HETEROGENEITY IN RISK AVERSION AND RISK SHARING REGRESSIONS

Pierfederico Asdrubali [*] John Cabot University

Simone Tedeschi [†] Third University of Rome

Luigi Ventura [‡] Sapienza University of Rome

Abstract

By using data on consumption and income of Italian households for a recent time spell, including a substantial portion of the financial crisis, we show that heterogeneity of risk aversion coefficients may induce a bias on the income coefficient of standard consumption insurance regressions, whose direction is, *a priori*, not certain. We show that, extending the theoretical analysis and empirical findings in Schulofer-Wohl (2011), the sign of the bias is ambiguous, and depends on average income (growth) and on the covariances of both aggregate and idiosyncratic risk with individual risk aversion. While confirming that aggregate risk is negatively associated to risk aversion, our results further show that the latter may co-vary positively with idiosyncratic risk, thus implying a positive bias. Nevertheless, the role of average income may well still bring about a negative bias in standard consumption insurance tests.

Keywords: risk sharing; risk aversion; heterogeneity; endogeneity.
JEL classifications: C23; C26; E21.

1 INTRODUCTION

Standard tests of risk sharing, both from a macro and a micro perspective, have been performed under the more or less explicit assumption of homogeneous risk attitudes and homogeneous time preferences. In a recent paper by Schulhofer-Wohl (2011) (henceforth SW11) these assumptions have been brought under scrutiny, and shown to bring about a positive bias in risk sharing coefficients, implying an underestimation of insurance against idiosyncratic shocks. In this paper we complement and generalize SW11's arguments and show that:

1. theoretically, heterogeneity in risk aversion coefficients might lead to either an underestimation or an overestimation of risk sharing, depending on key co-variances of risk aversion with both aggregate and idiosyncratic risks, as well as on the sign and the relative importance of income first moments;

2. empirically, there is no evidence of an upward bias in risk sharing coefficients related to the omission of heterogeneous risk preference in recent Italian micro-data;

3. aggregate risk is positively related to the level of risk tolerance (as shown in other recent contributions) and, in addition, idiosyncratic risk is negatively related to risk tolerance (which is novel).

[*] Department of Economics and Law
[†] Department of Economics
[‡] Department of Economics and Law

2 THE THEORETICAL FRAMEWORK

Following SW11 and standard contributions in the literature, under the assumption of CRRA instantaneous utility functions we can test consumption risk sharing by running the following regression:

$$\log c_{it} = \frac{\log \alpha_i}{\gamma_i} + \frac{-\lambda_t}{\gamma_i} + \beta \log X_{it} + \varepsilon_{it} \tag{1}$$

where c_{ti} is household i's consumption at time t, α_i is its Pareto weight in the welfare maximization problem, γ_i its risk aversion coefficient, λ_t is the Kuhn-Tucker multiplier of the resource constraint and $\log X_{it}$ the log of household i's income at time t, representing an idiosyncratic shock - and then testing the statistical significance of β. However, standard consumption insurance tests do not estimate 1, but rather:

$$\log c_{it} = \frac{\log \alpha_i}{\gamma} + \frac{-\lambda_t}{\gamma} + \beta \log X_{it} + \hat{\varepsilon}_{it} \tag{2}$$

where the coefficients of relative risk aversion have been assumed equal across households, and the second term on the RHS reduces to a simple time indicator variable.

To get rid of the (household specific) Pareto weights, we can take the first difference of 1 and 2,[1] but even this will not remove heterogeneity in risk aversion from the corresponding equations. Therefore, it becomes important to establish what kind of bias on the coefficient can arise when the true DGP is 1, whilst 2 is estimated.

To answer this question, we follow SW11 in assuming that household income can be expressed as: $\log X_{it} = q_i m_t + u_{it}$, where u_{it} is an idiosyncratic shock to household i's income, and q_i is household i's loading factor on aggregate shocks. Moreover, we assume that u_{it} and ε_{it} are i.i.d. and that idiosyncratic shocks to consumption and aggregate shocks are uncorrelated, *i.e.* $\mathrm{Cov}(q_i m_t, \varepsilon_{it}) = 0$. Finally, we assume that $E(\frac{1}{\gamma_i}) = \frac{1}{\gamma}$, *i.e.* that the common relative risk tolerance $\frac{1}{\gamma}$ is the mean of households' relative risk tolerances, and that $\mathrm{Cov}\Big(u_{it}, (\frac{1}{\gamma_i} - \frac{1}{\gamma})\Big) = \mathrm{Cov}\Big(u_i, (\frac{1}{\gamma_i} - \frac{1}{\gamma})\Big)$ *i.e.* it is constant over time.

When 2 is erroneously estimated instead of 1, the error term in 2 becomes:

$$\hat{\varepsilon}_{it} = \Big(\frac{1}{\gamma_i} - \frac{1}{\gamma}\Big)(-\lambda_t) + \varepsilon_{it} \tag{3}$$

As is well known, the OLS estimator for the coefficient β in equation 2 is unbiased *iff* $\mathrm{Cov}\Big(\log X_{it}, \hat{\varepsilon}_{it}\Big) = 0$ and biased upwards (downwards) *iff* $\mathrm{Cov}\Big(\log X_{it}, \hat{\varepsilon}_{it}\Big) > 0$ (< 0).

The key issue is therefore that of assessing the sign of that covariance. By using the law of iterated expectations, it is possible to show that:

$$\begin{aligned}&\mathrm{Cov}\Big(q_i m_t, (\tfrac{1}{\gamma_i} - \tfrac{1}{\gamma})(-\lambda_t)\Big) + \mathrm{Cov}\Big(u_{it}, (\tfrac{1}{\gamma_i} - \tfrac{1}{\gamma})(-\lambda_t)\Big) = \\ &-\mathrm{Cov}\Big(m_t, \lambda_t\Big)\mathrm{Cov}\Big(q_i, \tfrac{1}{\gamma_i}\Big) - E(m_t)E(\lambda_t)\mathrm{Cov}\Big(q_i, \tfrac{1}{\gamma_i}\Big) - E(\lambda_t)\mathrm{Cov}\Big(u_i, \tfrac{1}{\gamma_i}\Big)\end{aligned} \tag{4}$$

Expression 4 is the generalization of expression (12) in SW11, where all the relevant interactions have been taken into account. If we only considered the first term on the RHS of 4, as in SW11, since the Kuhn-Tucker multiplier on the aggregate resource constraint is decreasing in aggregate consumption [implying $-\mathrm{Cov}\Big(m_t, (\lambda_t)\Big) > 0$], we would obtain a positive covariance under the hypothesis that $\mathrm{Cov}\Big(q_i, \frac{1}{\gamma_i}\Big) > 0$ *i.e.* that less risk averse individuals feature larger $q_i's$. If that were true, we would deduce an upward bias on the β coefficient in 2.

However, even conditioning on $\mathrm{Cov}\Big(q_i, \frac{1}{\gamma_i}\Big) > 0$, an upward bias won't necessarily ensue, if we take due account of all the other terms in 4. In particular, the signs of $E(m_t)$ and $\mathrm{Cov}\Big(u_i, \frac{1}{\gamma_i}\Big)$ are also key in determining the direction of the bias.

Regarding the former, in the case of equations written in levels, since $E(m_t)$ is always positive, the second term in

the RHS of 4 is negative, thus curbing or even reversing the upward bias reported in SW11.
Therefore, the direction of the overall bias is ambiguous. It depends on the signs of correlations between aggregate and idiosyncratic risk, on the one hand, and individuals' risk aversion, on the other, as well as on the relative importance of aggregate *vs.* idiosyncratic risk. Finally, the direction of the bias is influenced by the sign of income first moments. If the model is written in levels it is unambiguously negative while with a specification in differences it depends on the relative importance of negative over positive shocks on the cycle. In the following sections, we will present a consumption insurance analysis of a recent sample of Italian households, where we find no sign of a positive bias of β related to the (neglected) heterogeneity in the risk attitude, but instead (weak) signs of a negative bias. The full derivation of 4 is helpful to understand the results.

3 EMPIRICAL STRATEGY

We follow SW11's empirical strategy in bringing the analysis of section 2 to the data. After a brief description of our dataset, we proceed in two steps: first, we empirically establish the signs of the co-variances in 4 by measuring: *i)* the relationship between household income and aggregate shocks in a regression framework, to confirm that $\text{Cov}\left(q_i, \frac{1}{\gamma_i}\right) > 0$; *ii)* the relationship between risk attitude and idiosyncratic shocks, in a framework of two-stage least squares; *iii)* the sign of average aggregate income growth, in order to check the sign of $E(m_t)$, in the case of a model with differenced variables.
In a second step, we estimate a (set of) equation(s) for household nondurable consumption using 1 and 2 to analyze the degree of consumption insurance enjoyed by households in the sample. By using the answers to a hypothetical gambling question available in the 2010 wave of the Bank of Italy's Survey on household income and wealth (SHIW), we obtain a measure of individuals' risk aversion, which we assign to all the repeated observations for the same household, under the standard assumption - as in SW11 - of time-invariance of the risk attitude parameter (at least over relatively short time horizons). Unlike SW11, we are thus able to build a (quasi-)continuous measure of risk aversion and, most importantly, to be directly included in the risk sharing regressions to control for households' risk attitude.

3.1 Data and the measurement of risk aversion

We use data on households' non-durable consumption, and households' disposable income, from the biannual panel component of the Bank of Italy's SHIW 2000-2014.

This survey collects data on a large number of variables, namely: social and geographic characteristics of the household and of its members, as well as those of the spouses' parents; income data, consumption and saving, and gross and net wealth, in terms of real estate, financial assets, liabilities, just to cite some of the main items. The unit of analysis is the household, and the reference person for socio-demographic variables is the survey respondent.

A coefficient of absolute risk aversion (ARA), $A(w)$, is computed by using answers to a hypothetical gambling situation, question C38, asked to a 50% random subsample of the 2010 wave:

Question C38. Suppose you can participate in the following lottery: for each euro invested, you can double (win one more euro) or half (lose 50 cents) according to the result of a coin toss: head you win, tail you lose. How much would you invest?

By taking a first order Taylor expansion, we obtain the following measure of absolute risk aversion (details in the

Appendix):

$$A(w) = -\frac{u''(w)}{u'(w)} = \frac{1}{2.5x},$$

where x is the amount invested, and which confirms the intuition that the larger x, the lower the coefficient of absolute risk aversion.

A coefficient of relative risk aversion (RRA), *R(w)*, is also computed as the product of *A(w)* times the average monthly nondurable expenditure per adult equivalent. As is clear from the descriptives in Table I below, a few answers with very low amounts reported will drive up the mean of both ARA and RRA, whereas their median values belong to more familiar ranges. The pattern of risk aversion in our sample across economic, social and demographic variables closely tracks those in other contributions (see, for example, Dohmen et al. (2005)).

Table I: **ARA (left panel) and RRA (right panel) unconditional distributions**

	A(w)				*R(w)*		
Percentiles	Values			Percentiles	Values		
1%	0			1%	0.02		
5%	0			5%	0.15		
10%	0	Obs	2,383	10%	0.3	Obs	2,383
25%	0.002	Sum of Wgt.	1,991.80	25%	1.6	Sum of Wgt.	1,991.80
50%	0.008	Mean	0.07	50%	6.7	Mean	50.1
		Std. Dev.	0.12			Std. Dev.	101.6
75%	0.04			75%	42.6		
90%	0.2	Variance	0.01	90%	160	Variance	10332.8
95%	0.4	Skewness	2.09	95%	240	Skewness	3.6
99%	0.4	Kurtosis	6.11	99%	480	Kurtosis	20.8

Note: Sample of respondent households in the 2010 wave of the SHIW

3.2 Risk tolerance and the relationship between income and aggregate risk

As in SW11, to gauge the relationship between risk aversion and aggregate risk we run the equation:

$$x_{it} = q_0 + q_1 PCC_t + q_2 PCC_t * \gamma_i + \delta' \mathbf{w}_{it} + \eta_i + u_{it} \tag{5}$$

where x_{it} is (the log of) household i's (monetary) disposable income, and PCC_t is (the log of) per capita aggregate consumption; therefore, q_2, the coefficient of the latter interacted with relative risk aversion (γ_i), represents the differential impact of aggregate shocks for increasing levels of RRA; $\mathbf{w}_{it}$ is a vector of controls such as household head's education, years of experience, civil status, type of occupation (white/blue collar employee, manager, self-employed), sector (primary, industry, public administration, other) and the number of household earners. Finally, η_i is an individual fixed effect and u_{it} is the idiosyncratic disturbance.

As we can see from Table II, our findings are in line with SW11 and, in general, with previous results on the topic, suggesting that more risk tolerant (averse) agents tend to bear a larger (smaller) share of aggregate risk (see, for example, Lusardi (1997), Guiso et al. (2002), and Moskowitz and Vissing-Jørgensen (2002)). In particular, the interaction between the aggregate shock variable (PCC) and our measure of RRA is negative and significant, and more so when the variable γ_i is log-transformed (column 2).

Table II: **FE estimation of the household income process and the relationship between risk attitude and aggregate shock**

(1)	(2)
PCC 0.436**	PCC 0.618***
(0.2178)	(0.2057)
PCC*γ_i **-0.003****	**PCC***$\log(\gamma_i)$ **-0.178****
(0.0013)	**(0.0731)**
controls Yes	controls Yes
R^2 0.181	R^2 0.183
N. of cases 3,400	N. of cases 3,400
N. of groups 720	N. of groups 720
Avg n. of T 4.722	Avg n. of T 4.722
Prob > F (all η_i=0) 0.000	Prob > F (all η_i=0) 0.000

Note: FE estimates of equation 5. Standard errors in parenthesis are clustered at the household level.
$***$ p<0.01; $**$p<0.05; $*$ p<0.1.

3.3 Risk tolerance and the relationship between income and idiosyncratic risk

Another key term in 4 yet to be investigated is the correlation between the idiosyncratic disturbance and households' risk aversion.

To assess the sign of this correlation we use an indirect approach: we show that risk aversion as an explanatory variable of income has a statistically significant negative impact but it is also endogenous, and then assess the sign of the bias in the corresponding OLS model, thus uncovering the sign of the covariance between risk aversion and idiosyncratic shocks to income. Starting from equation 7 - which we re-specify in a cross-sectional fashion, as risk aversion is measured at one point in time only - we obtain:

$$x_i = q_0 + q_1\gamma_i + \delta^{'}\mathbf{w}_{it} + u_i \tag{6}$$

This specification allows us to directly test for the endogeneity of γ_i.

Our time-invariant instrument is information on the household head's parental age and education[2], while as timevarying instrument we use the presence of newborn children (aged 0 to 3)[3].

Table III compares an OLS model (column 1) with an instrumental variable (IV) estimation under the hypothesis of γ_i endogeneity only (column 2), and both γ_i and head's educational level endogeneity (column 3).

The main empirical result is that the IV estimator lowers the γ_i elasticity by 17 to 35 percentage points compared to the OLS, revealing that:

1. the true effect of the head's risk aversion on the average household income is likely negative (rather than nil) and, most importantly

2. the OLS estimator is upward biased, implying $\text{Cov}\left(u_i, \frac{1}{\gamma_i}\right) < 0$, and thus - according to the third term on the RHS of 4 - revealing the relevance of a second possible channel of upward bias in the β coefficient.

Table III: **Endogeneity analysis of RRA w.r.t the household income process**

	(1) OLS	(2) IV-1	(3) IV-2
$\log(\gamma_i)$	**-0.017**	**-0.183***	**-0.361***
	(0.0105)	**(0.0965)**	**(0.1352)**
Head educ.	0.073***	0.062**	0.362
	(0.2658)	(0.319)	(1.1924)
controls	Yes	Yes	Yes
R^2	0.358		
N. of cases	909	909	909
F test of excl. instruments		5.10	3.24/10.38
Sargan p-val		0.112	0.424
Endogeneity test p-val		0.055	0.032
		IV (First stage) for $\log(\gamma_i)$	*IV (First stage) for* $\log(\gamma_i)$
		Coef. $P>\lvert t\rvert$	Coef. $P>\lvert t\rvert$
Child03		.487 [0.047]	.509 [0.037]
Moth.educ(degree)		1.93 [0.002]	1.54 [0.021]
Log diff. parents age		.227 [0.043]	.218 [0.052]
			IV (First stage) for head's education
			Coef. $P>\lvert t\rvert$
Fath.educ			.168 [0.000]
Moth.educ			.113 [0.024]

Note: OLS (column 1) and IV (columns 2 e 3) estimates of equation 6. Standard errors in parenthesis. P-values in square brackets. $***$ p<0.01; $**$p<0.05; $*$ p<0.1. Among first stage significant predictors for RRA, in addition to the excluded instruments, female gender and age must be mentioned (with positive sign) while, in terms of correlations, head's self-employment (negatively) and working as employee in the public sector (positively).

3.4 Aggregate income growth

In view of the discussion in section 2, one element inducing a a negative bias in the slope coefficient in regression 2 is the average income $E(m_t)$. When equation 4 is re-written in differences:

$$\begin{aligned}&\mathrm{Cov}\Big(q_i\Delta m_t, (\tfrac{1}{\gamma_i}-\tfrac{1}{\gamma})\Delta(-\lambda_t)\Big)+\mathrm{Cov}\Big(u_{it}, (\tfrac{1}{\gamma_i}-\tfrac{1}{\gamma})\Delta(-\lambda_t)\Big)=\\&-\mathrm{Cov}\Big(\Delta m_t, \Delta\lambda_t\Big)\mathrm{Cov}\Big(q_i, \tfrac{1}{\gamma_i}\Big)-E(\Delta m_t)E(\Delta\lambda_t)\mathrm{Cov}\Big(q_i, \tfrac{1}{\gamma_i}\Big)-E(\Delta\lambda_t)\mathrm{Cov}\Big(u_i, \tfrac{1}{\gamma_i}\Big)\end{aligned} \tag{7}$$

the second term in the RHS is a function of $E(\Delta m_t)$ and thus can be either positive or negative, depending on the business cycle. Here we suppose that aggregate risk loading factors and risk aversion coefficients do not vary over the reference time frame, and all arguments and caveats illustrated in section 2 apply.

Aggregate disposable household income growth in 2010, 2012 and 2014 is negative in our data, therefore in case of equations written in differences, the contribution to the overall bias will be positive up to 2008 and then negative.

3.5 Risk sharing regressions accounting for heterogeneity in risk attitude

Since our dataset allows us to explicitly control for heterogeneity in risk attitudes - though in a randomized subsample only - we decided to exploit it by comparing different models and specifications under the alternative assumption of homogeneity and heterogeneity in risk preferences. All the following models exploit the panel structure of the dataset, which allows us to control for the role of time-invariant unobserved heterogeneity.

We also address other possible endogeneity problems by running two-stage least squares within-estimator regressions, with instruments which are plausibly exogenous to consumption and correlated with household income, like a contingent unemployment for a household member (provided that he is male, aged between 30 and 55 and with children) and - assuming preference separability between consumption and leisure - the head's yearly worked hours[4]. The first set of regressions - which assumes risk preference homogeneity - corresponds to the empirical implementation of 2 and is based on the specification:

$$c_{it} = \alpha_1 + \tau_1 PCC_t + \beta_1 x_{it} + \Theta_1^{'}\mathbf{w}_{it} + \mu_{1,t} + \varepsilon_{1,it} \tag{8}$$

The second set of regressions which assumes heterogeneity in risk preferences implements equation 1 and is based on the specification:

$$c_{it} = \alpha_2 + \tau_2\Big(\frac{PCC_t}{\gamma_{it}}\Big) + \beta_2 x_{it} + \Theta_2^{'}\mathbf{w}_{it} + \mu_{2,t} + \varepsilon_{2,it} \tag{9}$$

In the equations above, c_{it} is (the logarithm of) household nondurable consumption expenditure, and PCC_t is (the log of) average (per capita) total consumption, which is used as a proxy for λ_t in 1 and 2; $\mathbf{w}_{it}$ is a vector of socio-demographic controls (such as a polynomial in head's age, civil status, the number of household components, the presence of small children and the house ownership *vs.* tenancy); x_{it} is (the log of) household disposable income.

We restrict our sample to households with at least three consecutive repeated observations, ending up with 3,836 observations for 875 households.

Table IV compares the estimates from 8 and 9. In particular, columns (1-2) report on the implementation of a standard within estimator (FE), and columns (3-4) address more general endogeneity issues by using a two-stage least-squares within-estimator regression (IV-FE). Moreover, odd columns are related to equation 8, while even columns re-estimate the same models based on 9. A comparison between odd and even columns reveals that the overall bias related to the risk attitude, if at all present, leads to overestimate the extent of insurance through a downward bias of the coefficient β that measures the impact of an income shock on nondurable consumption, when heterogeneity in risk preference is erroneously omitted. The bias, at least in our sample, seems to be small and hardly statistically significant. However, different specifications and different estimators always correct the coefficient upwards rather than downwards as suggested by SW11. This is totally consistent with expression 4, as the first and the last terms are positive in view of the correlations described in the previous sections, but the second term is negative, and might well offset the former. Moreover, different sources of endogeneity seem to bias the estimates towards the hypothesis of perfect risk sharing for the model in levels. In fact, applying instrumental variables, the estimated income elasticity increases by about 40% (from about 0.15, column 2, to about 0.22, column 4). Finally, Table V shows that in the more standard risk sharing models, where variables are specified in differences, controlling for RRA heterogeneity still marginally increases the β estimates compared to the same models under the assumption of homogeneity in risk attitudes. This is also consistent with expression 7, as the first two terms are positive under the maintained hypotheses, and convexity of preferences, while the sign of the third depends on the expected value of the multiplier's growth rates. However, in this case, other sources of endogeneity seem to determine an upward bias with OLS estimates in differences[5] (columns 1 and 2 of Table V).

4 CONCLUDING REMARKS

We started from the recent contribution by Schulhofer-Wohl (2011), who shows that disregarding risk aversion (and other types of) heterogeneity could bias risk sharing coefficients upwards, and thus falsely induce a (much) lower estimate of consumption smoothing. We showed that an opposite result may well be found, if the model is specified in levels or, in case it is written in differences, positive average shocks prevail on the negative ones. We extended SW11's analysis both theoretically - by showing that the bias can occur in either direction - and empirically. In the empirical section, we present an example built on Italian micro data (which allows using a direct measure of relative risk aversion), exhibiting a *downward* bias, albeit weak, in the risk sharing coefficient. No upward bias from heterogeneity in the risk attitude was detected in our empirical analysis.

Table IV: **Risk sharing regressions in log-levels**

$Dep.Var.: c_{it}$	FE (1)	(2)	IV-FE (3)	(4)
$\widehat{\beta_{1,2}}$	**0.135***	**0.144***	**0.205***	**0.215***
	(0.0166)	**(0.0176)**	**(0.0411)**	**(0.0423)**
$\widehat{\tau_{1,2}}$	0.989***	0.012***	0.926***	0.011***
	(0.1166)	(0.0038)	(0.1193)	(0.0038)
Controls	Yes	Yes	Yes	Yes
R^2	0.195	0.163	0.180	0.148
N. of cases	3,836	3,836	3,836	3,836
N. of groups	875	875	875	875
Avg n. of T	4.386	4.386	4.386	4.386
F test of excluded instruments	-	-	25.76	25.81
Cragg-Donald Wald F	-	-	207.6	210.6
Hansen J p-val	-	-	0.116	0.228
Endogeneity test p-val	-	-	0.041	0.042

Note: Estimates of equations 8 (columns 1 and 3) and 9 (columns 2 and 4). Columns 1 and 2 apply FE estimator. Columns 3 and 4 apply IV-FE estimator with x_{it} as endogenous regressor. Standard errors in parenthesis are clustered at the household level. $***$ p<0.01; $**$p<0.05; $*$ p<0.1.

Table V: **Risk sharing regressions in log-differences**

$Dep.Var.: \Delta c_{it}$	OLS (1)	(2)	IV (3)	(4)
$\widehat{\beta}$	**0.206***	**0.214***	**0.160**	**0.211***
	(0.0232)	**(0.0240)**	**(0.0697)**	**(0.0702)**
$\widehat{\tau_{1,2}}$	1.234***	0.014***	1.256***	0.014***
	(0.1571)	(0.0047)	(0.1573)	(0.0047)
Controls	Yes	Yes	Yes	Yes
R^2	0.14	0.104	0.137	0.103
N. of cases	3,111	3,111	3,111	3,111
F test of excluded instruments	-	-	28.46	28.76
Cragg-Donald Wald F	-	-	151.3	154.4
Hansen J p value	-	-	0.769	0.887
Endogeneity test p value	-	-	0.492	0.955

Note: Estimates of equations 8 (columns 1 and 3) and 9 (columns 2 and 4) re-written in log-differences. Columns 1 and 2 apply OLS estimator. Columns 3 and 4 apply IV estimator with x_{it} as endogenous regressor. Standard errors in parenthesis. $***$ p<0.01; $**$p<0.05; $*$ p<0.1.

Notes

[1]Or, alternatively, with more than two repeated observations, apply a within estimator.

[2]For example, Dohmen et al. (2011) show that parental education is a significant predictor of the willingness to take risks but while father's education has a positive and significant impact on risk taking in all contexts, mother's education would be important in a more restricted domain, namely for risk taking in sports and leisure and career.

[3] Browne et al. (2016) in a still unpublished paper find - among others factors - substantial changes in risk attitudes over time with respect to giving birth to a child for the first time.

[4]We may well think that this type of worker/household member is characterized by a fairly rigid (full-time) labor supply.

[5]Please note that in the specification in levels, only, the endogeneity tests bring about a clear refusal of the null of household income exogeneity, while the same tests can not refuse the null in the differences specification cases.

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