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Causal inference methods without exclusion restrictions: an economic application

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1 Introduction

The paper considers an application of causal inference methods with non-compliance when the exclusion restrictions is relaxed. The exclusion restriction basically states that in an experiment with non-compliance the assignment to treatment has no direct effect on the outcome, but has only a treatment mediated effect. Recently the problems related to this assumption has been analyzed from theoretical and applicative point of views (Hirano et al., 2000; Imbens and Rubin, 1997; Jo, 2002). In particular Imbens and Rubin (1997) introduced a weak version of the exclusion restriction requiring that the assignment to treatment is unrelated to potential outcomes for non-compliers. Their is a first step towards less restrictive causal model; but relaxing completely the exclusion restrictions produces weakly identified models, in the sense of not having unique maximum likelihood points. Section 1 analyses the likelihood function of a randomized experiment with imperfect compliance, without exclusion restrictions and with a normally distributed outcome. In particular, a simulation based analysis suggests to impose appropriate restrictions on the parametric space in order to obtain unique maximum likelihood estimates. The proposed procedures will be used in Section 2 for evaluating the effect of the educational level on earnings using the cohort of birth as an instrumental variable and supposing the existence of direct effects of the instrumental variable on the outcome.

2 Theoretical framework

The likelihood function for a randomized experiment with imperfect compliance can be written as proposed by Imbens and Rubin (1997). Given the assumption of: SUTVA (Stable Unit Treatment Value Assumption), random assignment to treatment, monotonicity (Angrist et al., 1996), and supposing normal distributions for the outcomes, the likelihood function is (Mercatanti, 2002):

$$L(\boldsymbol{\theta}|\mathbf{y}_{obs}) \propto \prod_{i\in(D_i=1,Z_i=0)} \omega_a \cdot N\left(y_i|\mu_{a0},\sigma^2_{a0}\right) \times \prod_{i\in(D_i=0,Z_i=1)} \omega_n \cdot N\left(y_i|\mu_{n1},\sigma^2_{n1}\right) \times$$

$$\times \prod_{i\in(D_i=1,Z_i=1)} \left[\omega_a \cdot N\left(y_i|\mu_{a1},\sigma^2_{a1}\right) + \omega_c \cdot N\left(y_i|\mu_{c1},\sigma^2_{c1}\right)\right] \times$$

$$\times \prod_{i\in(D_i=0,Z_i=0)} \left[\omega_n \cdot N\left(y_i|\mu_{n0},\sigma^2_{n0}\right) + \omega_c \cdot N\left(y_i|\mu_{c0},\sigma^2_{c0}\right)\right]. \qquad (1)$$

where y_i is the individual outcome; ω_t is the probability of an individual of being in the t group, t=c (complier), n (never-taker), a (always-taker); μ_{tz} is the mean and σ_{tz} is the standard error of the outcome for individuals in the t-group and assigned to the z-treatment; $N\left(y_i|\mu_{tz},\sigma^2_{tz}\right)$ is the normal density function. The treatment is supposed to be a binary variable. Under this set of conditions, the three probabilities ω_a, ω_n and ω_c, can be consistently estimated respectively by (Imbens and Rubin, 1997):

- the proportion of treated in the group of units not assigned to the treatment, ϕ_a;
- the proportion of non-treated in the group of units assigned to the treatment, ϕ_n;
- the difference $\phi_c = 1 - \phi_a - \phi_n$.

The likelihood function (1) is characterized by the presence of two mixtures of densities; this complicates an analytical study of the likelihood and justifies a simulation based analysis, Mercatanti (2003). In order to address complications in making likelihood based inference from (1), we firstly consider artificial dataset #1; this is generated by drawing 10000 units from an hypothetical population and according to the function (1). Table 1 presents: the value of the parametric vector ($\boldsymbol{\theta}_1$) for this hypothetical population and the maximum likelihood points detected using the EM algorithm ($\hat{\boldsymbol{\theta}}_{1,.}$).

Table 1. *MLE for Dataset #1*

	$\boldsymbol{\theta}_1$	ϕ_t	$\hat{\boldsymbol{\theta}}_{1,1}$	$\hat{\boldsymbol{\theta}}_{2,1}$	$\hat{\boldsymbol{\theta}}_{3,1}$	$\hat{\boldsymbol{\theta}}_{4,1}$
ω_a	0.4	0.4	0.400	0.400	0.387	0.387
ω_n	0.25	0.25	0.252	0.323	0.250	0.323
ω_c	0.35	0.35	0.349	0.276	0.362	0.289
μ_{a0}	0		-0.014	-0.014	-0.014	-0.014
μ_{a1}	1		0.906	0.907	7.025	7.025
μ_{n0}	1		1.020	5.985	1.017	5.987
μ_{n1}	2		1.990	1.990	1.990	1.990
μ_{c0}	6		5.978	1.034	5.977	1.039
μ_{c1}	7		7.025	7.026	0.906	0.905
LogLik.			-30332	-30394	-30345	-30345

$(\sigma_{a0}, \sigma_{a1}, \sigma_{n0}, \sigma_{n1}, \sigma_{c0}, \sigma_{c1}) = (1, 1.2, 1.15, 1, 0.85, 0.7)$; $P(Z_i = 1) = 0.25$.

The choice of the EM algorithm (Dempster at al., 1977; Tanner, 1996) is suggested by considering its main peculiarity, that is the two steps, imputation-maximization, way for locating maximum likelihood points (a short illustration of it is in the Appendix). In particular the analysis of the matrix whose entries are the imputation probabilities calculated during the imputation step has been proved very helpful for our aim. Indeed an analysis of the imputation probabilities obtained at convergence of the EM algorithm, shows that the three solutions $\hat{\boldsymbol{\theta}}_{2,1}$, $\hat{\boldsymbol{\theta}}_{3,1}$ and $\hat{\boldsymbol{\theta}}_{4,1}$ are characterized by a wrong disentanglement of the mixtures (apart from sampling variabilities). The right imputations correspond to $\hat{\boldsymbol{\theta}}_{1,1}$ where, differently from the other three solutions, the estimates of ω_t are in a small spherical neighborhood of the vector $\boldsymbol{\phi} = (\phi_a, \phi_n, \phi_c)$. Then, for artificial dataset #1 a maximization procedure restricted to a neighborhood of $\boldsymbol{\phi}$ could be sufficient to obtain an unimodal likelihood function. In order to evaluate the performance of the proposed restricted procedure, repeated samples of 10000 units had been drawn from the same hypothetical population. Table 2 shows the operating characteristics for 100 replications of the maximum likelihood estimates restricted to a neighborhood of ϕ and in comparison to other standard methods (i.e. maximum likelihood estimation under the usual exclusion restriction and Instrumental Variable method). The proposed procedure is clearly superior in term of mean biases, mean squared errors, coverage rate and mean width of corresponding 95% confidence intervals. Only for the Instrumental Variable Estimator (IVE) the coverage rate of corresponding 95% confidence intervals is higher but at the cost of a dramatically higher mean width.

Table 2. *Operating characteristics of various procedures for replications from Dataset #1*

	estimator	Mean bias	Root MSE	95% interval Coverage rate	95% interval Mean width
μ_{c0}	MLE restr. neigh.of ϕ	0.002	0.079	0.947	0.312
	MLE under excl. rest.	0.204	0.220	0.240	0.306
μ_{c1}	MLE restr. neigh. of ϕ	0.002	0.024	0.991	0.072
	MLE under excl. rest.	0.256	0.272	0.237	0.377
σ_{c0}	MLE restr. neigh. of ϕ	0.004	0.041	0.947	0.163
	MLE under excl. rest.	0.042	0.088	0.846	0.156
σ_{c1}	MLE restr. neigh. of ϕ	-4.9×10^{-4}	0.054	0.940	0.224
	MLE under excl. rest.	-0.006	0.061	0.920	0.216
CACE*	MLE restr. neigh. of ϕ	1.1×10^{-4}	0.096	0.940	0.368
	MLE under excl. rest.	0.051	0.111	0.912	0.368
	IVE	-1.844	1.857	1.000	15.99

*: Complier Average Causal Effect

Another simulation analysis concerns artificial dataset #2; it is generated by drawing 10000 units from an hypothetical population whose parametric vector ($\boldsymbol{\theta}_2$) presents the same values of $\boldsymbol{\theta}_1$ apart from a smaller difference $(\mu_{n0} - \mu_{c0})$, Table 3. Now restricting the maximization to a spherical neighborhood of ϕ is not sufficient to get an unique solution. Table 3 shows that the restricted procedure produces two maximum likelihood points. The closeness of μ_{n0} to μ_{c0} reasonably confounds the two densities in the mixture of never-takers and compliers not assigned to the treatment, then obstructing its disentanglement. Extra restrictions on the parameter space are required for obtaining well behaved likelihood function; an intuitive restriction is in imposing an order for the two means μ_{n0} and μ_{c0}. For dataset #2 the appropriate assumption is $\mu_{n0} < \mu_{c0}$, that is $(\mu_{n0} - \mu_{c0}) < 0$.

Table 3. *MLE for Dataset #2 (restricted to a neighborhood of ϕ)*

	$\boldsymbol{\theta}_2$	ϕ_t	$\hat{\boldsymbol{\theta}}_{2,1}$	$\hat{\boldsymbol{\theta}}_{2,2}$
ω_a	0.4	0.4	0.401	0.401
ω_n	0.25	0.25	0.251	0.249
ω_c	0.35	0.35	0.349	0.349
μ_{a0}	0		0.020	0.020
μ_{a1}	1		0.988	0.988
μ_{n0}	1		0.938	1.241
μ_{n1}	2		2.059	2.059
μ_{c0}	1.2		1.200	0.983
μ_{c1}	7		7.015	7.015
LogLik.			-27260	-27261

$(\sigma_{a0}, \sigma_{a1}, \sigma_{n0}, \sigma_{n1}, \sigma_{c0}, \sigma_{c1}) =$
$= (1, 1.2, 1.15, 1, 0.85, 0.7)$;
$P(Z_i = 1) = 0.25$.

These two simulation based analyses have shown that restricting the likelihood maximization to a neighborhood of ϕ, and eventually imposing conditions about the order of the outcome means in the mixtures in (1), can be a feasible way for obtaining an unimodal maximum likelihood function.

3 An application: the evaluation of return to schooling using the cohort of birth as an instrumental variable

The fields of application for causal inference methods are huge and ranging for example from epidemiology, to economics and social sciences. In particular the statistical theory concerning randomized experiment with non-compliance is very important because not only limited to the analysis of experiments but extendible to observational studies. The randomized experiment with non-compliance is indeed a template, the situation involving so-called instrumental variables, Rubin (2000). In this sense the template is adopted to evaluate causal effects from non-experimental sources, and for which the instrumental variable has the role of a random assignment to treatment; under this theoretical structure the effects for compliers can be consistently evaluated.

An interesting microeconomic study concerning the evaluation of causal effects for compliers has been proposed by Ichino and Winter-Ebmer (2004) (IW) whose aim was to quantify the long run educational cost of World War Two (WWII) on earnings. Their idea is in the fact that people born during the period 1930-39 was damaged by WWII in their educational choices respect to people born during the immediately previous and subsequent cohort. Given date of birth can be reasonably supposed to be a random event, an Instrumental Variable Estimate (IVE) of education attained on earnings, using the cohort of birth as an instrument, was proposed to evaluate the long run effect of WWII. The identification of average causal effect for compliers by the IV method relies on the satisfaction of some assumptions and in particular on the exclusion restriction.

This section illustrates an evaluation of the effect of educational level on earnings for Germany and Austria using the cohort of birth as an instrumental variable, without exclusion restrictions and adopting the procedures proposed in the previous section. Definitions of the variables[1] are analogous respect to IW. In particular log hourly earnings are observed about 40 years after the end of WWII (at 1986 for Germany, at 1981 for Austria). In order to consider the increasing trend of earnings respect to age, the outcomes Y are defined as the residuals of a regression of log hourly earnings on a cubic polynomial in age. An increasing trend respect to age characterized the years of education too; for this reason the residuals of a regression of years of education on a cubic polynomial in age are calculated. But in order to apply the procedures previously proposed, the treatment has to be a binary variable. Then in defining the treatment, D, a comparison of residuals of the regression respect to their average is helpful. More precisely the treatment is defined equal to one if the residual is smaller than the average (low educated

[1] The data are from Mikrozensus 1981 and the Austrian Census 1981 for Austria and from the Socio-Economic Panel and the GSOEP for Germany. The population is defined as the individuals born between 1925 and 1949 (only males). The sample size is 3326 for Germany, and 15434 for Austria

people) and equal to zero if the residual is greater than the average (high educated people). The cohort of birth is used as an instrumental variable, Z, having the role of a random assignment to treatment. At this purposes, Z will have necessarily value one for people assigned to be low educated and zero for people assigned to be high educated. Table 4 shows that the average years of education and the average residuals are smaller for individuals in the cohort 1930-39 than for people in the cohort 1925-29 or 1940-49. These results suggests to define $Z_i = 1$ for individuals born during the period 1930-39, and $Z_i = 0$ for individuals born during the period 1925-29 or 1940-49.

Tab. 4. *Average years of education and average residuals per Country and Cohort of birth*

Country	Cohort of birth	Average years of education	Average residuals
Germany	1930-39	11.438	-0.1702
	1925-29 or 1940-49	11.941	0.1274
Austria	1930-39	9.116	-0.205
	1925-29 or 1940-49	9.480	0.036

Table 5 shows the results for Germany (calculations are performed by the EM algorithm); here the restricted likelihood maximization produces an unique solution. In particular the analysis are restricted to a spherical neighborhood of raw 0.015 around ϕ. Standard errors in parenthesis are calculated by the second derivatives of the likelihood at the maximum point. Last rows of the Table 5 shows significant effects of the cohort of birth for non-compliers, but an effect for compliers not significantly different from zero. The importance of relaxing the exclusions restriction in this application proved by a comparison of Table 5 results with the estimated effect for compliers produced by the IV method: -0.4029 (s.e.: 0.1465). The non-parametric local average treatment effect estimated by imposing the exclusion restriction is in absolute about five times the difference $\hat{\mu}_{c1} - \hat{\mu}_{c0}$. More reasonable results are instead produced by allowing different from zero effects for non-compliers. This is in line with microeconomic consideration of general equilibrium effects. Card and Lemieux (2001) using a model with imperfect substitution between similarly educated workers in different cohort of birth, argued that return to education reflects changes in the relative supply of highly educated workers across cohorts. The low average education in the 1930-39 cohort evidently raise the return to education for never-takers, that is for individuals always educated from a counterfactual point of view.

Table 5. *MLE for Germany (restricted to a neighborhood of ϕ)*

	ϕ_t	$\hat{\boldsymbol{\theta}}$
$\hat{\omega}_a$	0.725	0.715 (0.014)
$\hat{\omega}_n$	0.221	0.219 (0.009)
$\hat{\omega}_c$	0.054	0.064 (0.007)
$\hat{\mu}_{a0}$		-0.086 (0.016)
$\Delta\hat{\mu}_{a1}$		-0.063 (0.018)
$\Delta\hat{\mu}_{n0}$		0.287 (0.023)
$\Delta\hat{\mu}_{n1}$		0.436 (0.037)
$\Delta\hat{\mu}_{c0}$		0.465 (0.114)
$\Delta\hat{\mu}_{c1}$		0.381 (0.142)
LogLik.		-6051.5
$\hat{\mu}_{a1} - \hat{\mu}_{a0}$		-0.063 (0.018)
$\hat{\mu}_{n1} - \hat{\mu}_{n0}$		+0.148 (0.036)
$\hat{\mu}_{c1} - \hat{\mu}_{c0}$		-0.084 (0.1817)

For what concern Austria; the likelihood maximization restricted to a spherical neighborhood of ϕ (of raw 0.015) does not produces an unique solution. But imposing further rescrictions about the signs of the differences $(\mu_{c1} - \mu_{a1})$ and $(\mu_{c0} - \mu_{n0})$ achieve the aim. The results are presented in Table 6 where there are four restricted maximum likelihood points, one point for each specification of the outcome means order in the mixtures (calculations are performed by the EM algorithm). The choice of the particular solution depends on both likelihood performances and economic considerations; here the most reasonable solutions could be the first two $\hat{\boldsymbol{\theta}}_1$ and $\hat{\boldsymbol{\theta}}_2$. These solutions presents log-likelihood values appreciably greater than the others, and more plausible orders for the outcome means in the two mixtures. Indeed, compliers can be considered more motivated and able individuals respect to always-takers (these last are never educated from a counterfactual point of view). Then, it is reasonable to think that outcome mean for compliers are greater than outcome mean for always-takers in the relevant mixture: $\mu_{c1} >$

μ_{a1}. This is the restriction imposed for obtaining the solutions $\hat{\boldsymbol{\theta}}_1$ and $\hat{\boldsymbol{\theta}}_2$. The choice about the order of the outcome means in the other mixture is more problematic; compliers can be considered more motivated and able individuals also in this mixture. But never-takers are always high educated from a counterfactual point of view, so presumably in good social status and exploiting more advantages and opportunities in the labor market. For these reason the choice of the sign for the difference $(\mu_{c0} - \mu_{n0})$ is more questionable. Again the importance of relaxing the exclusions restriction proved by considering the estimated effect for compliers produced by the IV method: -0.3006 (s.e.: 0.0720). In particular solution $\hat{\boldsymbol{\theta}}_2$ presents a very close value of the compliers effect -0.3024, joined with a not significant effect for always-takers and with a small positive effect for never-takers (+0.0289). Solution $\hat{\boldsymbol{\theta}}_1$ presents more pronounced effect for compliers (-0.3832) and never-takers (+0.0696). In summary the two more reasonable solutions for Austria share a not significant effect for always-takers, a positive effect for never-takers and a negative remarkable effect for compliers.

Table 6. *MLE for Austria (restricted to a neighborhood of ϕ)*

	ϕ_t	$\hat{\boldsymbol{\theta}}_1 : \mu_{c1} > \mu_{a1}$ $\mu_{n0} < \mu_{c0}$	$\hat{\boldsymbol{\theta}}_2 : \mu_{c1} > \mu_{a1}$ $\mu_{n0} > \mu_{c0}$	$\hat{\boldsymbol{\theta}}_3 : \mu_{c1} < \mu_{a1}$ $\mu_{n0} < \mu_{c0}$	$\hat{\boldsymbol{\theta}}_4 : \mu_{c1} < \mu_{a1}$ $\mu_{n0} > \mu_{c0}$
$\hat{\omega}_a$	0.779	0.776 (0.007)	0.776 (0.007)	0.775 (0.007)	0.774 (0.007)
$\hat{\omega}_n$	0.151	0.148 (0.004)	0.148 (0.004)	0.147 (0.004)	0.146 (0.004)
$\hat{\omega}_c$	0.070	0.074 (0.006)	0.075 (0.005)	0.077 (0.005)	0.078 (0.005)
$\hat{\mu}_{a0}$		-0.074 (0.003)	-0.074 (0.003)	-0.074 (0.003)	-0.074 (0.003)
$\Delta\hat{\mu}_{a1}$		-0.006 (0.005)	-0.006 (0.005)	0.007 (0.005)	0.007 (0.005)
$\Delta\hat{\mu}_{n0}$		0.354 (0.013)	0.395 (0.015)	0.353 (0.013)	0.395 (0.015)
$\Delta\hat{\mu}_{n1}$		0.424 (0.012)	0.424 (0.012)	0.424 (0.012)	0.424 (0.012)
$\Delta\hat{\mu}_{c0}$		0.413 (0.028)	0.332 (0.021)	0.412 (0.027)	0.333 (0.021)
$\Delta\hat{\mu}_{c1}$		0.030 (0.032)	0.030 (0.032)	-0.102 (0.016)	-0.101 (0.016)
LogLik.		-20798.0	-20799.4	-20838.3	-20839.6
$\hat{\mu}_{a1} - \hat{\mu}_{a0}$		-0.006 (0.007)	-0.006 (0.005)	+0.007 (0.008)	+0.007 (0.008)
$\hat{\mu}_{n1} - \hat{\mu}_{n0}$		+0.069 (0.018)	+0.028 (0.019)	+0.070 (0.018)	+0.028 (0.019)
$\hat{\mu}_{c1} - \hat{\mu}_{c0}$		-0.383 (0.043)	-0.302 (0.038)	-0.514 (0.032)	-0.435 (0.026)

4 Conclusions

The problem of relaxing the usual exclusion restrictions in causal model with imperfect compliance has been considered. From a theoretical point of view, the simulations in Section 1 show that restricting the analysis to a spherical neighborhood of ϕ (estimated vector of the probabilities of being in a particular compliance-status), and eventually imposing conditions about the order of the outcome means in the mixtures involved in the likelihood function (1), can be appropriate for obtaining unique maximum likelihood points. An application concerning the evaluation of the effect of educational level on earnings (in Germany and Austria) using the cohort of birth as an instrumental variable has been presented in Section 2. Results shows the usual exclusion restriction can be successfully relaxed for Germany. For Austria an extra assumption about the order of the outcomes in the mixtures involved in (1) is necessary. The estimated effects for these two countries proved to be reasonably in line with the recent literature in labor economics.

APPENDIX

The EM algorithm works in general as follows (at the t iteration):

- in the E-step, by calculating the expected log-likelihood $l(\boldsymbol{\theta}|\mathbf{X})$ with respect to $f\left(\mathbf{X}_{mis}|\mathbf{X}_{obs};\boldsymbol{\theta}^{(t-1)}\right)$:

$$Q(\boldsymbol{\theta},\boldsymbol{\theta}^{(t-1)}) = \int l(\boldsymbol{\theta}|\mathbf{X})\, f\left(\mathbf{X}_{mis}|\mathbf{X}_{obs};\boldsymbol{\theta}^{(t-1)}\right) d\mathbf{X}_{mis}.$$

  where: $\mathbf{X}_{obs}$ is the observed data, $\mathbf{X}_{mis}$ is the unobserved data, and $\mathbf{X}$ is the complete dataset that is the dataset created by merging the observed and unosberved data.

  If $l(\boldsymbol{\theta}|\mathbf{X})$ is linear in $\mathbf{X}_{mis}$ this is equivalent to augmenting the dataset by the quantities $\mathbf{X}_{mis}^{(t)}$:

$$\mathbf{X}_{mis}^{(t)} = E\left(\mathbf{X}_{mis}|\mathbf{X}_{obs};\boldsymbol{\theta}^{(t-1)}\right);$$

- and in the M-step, by maximizing $Q(\boldsymbol{\theta}, \boldsymbol{\theta}^{(t-1)})$ with respect to $\boldsymbol{\theta}$ to obtain $\boldsymbol{\theta}^{(t)}$. If $l(\boldsymbol{\theta}|\mathbf{X})$ is linear in $\mathbf{X}_{mis}$ this is equivalent to maximize the log-likelihood based on the augmented dataset, $l\left(\boldsymbol{\theta}|\mathbf{X}_{mis}^{(t)}, \mathbf{X}_{obs}\right)$, to obtain $\boldsymbol{\theta}^{(t)}$.

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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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- ☐ **n. 90** Riccardo Cambini, Funzioni scalari affini generalizzate.
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Anno: **1996**

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- ☐ **n. 112** Riccardo Cambini, Generalized concavity for bicriteria functions.
- ☐ **n. 113** Vincenzo Bruno, Variazioni dinamiche (1971-1981-1991)dei fenomeni demografici dei comuni (urbani e rurali) della Lombardia, in relazione ad alcune caratteristiche di mobilità territoriale.

Anno: **1997**

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Anno: **1998**

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Anno: **1999**

- ☐ **n. 130** Alessandro Valentini, Impatto delle immigrazioni sulla popolazione italiana: confronto tra scenari alternativi.
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- ☐ **n. 132** Alessandro Valentini, Proiezioni demografiche multiregionali a due sessi, con immigrazioni internazionali e vincoli di consistenza.
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- ☐ **n. 141** Alessandro Valentini, Proiezioni demografiche con algoritmi di consistenza per la popolazione in Italia nel periodo 1997-2142: presentazione dei risultati e confronto con metodologie di stima alternative.
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- ☐ **n. 147** Alessandro Valentini - Francesco Billari - Piero Manfredi, Utilizzi empirici di modelli multistato continui con durate multiple
- ☐ **n. 148** Francesco Billari - Piero Manfredi - Alberto Bonaguidi - Alessandro Valentini, Transition into adulthoold: its macro-demographic consequences in a multistatew stable population framework
- ☐ **n. 149** Francesco Billari - Piero Manfredi - Alessandro Valentini, Becoming Adult and its Macro-Demographic Impact: Multistate Stable Population Theory and an Application to Italy
- ☐ **n. 150** Alessandro Valentini, Le previsioni demografiche in presenza di immigrazioni: confronto tra modelli alternativi e loro utilizzo empirico ai fini della valutazione dell'equilibrio nel sistema pensionistico
- ☐ **n. 151** Emilio Barucci - Roberto Monte, Diffusion processes for asset prices under bounded rationality
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Elenco dei report pubblicati

Anno: 2000

- ☐ **n. 161** Piero Manfredi and Luciano Fanti, Labour suppley, population dynamicics and persistent oscillations in a Goodwin-type growth cycle model
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- ☐ **n. 169** A. Marchi, Theorems of the Alternative by way of Separation Theorems
- ☐ **n. 170** Emilio Barucci - Maria Elvira Mancino, Asset Pricing and Diversification with Partally Exchangeable random Variables
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- ☐ **n. 179** E. Barucci - M.E. Mancini - Roberto Renò, Volatility Estimation via Fourier Analysis
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- ☐ **n. 196** M.Bottai, M.Bottai, N. Salvati, M.Toigo, Le proiezioni demografiche con il programma Nostradamus. (Applicazione all'area pisana)
- ☐ **n. 197** A. Lemmi - B. Cheli - B. Mazzolli, La misura della povertà multidimensionale: aspetti metodologici e analisi della realtà italiana alla metà degli anni '90
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Anno: 2001

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- ☐ **n. 205** Vincenzo Bruno, I Consumi voluttuari, nell'ultimo trentennio, in Italia
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- ☐ **n. 209** Riccardo Cambini - Claudio Sodini, A Finite Algorithm for a Class of Non Linear Multiplicative Programs.
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- ☐ **n. 215** Michele Longo - Vincenzo Valori, Existence end stability of equilibria in OLG models under adaptive expectations.

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- n. 218 Anna Marchi, On the connectedness of the efficient frontier: sets without local efficient maxima
- n. 219 Laura Lecchini - Odo Barsotti, Les disparités territoriales au Maroc au travers d'une optique de genre.

Anno: 2002

- n. 220 Gilberto Ghilardi - Nicola Orsini, Sull'uso dei modelli statistici lineari nella valutazione dei sistemi formativi.
- n. 221 Andrea Mercatanti, Un'analisi descrittiva dei laureati dell'Università di Pisa
- n. 222 E. Barucci - C. Impenna - R. Renò, The Italian Overnight Market: microstructure effects, the martingale hypothesis and the payment system.
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- n. 225 Andrea Mercatanti, Effetto dell'uso di carte Bancomat e carte di Credito sulla liquidità familiare: una valutazione empirica
- n. 226 Piero Manfredi - John R. Williams, Population decline and population waves: their impact upon epidemic patterns and morbidity rates for childhood infectious diseases. Measles in Italy as an example.
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