Stability of risk attitude, agricultural policies and production shocks: evidence from Italy

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Abstract

This article investigates the stability of farmers' risk attitude over time. To this end, we estimate responses to changes in agricultural policies and production shocks. We use a unique panel data of over 36,000 Italian farms specialised in cereals, during the period 1989–2009. We find evidence of risk preference changes over time in response to changes in the European Union Common Agricultural Policy and possibly after a drought-induced production shock.

JEL classification: D1, Q12, Q18

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1. Introduction

Temporal stability of risk preferences can mean that subjects exhibit the same risk attitudes over time or that their risk attitudes are a stable function of states of nature that change over time (Andersen *et al.*, 2008). Risk preferences, in turn, should be disentangled into risk attitudes and risk perception (Pennings and Garcia, 2001; Pennings and Wansink, 2004; Just, 2008).[1] Risk attitude,

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1 Debates surrounding the nature of risk preference and its measurement have a long history in economics and psychology, and alternative definitions of risk preferences are suggested in the literature (Mata *et al.*, 2018).

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risk perception and their interaction are important determinants of the adoption and use of specific risk management tools and practices (Pennings, Wansink and Meulenberg, 2002; Pennings and Garcia, 2004; Pennings and Wansink, 2004; Just and Just, 2016). However, measuring these concepts in separation in non-experimental settings is very difficult (Iyer *et al.*, 2020).

In this article, we attempt to capture risk attitude recovering Arrow–Pratt (AP) and downside (DS) risk coefficients from secondary data of farmers' observed behaviours. We focus on risk attitude that reflects an agent's general predisposition to risk in a consistent way, whereas risk perceptions may be defined as her assessment of the uncertainty of the risk content inherent in a particular situation (Pennings, Wansink and Meulenberg, 2002). Using panel data of over 36,000 Italian farms during the period 1989–2009, we test whether the assumption of risk attitude stability over time holds in the context of major changes in the environment in which farmers operate. In particular, we investigate changes in farmers' risk attitude in response to multiple shocks coming from policy reforms over a long period of time and production shocks triggered by droughts, a type of extreme climate event.

Several studies in the general economic literature have shown that experiencing extreme events such as droughts, floods and earthquakes (Eckel, El-Gamal and Wilson, 2009; Page, Savage and Torgler, 2014; Hanaoka, Shigeoka and Watanabe, 2018) or a major traumatic experience such as a civil war (Kim and Lee, 2014) can affect risk attitude over time. Studies conducted after the onset of the financial crisis that started in 2008 address the possibility that a (negative) shock can trigger large increases in agents' risk aversion over a relatively short period of time (Malmendier and Nagel, 2011; Guiso, Sapienza and Zingales, 2018). Previous literature has investigated these effects based on controlled experimental research (Cohn *et al.*, 2015) or exploited specific potentially exogenous shocks (Kahsay and Osberghaus, 2018; Hanaoka, Shigeoka and Watanabe, 2018). The general idea is to investigate decision makers' risk-taking behaviour in a field setting. Kahsay and Osberghaus (2018) use two rounds of survey data of German households and find evidence of an increase in risk-seeking behaviours after people experience storm damage. The authors highlight the fact that there is scarcity of research looking at the effects of natural hazards on risk preferences in the European context, unlike other developed regions of the words such as the USA (Eckel, El-Gamal and Wilson, 2009), Australia (Page, Savage and Torgler, 2014) and East Asia (Hanaoka, Shigeoka and Watanabe, 2018). Furthermore, only few applications investigate the temporal stability of risk attitude with focus on the agricultural sector (Love and Robison, 1984 and Koundouri, Laukkanen and Myyra, 2009).[2] The focus on agricultural applications is highly promising because farmers face a high-risk exposure and extreme events from different risk sources (Moscardi and de

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2 The article by Love and Robison (1984) is one of the early contributions to this field. The authors examined the intertemporal stability of risk preferences eliciting them from a small sample of 23 American farmers in 1979 and then again in 1981 and concluded that risk preferences were not stable over time.

Janvry, 1977; Leathers and Quiggin, 1991; Chavas and Holt, 1996; Hennessy, 1998; Moschini and Hennessy, 2001; Roe 2015; Trujillo-Barrera, Pennings and Hofenk, 2016). Yet, the role of different sources of shocks on farmer risk attitudes remained unexplored so far. More specifically, public policies have been shown to affect risk attitude but this side effect of public policies has been mostly neglected in the academic and policy debate (Aragón, Molina and Outes-Leon, 2017).

We aim to contribute filling these gaps by investigating the (in)stability of farmer risk attitude over time looking at the effect of both policy changes and production shocks driven by weather. To this end, we recover farmer risk attitude from a large and temporally long panel data set that includes farm-level information on input expenditures and output realisation and test for changes in risk attitude after facing major policy and climate shocks. More specifically, we use a unique data set for Italian cereal farmers covering a period of 21 years from 1989 to 2009 and containing more than 36,000 farm observations. Data are unbalanced. The total number of farm-year observations is about 116,700. We investigate farmers' adjustments in input choices based on the *method of moments* approach first proposed by Antle (1983, 1987). The use of panel data techniques allows to control to a certain extent for unobserved heterogeneity or path dependency. However, the use of secondary data does not allow us to control for all unobserved variation in the environments in which farmers operate, or due to environmental (Saastamoine, 2015) or individual characteristics not recorded in our data set such as age, gender and parental background (Eckel and Grossman, 2002; Dohmen *et al.*, 2011; Kim and Lee, 2014). As thoroughly discussed in a recent review by Iyer *et al.*, (2020), the here chosen approach does not allow to observe the range of attributes that we would observe in experimental data. This shortcoming substantially increases the chance that the estimated risk aversion coefficients will be biased. Lence (2009) warns that typical production data rarely contain enough information to allow identification of the structure of risk aversion, but seems to accept estimations of risk coefficients in large samples (Lence, 2009; Foudi and Erdlenbruch, 2012). Just and Just (2011) criticise more harshly the method arguing that separate identification of (production) risk and risk preferences from behavioural equations estimated on observed data is impossible, hence restrictions have to be imposed either on the technology or on the form of the utility function for parameters to be identified, which in turn undermines the possibility of global identification of risk preferences.

The use of methods based on multi-item scales and methods based on lottery-choice tasks, which become more frequent in the last decade (Iyer *et al.*, 2020), partly overcomes the limitations faced by empirical studies from observed behavioural secondary data, like the analysis suggested in this article. However, these methods elicit preferences from primary data, and although we recommend them for future studies, they would not allow us to recover risk attitude and changes related to major events that happened in the past. Furthermore, none of the previous studies using econometric estimations based on secondary farm-level data to attempt recovering risk attitude parameters

relies on a large countrywide panel data set, covering such a long period of time as the one used in this article. Italy is also an excellent case study to investigate the impact of changing policy and climate shocks on European farms, due to the highly heterogeneous climatic, soil, socio-economic and topographical features of the Italian peninsula (Bozzola *et al.*, 2018). In the 20 years covered by our data, the Italian farming sector experienced various shocks. This allows us to investigate changes in farmers' risk attitude in response to multiple shocks coming from policy reforms over a long period of time and extreme climate events. First, we look at shocks related to major changes in the European Union (EU) Common Agricultural Policy (CAP), i.e. the agricultural policy of the EU. These reforms started in 1992 and introduced main policy changes in European agriculture. Second, we look at climate shocks, focusing on the effects of two large-scale drought events in 2003 and 2007.

We find evidence of risk preference changes over time in response to major changes in the agricultural policy and to the droughts. Thus, we provide new evidence that the CAP and changes therein as well as exposure to climate shocks influence risk attitudes.

The remainder of the paper is organised as follows. In the next section, we provide a policy and historical background of the main events in the Italian agriculture covered in this article and we formulate some hypothesis on farmers' risk attitude based on the existing literature. In Section 3 we describe the data set. In Section 4, we outline the conceptual and empirical model. Section 5 presents the main results. In Section 6, we draw some conclusions and outline new research avenues.

2. Policy and historical background

In this section, we provide background information about the events that could have led to temporal instability of Italian farmers' risk attitude during the 20-year period covered by our data. We offer an overview in Table 5, with a focus on the influence of these events on the Italian farming sector. We refer to the relevant literature to formulate hypothesis on farmers' risk attitude.

2.1. The reforms of the EU CAP

Farms specialised in cereals are of outmost importance for Italian agriculture and were particularly affected by the CAP reforms of the 1990s (Sckokai and Moro, 2006; Platoni, Sckokai and Moro, 2012) and by the Fishler reform (Swinnen, 2008; Moro and Sckokai, 2013). The process towards a (partially) decoupled support to farmers started in 1992 with the Mac Sharry reform. The 1992 reform constituted a major shift from product support (through prices) to producer support (through income support) and marked the beginning of a series of CAP reforms. The Mac Sharry reform transitional period ended in July 1995 and the main implementation phase lasted until 1998.[3] In 1999 the

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3 In November 1993, the European Commission (EC) presented an Agricultural Strategy Paper in which it examined three different options for reforming further the CAP. These ranged from

Table 1. Selection of sub-periods for the empirical analysis

Relevant events and time period	Background information and hypothesis regarding risk attitude
[1] 1989–1991 Pre-CAP reform period	Farmers are expected to be risk averse or risk neutral (Roche and McQuinn, 2004; Koundouri, Nauges and Tzouvelekas, 2006; Groom *et al.*, 2008; Koundouri, Laukkanen and Myyra, 2009).
[2] 1992–1999 Mac Sharry reform	The process towards a (partially) decoupled support to farmers started in 1992 with the Mac Sharry reform. The CAP shifts from market support to producer support. This 7-year reform period can be broadly divided into two phases: a transition phase (1993–1995) and the implementation phase (1996–1999).
[2a] 1993–1995 Mac Sharry—transition phase The Uruguay Round Agreement in Agriculture entered into force In 1993 negotiations for accession of Austria, Finland and Sweden to the EU began. These countries entered the EU in 1995	Ambiguous impact on cereal farmers' risk attitudes: Risk aversion may decrease because of the reduction in the random component of farmers' income and the effect of direct payments on wealth effect (Koundouri, Laukkanen and Myyra, 2009; Femenia, Gohin and Carpentier, 2010). Farmers' risk aversion may increase due to policy uncertainties (Moschini and Hennessy, 2001; REAS, 2010).
[2b] 1996–1999 Mac Sharry—implementation phase	The uncertainty stemming from the new CAP reform and the other policy changes that happened at the beginning of the 90s decreased.
[3] 2000–2002 Agenda 2000 CAP reform, up to the 2003 climate shock	This CAP reform did not change the basic structure of the new regime, but further reduced the intervention prices and increased the cereal area payments (Platoni, Sckokai and Moro, 2012). Hence, the two effects described in row 2a may still apply, but we expect the drivers of increasing risk attitude to be less severe.

Continued

Table 1. Continued

Relevant events and time period	Background information and hypothesis regarding risk attitude
[4] 2003–2004 The Midterm Review of Agenda 2000 (a.k.a. the 'June 2003 Fischler reform'), and the 2003 climate shock	The June 2003 Fischler reform introduced the single farm payment decoupling, a large share of CAP support from production from 2005 onwards. The switch to decoupling was a major structural change in farmers' income support. In 2003 the whole Italian peninsula was hit by a major drought. After individuals experience, a main shock risk attitude may change but the sign of the change is ambiguous a priori. (Guiso, Sapienza and Zingales, 2018) conclude that a (negative) shock can trigger large increases in agents' risk aversion over a relatively short period of time. Kahsay and Osberghaus (2018), however, find evidence of an increase in risk-seeking behaviours after people experience storm damage. None of these studies focuses on the agricultural sector. In agriculture, experiencing extreme events may lead to sudden changes in farmers' behaviour, investment and productive decision therein (Carey and Zilberman, 2002; Ding, Schoengold and Tadesse, 2009; Nauges, Wheeler and Zuo, 2016).
[5a] 2005–2007 Fischler reform—implementation phase	From 2005 to 2009 the CAP went through the implementation phase of the Fischler reform. If farmers' risk preferences are consistent with DARA, direct payments should reduce their risk aversion (Saha, Love and Schwart 1994; Chavas and Holt, 1996; Koundouri *et al.*, 2009). The new policy may have induced those farmers who choose to produce to allocate more land to riskier products than previously (Roche and McQuinn, 2004).
[5b] 2008–2009 Post-2007 climate shock	Repeated production shocks, in the article captured by recurring droughts (i.e. a type of climate shock), may exacerbate the effects described in row 4.

'Agenda 2000' reform was introduced. This second CAP reform did not change the basic structure of the new regime, but further reduced the intervention

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maintaining the status quo to proposing a new radical reform, drastically reducing EU prices to world market levels and abolishing production quotas and other supply management measures (EC, 1995).

prices and increased the cereal area payments (Platoni, Sckokai and Moro, 2012). Following the 1992 reform, the level of support for cereal farmers reduced by about 30 per cent, with guaranteed prices of cereals lowered by 35 per cent. However, volatility of agricultural prices in Italy remained limited and below the European average (Visciaveo and Rosa, 2012). Italian cereal farmers anticipated the fall in prices. Koundouri, Laukkanen and Myyra (2009) argue a similar case about Finnish cereal producers: although output prices decreased significantly at the time Finland entered the EU, the fall in prices was completely anticipated by producers at the time cereal production decisions for 1995 were made. Furthermore, the authors argue that in the period 1995–2003 yield variability was the dominant determinant of variability in wheat and barley revenues, whereas the variability of prices and acreage was less relevant.

Compulsory set-aside measure and other accompanying measures were also introduced, together with two components to farm income: an *area payment* component and a *market* component. The first depends on the land allocation decision, whereas the second is the one that might generate more uncertainty, which can come from both market prices and yields. Some authors argue that the increase in the cereal area payment substantially increased the share of non-random income, which in turn led to riskier behaviour (Koundouri, Laukkanen and Myyra, 2009). Direct payments can affect the incentives to produce for risk-averse farmers through their impact on farmers' wealth (Hennessy, 1998; Femenia, Gohin and Carpentier, 2010). Other authors, on the contrary, stress that the reform increased uncertainty in the sector because farmers felt uncertainty about the stability of political support for direct payments, which was a major cause of farmers' resistance to lower prices (Bernstein *et al.*, 1999). Moreover, the Single Payment Scheme providing farmers with a fixed amount was introduced only after 2003.

During the implementation period of the Mac Sharry reform, the European agricultural sector went through other major policy changes: the Uruguay Round Agreement in Agriculture (The [General Agreement on Tariffs and Trade (GATT)] entered into force and the EU enlarged with three new members: Austria, Finland and Sweden). Subsequently, the EC started examining different options for the future development of the CAP. The Mac Sharry reform and the Agenda 2000 are discussed and implemented before the 2003 climate shock, whereas the Fischler reform started in June 2003 with the so called 'The Mid-term Review of the Agenda 2000'. The Fischler reform, with the switch to decoupling, has been another major structural change in farmers' income support. The main element of the June 2003 Fischler reform was the introduction of the single farm payment, decoupling a large share of CAP support from production. While this reform was announced and implemented two major droughts hit the Italian peninsula, causing severe damages to the agricultural sector (production shocks).

2.2. The 2003 and 2007 droughts

In the summer of 2003, the whole Italian peninsula and many other areas of Europe faced the most severe drought event and heat wave in decades and the hottest summer in Europe since the 16th century (World Meteorological Organization, 2003). These severe climate conditions began in Europe in June 2003 and continued until mid-August, with summer temperatures 20–30 per cent higher than the seasonal average. The agricultural sector was amongst the most severely hit by the persistent droughts (Fink *et al*., 2004; Ciais *et al*., 2005; García-Herrera *et al*., 2010).

A second severe drought hit the Italian agricultural sector in 2007. This second drought was particularly severe in the north of the Italian peninsula: it started by the end of 2006 and the poor weather conditions extended through 2007, with a spring and summer drier and hotter than seasonal long-term averages in many parts of Italy (Italian National Institute for Environmental Protection and Research (ISPRA), 2008).

3. Data

Our data set comprises farm-level data from the Italian Farm Accountancy Data Network (FADN) and includes more than 36,600 farms specialised in cereals and various field crops covering the 21-year period, from 1989 to 2009. Farms are located across the Italian territory. Data are unbalanced and on average each farm is included in the data set for 3.2 years, and the total number of farm-year observations is about 116,700. The Italian authorities in charge of data collection had improved the sampling methodology through the years, but the Italian cereal farmers included in the analysis can be considered a representative sample through time. For this reason, we perform the analysis using the entire available data set instead of relying on a much smaller data set weakly unbalanced, for which selection biases would be expected.[4]

The data set provides detailed information about the main production orientation, the value of the fixed assets, variable input expenditure by crop for key inputs such as fertilisers, labour, seeds and crop protection products, some other input information (such as average work units), income and structural characteristics of the farms. Table 2 provides the descriptive statistics and short definitions for each variable used in this study.

4. Conceptual and empirical model

We use a flexible estimation approach where we consider uncertainty by using moments of the profit distribution as determining farmers' decisions regarding the input mix. This flexible approach has been developed by Antle (1983,

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4 For robustness analysis, we also obtained the AP and DS risk aversion parameters for the entire period from a smaller sample of 1,779 cereal farms, including in the estimations only those farms appearing in the data set for at least 10 years. The estimated AP and DS risk aversion parameters are 0.6193 and 0.0088, respectively, and we strongly reject the null hypothesis of these coefficient being zero, coherently with the results presented in Table 3.

Table 2. Descriptive statistics and variables definitions

Variables	Mean	SD	Min	Max
Net profit/loss	22.5	78.8	−3207	5965
Farm net income ('000€)				
Fertilisers	3,300	9,335	0	1,172,010
Total fertiliser and soil improvers expenditure (€). It includes purchased lime, compost, peat, manure				
Labour	1.61	1.91	0.02	140.16
Total labour input expressed in annual work units (AWU) = full-time person equivalents				
Seeds/seedlings purchased	3,443	18,202	0	2,597,102
Total seeds/seedlings purchased (€)				
Crop protection	2,507	10,842	0	1,098,953
Total crop protection products expenditure (€)				
Total assets ('000€)	446.5	1,547	0	90,600
Value of fixed assets (€).				
Land	28.6	64.5	0.1	4290
Total UAA				
Share of irrigated land	0.11	0.27	0	1
Total UAA under irrigation/total UAA (hectares/hectares)				
Share of rented land	0.31	0.39	0	1
Total UAA rented/total UAA (hectares/hectares)				
Share fallow land	0.02	0.07	0	0.96
Agricultural land not cultivated for agricultural reasons/total UAA (hectares/hectares)				
Dummy LFA	0.42	0.49	0	1
Dummy variable, 1 if the majority of the UAA is in a mountain or non-mountain LFA				
Family	0.26	0.44	0	1
Dummy variable, 1 if organisational form is registered as an individual (family) farm				

Total farms: 36,691, including: specialist cereals, general field cropping and mixed cropping. More details on the data used in this study are provided in the Table A1. SD = standard deviation; Min. = minimum; Max. = maximum.

1987); Antle and Goodger (1984) and Kumbhakar (2002) and has been used to estimate risk attitude (Groom *et al.*, 2008; Koundouri, Laukkanen and Myyra, 2009). This model is ideally suited to analyse responses to interventions in uncertain environments (Groom *et al.*, 2008 pp. 316) and is particularly appropriate to be applied when agents are exposed to various type of risks, such as production, market and policy risks (Gardebroek, 2006).

We assume that farmer's behaviour is consistent with expected utility theory.[5] Hence, cereal farmers maximise expected utility, which is depicted as a function of moments of the profit distribution by estimating the distribution

5 Just and Peterson (2010) argue that the Expected Utility Theory is not applicable to every situation and suggest a calibration procedure to compute the degree of minimum concavity that rationalises observed choice behaviour, which is directly comparable with the estimates of risk-aversion coefficients based on the same (Just and Peterson, 2010).

of the error term. Those moments have the same vector X as an argument, so that the farmer's program becomes:

$$\max_{X} E\left[U\left(\pi\right)\right] = F\left[\mu_1(X), \mu_2(X), \ldots, \mu_m(X)\right] \tag{1}$$

where U is the *von Neumann–Morgenstern* utility function representing the farmer's risk preferences and π is the farmer's profit. $\mu_j = E\left[\Pi - E\left(\Pi\right)\right]^j$ i.e. $\mu_j = E\left[\Pi - \mu_1\right]^j$ and $j = 2, \ldots, m$ is the central m^{th} moment of profit.[6] These are obtained following the sequential procedure described below, which follows Kim and Chavas (2003).

First, we estimate the conditional expectation of profit (mean effect regression):

$$\pi_{it} = f\left(\boldsymbol{x}_{it}, \boldsymbol{q}_{it}, z_t; \beta\right) + \alpha_i + \gamma_{it} \text{ where } \alpha_i + u_{it} = \gamma_{it} \tag{2}$$

We use a large data set of Italian agriculture that comprises farm-level data for more than 36,600 farms specialised in cereals and other open field crops production, referring to the period 1989–2009. We regress total observed profit from production (π_{it}) on all levels, squares and interaction of variable input expenditure/annual units, captured by the vector x_{it} (fertilisers, seeds, crop protection products and labor).[7] We use a quadratic functional form following earlier studies (Kumbhakar and Tveteras, 2003; Koundouri, Nauges and Tzouvelekas, 2006; Groom *et al.*, 2008; Zuo, Nauges and Wheeler, 2015). We also include a vector q_{it} of other important covariates, such as the area of the utilised agricultural area (UAA), the share of land rented and under irrigation, the value of fixed assets and a dummy variable indicating if the farm is registered as a family farm. The subscripts $i = 1, \ldots, N$ and $t = 1, \ldots, T$ denote, respectively, individual farm units and the time periods (years).

We also include year dummies (vector z) to remove any general time trend affecting all farmers identically. Moreover, the year dummies are important in our setting, to mitigate the confounding effect that inflation may have on expenditure data, since we use a very long panel data set (1989–2009).[8] Other variables such as weather conditions are not directly under the farmer's control and contribute to profit variability (Zuo, Nauges and Wheeler, 2015).

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6 (Chavas, 2004) shows that the AP risk premium can be approximated by a linear function of the first m moments of the distribution of profit.

7 All variables are scaled by their standard deviation, as in (Groom *et al.*, 2008) and (Vollenweider, Di Falco and O'Donoghue, 2011).

8 FADN data for input quantities are not available for the time period under study, with the exception of labour. Hence, the use expenditure data might include price developments, especially given the long period of time under scrutiny in this paper. The here used FADN has already applied rates to convert monetary values from the Italian lira to EUR that change over time for the period 1989–1999 to reflect inflation. Moreover, we include year fixed effects in the estimations and we conduct several sensitivity analyses for sub-periods. Beyond this, the empirical case for deflating may be weak compared with the risk of introducing severe measurement errors that could be due, amongst other, to the need of compiling longer time series from the price indices released for specific sub-periods, and with different base years. These indices, in turn, present variation in the type and number of items included in their underlying baskets through time.

The panel structure of our data set allows to estimate equation (2) using a fixed-effect estimator. This provides consistent parameters even if there is correlation between the independent variables and time-invariant unobserved heterogeneity.[9] Hence, the error term associated to equation (2) is disentangled as follows: $\gamma_{it} = \alpha_i + u_{it}$. We control for unobserved farm-specific effects through α_i while u_{it} is the idiosyncratic error term. Hynes and Garvey (2009) highlight the advantages of panel data techniques over static frameworks and warn that where no attempt is made to control for unobserved heterogeneity or path dependency, the effects of the farm-specific characteristics may be overestimated (Hynes and Garvey, 2009 pp. 546). The use of fixed-effects estimation methods, the inclusion of farm size and other relevant explanatory variables and the fact that we focus on a group of farms relatively homogeneous in term of technologies adopted, i.e. cereal farmers[10], limit our concerns of cross-sectional heterogeneity in risk attitude. This can be due to producer characteristics such as education and parental background or environment characteristics (Saastamoinen, 2015). In this setting, although cross-individual heterogeneity cannot be fully accounted for, we attempt to obtain information about changes in risk aversion in the farm population following aggregate shocks, but we cannot trace this back to each individual farm. We acknowledge the fact that even if we focus on a certain production type, heterogeneity in risk attitude cannot be fully captured. Pennings and Garcia (2004) for example, found that the impact of risk attitudes on behaviour may change across farmers even for homogeneous farmers.

In the second step of our empirical procedure, the estimates of the first moment of the profit distribution $(\hat{u}_{it})$ are then raised to the power two and three to compute conditional higher moments (variance and skewness, respectively). We opted against consideration of additional moments as Antle (1983) and Groom *et al.* (2008) found that estimation coefficients associated to the kurtosis aversion were not significant. Along these lines, Koundouri, Nauges and Tzouvelekas (2006) and Zuo, Nauges and Wheeler (2015) recognised that distributions of profits in agriculture are well approximated by their first three moments.

The predicted values raised to the power two and three are then regressed, using the fixed-effects estimator, on the same explanatory variables included in the estimation of the mean effect, consistently with previous empirical work (Koundouri, Nauges and Tzouvelekas, 2006; Groom *et al.*, 2008).

$$\hat{u}_{it}^2 = w\left(\boldsymbol{x}_{it}, \boldsymbol{q}_{it}, \boldsymbol{z}_t; \delta\right) + \check{\gamma}_{it} \tag{3}$$

$$\hat{u}_{it}^3 = s\left(\boldsymbol{x}_{it}, \boldsymbol{q}_{it}, \boldsymbol{z}_t; \phi\right) + \tilde{\gamma}_{it} \tag{4}$$

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9 Besley and Case (1993); Koundouri, Nauges and Tzouvelekas (2006); Kahsay and Osberghaus (2018) and Hanaoka, Shigeoka and Watanabe (2018) discussed limits in using simple cross-section data.

10 Some authors stressed that conclusions on risk preference of one group of farmers (e.g. cereal farmers) cannot be extended to other producers (e.g. vegetable farmers) (Iyer *et al.*, 2020).

We use the estimated coefficients from equations (2, 3, 4) (i.e. the vectors $\hat{\beta}$, $\hat{\delta}$, $\hat{\phi}$) to compute twelve analytical expression for derivatives of the each moment with respect to each variable input (fertilisers, plant protection, seeds and labour).

Finally, because the first-order conditions of the farmer's program can be approximated using a Taylor expansion in matrix form, we can estimate system (5) of K equations (Antle 1983, 1987; Chavas, 2004). Each equation is itself derived from the first-order condition for the kth input. Hence, system (5) has four equations, as the number of variable inputs included as explanatory variables, and twelve unknowns, determined by the number of variable inputs and the number moments considered.

$$\frac{\partial \mu_1 (\boldsymbol{x})}{\partial x_k} = \theta_{1k} + \theta_{2k}\frac{\partial \mu_2 (\boldsymbol{x})}{\partial x_k} + \theta_{3k}\frac{\partial \mu_3 (\boldsymbol{x})}{\partial x_k} + u_k \qquad (5)$$

where $\theta_{jk} = -1/j!*\partial F (\boldsymbol{x}) /\partial \mu_j (\boldsymbol{x}) /\partial F (\boldsymbol{x}) /\partial \mu_1 (\boldsymbol{x})$; $(j = 1,2,3)$ and u_k is the econometric error term.

We estimate for each input k its marginal contribution to the expected profit (given by $\partial \mu_1 (\boldsymbol{x}) /\partial \boldsymbol{x}_k$, which can be expressed as a linear combination of the marginal contributions of each input to the variance ($\partial \mu_2 (\boldsymbol{x}) /\partial \boldsymbol{x}_k$) and the skewness ($\partial \mu_3 (\boldsymbol{x}) /\partial \boldsymbol{x}_k$).

The most important feature of this model is that the estimated parameters θ_{2k} and θ_{3k} are interpretable as the AP and DS risk aversion coefficients, respectively (Antle 1987; Chavas, 2004; Groom *et al.*, 2008).

More specifically, the AP absolute risk aversion coefficient is approximated by:

$$AP = -\frac{E \left(\partial^2 U/\partial \pi^2\right)}{E \left(\partial U/\partial \pi\right)} \cong -\frac{\partial F (\boldsymbol{x}) /\partial \mu_2 (\boldsymbol{x})}{\partial F (\boldsymbol{x}) /\partial \mu_1 (\boldsymbol{x})} = \boldsymbol{2\theta_2} \qquad (6)$$

whereas the DS risk coefficient is approximated by:

$$DS = \frac{E \left(\partial^3 U/\partial \pi^3\right)}{E \left(\partial U/\partial \pi\right)} \cong \frac{\partial F (\boldsymbol{x}) /\partial \mu_3 (\boldsymbol{x})}{\partial F (\boldsymbol{x}) /\partial \mu_1 (\boldsymbol{x})} = \boldsymbol{-6\theta_3} \qquad (7)$$

We recover the population average AP and DS risk aversion measures estimating system (5) through a three-stage least square (3SLS) model. We used as instruments (i) the share of farmland *not cultivated for agricultural reasons* (i.e. fallow land without subsidies) during the previous year and (ii) a dummy variable indicating if the majority of the UAA is in an area designated as 'less favoured' by the relevant regulation. Fallow land is commonly defined as a farmland ploughed but left for a period without being sown to restore its fertility and increase biodiversity. Less-favoured areas (LFA) are defined by the EC as areas where agricultural production or activity are more difficult because of *natural handicaps*, such as difficult climatic conditions or low soil productivity. These exogenous variables are assumed correlated with fertilisers, plant protection, seeds and labour choices in the current growing season

but exogenous to risk attitude. After selecting the instrumental variables, we estimated system of equation (5) exploiting interequations correlation of errors through the 3SLS procedure, as we believe that errors across the three equations of the system are correlated.[11] We acknowledge that the use of instrumental variables in a risk management analysis is still a controversial issue in empirical applications (Staiger and Stock, 1997; Heckman and Urzua, 2010; Larcker and Rusticus, 2010).[12] Our central research question pertains to the change in risk attitude through time, rather than to the estimation of the magnitude of risk attitude and their impacts on farm managerial decisions whether to adopt a specific technology such as irrigation (Koundouri, Nauges and Tzouvelekas, 2006) or on the outcomes of specific policies on input use. Groom *et al.*, 2008; for example, provide an analysis of the importance to consider risk attitude, in particular whether farmers are risk neutral or risk averse, when predicting the impact of water quotas on production decisions. For this reason, we call for caution in interpreting the magnitude of the estimated parameters.

We also bootstrap the standard errors, resampling them over individuals to obtain heteroscedasticity-robust standard errors (Cameron and Trivedi, 2010).

The risk aversion parameters are constrained such that $\hat{\theta}_{2k} = \hat{\theta}_2$ *and* $\hat{\theta}_{3k} = \hat{\theta}_3$. Although each input can affect the moments of profit in different ways, the AP and DS coefficients are related to the attitude over the moment of profit and thus not assumed to be input specific.

A positive AP coefficient indicates that the decision maker is risk averse. This implies that an agent with a positive and significant AP coefficient has the incentive to reduce its risk exposure. Any increase in the variance of profit would in fact increase the private cost of risk bearing.

A positive DS coefficient indicates that the decision maker is averse to DS risk, that is, he is averse to risk distribution towards low outcomes (such as crop failure), holding both the mean and the variance constant (Menezes, Geiss and Tressler, 1980; Antle, 1983; Di Falco and Chavas, 2009). This implies that an agent with a positive and significant DS coefficient is prone to implement management strategies that affect positively the skewness of the distribution of profits (e.g. by reducing the probability of crop failure). These considerations are related to our understanding of farmer's behaviour in a risky environment. Risk-averse farmers would be more willing to adopt strategies to reduce the variance of profit and/or exposure to DS risk.

The empirical procedure outlined in this section is implemented to estimate system of equation (5) for pooled data of relevant sub-periods and for each year separately as a robustness analysis. Stability of risk attitude implies that,

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11 We computed the correlation matrix for the fitted residuals, and used it to perform the test of independence of the errors in the three equations. The Breusch-Pagan Lagrange multiplier test for error independence indicates a statistically significant correlation between the errors in the three equations. This should be expected as the marginal contribution of fertilisers, plant protection, seeds and labour to the expected profit may have similar underlying determinants.

12 For robustness, we also obtained the risk aversion parameters using a seemingly unrelated regression model (SUR) and they were in line with those obtained using the 3SLS estimation procedure.

Table 3. Estimated risk aversion: entire period 1990–2009

Years	AP	DS
1990–2009	0.4284***(0.0111)	0.0095*** (0.0005)

Notes: Estimation method: 3SLS. Bootstrapped standard errors of the underlying estimated coefficients ($\hat{\theta}_2$ *and* $\hat{\theta}_3$) in parenthesis. Bootstrap replications: 500. Number of farms: 34,825. *** indicates that the null hypothesis of coefficients of risk aversion being zero can be rejected at the 1 per cent level. We do not include 1989 in the reference period because of the use of a lagged variable as instrument for the 3SLS estimation procedure.

in the absence of measurement error, one should obtain the same estimate of the parameter of interest when measuring an individual's risk attitude repeatedly (Schildberg-Hörisch, 2018). Based on these results on farmers' risk attitude (i.e. AP and DS coefficients), various hypotheses are tested that concern whether this attitude (i) remains stable over time and (ii) whether risk attitude is influenced by major agricultural policy regime shifts and climatic extreme events. The here considered events were presented in more detail in Section 2. Wald test is used to test for equality of these parameters across the relevant time periods. The null hypothesis is that attitude parameters across time periods are stable. We also perform several sensitivity analyses (Table A1) to compare the robustness of results when we vary marginally the length of the time under consideration.

5. Results

The AP and DS risk aversion parameters obtained estimating the system of equations (5) for the whole period under analysis, 1990–2009, are positive, equal to 0.4284 and 0.0095, respectively, and statistically significant (Table 3).[13],[14] Our results suggest that on average Italian cereal farmers have been averse to risk (profit variance) and to a lesser degree also to DS risk (negative profit skewness). This suggests that farms in our sample display some evidences of decreasing absolute risk aversion (DARA) preferences. Previous literature found that DARA preferences are a common property of farmers' risk preferences (Chavas, 2004; Zuo, Nauges and Wheeler, 2015).[15] As a robustness analysis, we also report in Table A3 the year-by-year risk aversion coefficients, obtained estimating system (5) for each cross-section (year).

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13 We provide in Table A2 the fixed-effects estimation results for the profit, variance and skewness equations, corresponding to equations 2, 3 and 4, respectively.

14 The results are robust if we exclude from the estimation of system (5) the 2 years hit by a major climate shock (the 2003 and 2007 droughts). In this case, the estimated AP and DS risk aversion parameters are 0.4291 and 0.0090, respectively, and we strongly reject the null hypothesis of these coefficients being zero.

15 A more robust conceptualisation and prediction of farmers' reactions would require to be able to assess both risk attitude and risk perception (Pennings, Wansink, and Meulenberg, 2002). This is not possible with our data. Risk perceptions reflect the farmers' assessment of the uncertainty of the risk content inherent in a particular situation, and this might also change after exposure to a climate shock or a significant policy change.

We further look at the results obtained estimating system (5) for the relevant sub-periods outlined in Table 5. These reveal that the parameters capturing risk attitude might have been unstable through time. The results presented in this section provide an indication of the general tendency in producers' risk attitude changes through time, in responses to major policy changes and exogenous climate shocks, but we caution the reader to keep in mind that real world changes in inputs will be the result of all the relevant market adjustments, as well as some technological constraints, which cannot be fully accounted for in our simple empirical set-up (Moro and Sckokai, 1999).

5.1. Policy-driven changes in risk attitude

We report in Table 4 the AP and DS risk aversion parameters for sub-periods associated to relevant policies changes. The AP coefficients change in magnitude but remain positive and statistically significant over all sub-periods analysed. This suggests risk-averse behaviour, characterised by a concave utility function. The DS coefficient for the different sub-periods are either positive and statistically significant or not statistically different than zero, indicating, on average, risk aversion or risk neutrality over DS risk. We now analyse in more detail if and how the parameters of risk aversion have changed through time, particularly in conjunction with major policy changes and major climate shocks. Previous work indicated that the impact of the changes in agricultural support on cereal farmers' risk attitudes, brought about by the application of the CAP, is ambiguous a priori (Koundouri, Laukkanen and Myyra, 2009). For this reason, we complement the estimation showing the Wald test statistic to test for the temporal stability of risk preference, using the null hypothesis of equality between the AP and the DS parameters across different time periods.

First, we analyse if and how the CAP reforms triggered a change in risk attitude of cereal farmers. We report in Table 4 the AP and DS risk aversion parameters for the periods encompassing the pre-CAP reform period, as well as for the Mac Sharry reform, the Agenda 2000 implementation period up to the 2003 climate shock, and for the Fischler reform, as outlined in Table 5. We also conducted sensitivity analysis by excluding from the estimation of system of equation (5) the final year preceding the implementation phase of a new reform.

We find cereal producers to be risk averse over the variance of profit (AP coefficient) in the pre-CAP period (1990–2002), whereas the DS risk aversion coefficients are not statistically significant, which indicate risk neutrality over the third moment of profit distribution. The latter result is easily expected in the pre Mac Sharry reform's context, where the CAP achieved income stabilisation indirectly, through price support mechanisms.

The AP coefficient increases and remains statistically significant during the Mac Sharry reform (1993–1999), and the DS risk coefficient becomes positive and statistically significant. This suggests that farmers display clearer evidences of risk-averse behaviour after the introduction of the first CAP reform

Table 4. Estimated risk aversion and Wald test statistic for equality of coefficients

Estimated Coefficients	Wald test statistic for equality across periods	
	H_0: $\hat{\theta}_2^{(t1)} = \hat{\theta}_2^{(t2)}$	H_0: $\hat{\theta}_3^{(t1)} = \hat{\theta}_3^{(t2)}$
[1] **Pre-CAP reform** (1990-1992)		
AP: 0.3400*** [0.0097]		
DS: 0.0001 [0.0002]	$\chi^2(1) = 30.72$	$\chi^2(1) = 36.99$
[2] **Mac Sharry Reform** (1993-1999)	Pr > chi2 = 0.0000	Pr > chi2 = 0.0000
AP: 0.4866*** [0.0096]		
DS: 0.0127*** [0.0003]		
[2a] **Transition phase** (1993-1995)		
AP: 0.4489*** [0.0232]		
DS: 0.0082* [0.0007]	$\chi^2(1) = 21.36$	$\chi^2(1) = 31.26$
[2b] **Implementation phase** (1996-99)	Pr > chi2 = 0.0000	Pr > chi2 = 0.0000
AP: 0.4111*** [0.0113]		
DS: 0.0057*** [0.0003]		
[3] **Agenda 2000 pre mid-term review & 2003 climate shock** (2000-2002)	$\chi^2(1) = 7.38$	$\chi^2(1) = 4.93$
AP: 0.4059*** [0.0149]	Pr > chi2 = 0.0066	Pr > chi2 = 0.0263
DS: 0.0076** [0.0006]		
[4] **Fischler reform implementation phase (up to 2007 climate shock)** (2005-2006)[16]	$\chi^2(1) = 3.04$	$\chi^2(1) = 0.21$
AP: 0.4044*** [0.0222]	Pr > chi2 = 0.0815	Pr > chi2 = 0.1357
DS: 0.0074 [0.0009]		

Notes: Bootstrapped standard errors of the underlying estimated coefficients ($\hat{\theta}_2$ and $\hat{\theta}_3$) in parenthesis. Bootstrap replications: 500. ***, ** and * denote significance at 1, 5 and 10 per cent, respectively. The suffix t1 and t2 indicate the periods compared in the Wald test statistics as shown in each row. Dotted brackets indicate that we reject the null hypothesis of stable risk attitude for both AP and DS coefficients at 1, 5 or 10 per cent level. If only one bracket is drawn across two periods, we reject/accept the hypothesis for both AP and DS coefficients.

in the early 1990s. The Wald test for the stability of the AP the DS parameters between the pre-CAP reform period and the Mac Sharry reform period strongly reject the hypothesis of stable risk attitude through the decade 1990–1999. We also strongly reject the null hypothesis of equality of estimated AP and DS risk parameters between the first triennium of the Mac Sharry period (1993–1995), which roughly corresponds to the formal *transitional phase* of the reform, compared with the *implement*ation phase of this reform (1996–1999). This suggests that farmers incur an implicit higher cost of risk bearing during the *transition phase*, possibly due to the uncertainties brought by the introduction of the CAP reform, the GATT's entrance into force and the EU enlargement

Table 5. Estimated risk aversion and Wald test statistic for equality of coefficients

Estimated Coefficients	Wald test statistic for equality across periods	
	$H_0: \hat{\theta}_2^{(t1)} = \hat{\theta}_2^{(t2)}$	$H_0: \hat{\theta}_3^{(t1)} = \hat{\theta}_3^{(t2)}$
[3] Agenda 2000 pre-shock (2000-2002)		
AP: 0.4059*** [0.0149]		
DS: 0.0076** [0.0006]	$\chi^2(1) = 7.38$	$\chi^2(1) = 4.93$
[4] post 2003 shock (2004-2006)	Pr > chi2 = 0.0066	Pr > chi2 = 0.0263
AP: 0.3202*** [0.0127]		
DS: 0.0011 [0.0005]		
[4] Fischler reform implementation phase (up to 2007 climate shock) (2005-2006)	$\chi^2(1) = 3.95$	$\chi^2(1) = 2.58$
AP: 0.4044*** [0.0222]	Pr > chi2 = 0.0468	Pr > chi2 = 0.1082
DS: 0.0074 [0.0009]		
[5] post 2007 shock (2008-2009)		
AP: 0.3200*** [0.0192]	$\chi^2(1) = 9.76$	$\chi^2(1) = 5.27$
DS: 0.0051 [0.0010]	Pr > chi2 = 0.0018	Pr > chi2 = 0.0217

Notes: Bootstrapped standard errors of the underlying estimated coefficients ($\hat{\theta}_2$ *and* $\hat{\theta}_3$) in parenthesis. Bootstrap replications: 500. *** and ** denote significance at 1 and 5 per cent, respectively. The suffix t1 and t2 indicate the periods compared in the Wald test statistics as shown in each row. Dotted brackets indicate that we reject the null hypothesis of stable risk attitude for both AP and DS coefficients at 1 or 5 per cent level. If only one bracket is drawn across two periods, we reject/accept the null hypothesis for both AP and DS coefficients.

by three new member States. During the implementation phase in of the Mac Sharry reform (1996–1999), farmers risk aversion gradually decreased again.[16]

These results suggest the existence of two opposite effects. First, it is frequently difficult to foresee changes in government policies, particularly where decisions are influenced by social and political considerations, and this situation may trigger an increase of farmers' risk aversion (REAS, 2010). Moschini and Hennessy (2001) argue that changes in policy interventions, such as the CAP reforms in the 1990s, can become sources of uncertainty that can create risk for agricultural investment (Gardner, 2002, for example from US agricultural policies). This might happen despite the fact that policy interventions in the agricultural sector are often intended to reduce the level of risk faced by farmers. This consideration brings us to the second possible effect: the policy reforms reduced the random component of farmers' income, which could lead to a decrease in farmers' risk aversion compared with the pre-CAP reform period, as discussed in (Koundouri, Laukkanen and Myyra, 2009).

16 These results are robust if we exclude from the estimation of (5) the year preceding a reform, that is, 1992 from the pre-CAP reform estimation period and 1999 in the Mac Sharry reform estimation period. We present these sensitive analyses in Supplementary Table A4 in the Appendix.

This effect seems to dominate in the implementation phase of the Mac Sharry reform as well as during the implementation of the Fischler reform. During the latter, the switch to decoupling, with the introduction of the single farm payment, led to farmers being less risk averse over the variance of distribution of profit, and become risk neutral over DS risk. This is coherent with the fact that Italian farmers participation rates to insurances programs is exceptionally low (Mahul and Stutley, 2010; Di Falco *et al.*, 2014; Santeramo *et al.*, 2016) and it is also in line with the findings of studies estimating farmers' risk attitude in other European countries (Gardebroek, 2006; Groom *et al.*, 2008).

5.2. Changes in risk attitude after the climate shocks

We report in Table 4 the estimated risk parameters for the post 2003 and 2007 droughts sub-periods.

The results of Table 4 offer further evidences of instability of risk aversion through time. We perform the Wald test for equality of risk parameters before and after the droughts, within relevant sub-periods identified through the analysis presented in Table 5. The null hypothesis of equality of AP risk aversion parameters before and after the 2003 and 2007 climate shocks is strongly rejected. The AP risk aversion parameter slightly decreases after the 2003 and 2007 droughts. The null hypothesis of equality of DS risk aversion parameters before and after the 2003 and 2007 climate shocks is rejected after the 2003 but not after the 2007 drought.

As discussed, the 2003 drought hit the Italian economy the same year of the midterm review of the Agenda 2000. Hence, these results might indicate that the risk aversion mitigating effects triggered by the described policy changes dominated a possible increase in risk aversion, which could be expected after farmers experience a major drought.

We warn the reader that this is not the only confounding effect, for example while the Italian agricultural sector was counting the damages caused by the 2007 drought, the whole Italian economy was put under stress by the onset of the 2008 financial crisis. Disentangling these multiple effects is very difficult or impossible, especially through econometric applications on observed economic behaviour from secondary data, as the one adopted in this article. This makes it impossible to draw conclusive comparison on the impacts that climate shocks may have on risk attitude compared with those triggered by major policy changes. The result that the DS risk aversion parameters remain very small or not statistically significant is interesting and deserves further investigation, because of its implication on farmers' vulnerability and need for public support to cope with climate shocks, which can become more frequent under climate-change scenario.

5.3. Concluding remarks

Agents' risk attitude, along with risk perception and risk exposure, are crucial for explaining their production and investment decisions. The question

whether structural parameters of human behaviour such as risk aversion and risk perceptions are stable through time become popular in macroeconomic modeling since the publication of the classic Lucas critique (Lucas, 1976). The question is also debated in studies that focus on the agricultural sector, which is particularly exposed to climate-related shocks and policy changes (Louhichi *et al.*, 2010; Lehmann, Briner and Finger, 2013).

This article adds to the existing literature, as we attempt to identify the impact of a range of diverse shocks on risk attitude. Using a rich panel data structure, we capture the time dimension in risk attitude's evolution. Notably, we analyse both the impact of policy changes and production shocks driven by droughts, recovering risk attitude looking at agents' revealed decision through a long period of time. We find evidence for temporal instability of risk attitude using the example of Italian agriculture over the period 1990–2009. These changes in risk attitude can be associated to major changes in agricultural policy regimes and to some extent to shocks in climate conditions.

More specifically, we find temporal instability for farmers' aversion to the variance of profit, the AP coefficient of risk aversion. In contrast, farmers' aversion to extreme (DS) events, expressed with the skewness of profits, exhibits more stability through time.

Other authors provided general insights that the changes in policy interventions, such as the CAP reforms in the 1990s, can become sources of uncertainty that can create considerable risk for agricultural investment (Moschini and Hennessy, 2001). In general, our analysis indicates that risk aversion coefficients tend to increase at each introduction of a policy change. We suggest that the build-up of progressive uncertainty through the complex system of government interventions that characterises the European policy framework of the late 1990s and early 2000 led at first to an increase of risk aversion. Instead, policies that aim to stabilise farmers' income and that are implemented for long enough without undergoing a continued reform process tend to trigger more risk neutral or loving behaviours, a result also found in other studies (Koundouri, Laukkanen and Myyra, 2009). As new policies are introduced, farmers tend to become more risk averse, but this effect fades through the implementation period of the policy, and eventually the mitigating effect on risk aversion of the policy dominates. This factor seems to be particularly strong after the introduction of the Fischler reform, with the switch to the single farm payment, decoupling a large share of CAP support from production from 2005 onwards. It may also have mitigated the increases in risk aversion expected after the 2003 and 2007 droughts. More specifically, we do not find large effects of these drought events on farmer risk attitudes. This contrasts findings in non-agricultural domains (Eckel, El-Gamal and Wilson, 2009, Page, Savage and Torgler, 2014, Hanaoka, Shigeoka and Watanabe, 2018) but may be also due to the confounding occurrence of climate and policy shocks.

In our context, this means that key behavioural parameters such as risk aversion coefficients may change through time in relationship to significant changes in policies, and we shall take this into account if we are going

to be concerned about the policy implications of our models or about our ability to forecast when there are changes in policy regimes, such as the CAP reforms. Since the publication of the seminal article by Lucas (1976), several studies looking at investors' behaviours also debated whether agents react differently to policy shocks perceived as permanent or transitory. Our findings suggest the need for such better understanding also in the specific context of decision-making in the agricultural sector.

These findings open up future avenues of research, as it is often assumed that individual risk aversion is time invariant. The instability of risk attitude is potentially a serious barrier to our understanding of farmers' decision-making under uncertainty, particularly relevant also in the context of risk exposure in agriculture.

In this article, we use the method of moments on a much larger panel data set of those used in previous similar studies. A limitation of this study is that we can only recover risk attitude from past farm-level accounting data, but we cannot provide a full account on how risk preferences changed through time. To analyse the latter, we would need to obtain separately risk attitude and risk perception, but unfortunately our data set does not allow us to recover information on farmers' risk perceptions in the past. This, in turn, would allow a more robust conceptualisation and prediction of farmers' reactions and resilience to shocks (Meuwissen *et al.*, 2019; Pennings, Wansink and Meulenberg, 2002). To date, there is only a limited number of studies measuring risk preferences with the same individuals' multiple times over time (Mata *et al.*, 2018). Yet, these are virtually inexistent for farmers. Future research should account for this shortcoming. A recent article by Schildberg-Hörisch (2018) stressed that once we empirically show the possibility of systematic change in risk preferences, an array of fundamental questions arises, for example on how can we evaluate alternative policy options or perform welfare analyses when individuals' preferences lack complete (time) stability. European agriculture is increasingly exposed to climate change and possibly significant policy changes, and the support of risk management is of increasing relevance in recent CAP reforms (El Benni, Finger and Meuwissen, 2016). For these reasons, researchers and policymakers alike need to enhance their understanding of the implication of risk preferences that are unstable over time.

Despite the discussed potential issues related to the challenge of identifying separately (production) risk and risk preferences from behavioural equations (i.e. first-order conditions on input choices) on observed data (Lence, 2009; Just and Just, 2011), this method provides an empirical application that allows retrieving agents' attitude for risk even when survey data with risk preferences elicitation nor the results of *ad hoc* field and lab experiments are available. Our approach does not replace the methods based on primary data. On the contrary, such methods shall provide further tests for the robustness of our findings and enhance understanding of time-invariant risk aversion (and risk preferences) in agriculture. Such methodological comparison is beyond the scope of this article and left for future research.

Supplementary data

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