Università degli Studi di Pisa
Dipartimento di Statistica e Matematica
Applicata all'Economia

Report n. 279

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Pisa, Gennaio 2006
- Stampato in Proprio –

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Spatial M-quantile Models for Small Area Estimation

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Riassunto: In questo lavoro si inserisce l'informazione spaziale nell'approccio basato sulla regressione M-quantilica per la stima per piccole aree, proponendo un'alternativa semiparametrica allo stimatore di tipo SEBLUP (Spatial Empirical Best Linear Unbiased Predictor). Il metodo è applicato alla stima della produzione media e mediana di olive per azienda in Toscana con dati provenienti dall'indagine campionaria sulla Struttura e Produzione delle aziende agricole (2003).

Keywords: Small area estimation, Spatial EBLUP, M-quantile regression.

1. Introduction

In the context of Italian agricultural surveys the term "small area" generally refers to a local geographical area, such as province or municipality. Predicting the average or the total crop area for municipalities is of interest. However, many difficulties are paused due to the lack of available survey data. Thus, previous work has focused on estimating at higher geographical levels such as Italian provinces (Benedetti *et al.*, 2004).
When the traditional municipality-specific direct estimator does not provide adequate precision, it is possible to employ indirect estimators that "borrow strength" from related areas. These indirect estimators incorporate area-specific random effects that account for between area variation beyond what is explained by auxiliary variables included in the model. Although it is custom to assume that the random area effects are independent, in most applications with environmental data it is reasonable to assume that the random area effects of neighbouring areas (for instance the neighbourhood could be defined by a contiguity criterion) are correlated with the correlation decaying to zero as the distance increases (Pratesi and Salvati, 2005). However, this type of modelling also depends on strong distributional assumptions, requires a formal specification of the random part of the model and does not easily allow for outlier robust inference.
A new approach to small area estimation that is based on modelling quantile-like parameters of the conditional distribution of the variable of interest y given the covariates $\mathbf{x}$ has been recently proposed by Chambers and Tzavidis (2006). Unlike

traditional random effects models, this technique does not depend on strong distributional assumption, and it is robust against outlying values. Moreover, this approach overcomes the important problem of small area estimation with changing small area geographies. The aim of this work is to extend the M-quantile approach to account for spatial correlation between small areas. The application will be on the estimation of the average and median production of olive per farm in each Local Economy System (LES) in Tuscany. The LES are defined as aggregations of municipalities but they are different from provinces.
The paper is organized as follows: In section 2 M-quantile models for small area estimation are reviewed and an extension of the M-quantile approach to account for spatial correlation between the small areas is proposed. In Section 3 the performance of the proposed M-quantile estimator is compared to the performance of alternative, traditional small area estimators obtained via random effects models using Monte Carlo simulation. In Section 4 we discuss the results of an application of the newly proposed method for estimating the average and median production of olives per farm at LES level.

2. M-quantile models with spatial information

Chambers and Tzavidis (2006) proposed a new approach to small area estimation that is based on modelling quantile-like parameters of the conditional distribution of the study variable (y) given the covariates (Breckling and Chambers, 1988). Mixed effects models assume that variability associated with the conditional distribution of y given $\mathbf{x}$ can be at least partially explained by a pre-specified hierarchical structure, such as the small areas of interest. An alternative approach to modelling the variability in this conditional distribution is via M-quantile regression, which does not depend on a hierarchical structure. Instead, we characterise the conditional variability across the population of interest by the M-quantile coefficients of the population units.
The first target in M-quantile models for small estimation is to estimate the M-quantile coefficients, $\{q_{js}; j \in s\}$, of the units in the sample without reference to the small areas of interest. To do this we define a fine grid on the (0,1) interval and, using the sample data, we fit the M-quantile regression line at each value q on this grid. Let us define by ψ an influence function and by $\mathbf{x}$ the matrix of covariates $n \times p$, with $n = \sum_{i=1}^{m} n_i$ and m the number of small areas. Estimation of the regression parameters of the linear M-quantile regression model

$$Q_q = (\mathbf{x}, \psi) = \mathbf{x}^T \boldsymbol{\beta}_\psi(q) \tag{1}$$

is achieved by solving the estimating equations below using iteratively reweighted least squares (IRLS)

$$\sum_{j=1}^{n} \psi_q(r_{jq\psi}) \mathbf{x}_j = 0 \tag{2}$$

where $r_{jq\psi} = y_j - \mathbf{x}_j^T \boldsymbol{\beta}_\psi(q)$, $\psi_q(r_{jq\psi}) = 2\psi(s^{-1} r_{jq\psi})\{qI(r_{jq\psi} > 0) + (1-q)I(r_{jq\psi} \le 0)\}$ and s is a robust measure of spread. Chambers and Tzavidis (2006) assume that $s = MAD/0.675$, where MAD is the Median Absolute deviation, and ψ is a Huber proposal 2 influence function.

If a data point lies exactly only a fitted M-quantile regression line, then the M-quantile coefficient of the corresponding sample unit is equal to q. Otherwise, if a data point lies between two fitted M-quantile regression lines, the M-quantile coefficient of the corresponding sample unit is derived by linear interpolation.
If a hierarchical structure does explain part of the variability in the population data, then we expect units within clusters defined by this hierarchy to have similar M-quantile coefficients. Therefore, we can compute the average value of the sample M-quantile coefficients of units j in area i, $\hat{\bar{q}}_i = \sum_{j \in i} q_{js}$. A small area estimator of the mean, $\hat{\bar{y}}_i$, is then derived by fitting an M-quantile regression model at $\hat{\bar{q}}_i$

$$\hat{\bar{y}}_i = \frac{1}{N_i}\left(\sum_{j \in s_i} y_j + \sum_{j \in r_i} \mathbf{x}_i^T \hat{\boldsymbol{\beta}}_\psi(\hat{\bar{q}}_i) \right) \tag{3}$$

where s_i and r_i respectively denote the sampled and non sampled units in area i and N_i is the number of population units in area i. The unobserved value y_j for population unit $j \in r_i$ is predicted using $\mathbf{x}_i^T \hat{\boldsymbol{\beta}}_\psi(\hat{\bar{q}}_i)$.
In this paper we propose an extension to the M-quantile model of Chambers and Tzavidis (2006) to account for spatial correlation between the small areas. This approach defines an alternative, to the traditional Spatial EBLUP estimator (Pratesi and Salvati, 2005a; Petrucci and Salvati, 2005; Saei and Chambers, 2003), small area estimator.
To account for spatial correlation in the M-quantile approach, we propose modelling the M-quantile coefficients, derived from the first step of the Chambers and Tzavidis (2006) approach using a model of the following form

$$log\left(\frac{\mathbf{q}_s}{1 - \mathbf{q}_s} \right) = \mathbf{x}\boldsymbol{\beta} + \mathbf{Z}(\mathbf{I} - \rho\mathbf{W})^{-1}\mathbf{u} + \boldsymbol{\varepsilon} \tag{4}$$

where $\mathbf{x}$ be the matrix of covariates $n \times p$, $\mathbf{Z}$ is the incidence matrix $n \times m$ for the random effects vector, the deviations from the fixed part of the model $\mathbf{x}\boldsymbol{\beta}$ are the result of an autoregressive process with parameter ρ (spatial autoregressive coefficient) and proximity matrix $\mathbf{W}$ $m \times m$ and $\boldsymbol{\varepsilon}$ the error vectors. $\mathbf{I}$ is an identity matrix $m \times m$. Under the model, the Spatial EBLUP estimator of $\hat{\bar{q}}_i$ is:

$$\hat{\bar{q}}_i(\hat{\sigma}_u^2, \hat{\sigma}_\varepsilon^2, \hat{\rho}) = \frac{exp(p_i)}{1 + exp(p_i)} \tag{5}$$

where $p_i = \bar{\mathbf{x}}_i\hat{\boldsymbol{\beta}} + \mathbf{b}_i^T \hat{\sigma}_u^2 \left[(\mathbf{I} - \hat{\rho}\mathbf{W})(\mathbf{I} - \hat{\rho}\mathbf{W}^T) \right]^{-1} \mathbf{Z}^T \times \left\{ \hat{\sigma}_\varepsilon^2 \mathbf{I}_n + \mathbf{Z}\hat{\sigma}_u^2 \left[(\mathbf{I} - \hat{\rho}\mathbf{W})(\mathbf{I} - \hat{\rho}\mathbf{W}^T) \right]^{-1} \mathbf{Z}^T \right\}^{-1} (\mathbf{q}_s - \mathbf{x}\hat{\boldsymbol{\beta}})$, $\bar{\mathbf{x}}_i$ are the known population mean, $\mathbf{q}_s$ is the $n \times 1$ vector of q_{js}, $\hat{\sigma}_u^2, \hat{\sigma}_\varepsilon^2, \hat{\rho}$ are asymptotically consistent estimators of the parameters and $\mathbf{b}_i^T$ is a $1 \times m$ vector $(0,0 \ldots 0,1,\ldots 0)$ with value 1 in the *i-th* position.
Similarly to (3), an M-quantile small area estimator of the mean, that accounts for spatial correlation, is then defined by fitting an M-quantile model at $\hat{\bar{q}}_i(\hat{\sigma}_u^2, \hat{\sigma}_\varepsilon^2, \hat{\rho})$. The drawback of the previous approach is that although we use the M-quantile approach in

order to avoid specifying a parametric model, we still use parametric model (4) to account for spatial correlation in the M-quantile coefficients. Ideally, we would like to employ a non-parametric approach to account for spatial correlation in the M-quantile coefficients. However, this is beyond the scope of this paper.

Estimator (3) can be more generally defined by appropriately integrating the following distribution function

$$\hat{F}_i(t) = \frac{1}{N_i}\left(\sum_{j \in s_{n_i}} I(y_{ij} \le t) + \sum_{j \in r_i} I\left(\mathbf{x}_{ij}^T \hat{\boldsymbol{\beta}}_\psi(\hat{\bar{q}}_i) \le t\right)\right) \tag{6}$$

As Tzavidis and Chambers (2006) noticed, the M-quantile estimator (3) is biased particularly when small areas contain outliers. Tzavidis and Chambers (2006) proposed a bias-adjusted M-quantile estimator of the mean. Their proposal is based on an adaptation of the Chambers-Dunstan (denoted by a subscript CD) (1986) estimator of the distribution function defined by

$$\hat{F}_{CD,i}(t) = \frac{1}{N_i}\left(\sum_{j \in s_{n_i}} I(y_{ij} \le t) + \frac{1}{n_i}\sum_{j \in r_i}\sum_{k \in s_{n_i}} I\left\{\left[\mathbf{x}_{ij}^T \hat{\boldsymbol{\beta}}_\psi(\hat{\bar{q}}_i) + \left(y_{ik} - \mathbf{x}_{ik}^T \hat{\boldsymbol{\beta}}_\psi(\hat{\bar{q}}_i)\right)\right] \le t\right\}\right) \tag{7}$$

The bias-adjusted mean estimator of Tzavidis and Chambers (2006) for small area j is given by

$$\hat{\bar{y}}_i = \int t d\hat{F}_{CD,i}(t) = \frac{1}{N_i}\left(\sum_{j \in s_{n_i}} y_{ij} + \sum_{j \in r_i} \mathbf{x}_{ij}^T \hat{\boldsymbol{\beta}}_\psi(\hat{\bar{q}}_i) + \frac{N_i - n_i}{n_i}\sum_{j \in s_{n_i}}\left(y_{ij} - \mathbf{x}_{ij}^T \hat{\boldsymbol{\beta}}_\psi(\hat{\bar{q}}_i)\right)\right) \tag{8}$$

Other quantiles of the distribution function can be obtained by appropriately integrating the CD estimator of the distribution function.

An MSE estimator of (3) has been proposed by Chambers and Tzavidis (2005). The main limitation of this estimator is that it does not account for the variability introduced in estimating the area specific q's. Thus it may underestimate the true MSE. As an alternative approach Pratesi and Salvati (2005) propose a bootstrap estimator. In the economy of this work we focus our attention on the performance of M-quantile point estimator (8), obtained with and without spatial information, in comparison with the EBLUP and the Spatial EBLUP performances. In Section 3 we present the results of a simulation experiment where the MSE of the considered estimators is only empirically evaluated. In the application of the M-quantile estimator to real life data (Section 4) we limit to the calculation of mean median and distribution function without proposing an estimator of the variance of the estimators.

3 Simulation experiments

In order to assess the use of the spatial small area methodology, simulated experiments were carried out. The EBLUP estimator, the Spatial EBLUP estimator, the M-quantile method, the Spatial M-quantile estimator were compared. We generate a synthetic population of y_{ij} -values, for the unit j in the small area i using the spatial nested error regression model with random area effects of neighbouring areas correlated according to the SAR dispersion matrix with established spatial autoregressive coefficient.

$$y_{ij} = x_{ij}\boldsymbol{\beta} + v_i + e_{ij}x_{ij}^{1/2} \tag{9}$$

where x_{ij} is the value of auxiliary variable x, v_i is the random area specific effect and e_{ij} is the individual error.

The experiment is designed following Rao and Choudry (1995, Section 27.2.3). We put $\beta = 0.21$, $\sigma_u^2 = 100$ and $\sigma_e^2 = 1.34$, and used a fixed number of small areas $m = 42$. We generated independent random variables $\mathbf{v} = [v_1,...v_m]^T$ and $\mathbf{e} = [e_{11}, e_{12},...e_{ij},..., e_{mN_m}]^T$ from $MVN(0, \sigma_u^2[(\mathbf{I} - \rho\mathbf{W})(\mathbf{I} - \rho\mathbf{W}^T)]^{-1})$ and $N(0, \sigma_e^2)$ respectively while x_{ij} values were generated from a uniform distribution between 0 and 10. Values of the study variable, y_{ij}, are then obtained for each x_{ij} using

$$y_{ij} = 0.21x_{ij} + v_i + e_{ij}x_{ij}^{1/2}.$$

The SAR dispersion matrix was generated with ρ equal to $\pm 0.25, \pm 0.50, \pm 0.75$ and the neighbourhood structure (**W**) is obtained by randomly assigning neighbours for each area as follows: The value 1 is assigned to the spatial weight w_{ij} if the value drawn from a uniform distribution $[0,1]$ is greater than 0.5, 0 otherwise. The maximum number of neighbours for each area was 5, and the W matrix was standardized by row, that is, the row elements sum to one. Thus, we can refer to ρ as an autocorrelation parameter. The **W** matrix was kept fixed for all simulations. We conduct a total of T=100 simulations. For each sample drawn, the small area mean is estimated using (a) the direct estimator, (b) the EBLUP estimator, (c) the SEBLUP estimator, (d) the M-quantile estimator and (e) the spatial M-quantile estimator

For each estimator we computed the Average Absolute Relative Bias ($\overline{ARB}$), the Average Relative Root MSE ($\overline{RRMSE}$) and the percentage relative bias $\overline{PRB}$ defined as follows:

$$\overline{ARB} = \frac{1}{m}\sum_{i=1}^{m}\left|\frac{1}{T}\sum_{t=1}^{T}\left(\hat{Y}_{it}/Y_i - 1\right)\right|$$

$$\overline{RRMSE} = \frac{1}{m}\sum_{i=1}^{m}\frac{\left[MSE(\hat{Y}_i)^{1/2}\right]}{Y_i}$$

$$\overline{PRB} = \frac{1}{m}\sum_{i=1}^{m}\frac{B_i^2}{MSE(\hat{Y}_i)}$$

$$\text{where } B_i^2 = \frac{1}{T}\sum_{t=1}^{T}\left(\hat{Y}_i - Y_i\right)^2$$

Tables 1 and 2 report the results from the simulation study. The main findings from the simulation study are:

i) M-quantile estimator reduces the bias in comparison with EBLUP and SEBLUP especially when we compute the CD estimator (see ARB values). This happens for

Table 1. *Comparison of small area estimators* $\rho > 0$.

	Estimator	ARB(%)	RRMSE(%)	PRB(%)
$\rho = 0.75$	*SEBLUP*	2.93	5.64	32.52
	EBLUP	2.38	6.11	18.44
	M quantile	2.93	6.18	29.67
	M quantile CD	1.87	6.03	13.42
	Spatial M.quantile	2.51	5.72	27.37
	Spatial M.quantile CD	1.55	5.81	10.96
$\rho = 0.5$	*SEBLUP*	2.66	4.53	46.55
	EBLUP	2.78	4.65	47.61
	M quantile	1.73	4.88	22.97
	M quantile CD	1.12	5.36	12.68
	Spatial M.quantile	1.98	4.40	27.92
	Spatial M.quantile CD	1.12	4.94	11.10
$\rho = 0.25$	*SEBLUP*	2.72	4.48	47.04
	EBLUP	2.69	4.39	47.81
	M quantile	1.78	4.59	21.83
	M quantile CD	1.11	4.89	10.04
	Spatial M.quantile	2.12	4.39	31.06
	Spatial M.quantile CD	1.17	4.74	11.03

Table 2. *Comparison of small area estimators* $\rho < 0$.

	Estimator	ARB(%)	RRMSE(%)	BIAS(%)
$\rho = -0.25$	*SEBLUP*	2.76	5.22	39.01
	EBLUP	2.78	5.20	39.31
	M quantile	2.01	5.34	21.61
	M quantile CD	1.32	5.64	9.39
	Spatial M.quantile	2.03	5.25	21.53
	Spatial M.quantile CD	1.23	5.62	7.99
$\rho = -0.5$	*SEBLUP*	3.25	5.80	38.03
	EBLUP	3.18	5.73	37.22
	M quantile	2.50	5.70	23.34
	M quantile CD	1.73	6.02	10.31
	Spatial M.quantile	2.41	5.72	20.16
	Spatial M.quantile CD	1.62	5.99	7.74
$\rho = -0.75$	*SEBLUP*	2.36	4.10	41.44
	EBLUP	2.40	4.23	41.92
	M quantile	1.68	4.33	20.92
	M quantile CD	1.19	4.66	10.51
	Spatial M.quantile	1.73	4.00	22.53
	Spatial M.quantile CD	1.05	4.40	9.56

every level of spatial correlation. When the semiparametric model is used this reduction is even more evident. Obviously the amount of reduction is higher for high level of spatial correlation (for $\rho = 0.75$ the $\overline{ARB}$ value is 1.55 for Spatial M quantile CD versus the 1.87 of the M quantile CD);
ii) the variability of M quantile estimator and of Spatial M quantile estimator is not always a good competitor of the variability of EBLUP and SEBLUP. This is probably due to the nature of the simulation study: we have assumed a normal distribution of random part in the model (9). Under this assumption it is well known that EBLUP and SEBLUP have good performance;
iii) anyway, the Spatial M quantile estimator based on a suitable spatial model (i.e. one that adequately reflects the between area variability in the population) work well when high or medium spatial correlation exists.

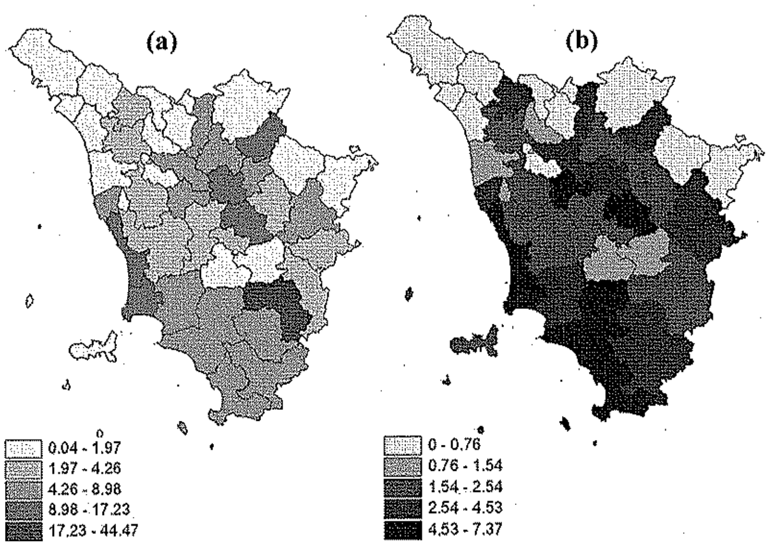
4 Results

The Farm Structure Survey is carried out once each two years and collects information on farms land by type of cultivation, the amount of breeding, the kind of production, the structure and the amount of farm employment. The sample is selected by a stratified one stage design with self-representation of larger farms (agricultural holdings). The sample size is 55,030 farms: 52,713 of them are drawn from the 2,150,000 firms of the so-called European Community target, while the additional 2,317 are selected from the 440,000 firms of the so-called Italy target[1]. The stratification is done in three phases. In the first one, the 6,972 self-represented farms are included in the sample on the basis of their economic dimension and/or their utilized surface area and/or number of bovines. In the second one, the residual EC targeted farms are divided into 407 strata utilizing dimensional, geographical and gross income parameters. Finally, the farms of the Italy target are stratified into 21 regional strata. The optimal allocation of sample size to the strata is obtained minimizing the sampling error at regional and national level.
Accurate estimates at sub-regional level require either the enlargement of the sample in provinces or municipalities or the application of small area estimation models. The Tuscany region is divided in 42 LES. The objective of inference is the average production of olives per farm ($\vartheta = \bar{y}$) for each of the 42 small areas (LES). The proposed method is implemented on this data using the expressions (4) and (5). The major determinant of the average production of olives per farm at area level is utilized surface area for olive production (ha). The neighbourhood structure **W** is defined as follows: spatial weight, w_{ij}, is 1 if area i shares an edge with area j and 0 otherwise. For an easier interpretation, the general spatial weight matrix is defined in row standardized form, in which the row elements sum to one.
The value of the estimated spatial autoregressive coefficient $\hat{\rho}$ is 0.441 ($s.e. = 0.183$) with REML (Restricted Maximum Likelihood) procedure and suggests a spatial relationship of a medium strength.
The maps in Figure 1 show two different aspects of the spatial distribution of the production per farm across the small areas (LESs): the mean production per farm

[1] The EC target is constituted by farms with more than 1 ha. of utilized surface area or less than 1 ha. if they produce for the market or when their production is over defined thresholds. The Italy target is constituted by the all the active censused farms in 2000.

(Figure 1a) highlights the LESs with relevant levels of production per farm (Val d' Orcia and the Chianti). The median (Figure 1b) is insensitive to the presence of few big farms that raise the medium level of production and, as a consequence, the spatial distribution of the median production seems to be more homogenous. This fact stresses the importance of maps that represent not only the spatial distribution of mean but also of other quantiles, emphasizing the importance of the estimation of the cumulative distribution function in each small area area. This makes it possible to know the concentration of the agricultural production at local level, giving a crucial information both for Government intervention on agricultural policies and for guidance activities for the final users of the data, the research institutes, the citizens.

Figure 1. (a) *Mean and (b) median production of olives per farm per LES*



5. Conclusions

We propose an extension to the Chambers and Tzavidis (2006) small area M-quantile model to allow for spatially correlated random effects. Spatial information is incorporated into the M-quantile model by modeling the M-quantile coefficients, obtained from the first step of the Chambers and Tzavidis (2006) approach, using a parametric model that allows for spatially correlated random effects. Small area estimates are then obtained by fitting an M-quantile model at the average area specific M-quantile coefficient predicted under the parametric model. Results from a simulation study indicate that this approach works well and competes the conventional spatial EBLUP estimator. The drawback of this approach is that we still need to specify a fully parametric model for modeling the unit-specific M-quantile coefficients. Currently, we investigate the use of non-parametric methods to incorporate spatial correlation in the M-quantile approach.

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Anno: **1991**

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Anno: **1992**

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Anno: **1993**

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Anno: **1994**

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Anno: **1995**

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Anno: **1996**

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Anno: **1997**

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Anno: **1998**

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Anno: **1999**

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Elenco dei report pubblicati

Anno: **2000**

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Anno: **2001**

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Anno: **2002**

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Anno: **2003**

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Anno: **2004**

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Anno: **2005**

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Anno: **2006**

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