dots, n$ $j=1, \dots, m$) is overfitted choosing the highest number of clusters between those admissible according to a selection criterion; let L be the set of clusters with size 1 having elements l_s ,
- b) each simulated record (y_{ij} , $i=1, \dots, n$ $j=1, \dots, m$) is assigned to the closest centroid identified in (a); let H be the set of non empty clusters having centroids in L and H_{l_s} the cluster around l_s ,
- c) for each $l_s \in L$, if l_s is the s^{th} element of L , $D_g \equiv \max_j |x_{sj} - \text{median}(x_j)|$ is retained (that is, the worst dimension in term of identifiable units is considered),
- d) for each $l_s \in L$, if y_u is the u^{th} record of H_{l_s} , $d_s \equiv \min_u |x_{sg} - y_{ug}| / D_g$ is calculated and a threshold $q \in [0, 1]$ is defined,
- e) $ns \equiv |H|/|L|$ expresses the proportion of simulated records potentially not secured; $ans \equiv ns \cdot E_s[I(d_s < q)]$ adjusts ns for the proportion of records subjectively secured because of the proximity of simulated and median values for the less favourable variable detected in (c),
- f) the global protection indicator is defined as $gp \equiv 1 - ans$.

About points (c) and (d), alternative ways for associating a unique measure to each record could be obtained choosing some synthesis of gaps; standardizations different from the proposed one can be applied but it is to notice that the interquartile range is not always useful to this aim due to its equality to zero for some variables; other possibilities are related to variable standardizations or rotations of Cartesian axes. For simplicity, the first step has been implemented resorting to the *k-means* algorithm, choosing the highest number of clusters, 3981, which locally maximizes the Calinski-Harabasz *pseudo-F* (Sugar and James, 2003) and minimizes the averaged *within* sum of squares; through overfitting, lack of detection for simulated units far from the region where the multivariate density is predominantly allocated should be avoided. From table 3.8, if $q = 0.1$, roughly 40% of simulated records belonging to H are considered critical in the first setting and the proposed protection measure can be calculated as $1 - 0.4 \cdot 0.1185 = 0.95$; in the second setting, q is less than 0.1 for the 90% of instances and gp is $1 - 0.9 \cdot 0.293 = 0.74$. Such behaviours are outcomes of (different) simulation abilities to mimic distribution details induced by data generation settings.

3 strata, 20% s.u.	ns	q	0.00	0.03	0.08	0.14	0.21	0.28	0.35	0.49	0.60
	0.1185	$E_s[I(d_s < q)]$	0.10	0.20	0.30	0.40	0.50	0.60	0.70	0.80	0.90
100 strata, 50% s.u.	ns	q	0.00	0.01	0.01	0.02	0.03	0.04	0.06	0.08	0.11
	0.2930	$E_s[I(d_s < q)]$	0.10	0.20	0.30	0.40	0.50	0.60	0.70	0.80	0.90

Table 3.8 Elements for estimating the global protection indicator

4 Some considerations

Within a sampling setting, a data simulation method has been proposed. The trade-off between accuracy and protection against disclosure is controlled through compression of univariate supports. Calibration weights obtained by means of moment conditions partially accommodate the information loss, so that samples drawn from artificial supports approximately retain univariate statistical features. Multivariate dependence relationships are maintained ordering simulated samples according to the matrix of ranks. The irrelevance of explicit conditional independence assumptions regarding subsets of variables provides robustness against misspecifications. Moreover, compression of supports and empirical copula expression of multivariate dependence jointly reduce the computational burden: while the former decreases the number of support points, the latter ensures a growth only linear w.r.t. the number of variables. Exercises conducted on Survey Household Income data of year 2006, collected from the Bank of Italy, have shown encouraging results. A global protection indicator has been proposed unifying two aspects: detection of statistical units remote to the multivariate density core and evaluation of the exchangeability between simulated ones. Extreme settings for data generation let understand the existence of tuning opportunities for accuracy and protection preferences. It remains to assess the behaviour of non linear combinations of simulated data when records are grouped according to known homogeneity features: further capability studies seem necessary.

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References

- Nelsen, R.B. (2006). *An Introduction to Copula*. Springer-Verlag.
- Owen, A.B. (2001). *Empirical Likelihood*. Chapman & Hall.
- Rueschendorf, L. (2009). On the distributional transform, Sklar's theorem, and the empirical copula process. *Journal of Statistical Planning and Inference*. 139(11). 3921-3927.
- Sugar, C., James, G. (2003) Finding the Number of Clusters in a Data Set: An Information Theoretic Approach. *Journal of the American Statistical Association*. 98. 750-763.
- Székely, G. J., Rizzo, M. L. (2004). Testing for Equal Distributions in High Dimension. *InterStat*. Nov. (5).
- Vinod, H. (2004). Ranking mutual funds using unconventional utility theory and stochastic dominance. *Journal of Empirical Finance*. 11(3). 353-377.
- Wu, C. (2004). Weighted Empirical Likelihood Inference. *Statistics and Probability Letters*. 66(1). 67-79.
- Wu, X. (2003). Calculation of maximum entropy densities with application to income distribution. *Journal of Econometrics*. 115(2). 347-354.