PEF, **13** (4): 389–419, October, 2014. © Cambridge University Press 2014
doi:10.1017/S1474747213000395 First published online 17 January 2014

*The effect of age on portfolio choices: evidence from an Italian pension fund**

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Abstract

According to optimal portfolio theories, investors should reduce their exposure to stock market risk as they grow old. Indeed, older workers, with only a few years left before retirement, are particularly vulnerable to unexpected falls in stock prices. Despite the theoretical and – as shown by the recent financial crisis – policy relevance of the issue, empirical evidence on this topic has been scant and inconclusive. The aim of the present paper is to assess the effect of age on portfolio choices, using a new panel dataset from an Italian defined-contribution pension plan. We find that on average holdings of risky assets do indeed significantly decrease with age. However, the effect is non-linear, being much stronger in the last part of one's career. Moreover, we also document that inertial behaviour is quite widespread, and can be very costly. Results are confirmed when we control for individual fixed effects and cohort effects.

JEL CODES: G11, G23

Keywords: Pension funds, portfolio choice.

1 Introduction

In recent years, many countries have reformed their public pension system, tightening the eligibility rules and reducing the generosity of benefits (Feldstein and Siebert, 2002). Partly as a result, private pension plans have grown in terms both of assets

* The views expressed in this paper are those of the authors and do not necessarily reflect the position of the Bank of Italy. We thank Giorgio Gobbi, Sandro Momigliano, Valentina Nigro, Domenico Depalo, two anonymous referees, and seminar participants at the Bank of Italy, the 10th Anniversary Conference of CERP (Turin, September 2009), the Netspar Pension Workshop (Amsterdam, February 2010) and LUISS (Rome, April 2010) for useful comments.

under management and of number of participants (OECD, 2009), becoming increasingly relevant as a means to provide adequate retirement income.

Contrary to Social Security schemes, private pension investment requires the worker to make several choices. She has to decide whether and how much to contribute, select the most appropriate portfolio, and possibly decide when to withdraw. These decisions are even more difficult in a time of financial turmoil, when both the probability and the cost of errors are magnified. In order to design policies, which help workers to get the most out of their pension investments, it is important to understand how their behaviour is systematically affected by individual characteristics such as age, sex, financial education and income.

In particular, the aim of the present paper is to assess the effect of age on portfolio choices, using a new panel dataset from a defined-contribution (DC) pension plan for employees of a medium-size Italian bank.

Recent optimal portfolio theories suggest that investors should reduce their exposure to stock market risk as they get near to retirement (see, e.g., Campbell and Viceira, 2002), and failure to do so can be quite costly, especially for elderly workers. They might find themselves overexposed to stock market risk, with a few years left before retirement to recover from an unexpected fall in stock prices.

Despite the theoretical and policy relevance of the issue, very few papers have studied the relationship between age and portfolio choices in a panel data framework, reaching different conclusions.

In particular, Agnew *et al.* (2003) use a four-year panel of participants in a large 401(k) plan. They include age and time effects in their regression specification together with demographic variables. They find a statistically significant and economically sizable negative relationship between age and equity holdings: each extra year translates into a 93 basis points reduction in the share of stocks in the average portfolio. On the contrary, Ameriks and Zeldes (2004), using a similar specification[1] and a very long panel dataset from TIAA-CREF (the large US pension plan, open to public sector teachers and university professors), find a flat age-equity profile[2].

While we see both contributions as important steps forward, we believe that some characteristics shared by both data sets make the interpretation of their results problematic.

First, the plans they study offer a wide array of funds, with the investors being allowed to choose whatever any combination of funds. Most of them are pure equity funds, whereas some are bond funds (in the case of Agnew *et al.* (2003), there are also some pre-mixed 'balanced' portfolios). As a growing body of evidence suggests (see e.g., Benartzi and Thaler, 2001, 2002, 2007; Choi *et al.*, 2006; Huberman and Jiang, 2006), investors allocate their wealth by using simple rules of thumb, often

[1] However, their regressions do not include any individual-specific, time-invariant characteristic.

[2] There are also a few other studies that investigate the question in a cross-section context. Even in this earlier literature, results are inconclusive (for a survey of this literature, see Guiso *et al.*, 2003; Ameriks and Zeldes 2004). Another interesting approach, in which the unit of analysis is not the individual investor but the pension fund, is pursued by Bikker *et al.* (forthcoming). They find that, in a large cross-section of Duch pension funds, an increase of one-year in the age of the average active participant reduces the fund equity exposure by 0.5%.

resulting in a suboptimal portfolio. For example, Benartzi and Thaler find that many workers simply allocate an equal fraction of their wealth to each available investment option: in the Agnew *et al.* (2003) setup, the observed strong effect of age on portfolio decisions could be due to investors *a' la* Benartzi and Thaler (2001) trying to correct their initial excessive stock market exposure.

In our setting, instead, plan participants can choose just one among five 'pre-mixed' funds, unambiguously ranked in terms of their risk profile.[3] Most importantly, at any given moment all the investor's wealth must be invested in a single fund. Moreover, in our setup the set of alternatives has remained unchanged during the whole sample period. In the Ameriks and Zeldes (2004) study, instead, the set of funds offered by the plan changes over time, with new equity funds becoming available. Poterba *et al.* (2007) show that such changes strongly influence investors' behaviour, and the lack of age effect in Ameriks and Zeldes (2004) might be due to the workers being enticed by the new riskier (and fancier) options.

Second, in the plans considered by the two previous studies workers have an almost unlimited possibility to change their allocation at any moment. Hence, it is not clear whether the observed portfolio reallocation is driven by long-term risk-rebalancing considerations, as the ones we are interested in, or by short-term 'return chasing' (the latter is documented, among others, by Odean, 1999 and Barber and Odean, 2001*a*, *b*). In our setting, individuals are instead allowed to change their allocation only once a year, in which case they are obliged to move all the previously accumulated wealth, not only the new contributions.[4]

It is important to remark that this very simple structure, while different from the one common in other countries (e.g., the USA) is not peculiar to our plan. It was imposed until recently by the Italian law to *all* the Italian pension plans (these restrictions were lifted only in 2007, and even today most pension plans stick to the old set of rules).[5]

Third, our data go from 2002 up to December 2008, one year after the beginning of a sharp and disorderly drop in share prices, allowing us to observe, at least to a certain extent, how investors reacted to such event.

Finally, from a methodological point of view, we apply recent econometric techniques that allow us to account for individual fixed effects and cohort effects, which are instead not considered in earlier contributions.

Of course, our empirical exercise also comes with some limitations. We use administrative data: while administrative records are about actual choices, thereby reducing the risk of measurement errors, they also contain relatively few variables for each individual. In particular, a limitation of our dataset – which is shared with the

[3] This set up, which is typical in Italian employer-sponsored pension plans, is also common in other countries. For example, mandatory individual accounts systems in Chile and other Latin American countries allow workers to choose among a limited number of 'lifestyle' funds. The same is true for the mandatory systems of Central and Eastern European countries. Other countries (e.g., Sweden and Australia) allow a much wider variety of choices (Tapia and Yermo, 2007).

[4] The switch has to take place at a pre-specified date, which is the same for all participants.

[5] Of course, we do not mean that these characteristics of our set-up represent an unambiguous advantage. Different kinds of data are more useful for answering other questions (for example the determinants and consequences of higher frequency trading).

above-mentioned studies – is the absence of information on non-retirement wealth. However, differences in unobserved wealth are unlikely to explain the observed age-effects: first, in the data for the Italian population at large there is no increasing relationship between age and the fraction of wealth invested in shares; second, prior literature does not find strong effects of wealth on the share of risky assets (conditional on the fact that some wealth was initially invested in shares: see Sahm, 2007; Brunnermeier and Nagel, 2008); third, our use of individual fixed effects should limit the omitted variable bias due to the unobservability of wealth.

Another limitation of our sample is that it is definitely not representative of the whole Italian population. Nevertheless, as it is made up of a group of agents characterized by a high degree of financial education (they are mostly clerical and managerial workers in the banking sector), one could at least argue that any deviation from optimizing/rational behaviour observed in our sample should be even more pronounced at the population level.

As it is well known, there is an inherent difficulty in simultaneously identifying cohort, time and age effects without imposing some a priori restrictions. In our baseline estimates, also for the sake of comparability, we follow Agnew *et al.* (2003) and assume the absence of cohort effects.[6] However, we also perform robustness checks in which cohort effects are included in the model using the identification strategy proposed by the recent paper by Malmendier and Nagel (2011).

To give a quick preview of our main results, we find that, contrary to Ameriks and Zeldes (2004), age induces investors to sensibly reduce their exposure to equities, broadly in line with the prescription of optimal portfolio theory. However, differently from Agnew *et al.* (2003) the equity share starts to decline quite late in one's career, dropping quickly in the final years prior to retirement. We also find that some workers never change their asset allocation and more generally there is a significant tendency to stick to the previous allocation. While we cannot exclude that this inertia might be rational, behavioural factors are also likely to play a role (the role of inertia has been emphasized, among others, by Madrian and Shea, 2001; Papke, 2004; Mitchell *et al.*, 2006; Bilias *et al.*, 2010).

The rest of the paper is organized as follows. In Section 2, we provide a brief outline of the Italian pension system, which can be helpful to put our results in perspective. In Section 3, we describe the structure of the pension plan under examination. In Section 4, we outline the characteristics of our dataset and present some summary statistics concerning investment choices and fund performance. In Section 5, we study the portfolio choices of the workers, and in particular the impact of age, controlling for several other possible determinants. In Section 6, we present several extensions and robustness checks. In Section 7, we draw some tentative conclusions and policy implications, and point to some avenues for possible further research.

[6] Ameriks and Zeldes (2004) present two sets of estimates: in the first one, they also assume no cohort effect; in the second, they allow for cohort effects but assume no age effect.

2 A short overview of the Italian pension system

Retirement income in Italy mainly comes from the public pay-as-you-go pension system, which is in turn based on two main schemes. First, there is a relatively small non-contributory scheme, granting a minimum benefit to any person with at least 65 years and with income below a given threshold. The benefit is linearly decreasing with income and becomes zero for income levels at or above the threshold, which was equal to 430 euros (€) per month in 2008.[7] In the second scheme, which is contribution-based, the right to get a pension is conditional to a minimum amount of contributions and/or a minimum eligibility age. The size of the benefit increases with the amount of contributions paid by the worker during the career. In 2008, a worker could qualify for a contributory pension with at least 65 years of age (60 years for women) and 20 years of contribution[8] and the average monthly benefit was about 1,000 €.[9]

Besides State-provided pensions, there are several private pension plans.[10] Enrolment in these plans is on a voluntary base,[11] even if there are fiscal incentives for those joining, and in the case of employer-sponsored plans most employers grant matching contributions.

The public pay-as-you-go system has been reformed many times in the past, starting from the early nineties, reducing significantly its generosity (see, e.g., Franco, 2002 and Franco and Sartor, 2006). Furthermore, after a reform passed in 1995, a significant fraction of the pension benefits of the contributory scheme depends not only on the amounts contributed by the worker during his active life, but also on GDP growth (positively) and on increases in longevity (negatively).[12] This means that public pensions cannot be considered as a riskless investment vehicle. Not only they are subject to 'political risk' (e.g., in 1992, in the midst of a difficult fiscal and exchange-rate crisis, a major pension reform cut public pension wealth of Italian workers by an estimated 30% almost overnight), but also to considerable macroeconomic and demographic risks.

Given the reduced generosity and the higher riskiness of public pensions, it is not surprising that assets and enrolment rates of private pension funds, have been

[7] The threshold itself is indexed to inflation (as of 2012, it is equal to 464 €).

[8] Alternatively, he/she needed at least 40 years of contributions, or 58 years of age (59 for the self-employed) and 35 years of contributions. Overall, about 90% of old-age public pensions are paid under the contributory scheme.

[9] The computation of benefits in the contribution-based pillar is quite complex. However, focusing on those with more than 15 years of contributions in 1992, which have constituted in the past (and will still constitute for several years to come) the vast majority of those entering retirement, in 2008 pension benefits were computed as a fraction of the average of their wages of the last ten working years. In particular, the pension/wage ratio was proportional to the years of contributions, reaching a maximum of 80% with 40 years of contributions.

[10] Differently from what happens in other countries, Italian pension funds are only concerned with the accumulation phase, and do not provide annuity products. At retirement, workers can use their pension wealth to buy an annuity from an insurance company.

[11] Differently from what happens, for example, in the UK, workers cannot opt-out from the public pension system to join a private plan.

[12] This kind of pay-as-you-go schemes is often called 'notional defined contribution' as they mimic quite closely the functioning of a funded defined-benefit pension plan (see the contributions in Holzmann *et al.* (2012), for a thorough discussion).

constantly increasing, even if they are still below those in other advanced countries. At present, assets under management of Italian private pension plans amount to 5.7% of GDP, and 28.9% of private sector workers are enrolled.

3 The dataset

We draw our data from an Italian DC pension plan. Our dataset includes information on yearly individual investment choices for all 3,820 retirement accounts – outstanding for at least 1 year – from December 2002 to December 2008, for a total of 20,505 year-investor data points. The plan is sponsored by a medium-size Italian bank operating mainly in Northern Italy and it is open to all the bank's employees, and it was set up in 2001.[13] At the end of 2008, the plan covered about 97% of the workforce. The plan does not envision default options concerning the decision to enrol or the fund chosen upon enrolment. Employees have the option to join the plan only at the moment in which they join the firm. They can leave the plan in any moment, but the fraction of workers that leaves the plan before retirement is negligible.[14] Upon enrolment, participants choose one of the five funds, characterized by different asset allocations. Once a year, at the end of November, participants can change the fund and the level of their monthly contributions. They receive a letter that reminds them of the deadline; an advisory service (internet and telephone-based) is active throughout the year, helping participants to self-assess their risk preferences and to choose the appropriate fund. There is no monetary cost of switching. If they choose to switch, the change is effective from January 1 of the following year. Participants can choose only one fund among those offered by the plan; that is, they cannot split their accumulated wealth among more funds. When a participant chooses to switch, her entire wealth is disinvested from the previous fund and moved into the new one. Our dataset includes information on yearly individual choices and on demographic and employment characteristics, such as gender, age, marital status, position and seniority of service. As is often the case with administrative data, our dataset contains no information on non-retirement wealth.

The plan offers five funds: guaranteed returns, money-market, bond, balanced and equity. The guaranteed returns fund is managed through insurance products. Each of the other four funds fund has a target asset allocation, which the portfolio manager maintains during the year, rebalancing the portfolio when necessary. The money-market fund is invested in euro-denominated money market instruments (at least 80%) and other debt securities (up to 20%); the bond fund is invested in euro-denominated money market instruments (up to 20%) and other debt securities (at least 80%); the balanced fund is invested in money market instruments (up to 20%), other debt securities (up to 80%), and equities (up to 40%); the equity

[13] The plan sponsor contributes a fraction equal to 3.25% of the salary for each contributing worker. The amount of this contribution does not vary with the amount of the worker's contribution.

[14] Moreover, the amount of workers leaving the firm before retirement during our sample period is also very small as it is often the case, in Italy, for big, established enterprises.

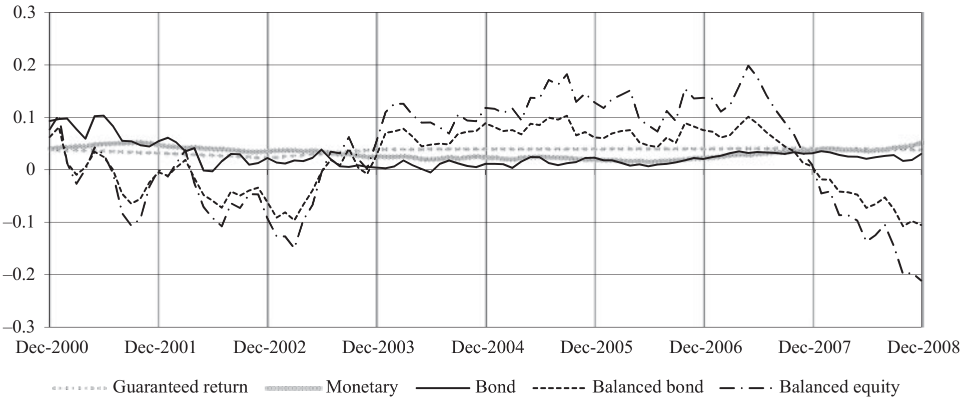

Figure 1. Fund performance net of management fees (annualized monthly returns).

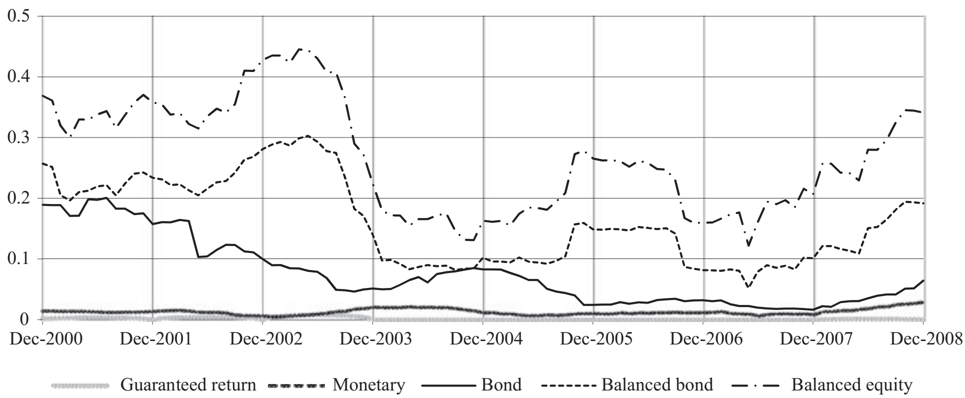

Figure 2. Standard deviation of annualized monthly fund returns.

fund is invested in money market instruments (up to 20%), other debt securities (up to 50%), and equities (up to 70%). The precise asset allocation of each fund in a given year is communicated to participants every year before they can choose their fund.[15] Each fund's return and that of its benchmark are published on a monthly basis.

As it is apparent from Figures 1 and 2, the funds can be naturally ordered according to their returns' volatility. The guarantee and monetary funds are by far the least volatile (with an average standard deviation respectively of 0.2 and 1.3% over the sample period), followed by the bond fund (7.5%). The two balanced funds exhibit a much higher volatility (16.1% for the balanced-bond and 26.8% for the balanced-equity).

[15] The guaranteed fund was introduced at the beginning of 2002, a year after the start of the plan. The exposure to equity of the balanced and equity funds has been slightly increased at the end of 2004, respectively, from 20% to 30% and from 40% to 60%. No other changes in the option offered have taken place during the sample period.

4 Summary statistics

4.1 Participants' characteristics

In Table 1, we present some statistics on the demographic characteristics of plan participants (information on salary, marital status and job position, as of December 2008) and compare them with those of Italian private sector workers at large, taken from the 2008 of the Bank of Italy survey on household income and wealth (SHIW).[16] Our sample differs from the Italian population in several respects.

Workers in our sample have, on average, higher earnings than private sector workers in general and a higher level of education (94% have completed high school or college, compared with 44% of private sector employees). They are almost all clerical or managerial workers (98% of the total); mostly male (68%); relatively young (24% are <30 years old) and with relatively short job tenure (43% have <5 years of tenure). About 40% of the participants have been in the sample for all the 7 years.

We can rely on SHIW data also for further information concerning Italian financial sector workers (Table 2). On average, they earn a net salary of about 41,000 € and have a financial wealth of about 53,000 €, of which 18,000 invested in private pension funds. Most of them are homeowners (with a real estate of about 370,000 €). Both earnings and wealth increase with age. Even if they expect most of their retirement income to come from the public pillar (replacement rate from this pillar is about 70%), the role of private pension funds is not negligible at all (with an expected replacement rate of about 20%), especially for younger workers.[17]

4.2 Investment choices

At the end of December 2008, 30% of plan participants had their wealth invested in the riskiest portfolio; 36% in the balanced one, 34% in the three remaining portfolios (Table 3).[18] Through time, there has been a shift in the relative importance of the two riskiest portfolios, which are the only ones which invest in shares: in 2002 they were chosen by 75% of participants, in 2008 this proportion drops to 65%. This is probably related to the disappointing stock market performance during the observed period.

Switches only account for about 9% of all the investor-year observations: most participants confirm their previous portfolio choices most of the time (Table 4).[19] However, during our 7-years period, 25% of the 3,820 individuals observed

[16] The survey provides a representative sample of the Italian population. More information is available in Bank of Italy (2008).

[17] The replacement ratio is defined as the ratio of the last salary to the first pension benefit. Cappelletti and Guazzarotti (2010) show that individuals in the SHIW survey compute their future replacement rate quite accurately. These figures also track quite closely the official projections (see, e.g., those included in Ministry of Economy and Finance 2012).

[18] At the end of 2008 the total wealth accumulated in the fund amounted to 108 million.

[19] This may be due, at least partly, to the fact that the intention to shift has to be notified to the fund while the choice to remain in the same line is made tacitly.

switched at least once. The percentage rises to 48% among those that joined the plan from the start.

Male and female workers do not differ much in their portfolio choices, even though females switch slightly less than males (8.5% and 9.9%, respectively).[20] With respect to education, the main difference is between the least educated group (people with only a primary school certificate) and the others. Indeed, less than 60% of the former invest in shares, compared with more than 70% in the other groups. More educated switchers are more likely to switch towards more risky funds than less educated ones. There are no clear patterns with regard to job position.

Sizable differences are apparent across age groups. In particular, while the share of workers who choose the two funds exposed to stock market risk is above 75% for those younger than 50, it drops to about 50% for those over 50. Moreover, the propensity to switch is higher for older workers, and in particular the elderly are relatively much more likely to switch towards less risky funds.

Finally, the average plan balance amounted to 32.600 € (with a median of about 20,000 €; Table 5). Not surprisingly, it increases monotonically with age, job tenure, job position and salary. Less intuitively, it decreases with education: this is due to the fact that older cohorts (which have more money in the plan) also have a lower education.

4.3 Performance

Looking at monthly annualized returns from 2002 to 2008, we can notice that our sample is characterized by two periods of low returns and high volatility in stock markets. The first started at end-2001 and lasted until mid-2003 and the second started in the summer of 2007, with the recent financial turmoil (Figures 1 and 2).

In particular, in 2008 the annual return of the balanced equity fund was equal to −20% while that of the balanced bond fund was −10%. Investing in one of these portfolios would have implied a severe loss in investors' plan balances, especially harmful for older workers, given their shorter investment horizon. In this section, we try to evaluate the effects of the decision to change fund on realized returns.

First, we look at returns in the year following a switch. In the short term, changing fund has been profitable, allowing the investor to gain on average more than 1% with respect to a passive conduct.

As one-period gains or losses are more important for workers approaching retirement, which do not have the option to wait for market values to recover, we made separate computations for older investors. Workers older than 50 years who changed asset allocation at least once earned on average a return 2.9% higher than those who did not. Moreover, in 2008 older workers who switched fund avoided considerable losses which amounted on average to 25% of their plan balance, i.e., more than 22,000 €.[21]

While looking at one-period-ahead returns might be a sensible approximation for older workers, this is not true for younger ones, who have a longer investment

[20] On gender differences in portfolio choices, see Barber and Odean (2001*a*).

[21] As we remarked above older participants tend to switch to safer funds in case they switch.

Table 1. *Statistics on plan participants (number of workers and percentages)*

	Statistics on all plan participants		Statistics on participants in the panel		Survey statistics on private sector employees	
	Number	%	Number	%	Private sector (%)	Financial sector (%)
Gender						
Female	1,216	31.8	496	30.5	38.7	49.4
Male	2,604	68.2	1130	69.5	61.3	50.6
Age						
Under 30	920	24.1	55	3.4	22.1	14.4
31–40	1,232	32.3	563	34.6	32.4	28.3
41–50	956	25.0	592	36.4	30.5	34.5
Over 50	712	18.6	416	25.6	15.0	22.7
Marital status						
Unmarried	1,517	39.7	405	24.9	59.6	61.8
Married	1,881	49.2	1100	67.7	32.6	26.4
No longer married	148	3.9	100	6.2	7.9	11.8
Unknown	274	7.2	21	1.3	–	–
Education						
Elementary and middle school	188	4.9	117	7.2	56.3	12.5
High school	2,008	52.6	977	60.1	35.3	60.6
University	1,572	41.2	530	32.6	8.4	26.9
Unknown	52	1.4	119	7.3	–	–
Job position						
Blue collar	76	2.0	18	1.1	63.6	0.3
White collar	2,450	64.1	906	55.7	30.6	76.5
Middle management	1,221	32.0	651	40.0	4.0	18.1
Senior management	73	1.9	51	3.1	1.6	5.1

Table 1 (*cont.*)

	Statistics on all plan participants		Statistics on participants in the panel		Survey statistics on private sector employees	
	Number	%	Number	%	Private sector (%)	Financial sector (%)
Salary in 2008 (thousands of €)						
Up to 25	188	4.9	83	0.0	54.4	22.8
25–35	1,793	46.9	505	1.1	25.6	22.9
35–45	774	20.3	500	55.7	11.3	24.7
45–55	434	11.4	257	40.0	5.8	11.7
55+	631	16.5	281	3.1	3.0	17.9
Tenure (years)						
Less than 5	1,635	42.8	1	5.1	–	–
5–14	1047	27.4	723	31.1	–	–
15–24	449	11.8	369	30.8	–	–
25–34	562	14.7	479	15.8	–	–
35+	127	3.3	54	17.3	–	–
Entry-exit						
Enrolled for 8 years (panel)	1,626	42.6	–	–	–	–
Enter late	1,922	50.3	–	–	–	–
Exit before December 2008	476	12.5	–	–	–	–
Enter late and exit early	130	3.4	–	–	–	–
Unknown	18	0.5	–	–	–	–
Total	3,820	100	1626	100	–	–

Table 2. *Income and wealth composition for workers in the Italian financial sector (Euro and percentages)*

	Financial sector employees (averages)						For comparison: Private sector employees (averages)					
		Financial assets			Expected replacement rate from:			Financial assets			Expected replacement rate from:	
	Salary		*of which: Private pension plan*	Real assets	Public pension	Pension funds	Salary		*of which: Private pension plan*	Real assets	Public pension	Pension funds
Age												
Under 30	23,519	17,578	*3,000*	161,231	–	–	17,694	8,179	*1,070*	70,395	63	16
31–50	41,203	49,903	*14,471*	402,007	65	22	24,337	17,397	*3,202*	138,870	65	15
Over 50	46,418	62,982	*23,031*	360,301	75	18	27,625	43,439	*9,269*	213,070	69	11
Job position												
Blue collar	–	–	–	–	–	–	20,193	9,265	*1,312*	98,268	64	13
White collar	33,927	39,829	*15,290*	359,703	68	19	29,252	34,087	*6,620*	216,903	68	15
Management	56,656	78,238	*21,397*	396,192	70	24	64,816	133,059	*7,424*	468,512	69	17
Total	41,464	53,084	*17,832*	371,804	69	20	24,263	21,494	*4,205*	145,899	65	14

Source: SHIW 2008.

Table 3. *Statistics on choices among funds (number of observations and percentages)*

	Statistics on fund choices					
	Composition of observations by fund (per cent)					
	Guaranteed return	Monetary	Bond	Balanced	Equity	Observations
Total	1,352	1,760	2,154	8,014	6,739	20,019
Year						
2001	0.0	6.1	4.4	10.7	10.7	1,781
2002	3.0	8.6	13.6	9.7	10.3	1,962
2003	5.7	9.9	15.2	9.5	10.3	2,040
2004	10.7	9.9	13.0	11.2	10.8	2,223
2005	13.8	8.5	10.7	12.0	12.6	2,376
2006	19.9	9.9	10.8	15.0	15.0	2,890
2007	23.1	19.9	15.8	16.8	15.6	3,403
2008	23.9	27.0	16.4	15.0	14.7	3,344
Gender						
Female	19.7	28.2	35.9	34.5	28.4	6,217
Male	80.3	71.8	64.1	65.5	71.6	13,802
Age						
Under 30	13.7	18.0	23.2	21.9	27.3	4,595
30–39	28.0	26.9	30.2	33.6	36.7	6,663
40–49	28.0	22.0	23.1	28.5	28.5	5,462
Over 50	30.4	33.1	23.5	16.1	7.6	3,299
Marital status						
Unmarried	25.3	6.9	11.1	39.6	37.5	6,888
Married	70.6	9.4	9.8	40.6	32.1	11,521
No longer married	4.1	4.3	11.1	43.0	36.0	954

Table 3 (*cont.*)

	Statistics on fund choices					
	Composition of observations by fund (per cent)					
	Guaranteed return	Monetary	Bond	Balanced	Equity	Observations
Education						
Elementary and middle school	4.8	10.2	8.8	5.7	3.6	1,128
High school	57.0	51.7	58.9	56.4	54.6	11,103
University	38.2	38.1	32.3	37.8	41.9	7,704
Job position						
Blue collar	0.3	2.0	1.9	1.8	0.7	277
White collar	51.0	54.9	63.7	63.3	62.9	12,346
Middle management	44.1	40.5	32.4	32.7	34.0	6,922
Senior management	4.6	2.5	2.0	2.1	2.3	474
Salary in 2008 (thousands of euros)						
Up to 25	4.2	5.1	6.3	4.9	4.6	986
25–35	30.8	35.1	40.0	43.8	45.3	8,463
35–45	25.6	23.9	28.6	23.4	22.9	4,803
45–55	14.1	16.1	12.7	12.3	12.5	2,572
55 +	25.2	19.8	12.3	15.5	14.7	3,195
Tenure (years)						
Less than 5	30.3	38.5	37.8	33.5	40.3	7,307
5–14	29.0	22.8	24.3	28.9	30.4	5,682
15–24	13.5	14.1	18.7	20.1	19.5	3,755
25–34	22.3	20.3	17.1	16.1	9.3	2,950
35 +	4.9	4.3	2.1	1.3	0.5	325

Table 4. *Statistics on switches between funds (number of decisions and percentages)*

	Statistics on switches			
	Switches over total decisions (%)	Switches to a safer fund over total switches (%)	Switches to a riskier fund over total switches (%)	Total investment decisions
Year				
2002	*17.8*	*85.3*	*14.7*	1,758
2003	*5.4*	*84.3*	*15.7*	1,887
2004	*14.1*	*64.3*	*35.7*	1,981
2005	*8.9*	*44.0*	*56.0*	2,152
2006	*10.6*	*64.2*	*35.8*	2,327
2007	*6.2*	*72.0*	*28.0*	2,805
2008	*7.3*	*87.1*	*12.9*	3,300
Gender				
Female	*8.5*	*66.7*	*33.3*	5,002
Male	*10.0*	*73.7*	*26.3*	11,208
Age				
Under 30	*8.4*	*56.4*	*43.6*	5,662
30–39	*9.9*	*71.5*	*28.5*	5,020
40–49	*9.5*	*81.0*	*19.0*	4,026
Over 50	*12.7*	*92.1*	*7.9*	1,502
Marital status				
Unmarried	*8.8*	*60.9*	*39.1*	5,373
Married	*10.1*	*76.8*	*23.2*	9,641
No longer married	*7.9*	*79.7*	*20.3*	807
Education				
Elementary and middle school	*10.0*	*87.2*	*12.8*	940
High school	*9.4*	*76.8*	*23.2*	9,099
University	*9.7*	*62.3*	*37.7*	6,133
Job position				
Blue collar	*8.9*	*83.3*	*16.7*	203
White collar	*9.1*	*65.2*	*34.8*	9,901
Middle management	*10.1*	*81.3*	*18.7*	5,705
Senior management	*11.7*	*76.6*	*23.4*	401
Salary in 2008 (thousands of euros)				
Up to 25	*8.4*	*59.7*	*40.3*	798
25–35	*8.6*	*59.5*	*40.5*	6,672
35–45	*10.6*	*77.0*	*23.0*	4,031
45–55	*10.1*	*84.7*	*15.3*	2,138
55 +	*10.4*	*82.0*	*18.0*	2,571
Tenure (years)				
Less than 5	*7.6*	*56.1*	*43.9*	4,876
5–14	*9.9*	*67.8*	*32.2*	5,180
15–24	*10.3*	*76.5*	*23.5*	3,179
25–34	*11.5*	*90.3*	*9.7*	2,692
35 +	*10.2*	*89.7*	*10.3*	283
Total	*9.5*	*71.8*	*28.2*	16,210

Table 5. *Statistics on plan balances (€)*

	Mean	Median	10th percentile	90th percentile
Gender				
Female	23,159	16,292	2,450	56,123
Male	37,499	22,325	3,460	87,543
Age				
Under 30	6,932	5,258	1,266	15,574
31–40	21,022	19,073	5,079	41,499
41–50	45,852	45,057	6,699	82,507
Over 50	74,629	68,370	6,764	144,471
Marital status				
Unmarried	16,399	9,087	1,730	41,664
Married	46,274	35,918	7,388	93,088
No longer married	45,449	40,175	9,019	84,205
Education				
Elementary and middle school	47,527	47,975	6,267	82,359
High school	38,736	26,372	3,288	87,000
University	23,196	12,915	2,607	52,131
Job position				
Blue collar	21,280	11,918	2,832	54,041
White collar	20,019	12,648	2,288	51,292
Middle management	51,853	46,246	7,148	101,597
Senior management	153,277	131,132	52,105	268,123
Salary in 2008 (thousands of €)				
Up to 25	19,317	17,111	2,887	39,535
25–35	13,412	9,284	1,886	28,280
35–45	39,749	40,479	5,301	72,924
45–55	49,990	51,259	5,414	91,424
55+	77,084	63,095	10,739	157,170
Tenure (years)				
Less than 5	14,792	6,064	1,624	38,280
5–14	29,438	23,242	11,850	50,267
15–24	47,121	46,739	15,769	71,255
25–34	75,658	68,679	34,448	113,674
35+	95,674	90,427	13,157	189,321
Total	32,647	20,053	3,069	77,494

horizon. So, we also compute gains and losses for the whole sample period. We consider the individuals that were present from the start to the end of the sample and decided to change,[22] and then compare their 7 years returns at the end of 2008 to what they would have earned if they had not switched. On average, the cumulative gains from switching amount to more than 18%.

[22] Only those which switched only once are considered in the computation. They represent however the overwhelming majority among those who switched.

5 Multivariate analysis

5.1 The empirical model

In this section, we describe and motivate the empirical model that we bring to the data. Consider a set-up in which the indirect utility of an investor i at time t depends on the share of risky assets α_{it} in his portfolio, on a vector of observable individual characteristics X_{it}, and on a stochastic element ε_{it}. More specifically, let us assume that the indirect utility function is given by $U(\alpha_{it};\ \phi_{it})$, where U is continuous in α, has the single-crossing property and $\phi_{i,t}=\beta X_{it}+\varepsilon_{it}$.[23]

Suppose also that investors can choose among three types of funds (labelled 0, 1 and 2), which differ in the fraction of risky assets (α_f) in their portfolios (without loss of generality, let their α_f be increasing: $0=\alpha_0<\alpha_1<\alpha_2$). It is straightforward to show that the fund chosen by the investor will be: fund 0 if $\beta X_{it}+\varepsilon_{it}\leqslant K_1$; fund 1 if $K_1<\beta X_{it}+\varepsilon_{it}\leqslant K_2$; fund 2 if $\beta X_{it}+\varepsilon_{it}>K_2$, with $K_1<K_2$.

If ε_{it} is distributed according to an $N(0,1)$ distribution, the conditional distribution of α_{it} given X_{it} is given in turn by:

$$\begin{aligned} P(\alpha_{it}=\alpha_0|X_{it}) &= P(\beta X_{it}+\varepsilon_{it}<K_1)=\Phi(K_1-\beta X_{it}),\\ P(\alpha_{it}=\alpha_1|X_{it}) &= \Phi(K_2-\beta X_{it})-\Phi(K_1-\beta X_{it}),\\ P(\alpha_{it}=\alpha_2|X_{it}) &= 1-\Phi(K_2-\beta X_{it}), \end{aligned} \tag{1}$$

where Φ is the cumulative density function of the standard normal. Equation (1) represents a multinomial ordered probit, which can be estimated using the standard maximum-likelihood techniques (Wooldridge, 2010), and this is the model that we adopt in the present paper.

5.2 Empirical results

We estimate the model described in the previous section on the pooled set of workers' choices for the 2002–2008 period. Besides age (summarized by four age dummies), we control for gender, marital status, education and job position. In all our regressions, we also include a full set of year dummies to capture unobserved time-specific effects, among which (perceived) changes in the process driving share prices. These time dummies are also interacted with the four age dummies, to check for possible changes over time of the age effect. Our baseline specification assumes no cohort effects (however, we relax this assumption in Section 5.1).

For the sake of clarity, we merge together the guaranteed return, the money-market and the bond funds (however, we checked that results do not vary if we consider each of the five funds separately).

Table 6 (columns 1–3) gives our baseline estimation results. It reports the average marginal effects of a change in the independent variables on the probability to choose each fund. These figures are computed as the average effects over all individuals of our population.

[23] U satisfies the single-crossing property if, for all $\alpha'<\alpha''$ and $\phi'<\phi''$ (and for all $\alpha'>\alpha''$ and $\phi'>\phi''$), $U(\alpha';\phi')\leqslant U(\alpha'';\ \phi')$ implies $U(\alpha';\phi'')\leqslant U(\alpha'';\ \phi'')$ (see, e.g., Mas-Colell *et al.*, 1995, for a discussion).

Table 6. *Ordered probit model: pooled regression (average marginal effects)*

	Baseline			Cohort effect		
Variable	Zero-share fund	Balanced fund	Equity fund	Zero-share fund	Balanced fund	Equity fund
Gender						
Male	−0.0346***	−0.0081***	0.0427***	−0.0348***	−0.0068**	0.0416***
Education						
High school	−0.0503***	−0.0232***	0.0736***	−0.0528***	−0.0221***	0.0749***
University	−0.0495***	−0.0227***	0.0723***	−0.0542***	−0.0229***	0.077***
Job position						
White collar	−0.0476**	−0.0098**	0.0574**	−0.0499**	−0.0083*	0.0581**
Middle management	−0.0551***	−0.0123**	0.0674***	−0.0557**	−0.0099*	0.0656***
Senior management	−0.077***	−0.0215**	0.0984***	−0.0832***	−0.0203**	0.1036***
Marital status						
Married	0.0026	0.0009	−0.0035	0.0031	0.0009	−0.004
No longer married	−0.0332***	−0.0141**	0.0473***	−0.0311***	−0.0116**	0.0426***
Age						
From 30 to 40 years old	−0.0359**	−0.015**	0.0509**	−0.0249	−0.0084	0.0333
From 40 to 50 years old	−0.0571***	−0.0277***	0.0847***	−0.0508***	−0.0209**	0.0718***
More than 50 years old	0.1373***	0.006	−0.1433***	0.1129***	0.006	−0.1189***
Time of the choice						
2003	0.154***	0.0028	−0.1568***	0.1442***	0.0008	−0.1449***
2004	0.0778***	0.0114**	−0.0891***	0.0757***	0.0089*	−0.0846***
2005	0.0085	0.0025	−0.011	0.0095	0.0024	−0.012
2006	−0.039**	−0.0166**	0.0556**	−0.0383**	−0.0144**	0.0527**
2007	0.0101	0.0029	−0.0131	−0.0025	−0.0007	0.0032
2008	0.0416**	0.0091*	−0.0508**	−0.0146	−0.0045	0.0191
Cohort effect	N	N	N	Y	Y	Y

Note: Changes in the average probability of choosing one of the three asset allocations when the value of each dummy variable changes from zero to one for every individual.
The reference is a female, primary and middle school, blue collar worker, unmarried, in 2001. Full set of interaction terms between the age and the year dummies are included in the estimated model (omitted in the table).
Significance levels: 1% (***); 5% (**), 10% (*). Standard errors are robust to heterogeneity and autocorrelation.

Table 7. *Model-based probabilities of choosing a given fund (percentage points)*

	2002	2003	2004	2005	2006	2007	2008
Zero-share fund							
Under 30	34.3	36.9	32.1	27.6	24.3	27.7	29.8
From 30 to 40	29.4	31.1	32.1	30.3	30.5	31.4	33.1
From 40 to 50	27.0	28.0	30.7	30.4	32.8	35.2	37.3
Over 50	42.8	45.3	45.7	44.3	48.1	50.7	51.2
Balanced bond fund							
Under 30	27.2	27.3	27.0	26.3	25.3	26.3	26.7
From 30 to 40	26.7	26.9	27.0	26.8	26.8	27.0	27.1
From 40 to 50	26.1	26.4	26.9	26.8	27.1	27.3	27.3
Over 50	26.8	26.4	26.3	26.6	25.9	25.2	25.0
Balanced equity fund							
Under 30	38.5	35.8	40.8	46.2	50.5	46.0	43.5
From 30 to 40	43.9	41.9	40.9	43.0	42.7	41.6	39.7
From 40 to 50	47.0	45.6	42.5	42.8	40.1	37.5	35.4
Over 50	30.4	28.3	28.0	29.1	26.1	24.1	23.7

Note: Estimated probabilities implied by the model. The reference is a male, white collar, high school and married worker.

Overall, the findings of the univariate analysis are confirmed. In particular, the reduction in equity-holding due to ageing is statistically significant (standard errors are clustered at the individual level). On average, the probability of being in a zero-share portfolio initially decreases and then increases with age, while the reverse is true for the probability to be in the riskiest fund.

The age–year interaction terms are statistically significant, starting from 2006.[24] Table 7 shows how the probability of choosing each fund for an individual with given characteristics changes through years and age classes.

In order to better assess the economic significance of the effects we can compute the expected fraction of equities in the chosen portfolio (α_{it}):

$$E(\alpha_{it}|X_{it}) = \alpha_0 P(\alpha_{it} = a_0|X_{it}) + \alpha_1 P(\alpha_{it} = a_1|X_{it}) + \alpha_2 P(\alpha_{it} = a_2|X_{it}).$$

According to our estimates, the relationship between age and the holding of stock changes across time, becoming stronger at the end of the sample; moreover, while in the first years of the sample the age-stockholding profile is hump-shaped, starting from 2005, it becomes monotonically negative (Table 8 and Figure 3).[25] This may be due to the fact that workers – observing the losses suffered by their colleagues who retired during periods of declining share prices – have learnt that being exposed to stock

[24] In non-linear models the significance of the coefficient of an interaction *term* does not necessarily imply that there is a significant interaction *effect*, which is the cross derivative of the probability with respect to age and time. However, we computed this cross-derivative and the associated standard error using the delta method as suggested by Ai and Norton (2003), and we verify their statistical significance.

[25] With a test on the related interaction effect (see footnote 25), we could verify that this differences are statistically significant.

https://doi.org/10.1017/S1474747213000395 Published online by Cambridge University Press

Table 8. *Expected asset allocation (percentage points)*

	2002	2003	2004	2005	2006	2007	2008
Under 30	31.26	29.68	32.61	35.57	37.86	35.50	34.11
From 30 to 40	34.34	33.24	32.64	33.81	33.65	33.07	31.97
From 40 to 50	38.81	36.00	35.27	33.53	33.73	32.21	30.70
Over 50	26.26	24.91	24.68	25.42	23.41	22.04	21.74

Note: Estimated probabilities implied by the model. The reference is a male, white collar, high school and married worker. Estimated shares of stocks implied by the ordered probit model assuming that the shares are equal to 0%, 30% and 60%, respectively for the zero-share, balanced bond and balanced equity funds.

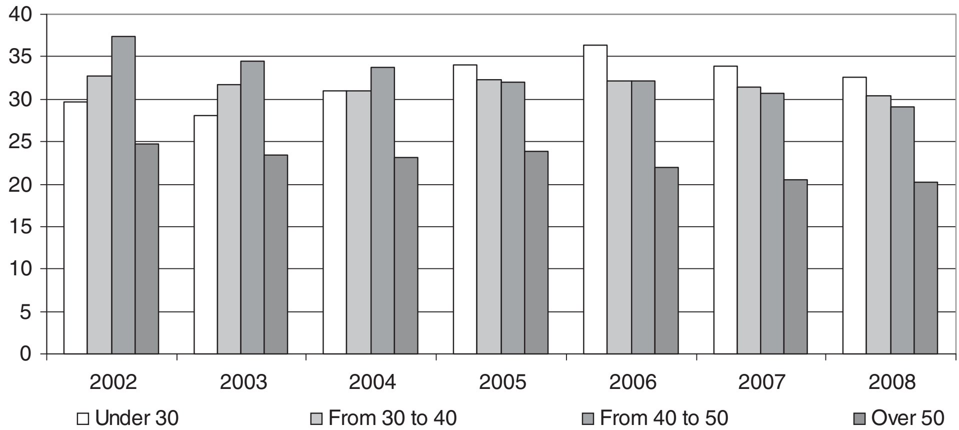

Figure 3. Model-based expected portion of equities by age and years (%).

market risk when they are near to retirement is very risky. In 2002, a married male with a white collar position and a high school degree can be expected to hold in equities a fraction of his portfolio equal to 31% if he is younger than 30, which rises to 39 if he is in his forties, then drops to 26% if he is older than 50. In 2008, these figures are 34%, 32% and 22%, respectively.

If we use age and age-squared instead of the age dummies then we can observe similar patterns. The relationship between stockholding and age turns negative at about 40, while in Agnew *et al.* (2003) the peak is reached at about 30.

Like the final decrease, the initial increase in the fraction of stocks is consistent with optimal portfolio theory, as the young are relatively more exposed to labour market risks (Campbell and Viceira, 2002).

Concerning controls, being male, having a better job position and a higher education decreases the probability of choosing a zero-share portfolio and increases the probability of choosing the riskiest portfolio in a statistically significant way.

Table 9a. *Expected asset allocation (percentage points)*

	2002	2003	2004	2005	2006	2007	2008
Subsample: female							
Under 30	30.4	28.7	32.1	35.3	36.7	34.9	33.9
From 30 to 40	32.8	32.0	31.1	31.9	32.5	32.4	31.0
From 40 to 50	38.1	35.0	33.6	32.6	32.0	31.9	30.6
Over 50	23.8	22.0	23.5	31.3	26.8	27.1	25.3
Subsample: male							
Under 30	31.0	29.5	32.2	34.9	37.9	35.1	33.4
From 30 to 40	34.3	33.0	32.6	34.1	33.5	32.7	31.8
From 40 to 50	38.5	35.8	35.2	33.2	33.7	31.6	30.0
Over 50	26.1	24.8	24.5	24.7	22.8	21.2	20.9

Note: Probabilities implied by the model estimated on different subgroups. The reference is a white collar, high school and married worker. Estimated shares of stocks implied by the ordered probit model assuming that the shares are equal to 0%, 30% and 60% respectively for the zero-share, balanced bond and balanced equity funds.

Table 9b. *Expected asset allocation (percentage points)*

	2002	2003	2004	2005	2006	2007	2008
Subsample: blue and white collar							
Under 30	31.6	30.0	32.8	36.0	38.4	36.1	34.8
From 30 to 40	33.4	32.6	32.6	33.8	34.0	33.6	32.6
From 40 to 50	38.3	35.2	34.3	33.1	33.9	33.4	32.2
Over 50	27.3	25.4	24.9	26.3	24.8	23.1	22.7
Subsample: managers							
Under 30	29.6	30.8	39.1	39.6	43.0	41.2	37.9
From 30 to 40	35.8	34.6	33.1	34.3	33.2	32.1	30.5
From 40 to 50	39.6	37.1	36.5	34.3	34.0	31.8	30.0
Over 50	26.2	25.1	25.0	25.4	23.2	21.9	21.7

Note: Probabilities implied by the model estimated on different subgroups. The reference is a male, high school and married worker. Estimated shares of stocks implied by the ordered probit model assuming that the shares are equal to 0%, 30% and 60%, respectively for the zero-share, balanced bond and balanced equity funds.

To investigate further the role of gender, we estimate separately the age-portfolio profiles for men and women.[26] It appears that, while in most years of the sample males start their career with a slightly riskier portfolio with respect to females, they are also relatively quicker (at least in the second half of the sample period) in moving towards safer assets as they age (Table 9a). This result is in line with the results by Barber and Odean (2001*a*), who show that men tend to trade more than women.

Interestingly, a similar pattern emerges if we estimate the model separately for managers and non-managerial workers (Table 9b), and if we distinguish between workers

[26] This is similar to what is done by Jianakoplos and Bernasek (1998).

Table 9c. *Expected asset allocation (percentage points)*

	2002	2003	2004	2005	2006	2007	2008
Subsample: not graduated							
Under 30	30.3	29.4	31.2	32.9	34.6	33.1	32.6
From 30 to 40	33.6	32.5	31.2	32.2	32.1	31.6	30.1
From 40 to 50	37.8	34.9	34.0	32.0	32.5	31.0	29.4
Over 50	25.1	23.9	23.7	24.3	22.3	21.1	21.0
Subsample: graduates							
Under 30	32.5	30.3	34.4	38.3	40.9	38.1	36.1
From 30 to 40	34.7	33.8	34.0	35.4	35.3	34.7	33.9
From 40 to 50	39.0	36.4	36.2	35.1	34.4	32.8	31.5
Over 50	26.1	24.3	24.4	25.8	24.0	22.2	21.1

Note: Probabilities implied by the model estimated on different subgroups. The reference is a male, white collar and married worker. Estimated shares of stocks implied by the ordered probit model assuming that the shares are equal to 0%, 30% and 60%, respectively for the zero-share, balanced bond and balanced equity funds.

with and without a university degree (Table 9c). Both managers and graduates start their careers with relatively more risky assets in their portfolio, but both categories are relatively faster in moving towards less risky assets. This behaviour might be related to the fact that they have easier access to financial information and/or a higher level of financial literacy (documented among others by Van Rooij *et al.*, 2011).

6 Extensions and robustness checks

6.1 Modelling cohort effects

As it is well known, due to collinearity it is impossible to separately identify cohort, time and age effects without imposing some additional identification restriction to the data-generating process. In the estimates of Section 5, we followed Agnew *et al.* (2003), therefore we controlled for time and age effects while assuming the absence of cohort effects. In this section, however, we relax this assumption and enrich our baseline model by adding cohort effects. To do this, we use the identification strategy pioneered by Malmendier and Nagel (2011). These authors argue that portfolio choices of people belonging to the same cohort are similar because they went through similar economic and financial experiences and in particular, they experienced the same stock market returns during their lifetime. Therefore, we enrich our specification to include the Malmendier and Nagel (2011) proxy for the cohort effect, namely the average yearly returns of the Italian stock market over the lifetime of each individual. Remarkably, this leaves unaffected the magnitude and significance of the age effect (Table 6, columns 4 to 6).

6.2 Fixed-effects estimation

As we remarked in the introduction, one limit of administrative records is that they do not contain much information about participants' characteristics. The possibility of a bias in our estimates due to omitted variables is therefore a source of concern. In the present section, we address this problem by estimating a fixed-effect model, which allows us to control for the effect of all the unobserved time-invariant characteristics of the participants.

We make two departures with respect to our baseline model: first, we use a logit instead of a probit distribution function; second, we model age as a quadratic function instead of a set of age-dummies. The first departure is needed because there is no way, to our knowledge, to account for unobserved individual heterogeneity within a multinomial probit model whereas in the multinomial logit case we can resort to some recent papers and in particular to Baetschmann *et al.* (2011). The second departure is needed because the Baetschmann *et al.* (2011) estimation technique does not exploit the information provided by all those individuals that never switch between funds. This entails a sharp decrease in the degrees of freedom that induced us to resort to a more parsimonious specification of the age effect.

In our setting, where the dependent variable can take three values (e.g., 0, 1, 2), the Baetschmann *et al.* (2011) procedure requires the creation of two dichotomous variables ($d^1_{i,t}$, $d^2_{i,t}$). The first is equal to one if and only if the chosen fund is the one with the highest share of risky assets; the second is equal to one if the chosen fund is not the safe one, i.e. (using the notation introduced in Section 5.1):

$$\begin{aligned} d^1_{i,t} &= I(\mu_i + \beta X_{it} + \varepsilon_{it} > K_1), \\ d^2_{i,t} &= I(\mu_i + \beta X_{it} + \varepsilon_{it} > K_2), \end{aligned} \tag{2}$$

where $I(\cdot)$ is the indicator function and μ_i is the individual fixed effect. The second step of the procedure requires the estimation of two binary logit-fixed effects models for $d^1_{i,t}$ and $d^2_{i,t}$ as in Chamberlain (1980). Indeed, Chamberlain (1980) shows that maximum-likelihood estimation of a binary logit-fixed effects model conditional on the sum of all the outcomes over time gives consistent estimates. In the final step of the procedure, a weighted sum of the two estimates obtained in the second step is computed.[27]

The second column in Table 10 shows that the estimated age effects obtained using the Baetschmann *et al.* (2011) procedure confirms our baseline results: the coefficients on age and age-squared are highly significant and the stock-holding/age profile has a hump-shaped pattern, with a peak at 40.

6.3 Fixed-effects estimation controlling for dynamics

The choice of the fund in period t might depend on previous choices. This would be the case if investors tend to stick to their status quo (inertia in financial decisions has been documented in a number of studies, among which Bilias *et al.*, 2010). Past choices would matter also if investors are relatively more responsive to the performance of their own fund. The specification used in the previous section cannot account for such dynamic effects, as lagged values of the dependent variable are not among the regressors. To allow for this possibility we extend the dynamic fixed-effect binary

[27] Weights are equal to the inverses of the respective variances.

Table 10. *Estimations of the age effect: robustness exercises (parameter estimates)*

	Pooled regression (ordered logit)	Individual fixed effect (BSW, 2011)	Dynamic individual fixed effect (BN, 2010)
Age	0.2827***	1.2468***	0.8923***
Age square	−0.0037***	−0.0197***	−0.0043***
Previous year choice	–	–	7.0805***
Age x year	Y	Y	Y
Individual heterogeneity	N	Y	Y
Dynamic structure	N	N	Y
Observations	17,627	7,563	3,665
Pseudo R-squared	0.03	0.18	–

Note: Baetschmann *et al.* (BSW, 2011) map the categorical response into a set of binomial variables, so that a fixed-effect binomial logit can be estimated. Bartolucci and Nigro (BN, 2010) propose a statistical model that allow dynamic fixed-effect, we apply it on the specification proposed by Baetschmann *et al.* (2011).

logit model recently developed by Bartolucci and Nigro (2010) to our trichotomous setting.[28]

Bartolucci and Nigro (2010) proposed the following binary model:

$$d_{i,t} = I(\mu_i + \beta X_{it} + d_{i,t-1}\gamma + e^*(\mu_i, X_i) + \varepsilon_{i,t} \geqslant 0), \tag{3}$$

where $d_{i,t}$ is a dichotomous variable, μ_i is the individual fixed effect, and $e^*(\mu_i, X_i)$ is a variable which can be interpreted as a measure of the effect of the present choice $d_{i,t}$ on future expected utility. We apply the Baetschmann *et al.* (2011) procedure described in the previous section to the model in equation (3) instead of applying it to a Chambelain-style static fixed effect binary logit as Baetschmann *et al.* (2011) do. Namely, we estimate equation (3) for the two dicothomous variables described in equation (2), and build an estimate of the impact of X_{it} on the original trichotomous variable using a weighted average of the estimates concerning the two dicothomous variables $d^1_{i,t}, d^2_{i,t}$.

The results of this procedure confirm the relation between asset allocation and age found in our baseline exercise (Table 10, column 3). Moreover, the dynamic factor turns out to be significant, suggesting that portfolio choices are indeed persistent.[29]

6.4 Conditional switching probabilities

In order to investigate the role of inertial behaviour in workers' choices in this section, we can also focus specifically on shifts from one fund to another.

To this aim, we run our baseline regression conditional on the fund chosen in the previous year. Besides being interesting per se, this exercise can be also seen a simple

[28] We thank Bartolucci and Nigro for suggesting us this extension of their own procedure.

[29] As suggested by one referee, we also estimated a couple of fixed-effect linear probability model, using both $d^1_{i,t}$ and $d^2_{i,t}$ as dependent variables. Among the independent variables, we included a full set of age dummies and the lagged dependent variable. Results are in line with those presented in Table 10.

way to control for both dynamic effects and unobserved heterogeneity, alternative to the one discussed in Section 6.2. In particular, it has the advantage that we can use our baseline specification, and that it does not ignore the information concerning the participants that never switch fund.

We proceed in two steps. First, we run our baseline regression on different subsamples, grouping people according to the fund that they chose in period $t-1$ (Table 11). As before, dependent variables include dummies for gender, education, job position, marital status, years and age.

Second, we use the estimated parameters to compute the conditional probability of switching from one fund to another. The probabilities are summarized in conditional transition matrices (Tables 12 and 13).[30] The elements on the main diagonal of each matrix give, for a particular participant (e.g., a male, middle manager, higher educated, unmarried participant choosing his retirement account asset allocation in 2008), the probability of remaining in the old fund; on the contrary, the elements off the main diagonal give the probability of switching from the fund on the row to the one on the column. We compute different matrices for alternative settings of the X variables in order to assess the impact of each covariates. All in all, this approach is analogous to that of Bertaut (1998).[31]

The age effect highlighted in the previous sections is again quite strong (Table 11). The probability of remaining in the riskiest fund is 96% for a less than −30 y.o. worker, falling to 85% for a 50+ y.o. worker. Moreover, the probability of switching towards less risky funds starting from the balanced one is much lower for the young than for the old participant (6% versus 18%).

The likelihood of switching towards less risky portfolios is higher at the beginning and at the end of the sample, when the returns from the stock market were particularly disappointing. In 2005 (a year of relatively bullish markets), the probability of not changing fund was 95% for those starting in the riskiest fund and 93% for those starting in the balanced fund. These probabilities were, respectively, 93% and 90% in 2008, and 87% and 86% in 2002 (Table 13). Most importantly, the probability of switching towards riskier funds for those in the zero-equities portfolios was much lower during the end-of-period and the beginning-of-period stock market crashes: indeed, for those starting from the no-shares funds, the probability was 18% in 2005, compared with 4% and 2%, respectively in 2002 and 2008. The effects of sex and job position on the probability of switching are not statistically significant (results not shown), as in the sample these are basically time-invariant characteristics. This result is different from what we found in the previous section, in which we studied the unconditional probability of choosing a particular fund.

So far, we focused on annual conditional probabilities. In the context of retirement saving we might want to evaluate probabilities over a longer horizon. For example, we might be interested in the model-based probability that a 35 years old plan participant who invests in the equity fund will end up 20 years later in the zero-share fund.

[30] A similar approach, applied to a different issue, is adopted by Nickell *et al.* (2000).
[31] The difference is that we have several years and not just two, and that, given our multinomial set-up, we have three-by-three instead of two-by-two matrices.

Table 11. *Ordered probit model: separate regressions (parameters estimates)*

	Zero-shares fund	Balanced fund	Equity fund
Male	0.0910	0.108**	−0.0614
	(0.0760)	(0.0449)	(0.0587)
Primary and middle school	–	–	–
	–	–	–
High school	0.534***	0.0482	−0.0198
	(0.206)	(0.0889)	(0.125)
University	0.660***	0.132	−0.0261
	(0.211)	(0.0956)	(0.132)
Blue collar workers	–	–	–
	–	–	–
White collar workers	0.433	−0.111	0.376
	(0.411)	(0.143)	(0.256)
Middle management	0.230	−0.143	0.304
	(0.418)	(0.148)	(0.257)
Senior management	0.476	−0.0696	0.535*
	(0.462)	(0.205)	(0.290)
Married	−0.101	0.0124	−0.138**
	(0.0776)	(0.0464)	(0.0559)
Under 30 y.o.	–	–	–
	–	–	–
From 30 to 40 y.o.	0.212**	−0.192***	−0.225***
	(0.0972)	(0.0637)	(0.0787)
From 40 to 50 y.o.	0.148	−0.204***	−0.317***
	(0.121)	(0.0703)	(0.0875)
Over 50 y.o.	−0.378***	−0.618***	−0.706***
	(0.142)	(0.0775)	(0.105)
2002	–	–	–
	–	–	–
2003	−0.0964	0.504***	0.590***
	(0.253)	(0.0815)	(0.106)
2004	0.847***	0.531***	0.0902
	(0.225)	(0.0883)	(0.0806)
2005	0.855***	0.827***	0.499***
	(0.226)	(0.0845)	(0.0957)
2006	0.686***	0.806***	0.0203
	(0.229)	(0.0826)	(0.0792)
2007	0.124	0.674***	0.229***
	(0.238)	(0.0772)	(0.0865)
2008	−0.239	0.266***	0.361***
	(0.249)	(0.0746)	(0.0919)
Cut 1	2.942***	−1.215***	−1.616***
	(0.474)	(0.165)	(0.288)
Cut 2	3.612***	2.439***	−1.206***
	(0.476)	(0.170)	(0.286)
Observations	3761	6565	5592
Pseudo R-squared	0.1183	0.0605	0.0454

Note: The table shows parameter estimates of ordered probit models run separately for participants starting from a zero-share, balanced bond and balanced equity funds.
The reference is a female, primary and middle school, blue collar, unmarried, under 30 y.o., 2002. Significance levels: 1% (***); 5% (**), 10% (*).

Table 12. *Model-based conditional transition matrix by age (percentages)*

Initial fund	Chosen fund 2008 Zero-shares (%)	Balanced (%)	Equity (%)
Under 30 years old			
Zero-shares	98.7	1.1	0.2
Balanced	6.2	92.1	1.7
Equity	1.6	2.6	95.8
From 30 to 40 years old			
Zero-shares	97.8	1.8	0.4
Balanced	**8.9**	**90.0**	**1.1**
Equity	**2.8**	**3.9**	**93.3**
From 40 to 50 years old			
Zero-shares	98.1	1.6	0.3
Balanced	**9.1**	**89.9**	**1.0**
Equity	**3.5**	**4.5**	**92.0**
Over 50 years old			
Zero-shares	99.5	0.4	0.1
Balanced	**17.9**	**81.8**	**0.3**
Equity	**7.7**	**7.8**	**84.6**

Note: The reference individual is a 30-to-40 years old male worker, white collar, with a high school degree and married. The percentages show model-based probabilities to switch form the initial fund (rows) to the chosen fund (columns). Probabilities in bold are statistically different from those of the reference matrix at the 5% significance level.

We can compute this probability using the transition matrices for 2008 in Table 12. For example, if we consider a married man, white-collar, with a high-school degree, and assume that his job-position does not change over time, there is 86% chance that by the end of the 20 years period he will be found in the zero-share fund. Alternatively, if we use the transition matrix for 2005, a year of positive returns, such probability goes down to 42%. We could be more accurate and allow for the fact that the years of negative and positive returns evolve stochastically over time; however, our simple exercise can be seen as giving reasonable lower and upper bounds.

7 Conclusions and suggestions for further research

We studied investors' portfolio choices in a very simple and clear-cut real-world setup. Some results prove quite robust across all the empirical exercises we performed. First, there is a tendency to choose safer funds as people age. This effect is still there after controlling for several demographic factors, for time effects, for the fund chosen in the previous period, for individual fixed effects (in a static as well as in a dynamic specification) and cohort effects (modelled as in Malmendier and Nagel, 2011).

Table 13. *Model-based conditional transition matrix by year (percentages)*

Initial fund	Chosen fund		
	Zero-shares (%)	Balanced (%)	Equity (%)
Year 2002			
Zero-shares	96.2	3.1	0.7
Balanced	14.0	85.5	0.5
Equity	6.1	6.7	87.2
Year 2005			
Zero-shares	**82.1**	**12.3**	**5.6**
Balanced	**2.8**	**93.1**	**4.0**
Equity	**2.0**	**3.1**	**94.9**
Year 2008			
Zero-shares	97.8	1.8	0.4
Balanced	**8.9**	**90.0**	**1.1**
Equity	**2.8**	**3.9**	**93.3**

Note: The reference individual is a 30-to-40 years old male worker, white collar, with a high school degree and married. The percentages show model-based probabilities to switch form the initial fund (rows) to the chosen fund (columns). Probabilities in bold are statistically different from those of the reference matrix at the 5% significance level.

However, the effect is non-linear, being much stronger in the very last years of the career. Moreover, not all elderly people in our sample reduced their exposure to risk. Looking at the ones present in the sample from the start, it turns out that more than 30% of the elderly workers, which were exposed to stock market risk in 2001 were still exposed to it in 2008. More generally, the significance of the persistence parameter in our dynamic estimates signals a possible status quo bias.

An elderly worker taking risk on the stock market could pay a high price if stocks fall. To the extent that this inertia is related to behavioural factors, there might be room for welfare-improving policy interventions. For example, the diffusion of life cycle funds could be promoted. This kind of investment vehicles automatically brings all the participants towards less risky allocations as they age (Viceira, 2007). In the Chilean system, for example, a lifecycle fund is the default option for all the workers. Moreover, the riskiest funds are closed to individuals older than a certain age.

We also document that the effect of age is more pronounced in the last years of the sample. This might be due to the fact that investors learn from the experience of their colleagues. Indeed, in our sample there have been periods of disappointing stock market performance. Having seen that people who retired during these bear market periods have been severely hit might have pushed investors towards a more active behaviour. A better understanding of this form of learning appears to be an interesting issue for further research.

We find that job position has an impact on portfolio choice (but not on the probability of switching): people with a higher position tend to take more risks. This tallies with previous empirical analyses and can be consistent with optimal portfolio allocation (Brunnermeier and Nagel, 2008; Chiappori and Paiella, 2008; Cappelletti, 2009). Finally, we find that education also increases both the share of stocks in the portfolio and the likelihood of switching for those in the zero-shares funds.

Before concluding, we would like to remark that the fact that the behaviour of most workers is by and large consistent with optimal portfolio theory does not rule out other possible interpretations of our results. For example, if investors pay more attention to retirement savings when they are older (because retirement is more salient or because the plan balance is bigger), and if attentive individuals tend to move towards saver portfolios (maybe because the sample period is overall a period of disappointing stock market performance), this could explain why workers leave risky investments as they age. An increased use of advisory services by older workers would be evidence in favour of this interpretation.[32]

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[32] Unfortunately we do not have data concerning this aspect. We thank one referee for suggesting this alternative reading of our results.

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