Can Security Design Foster Household Risk-Taking?

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ABSTRACT

This paper shows that securities with non-linear payoff designs can foster household risk-taking. We demonstrate this effect by exploiting the introduction of capital guarantee products in Sweden between 2002 and 2007. The fast and broad adoption of these products is associated with an increase in expected financial portfolio returns. The effect is especially strong for households with low risk appetite ex ante. These empirical facts are consistent with a life-cycle model in which households have pessimistic beliefs or preferences combining loss aversion and narrow framing. Our results illustrate how security design can mitigate behavioral biases and enhance economic well-being.

JEL Codes: I22, G1, D18, D12.

Keywords: Security design, household finance, capital guarantee product, behavioral biases, stock market participation, risk-taking.

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I. Introduction

A distinguishing feature of financial markets is that in every country, a sizable group of households only invest a small share of financial wealth in stocks and mutual funds (Calvet, Campbell, and Sodini, 2007).[1] This empirical fact is a challenge to canonical models of portfolio allocation (Campbell and Viceira, 2002; Cocco, Gomes, and Maenhout, 2005; Merton, 1971) because households with low equity holdings forfeit an important source of income over their lives (Haliassos and Bertaut, 1995; Mankiw and Zeldes, 1991), which reinforces wealth inequality (Bach, Calvet, and Sodini, 2019).[2]

Another, and potentially related, challenge to established finance theory is the impressive growth over the past two decades of the market for retail *capital guarantee products* (thereafter CGPs), a class of equity-linked contracts offering a capital protection. In 2015, CGPs total more than $4.5 trillion in global outstanding volumes and represent a significant share of household savings in major economies, such as the U.S., China, and the European Union.[3] In Sweden, where precise data on household portfolio composition is available, CGPs were adopted quickly and broadly, reaching 14% of the population within 5 years of their introduction. However, rational-choice portfolio theory does not provide a clear economic rationale for the success of these products. By contrast, several innovative financial assets with strong economic motivations, such as low-cost exchange-traded funds or inflation-indexed bonds, have experienced much slower speeds of adoption (Shiller, 2004).

Taken together, these major stylized facts raise a number of questions. Does the capital protection embedded in CGPs foster household financial risk-taking? If so, through which economic mechanism? Are households better off as a result? More generally, can security design mitigate behavioral biases preventing sizable groups of households from making efficient decisions?

In this paper, we take a first step in answering these questions by empirically studying

[1]See Table I.
[2]See also Gomes, Haliassos, and Ramadorai (2020) and the references therein.
[3]See Table II.

the effects of the introduction of CGPs on household risk-taking in Sweden in the 2000s. Our analysis exploits a unique administrative data set containing granular information on the demographics and exact portfolio composition of every Swedish resident (see Calvet et al. (2007)), which we merge with detailed information on all CGPs sold in Sweden (see Célérier and Vallée (2017)). The resulting panel offers a comprehensive coverage of the 2002-2007 period, the first five years of the development of the retail market for CGPs. In a second step, we investigate the theoretical mechanisms that can rationalize our empirical findings by augmenting the life-cycle model of portfolio allocation of Cocco et al. (2005). We include capital guarantee products in the set of financial assets available to households and span a series of preference and belief specifications for these agents. We calibrate each version of the model to the data. This theoretical exercise allows us to identify two possible economic explanations for our empirical findings and also to assess the implications of financial innovation for household well-being.

We begin our empirical analysis by showing that the CGPs sold in Sweden allow retail investors to earn a significant fraction of the equity premium. We conduct an asset pricing assessment of these products that accounts for all aspects of their design, including their exact payoff formula, disclosed fees, credit risk, and the ex-dividend nature of the final payoff. CGPs offer investors a risk premium amounting to 44% of the equity premium on average, even though they embed relatively high total markups that are on average equal to 1.6% per year. These expected excess returns and markups are comparable to the values obtained for equity mutual funds sold in Sweden over the same period.[4]

Among equity participants, households that adopt CGPs are found to increase their risk-taking significantly more than households that do not. We define a household's *risk-taking index* as the expected fraction of the yearly equity premium earned on their financial portfolio, net of fees.[5] Over the 2002-2007 period, the risk-taking index increases by 3

[4]Gennaioli, Shleifer, and Vishny (2015) report similar magnitudes for mutual funds in the U.S. once taking into account all types of fees.

[5]The literature usually measures risk-taking with the risky share, which is the weight of risky assets in the financial wealth, without adjusting for the heterogeneity in the risk premium that each risky asset might

percentage points (pp) for adopters and 1 pp for non-adopters, to compare with a median risk-taking index in 2002 of 17 pp for equity markets participants.

The relationship between CGP investing and an increase in risk-taking is significantly more pronounced for equity market participants that are initially less willing to take risk. This sizable group of households is wealthier on average that non-participants and therefore faces a significantly larger cost of opportunity of not investing along canonical predictions.[6] This group also represents a potentially important source of capital for the corporate sector. While the initial risk-taking index is only 2 pp for participating households in the bottom quartile of the willingness to take risk in 2002, the index increases by 13 pp for CGP adopters versus only 6 pp for non-adopters at the end of the sample period. This heterogeneity results from a higher demand for CGPs for household initially less willing to take risk, and from low substitution with traditional equity products.

To gain causal identification, we instrument household investments in CGPs by quasi-random shocks to the bank idiosyncratic supply of these products in a panel model with household and year fixed effects. We provide evidence that bank supply shocks drive an important share of the volumes of CGPs. Similar to Borusyak, Hull, and Jaravel (2018), our identification strategy relies on the exogeneity of the shock in the time series and does not require the exogenous matching of households and banks. We estimate the idiosyncratic supply shocks by regressing the quantity of CGPs a household holds in a given year on bank-year fixed effects, controlling for household characteristics, in a random half of the household population. We then use the other half to analyze the causal effects of the supply of CGPs on the household risk-taking index. We find that a 1 pp increase in the share of financial wealth invested in CGPs leads to a 0.69 pp increase in the risk-taking index.

We next examine the theoretical determinants of investments in CGPs. We develop a life-cycle model with stochastic labor income that extends standard models (e.g., Cocco et al.

offer based on the payoff design, beta and fees (see for example Calvet et al. (2007)). Another technique would be to use the delta coefficient used in derivatives pricing, but this continuous-time measure is ill-suited to analyze the portfolios of retail investors trading at lower frequencies.

[6]See Section II.A.

(2005)) along several dimensions. The investment set includes a bond, a stock, and a CGP with the exact same design, embedded markup, and illiquidity as the median product in our sample. We solve the life-cycle model across a set of utility functions and beliefs and relate the results to our empirical findings. This exercise provides a set of novel theoretical insights.

We show that preferences incorporating narrow framing on investment income with loss aversion (Barberis and Huang, 2009) explain why the introduction of CGPs fosters financial risk-taking, especially among households that are initially less willing to take risk. By contrast, Epstein and Zin (1989) preferences, general disappointment aversion (Gul, 1991; Routledge and Zin, 2010), and smooth forms of narrow framing cannot explain the data. The intuition is the following. When risk aversion is second-order, as is the case under Epstein-Zin preferences or smooth forms of narrow framing, the stock offers an attractive trade-off between risk and return, while the welfare benefits from CGPs and the demand for these products are weak. First-order risk aversion is therefore a natural avenue. However, as Barberis, Huang, and Thaler (2006) explain, the presence of other preexisting risks, such as labor income risk, makes a purely loss averse agent act in a second-order risk-averse manner toward independent, delayed gambles. The combination of narrow framing and loss aversion is therefore necessary to explain the empirical results in our life-cycle framework under rational expectations.

We demonstrate that pessimistic beliefs alone, for instance those captured by probability weighting (Prelec, 1998), can also explain the positive and heterogeneous response of risk-taking to financial innovation. Pessimistic households have a strong demand for CGPs because these contracts combine the upside potential of equity markets with a protection against adverse outcomes, which pessimistic households view as particularly likely. The increase in risk-taking is therefore the strongest for the most pessimistic households.

Building on these results, we assess the welfare gains associated with the introduction of CGPs. By revealed preference, a household should be strictly better off under the lens of its *decision* utility if it adopts the innovation, and we indeed observe large gains under this metric.[7] We estimate how the surplus created by the introduction of capital guarantee products is shared between financial institutions and households. We observe that, despite the comfortable markup that banks charge, households obtain a substantial share of the surplus. These results suggest that banks do not necessarily capture the entire surplus that they create when addressing a bias.

Last, we take a conservative approach by assessing household welfare through the lens of *experienced* utility (Kahneman, Wakker, and Sarin, 1997). Assuming that experienced utility exhibit less pronounced behavioral traits than the decision utility, we still find sizable welfare gains, except for households with high initial willingness to take risk. Financial institutions seeking to improve the well-being of their customers should target the sale of capital guarantee products to households with low levels of risk-taking.

This paper contributes to the strand of the household finance literature investigating low risk-taking by a sizable group of households. While the literature provides a long list of possible explanations for such behavior (e.g. Attanasio and Vissing-Jørgensen (2003); Barberis et al. (2006); Calvet et al. (2007); Guiso, Sapienza, and Zingales (2008); Haliassos and Bertaut (1995); Kuhnen and Miu (2017)), our work identifies specific preferences or beliefs as first-order mechanisms underlying such behavior by assessing the effectiveness of a targeted remedy.

In this respect, our study opens a new direction in the active debate on whether financial education (Bernheim, Garrett, and Maki, 2001), financial advisors (Gennaioli et al., 2015) or default options (Madrian and Shea, 2001) should be prioritized to address the frictions households face when making financial decisions. While the evidence on the effectiveness of financial education is mixed (e.g. Duflo and Saez (2003), Lusardi (2008)), Chalmers and Reuter (2020) show that in the context of U.S. retirement plans, introducing default options in target funds is more valuable to households than providing them with access to financial

[7] We make the simplifying assumption that the risk-free rate and the dynamics of equity are not impacted by financial innovation, which seems to be a reasonable approximation in our sample period.

advisors. Due to offsetting household behaviors at longer horizons however, extrapolating the short-run gains from default option introductions can significantly overstate their benefits at longer horizons, as Choukhmane (2019) documents. Our findings suggests that security design might be both more effective and more targeted than each of these alternatives by specifically addressing the bias distorting household financial decision-making.[8] In this sense, the security design solution we identify has the ability to provide customized efficiency, analogous to the decision process designs advocated by Thaler and Benartzi (2004) and others to encourage higher saving rates.

Our work also contributes to the literature on the cost and benefits of financial innovation. Several studies have underlined potential adverse effects of financial innovation, such as speculation (Simsek, 2013) or rent extraction (Biais, Rochet, and Woolley, 2015; Biais and Landier, 2018), particularly from unsophisticated agents (Carlin, 2009). The present paper illustrates how innovative financial products may also benefit unsophisticated market players by mitigating investor behavioral biases. This mechanism differs from and complements the more traditional role of financial innovation to improve risk-sharing and complete markets (Ross, 1976; Calvet, Gonzalez-Eiras, and Sodini, 2004).

This study adds to the literature examining how to tailor security design to investor preferences or beliefs. Célérier and Vallée (2017) document how banks design financial products to cater to yield-seeking investors, which allows them to charge larger markups.[9] The present paper further establishes that security design is a powerful tool for affecting economic decisions. In contrast to earlier work, however, we focus on the bright side and show that security design can foster actions beneficial to investors. Our paper therefore brings nuance to the prevailing negative view of tailored security design. In this respect, our findings expand the literature advocating contract design as a possible solution to behavioral biases (DellaVigna and Malmendier, 2004).

[8]The security design we study does not mitigate a bias by exploiting another one, such as inertia for default options or gambling propensity for lottery-saving accounts (Cole, Iverson, and Tufano, 2018).

[9]Henderson and Pearson (2011), Li, Subrahmanyam, and Yang (2018), and Vokata (2019) also focus on the dark side of non-linear products.

The paper is organized as follows. Section II provides background on retail capital guarantee products and presents the data for our empirical analysis. Section III describes the product design, and develops an asset pricing model to measure their expected returns and markups. In Section IV we test whether investing in capital guarantee products induces a causal increase in household risk-taking. Section V develops a theoretical life-cycle model of portfolio allocation to study the mechanisms that can explain the empirical effects we document. In Section VI, we measure the welfare gains from financial innovation and how they are divided between product providers and households. Section VII concludes. An Internet Appendix provides derivations and additional empirical results.

II. Background and Data

A. Background on Capital Guarantee Products and their Introduction in Sweden

Capital guarantee products are retail investments that offer exposure to the upside potential of risky assets and protect a substantial part of the invested capital, typically close to 100%.

Retail CGPs are widespread around the world. As of 2015, their total outstanding volumes exceed $4.5 trillion. Table II provides country-level outstanding volumes for the largest classes of CGPs. In the United States, guaranteed variable annuities represent a $1.7 trillion market (Ellul, Jotikasthira, Kartasheva, Lundblad, and Wagner, 2020). In France, Euro-life insurance contracts amount to $1.5 trillion, or 60% of GDP (Hombert and Lyonnet, 2020). In China, guaranteed wealth management products account for $854 billion. Finally, global outstanding volumes of retail structured products with a capital protection exceed $400 billion.[10] The pervasiveness and large volume outstanding of CGPs suggest that their design strongly appeals to retail investors.

[10]The risky assets covered by this list of products include public equities, bonds, and loans.

Financial institutions use three main approaches to structure a capital guarantee product. They can choose to design a synthetic product, implement a portfolio insurance strategy, or build reserves. Synthetic CGPs, also referred to as retail structured products with a capital protection, are passive, limited-horizon products with a non-linear payoff that depends on the performance of their underlying asset (Célérier and Vallée, 2017).[11]

The first synthetic CGPs were created in the United Kingdom in the early 1990s. These products initially targeted institutional investors. However, financial institutions quickly rolled-out the products to their retail client bases, as they discovered the popularity of the CGPs among retail investors. Then, the technology spread to other European countries over the decade and reached Sweden in the early 2000s.

We exploit the introduction of CGPs in Sweden over the 2002-2007 period as a laboratory to study the impact of security design on household risk-taking for the following reasons. First, CGPs were adopted quickly and broadly in this country, reaching 14% of the population within 5 years of their introduction. Figure IA.1 in the Internet Appendix illustrates the speed and depth of the adoption of CGPs in Sweden over the period. This choice is further supported by the unmatched quality and scope of Swedish data on household financial holdings and demographics, as described in the following section. Finally, the structure of the retail market for financial products in Sweden, where banks play a dominant role, allows us to develop an identification strategy aimed at establishing a causal claim.

In Table I, we document the low level of financial risk-taking by a substantial share of households across countries. We consider a selected set of countries for which data are available and provide summary statistics on household participation in equity markets. In Sweden, as of 2015, 17% of household total financial wealth is invested in equity (column 1), 68% of households that are 50 years or older participate in equity markets (column 2), and the median participating household invests 37% of its financial wealth in equity (column

[11] Portfolio insurance is a dynamic trading strategy aimed at managing downside risk. Reserves are built by the product provider to offset fluctuations in asset returns, as is the case for Euro life insurance contracts in France (Hombert and Lyonnet, 2020).

3).[12] While modest compared to predictions from standard portfolio allocation models, these levels are relatively high by international standards. Therefore, the relationships between CGPs and risk-taking that the present paper document in Sweden are likely to be stronger in other countries.

For expositional simplicity, one can conveniently consider three groups of households: (i) non-participants, (ii) participants with very low shares of financial wealth invested in equity, and (iii) households with substantial equity holdings in line with the prescriptions of canonical models (Campbell and Viceira, 2002; Cocco et al., 2005; Merton, 1971). The household finance literature devotes considerable attention to non-participants, which include a large group of households with very low levels of financial wealth and therefore limited ability to invest in stocks. For instance, in the U.S., 40% of nonparticipants older than 50 own less than $500 in financial wealth according to the Health and Retirement Survey.

Participants with low risky shares are much wealthier than non-participants and are therefore a sizable potential source of capital, which has received less attention in the literature. For instance, in the U.S., participants with an equity share lower than 20% represent 40% of all participants and have a median financial wealth equal to $272,000, compared to $5,000 for non-participants and $370,000 for participants with risky shares above 20%. Similar results hold in Sweden.[13] As a consequence, participants with low risky shares forego similar returns but much higher financial income compared to non-participants. As we will show in Section IV, this intermediate group is the most impacted by the introduction of CGPs in Sweden and is therefore central to our study.[14]

[12]Section 2 in the Internet Appendix provides details on the methodology used to obtain these statistics.

[13]In Sweden, participants with equity shares less than 20% represent 31% of all participants and have a median financial wealth of $30,000, compared to $3,000 for non participants.

[14]While nonparticipation is an important issue in many countries, Sweden is not the best country to study this topic because it has one of the highest participation rates in the world in 2002 and nonparticipants have very little wealth to invest.

B. Data

Our empirical analysis relies on a data set on all the synthetic CGPs and mutual funds sold to Swedish retail investors over the 2002-2007 period, merged with a data set on the portfolio composition and socio-demographic characteristics of all Swedish households over the same period.

1. Capital Guarantee Products and Equity Mutual Funds. The data set contains detailed information on the synthetic CGPs sold to Swedish retail investors between 2002 and 2007, which we retrieved from the Célérier and Vallée (2017) database of European retail structured products. It includes the underlying instrument, maturity, volumes, and disclosed fees of every CGP sold in Sweden, as well as text from which we obtain the payoff formula of each contract.[15] Panel A of Table III reports summary statistics. The sample contains 1,511 equity-linked contracts issued over the 2002 to 2007 period, for a total volume of $8 billion.[16]

For equity mutual funds, we obtain the historical fees, age, family, and geographical scope from each fund's fact sheet. The reported fees include transaction costs, operating costs, and management fees. The returns, volatility, and dividend distributions of mutual funds, and the historical returns and volatility of the instruments underlying CGPs are retrieved from Bloomberg, Datastream, and FinBas.[17]

2. Household Demographics, Income, and Wealth. The administrative household panel, described in Calvet et al. (2007), contains the demographics, income, and disaggregated financial holdings of every Swedish household between 2000 and 2007. Demographic and income variables include the age, gender, education level, parish of residence, and income of each member of a household. The panel's distinguishing feature is that it contains the comprehensive disaggregated financial holdings of each household, including the positions in

[15]See Célérier and Vallée (2017) for the description of the textual analysis involved.

[16]In Sweden, the large majority of CGPs offer equity exposure (87% of the products).

[17]FinBas is a financial database maintained by the Swedish House of Finance.

cash, equity mutual funds, stocks, and CGPs at the level of each account or security.[18] The security-level information is identified by the International Security Identification Number (ISIN). The panel also provides a unique identifier for the institution where each bank account is held.

The household panel covers the entire population of Sweden and provides the exact portfolio composition of each household. It is highly reliable because the wealth information is collected by Statistics Sweden for tax purposes and is incorporated in tax forms, which households then have an opportunity to correct in case of a mistake. Statistics Sweden collects this information from a variety of sources, including the Swedish Tax Agency, welfare agencies, and private employers. Financial institutions supply to the tax agency their customers' deposits, interest paid or received, security investments, and dividends.[19]

We construct the merged household panel as follows. We filter out households with a head younger than 25 years or with financial wealth lower than $200 in 2002. We then only keep households that are observable over the whole sample period, consistent with our aim to investigate the effects of capital guarantee products on household risk-taking over the 2002-2007 period.[20] Our final panel contains 3,107,893 households. We merge it with the CGP and equity fund data via the unique ISIN identifier. The high-quality panel covers the launch and subsequent high growth of the market for CGPs in Sweden.

C. *Summary Statistics*

Table IV reports demographic and financial characteristics for the full sample of 3.1 million households, the subsample of 2.1 million households that participate in equity markets in 2002 (68.5% of the total sample), and the subsample of 430,000 households that invest in

[18]Bonds and bond mutual funds, which we can also observe, are infrequent.

[19]The panel does not report defined contribution pension savings. These pension savings include assets in private pension plans and in public defined contribution accounts that were established in a 1999 pension reform. According to official statistics, defined contribution pension savings had an aggregate value of $25.6 billion in Sweden at the end of 2002, whereas aggregate household financial wealth invested outside pension plans amounted to $131.3 billion.

[20]In our data set, a household exits every time the composition of adults of the household changes, due to either death, divorce, marriage or change in partnership.

CGPs at least once over the sample period (13.9% of the full sample).

Panel A of Table IV focuses on 2002. While equity participation is relatively high in Sweden compared to other developed economies, the share of financial wealth invested in risky assets conditional on participation is 32.9% on average. Participants mostly take financial risk by investing in equity funds, which represent 22.9% of financial wealth on average (median = 16.9%), and individual stocks, which represent 9.3% of financial wealth on average (median = 1.4%). Moreover, household characteristics such as financial wealth, age, and income vary substantially across groups, which calls for using precise controls in the empirical analysis.

Panel B of Table IV illustrates that CGPs quickly gained traction within a few years. At the end of 2007, 13.9% of Swedish households had participated at least once in the new asset class, and participants allocated on average 11.9% (median = 7.3%) of financial wealth to these products.

III. Design, Expected Return, and Markup

In this section, we compute the risk premia that capital guarantee products provide to investors, and the gross markups earned by the financial institutions that market them. For this purpose, we develop a no-arbitrage pricing methodology that captures the specificities of these contracts, such as their option features, issue price, the dividend yield of the underlying instrument, and issuer credit risk. We document that CGPs offer a share of the equity premium that is slightly lower than the share offered after fees by equity mutual funds, the most popular form of household risky investments. In addition, the gross markups earned by sellers of CGPs are comparable to the gross markups earned by mutual fund companies.

A. Product Design

The majority of CGPs in our sample have the following design. The contract is sold at time $t = 0$ at the issue price P_0 and face value F, and reaches maturity at time M. The product offers upside potential by allowing the household to earn at maturity a fraction p of a benchmark return, R^*, applied to the face value F. The benchmark R^* is defined by the returns on an underlying asset, index, or basket of indexes. The contract also offers downside protection by offering a guaranteed net rate of return, g, on face value.

Capital guarantee products are typically structured as notes and therefore bear the credit risk of the bank structuring them. Let $\xi \in [0, 1]$ denote the random fraction of pledged cash flows that is paid at maturity, commonly called the payoff ratio (Jarrow, 2019; Jarrow and Turnbull, 1995). The gross return on the CGP is

$$1 + R_g = \frac{F}{P_0} \left[1 + \max(p\, R^*; g)\right] \xi \tag{1}$$

between issuance and maturity.

The benchmark return R^* is the average ex-dividend performance of the underlying measured at prespecified dates $t_1 < \cdots < t_n$:

$$1 + R^* = \frac{S_{t_1} + S_{t_2} + \ldots + S_{t_n}}{n S_{t_0}}, \tag{2}$$

where S_{t_0} is the initial reference level of an index or asset at t_0, which is typically the day of issuance or shortly thereafter. We call $t_n - t_1$ the length of the Asian option. If $n = 1$, the option is European and the length of the Asian option is equal to 0.

Panel B of Table III provides summary statistics. Contracts with this representative design account for 54% of CGPs issued in Sweden during our sample, and 60% of volumes. The average volume of an issuance is around \$5 million. The median maturity M is 4 years, the median net rate of guarantee g is 0%, the median issue price is 110% of face value, and the median participation rate p is 1.10.[21] We note that to this date, no default has occurred on CGPs sold to Swedish retail investors.

B. *Expected Return and Markup: Methodology*

We develop a no-arbitrage pricing method designed to compute the risk and return of CGPs. The model is based on the following assumptions. Under the physical measure $\mathbb{P}$, the underlying follows a geometric Brownian motion:

$$\frac{dS_t}{S_t} = (\mu - q)dt + \sigma dZ_t, \tag{3}$$

where μ is the drift, q is the dividend yield, and σ denotes volatility. The payoff ratio ξ is independent of the underlying, consistent with the view that credit risk is driven by operational risk. Let r_f denote the continuous-time interest rate. Under the risk-adjusted measure $\mathbb{Q}$, the drift of the underlying is $r_f - q$. We consider for simplicity that the payout ratio's distribution and independence from the underlying are not impacted by the change of measure.

The expected return on the CGP over the life of the contract is given by:

$$\mathbb{E}_0^{\mathbb{P}}(1 + R_g) = (1 - \kappa)\,\frac{F}{P_0}\,\mathbb{E}_0^{\mathbb{P}}[1 + \max(p\,R^*; g)], \tag{4}$$

where $1 - \kappa = \mathbb{E}_0^{\mathbb{P}}(\xi)$ denote the expected payoff on a \$1 promise. This approach provides conservatively low estimates of the expected return if default is more likely when the underlying is low.[22] In practice, we compute the expected return (4) as follows. We obtain

[21] A product can offer both substantial capital protection and a participation rate higher than unity because of the Asian option feature and the ex-dividend nature of the benchmark return, as Section III.B further explains.

[22] The expected return

$$\mathbb{E}_0^{\mathbb{P}}(1 + R_g) = (1 - \kappa)\,\frac{F}{P_0}\left\{1 + \mathbb{E}_0^{\mathbb{P}}[\max(p\,R^*; g)]\right\} + \frac{F}{P_0} Cov[\xi, \max(p\,R^*; g)]$$

is higher than (4) entails if the payoff ratio ξ and the benchmark return R^* co-move positively, that is if

$\mathbb{E}_0^{\mathbb{P}}[1 + \max(p\,R^*; g)]$ by Monte Carlo simulations of the underlying, as we explain in the Internet Appendix. We set κ equal to the CDS spread of the issuer.[23]

The fair issue price, $P_0^{\text{fair}} = (1-\kappa)\,F\,e^{-r_f\,M}\,\mathbb{E}_0^{\mathbb{Q}}[1+\max(p\,R^*; g)]$, is the price that equates the expected return of the contract under $\mathbb{Q}$ to the return on a riskless bond of same maturity. It is also conveniently computed by Monte Carlo.

The *gross markup* of the contract, $(P_0 - P_0^{\text{fair}})/P_0$, is the difference between the market issue price and the fair issue price divided by the market issue price. To compare it to the stream of fees generated by standard funds, consider a mutual fund company that charges a fraction φ of asset value at the beginning of each year. An initial investment of \$1 generates over M periods a flow of fees equal to $\sum_{t=0}^{M-1} \varphi\,(1-\varphi)^t = 1-(1-\varphi)^M$.[24] The gross markup on the CGP coincides with the fair value of the stream of fund fees if

$$\varphi_{CGP} = 1 - (P_0^{\text{fair}}/P_0)^{1/M}. \tag{5}$$

This formula allows us to convert a CGP's markup into its yearly mutual fund fee equivalent.

The baseline products in our sample cover 155 different underlying instruments, which can be a stock index, a basket of stock indices, or a basket of stocks. For each underlying, we estimate the risk premium at the monthly frequency, $\mathbb{E}(R_{i,t})$, by applying the World CAPM over the longest time-series available and a world market risk premium of 6%. We set the model's yearly drift μ_i to $12\,\ln[1 + \mathbb{E}(R_{i,t})]$, the volatility parameter σ_i to the historical volatility over the 1990-2007 period, and q_i to the latest dividend yield before issuance. We use the M-year SEK swap rate as the risk-free rate in the pricing model. The yearly expected excess return earned by an investor on a CGP is the difference between the product's annualized expected markup return and the annual yield on an M-year Swedish Treasury bond.

default is more likely in bad times than in good times.

[23] Since the CDS swap typically includes a risk premium, this choice produces conservatively low values of $\mathbb{E}_0^{\mathbb{P}}(1 + R_g)$

[24] This formula holds if the household invests \$1 at $t = 0$, keeps its investment in the fund until $t = M$, and makes no intermediate withholdings or contributions. See the Internet Appendix for the derivation.

C. Expected Return and Markup: Results

Panel A of Figure 1 displays the distribution of expected excess returns and yearly markups of CGPs sold in Sweden during the sample period. Corresponding key statistics are reported in Panel B of Table III. There are two take-aways. First, the expected excess return on CGPs is significantly positive and amounts to 2.7% per year on average, or close to half the premium on the world index.[25] More than 90% of products earn a positive risk premium. These results confirm that retail CGPs allow households to earn a significant part of the risk premium. Second, the average markups earned by banks on CGPs are equivalent to an annual fee $\varphi = 1.6\%$. In Table IA.1 of the Internet Appendix, we verify that these results are robust to alternative parameter choices.

For comparison purposes, Panel B of Figure 1 report the expected return and fees of equity mutual funds available to Swedish retail investors over the 2002-2007 period. We compute expected returns by applying the World CAPM and deducting fees. Beta coefficients are estimated from the historical returns of each fund over the longest period available. Equity funds have an average beta of 0.9 relative to the World Index and therefore a risk premium before fees of $0.9 \times 6\% = 5.4\%$ per year. Fees, which include transaction costs, operating costs, and management fees, amount to 2.1% per year on average during our sample period. The average expected excess return on equity funds is therefore 3.3% in annual units, or a fraction of about 55% of the world equity premium.

Overall, capital guarantee products and mutual funds exhibit comparable expected returns and similar markups on average. This finding suggests that banks have equivalent financial incentives to market equity funds and CGPs to retail investors.[26]

[25]Our share estimate exhibits little sensitivity to the value of the world index equity premium we assume.
[26]Discussions with practitioners also support this hypothesis.

IV. Measuring the Impact of Capital Guarantee Products on Household Risk-Taking

We have shown that the capital guarantee products marketed to Swedish households offer a substantial fraction of the equity premium even when accounting for embedded markups. In this section, we test whether the introduction of these products has an impact on household risk-taking.

A. *Measuring Household Risk-Taking*

The literature usually measures household risk-taking as the share of financial wealth invested in equity products (e.g. Calvet, Campbell, and Sodini (2009)). One limitation of this approach is that diverse equity products, such as stocks, mutual funds, allocation funds, and CGPs, tend to earn heterogeneous risk premia that vary with design, maturity, and fees. For this reason, we now develop a novel measure of equity market exposure.

We define the *risk-taking index* of a risky product p as the fraction of the market risk premium it provides investors:

$$\eta_p = \frac{[\mathbb{E}(1+R_p)]^{\frac{1}{M}} - e^{r_f}}{\mathbb{E}(1+R_m) - e^{r_f}}, \tag{6}$$

where M denotes product maturity, R_p the net arithmetic return on the product, R_m the net return on the world index, and r_f the average log yield on Swedish 1-year Treasury bonds.[27] We set $M = 1$ for a liquid product. The measure (6) intentionally focuses on the compensation for risk-taking, which motivates participation in risky assets markets, and not on downside risk.[28]

We obtain η_p for equity products in our sample as follows. The asset pricing results of

[27]The log yield satisfies $r_f = \ln(1 + R_f)$, where R_f is the yearly arithmetic yield on Swedish Treasury bonds. The yield R_f is 3.5% on average over the period.

[28]Our results are robust to using traditional measures of risk-taking such as the risky share or volatility. The index η_p is more conservative because it takes into account the design and fees of financial products.

Section III.C give the expected returns $[\mathbb{E}(1+R_p)]^{\frac{1}{M}}$ on CGPs and equity mutual funds. For the subsample of CGPs that we do not price, we use the average η_p in the sample of baseline CGPs. For stocks and exchange traded funds (ETFs), we assume that management fees amount to 0.2% and 0.5%, respectively, and that the World CAPM β is unity. We also assume that $\eta_p = 0.3$ for allocation funds, which represent around 2% of household financial wealth.[29]

Panels B and C of Table III provide summary statistics on the risk-taking index η_p of CGPs and equity mutual funds. As expected, CGPs offer a relatively lower fraction of the equity premium than equity mutual funds. The average risk-taking index is 0.44 for CGPs and 0.55 for equity funds. The gap is limited in part because the beta coefficient is on average higher for CGPs ($\beta = 1.1$) than for equity mutual funds ($\beta = 0.9$).

We define the *risk-taking index* of household h in period t by:

$$\eta_{h,t} = \sum_{p=1}^{n} Share_{p,h,t} \times \eta_p,$$

where $Share_{p,h,t}$ is the share of product p in the household's financial wealth in period t. The sum is taken over all CGPs, equity mutual funds, stocks, ETFs, and allocation funds.

Panel C of Table IV provides summary statistics on the risk-taking index of households. In 2002, the average index is 0.22 for stock market participants and 0.26 for CGP participants. Between 2002 and 2007, the proportional change in the index is 0.7% for stock market participants versus 17.6% for CGP participants, which suggests a positive correlation between risk-taking and CGP investing.

B. OLS Results: Capital Guarantee Products and Risk-Taking

1. Total Change in Risk Taking. We now investigate whether CGP investing is associated with an increase in household risk-taking. Panel A of Figure 2 plots the *risk-taking index*

[29] Allocation funds are hybrid funds that combine equity funds and money market funds.

in 2002 and in 2007 for: (i) households that participate at least once in capital guarantee products over the sample period, and (ii) a control group of equal size containing stock market participants matched based on their 2002 risk-taking index. The two groups exhibit diverging risk-taking indexes over the sample period. While by construction the gap between the two groups is close to zero in 2002, it increases to 2 pp, or more than 6% of the 2002 risk-taking index, by the end of the sample period.

In Panel B of Figure 2, we apply the same analysis to households in the bottom quartile of risk-taking index in 2002. The divergence in risk-taking index between CGP participants and the matched control group is significantly more pronounced than in Panel A, with a gap in risk-taking index of 8 pp in 2007. This gap is particularly large when compared to the baseline risk-taking index of 2 pp for this subsample in 2002. This finding suggests some heterogeneity across households in the extent of the relationship between CGP participation and change in risk-taking.

In column 1 of Table V, we confirm this result by running a cross-sectional regression of the evolution of the risk-taking index in the sample of 2002 equity market participants:[30]

$$\Delta_{2007,2002}(\eta_h) = \alpha + \beta_1 \mathbb{1}_{CGP,h} + \lambda' x_{h,2002} + \varepsilon_h. \tag{7}$$

In this regression, $\Delta_{2007,2002}(\eta_h)$ denotes the Davis and Haltiwanger (1992) growth rate of the index,[31] $\mathbb{1}_{CGP,h}$ is an indicator variable equal to unity if the household purchases a CGP at least once during the sample period, $x_{h,2002}$ is a vector of household characteristics in 2002, and ε_h is an error term. Characteristics include the percentage change in income and in financial wealth over the period, as well fixed effects for the number of children, household size, gender, locality, years of education, and deciles of financial wealth, income, age and

[30]We therefore estimate the effect at the intensive margin. Our results are robust to including the whole population. However, effects on the extensive margin are minimal, which could be due to the high level of stock market participation in Sweden, or to the existence of a fixed cost to participation, which would not be alleviated by CGPs.

[31]The Davis and Haltiwanger (1992) growth measure, $\Delta_{2007,2002}(\eta_h) = 2(\eta_{h,2007} - \eta_{h,2002})/(|\eta_{h,2007}| + |\eta_{h,2002}|)$, limits the extreme values created by low denominator values in a standard growth rate.

risky share. The coefficient of the variable $\mathbb{1}_{CGP,h}$ confirms that households that participate in CGPs increase their risk-taking index significantly more than households that do not. The percentage change in the index is 24 pp higher for CGP participants, while the average household increases its index by only 0.7 pp over the period. This magnitude is comparable to the increase in risk-taking resulting from having access to a financial advisor, as estimated in Chalmers and Reuter (2020). However, the effect we document applies to a larger base: the household's entire financial wealth instead of a single retirement investment account.

2. Active Change in Risk-Taking. We now show that the heterogeneous response of risk-taking to innovation is driven by active investment decisions and not simply by the mechanical effect of realized asset returns.[32] To do so, we measure the active change in the risk-taking index of household h between t and $t+n$, $\Delta^{A}_{t,t+n}(\eta_h)$, as the Davis and Haltiwanger (1992) growth rate between the initial index, $\eta_{h,t}$, and the *market-neutral risk-taking index* $\eta^{\text{MN}}_{h,t+n}$ in year $t+n$, which we define as follows. The market neutral index is the index that the household would achieve if all asset returns were equal to zero.[33] By construction, $\eta^{\text{MN}}_{h,t+n}$ only differs from $\eta_{h,t}$ as a result of active trading and saving decisions.

Figure IA.2 in the Internet Appendix reproduces Figure 2 using the market-neutral risk-taking index. As for the risk taking index, CGP participants and the matched control group exhibit diverging trends.

In column 3 of Table V, we regress the active change $\Delta^{A}_{2002,2007}(\eta_h)$ on CGP participation and household characteristics. The active change associated with CGP participation is

[32]Since active allocation decisions might be in part responses to passive performance, we view both exercises as complementary.

[33]The market-neutral risk-taking index is defined by $\eta^{\text{MN}}_{h,t+n} = \sum_{p=1}^{n} \eta_p \, Share^{\text{MN}}_{p,h,t+n}$, where $Share^{\text{MN}}_{h,p,t+n}$ is the share of product p in year $t+n$, adjusted for the mechanical changes due to realized asset returns from year t to $t+n$. Specifically,

$$Share^{\text{MN}}_{p,h,t+n} = \frac{X_{p,h,t} + \sum_{s=t+1}^{t+n}[X_{p,h,s} - (1+R_{p,h,s})\,X_{p,h,s-1}]}{FW_{h,t} + \sum_{s=t+1}^{t+n}[FW_{h,s} - (1+R_{h,s})\,FW_{h,s-1}]},$$

where $X_{p,h,s}$ is the amount invested in product p at date s, $R_{p,h,s}$ is the yearly realized return of product p from year $s-1$ to s, $FW_{h,s}$ is the total financial wealth, and $R_{h,s}$ is the return on financial wealth. Values are winsorized at the 1% level.

comparable to the result obtained with the total change in the index, which rules out that our results are purely mechanical.

3. Panel Model. The following panel specification allows us to measure the sensitivity of the risk-taking index to the purchased quantity of capital guarantee products:

$$\eta_{h,t} = \alpha + \beta_2 \, CGP \ Share_{h,t} + \lambda' x_{h,t} + \gamma_h + \mu_t + \varepsilon_{h,t}, \tag{8}$$

where $CGP \ Share_{h,t}$ is the share of CGPs in household h's financial wealth, $x_{h,t}$ is a vector of characteristics, γ_h is a household fixed effect, μ_t is a time fixed effect, and $\varepsilon_{h,t}$ is a stochastic error.

If a household fully funds CGP purchases from bank deposits, the linear coefficient β_2 is approximately equal to the average risk-taking index of CGPs. By contrast, if a household views CGPs as perfect substitutes for traditional equity products, it funds CGP purchases by selling traditional products and β_2 can be negative. We report the panel regression results in Table VI. The point estimate of β_2 is 0.21, around half of the average risk-taking index of CGPs. We find similar results when the market-neutral risk-taking index $\eta_{h,t}^{MN}$ is used as the dependent variable.[34]

C. Heterogeneity along Household's Willingness to Take Risk

1. Main Result. We now show that the increase in risk-taking associated with CGP investing tends to vary substantially with a household's initial willingness to take risk, as Panel B of Figure 2 suggests. We measure this willingness by filtering out household characteristics from the initial risk-taking index. That is, we write $\eta_{h,2002} = \bar{\eta}_h + b'(x_h - \bar{x}) + e_h$, where $\bar{\eta}_{2002}$ and $\bar{x}$ respectively denote the sample means of $\eta_{h,2002}$ and x_h. Hence $\eta_{h,2002}^{\mathrm{F}} = \eta_{h,2002} - b'(x_h - \bar{x})$ represents the household's initial willingness to take risk that is not

[34] The coefficient β is slightly stronger, consistent with the fact that capital guarantee products are valued at issuance price while traditional equity products are marked to market in our data.

captured by observable characteristics.

Figure 3 illustrates the relationship between a household's willingness to take risk and the change in risk-taking for adopters of CGPs. To construct the figure, we regress the household change in the risk-taking index over the 2002-2007 period on the indicator variable $\mathbb{1}_{CGP,h}$ interacted with the filtered risk-taking index in 2002:

$$\eta_{h,2007} - \eta_{h,2002} = \alpha + \beta_3 \, \mathbb{1}_{CGP,h} + \beta_4 \, \mathbb{1}_{CGP,h} \times \eta^{\mathrm{F}}_{h,2002} + \lambda' x_{h,2002} + \varepsilon_h, \tag{9}$$

where $x_{h,2002}$ includes fixed effects for deciles of wealth, income, and age, as well as the income change over the period. We then plot the proportional change in risk-taking for CGP participants, e.g. the ratio of the predicted incremental change in the risk-taking index for CGP participants vs. non-participants to their period-average of risk-taking index $(\eta_{h,2002} + \eta_{h,2007})/2$, as a function of their filtered risk-taking index in 2002.[35]

The incremental increase in risk-taking for CGP adopters monotonically falls with the initial willingness to take risk.[36] The magnitude is particularly large for households with a low initial willingness to take risk. For households with a filtered 2002 risk-taking index below 0.10, the adoption of CGPs result in an increase in the risk-taking index of more than 60%. By contrast, the effect is close to zero for households that have a filtered 2002 risk-taking index above the median, or 0.17.

In columns 2 to 5 of Table VI, we confirm these results by estimating equation (8) within each quartiles of filtered 2002 risk-taking index. The coefficient β_4 is a decreasing function of their initial willingness to take risk.

2. Mechanism. To better understand the mechanisms at play, we explore whether the demand for CGPs increases with household willingness to take risk, as is the case for stocks and mutual funds, or decreases with it, as is the case for bank deposits. We consider four

[35]We scale by $(\eta_{h,2002} + \eta_{h,2007})/2$ and not by $\eta_{h,2002}$ to reduce distortions when $\eta_{h,2002}$ is close to zero.

[36]We obtain a comparable result when using the ex-ante bank deposit share of the financial wealth as a proxy for household (un-)willingness to take risk.

asset classes: CGPs, bank deposits, stocks, and equity mutual funds. For each asset class j, we run the OLS regression of the share of financial wealth invested in the class at the end of 2007, $Share_{j,h}$, on the willingness to take risk:[37]

$$Share_{j,h} = \alpha_j + \beta_5 \, \eta^{\mathrm{F}}_{h,2002} + \lambda_j' x_{h,2002} + \varepsilon_{h,j}. \tag{10}$$

The vector of characteristics, $x_{h,2002}$, includes fixed effects for deciles of financial wealth, income, age, and years of education in 2002.

Figure 4 plots the predicted share of financial wealth invested in each asset class in 2007 as a function of the filtered 2002 risk-taking index. The share of stocks and mutual funds in 2007 is positively correlated with the initial willingness to take risk. This strong correlation is consistent with the persistence of household preferences and portfolio allocations, as the 2002 index is driven by stock and fund holdings. By contrast, the share of CGPs and the share of bank deposits are both negatively correlated with the initial willingness to take risk. The patterns of investment in CGPs are therefore similar to the patterns observed for bank deposits but opposite to the patterns of traditional equity products. These results suggest that households perceive CGPs to be closer to bank deposits than to traditional equity products, most likely because both protect the capital invested.[38]

D. *Instrumental Variable Analysis*

Our baseline result is a within-household positive correlation between risk-taking and CGP investing, controlling for a comprehensive set of time-varying household characteristics. Such correlation should be interpreted causally with caution. The share of capital guarantee products, $CGP\ Share_{h,t}$, and the error term of the structural equation (8), $\varepsilon_{h,t}$, may be driven by the same time-varying latent variables, such as the household's time-varying idiosyncratic willingness to take risk not predicted by characteristics. This endogeneity is-

[37] For each regression, we restrict the sample to participants in this given class.

[38] Bank deposits and capital guarantee products also have significantly different levels of liquidity.

sue could bias OLS estimates downward or upward. Therefore, we develop an instrumental variable estimation of the structural equation (8), which we implement by two-stage least squares (2SLS).

Design. We instrument the CGP share, $CGP\ Share_{h,t}$, by a measure of supply of capital guarantee products in year t from the banks with which household h has the strongest relationship at the beginning of the sample period. To do so, we exploit information on the identity of all the banks households receive interest income from. About two thirds of the sample of stock market participants declare an interest income.

The instrument is motivated by the evidence suggesting that bank supply largely drives CGP volumes. Figure IA.4 of the Internet Appendix illustrates the strong correlation of the CGP volumes issued inside and outside Sweden by Swedish banks. Table IA.2 takes a more systematic approach and documents that bank-year fixed effects have significantly more explanatory power than country-year fixed effects for explaining the volume of CGPs sold by a given bank in a given country in a given year. When we introduce bank-year fixed effects in addition to country-year fixed effects that should absorb local demand to a certain extent, the adjusted R^2 increases from 0.03 to 0.19. Possible explanations for strong supply effects include securing access to a structuring desk and marketing efforts.

Let $\theta_{h,b}$ denote the indicator variable equal to unity if bank $b \in \{1, \ldots, B\}$ is the bank where household h deposits the largest share of cash at the beginning of the sample period, and let $\theta_h = (\theta_{h,1}, \ldots, \theta_{h,B})'$. We instrument the CGP share of household h hold in year t by

$$Z_{h,t} = \hat{\Phi}_t' \theta_h,$$

where $\hat{\Phi}_t$ is a measure of bank supply shocks.

The instrument is valid if the following condition holds.

Identifying Restriction 1. *The exogeneity condition* $\mathbb{E}(\hat{\Phi}_t' \theta_h \, \varepsilon_{h,t}) = 0$ *holds for every* h *and* t, *where* $\varepsilon_{h,t}$ *is the error term of the structural equation* (8).

That is, supply shocks are exogenous to time-varying unobservable characteristics that might drive household portfolio decisions. Similar to Borusyak et al. (2018), our strategy does not require that the matching between households and banks be exogenous.

In the first stage of 2SLS, we regress the share of capital guarantee products on the instrument, household characteristics, and household and time fixed effects:

$$CGP\ Share_{h,t} = \alpha + \beta_6\, \hat{\Phi}_t' \theta_h + \lambda'\, x_{h,t} + \gamma_h + \mu_t + u_{h,t}, \tag{11}$$

where $x_{h,t}$ includes time-varying household characteristics that are driving the demand for CGPs, and $u_{h,t}$ is a stochastic error term. In the second stage, we estimate:

$$\eta_{h,t} = \alpha + \beta\, \widehat{CGP\ Share}_{h,t} + \lambda' x_{h,t} + \gamma_h + \mu_t + v_{h,t}, \tag{12}$$

where $\widehat{CGP\ Share}_{h,t}$ is the predicted share from the first stage.

Measuring the Banks' Time-Varying Supply Shocks. A first approach is to use banks' CGP issuance per depositor as a proxy for supply shocks. This approach is motivated by the previously described evidence that bank supply drives total volumes. While this first approach has the advantage of simplicity, it may not satisfy the identification restriction. Total volumes can also be driven by demand factors that vary heterogeneously across banks along with unobservable household characteristics, which may imply that $\mathbb{E}(Z_{h,t}\, \varepsilon_t) \neq 0$.

In a second approach, we address this issue by filtering out demand effects and trends from the volumes offered by banks, thereby focusing on idiosyncratic supply shocks at the bank level. We obtain $\hat{\Phi}_t$ by estimating the panel regression:

$$CGP\ Share_{h,t} = \alpha + \Phi_t' \theta_h + \lambda'\, x_{h,t} + \gamma_h + \mu_t + w_{h,t}, \tag{13}$$

where $x_{h,t}$ includes the same set of time-varying fixed effects interacted with year fixed effects as in the structural equation (8), and $w_{h,t}$ is a stochastic error term with zero mean.

Importantly, (11) is a random coefficients model, because the vector of linear coefficients Φ_t is allowed to vary randomly through time. We make the following assumption.

Identifying Restriction 2. *The error term $w_{h,t}$ of equation* (13) *satisfy* $\mathbb{E}(\theta_h\, w_{h,t}) = 0$ *for every h and t.*

This restriction is reasonable to the extent that θ_h is not time-varying.

To further ensure that the estimator $\hat{\Phi}_t$ produces a valid instrument $Z_{h,t} = \hat{\Phi}_t'\theta_h$, we randomly partition the household population into two sub-samples of equal size. We estimate the idiosyncratic supply shock $\hat{\Phi}_t$ on the first sub-sample, and run the second stage of 2SLS on the other sub-sample. By doing so, we reduce the likelihood that unobservable characteristics of households in the first sub-sample are correlated with the error terms $\varepsilon_{h,t}$ of households in the second sub-sample.[39]

In practice, households have multiple banking relationships. We also use as instruments the supply shocks of the second and third banks with which the household has the largest balances.

Results. In columns 1 and 2 of Table VII, we instrument a household's CGP share in year t by the issuance of CGPs per depositor from the household's main banks during the year. Standard errors are clustered at the bank $\times$ year level, the level of granularity of the instrumental variable. The regression coefficients are consistent with a positive causal effect of CGP investing on household risk-taking.

Columns 3 and 4 of Table VII report the regression coefficients of 2SLS estimated on the second half of the sample, the first half having been used to estimate the supply shocks $\hat{\Phi}_t$. Column 3 displays the coefficients of the first stage. A higher supply intensity of CGPs from a given bank significantly increases CGP investments by households in a relationship

[39] The estimator of the idiosyncratic supply shock at date t can be written as $\hat{\Phi}_t = \Phi_t + A\, w$, where w is the vector of yearly errors $w_{h,t}$ of households in the first subsample and the matrix A is a function of the observations $\theta_{h,t}$ and $x_{h,t}$. Since $\mathbb{E}(Z_t\, \varepsilon_{h,t}) = \mathbb{E}(\Phi_t'\, \varepsilon_{h,t}) + \mathbb{E}(w'\, A'\, \varepsilon_{h,t}) = 0$, the instrument is valid if supply shocks are uncorrelated with the error terms of the structural equation and if observations of households in different subsamples are independent.

with this bank, even when controlling for detailed time-varying household characteristics in a panel specification. The F-statistic of the first stage, at 569, is significantly above the threshold for strong instruments (Stock and Yogo, 2005).

Column 4 provides the coefficients of the second stage. The positive and significant coefficient on the instrumented quantity of CGPs confirms our central result and strengthens its causal interpretation: offering CGPs is associated with a significant increase in the risk-taking index of households. The larger magnitude of the coefficient in the instrumented specification suggests that sources of endogeneity are biasing our OLS results downwards. In columns 5 to 8, we restrict the sample to quartiles of filtered 2002 risk-taking index. Consistent with the OLS results, we find that the positive change in the risk-taking index is decreasing with household willingness to take risk, which provides for a causal interpretation of the cross-sectional result from the previous section. The sensitivity of the household risk-taking index with respect to the CGP share is on average equal to 0.69. Its confidence interval strongly overlaps with the distribution of the risk-taking index across CGPs (Table III), consistent with the weak substitutability of CGPs and equity mutual funds.

V. Can Economic Theory Explain the Impact of Capital Guarantee Products on Risk-Taking?

This section shows that two economic mechanisms can explain the increase in household risk-taking triggered by the introduction of capital guarantee products. In Section V.A, we develop a life-cycle model with stochastic labor income and three types of financial assets: a bond, an equity fund, and a CGP exhibiting the nonlinear payoffs and illiquidity of actual contracts. We use the life-cycle framework to assess how the introduction of CGPs impacts household portfolios under several specifications of preferences and beliefs. In Section V.B, we demonstrate that the causal impact of innovation on risk-taking is consistent with recursive preferences with loss aversion and narrow framing (Barberis and Huang, 2009), while other common preferences do not explain our empirical results. Pessimistic subjective beliefs are a powerful complementary explanation, which we investigate in Section V.C.

A. A Life-Cycle Model with Capital Guarantee Products

We develop a life-cycle model with stochastic labor income and CGPs. The model extends Cocco et al. (2005) by expanding the set of assets, alternative preferences, and beliefs.

1. Labor Income. The agent lives at dates $t = 1, \ldots, T$, and receives a stochastic labor income Y_t every period. Before retirement, labor income is specified by:

$$Y_t = Y_t^P \; Y_t^H,$$

where Y_t^P, is a persistent component of income and Y_t^H is a transitory component. The permanent component is specified by $Y_t^P = e^{f(t;\chi_t)+\nu_t}$, where $f(t;\chi_t)$ is a fixed effect driven by the vector of deterministic characteristics χ_t and ν_t follows a random walk with Gaussian increments: $\nu_{t+1} - \nu_t \sim \mathcal{N}(0, \sigma_u^2)$. The transitory components have identical lognormal distributions, are mutually independent, and are also independent from the permanent components. We denote by RA the retirement age. After retirement, income is $Y_t = \lambda \, Y_{RA}^P$, where λ is a replacement ratio.

2. Financial Assets. The agent can trade two liquid financial securities every period. The riskless asset has constant yield $1 + R_f = e^{r_f}$ on a 1-period investment. The equity fund has random return $R_{eq,t} = (1 - \varphi)(1 + R_{m,t})$ between $t - 1$ and t, where $R_{m,t}$ is the return on an equity index and φ is a per-period fee.

Before financial innovation, the agent can only trade these two liquid assets. After innovation, the agent can also invest in CGPs of staggered maturities. All CGPs are identical except for the issue date. A CGP issued at date t reaches maturity at date $t + M$, and we denote by $1 + R_{g,t+M}$ the return on the guaranteed product over the life of the contract.

We make several conservative assumptions: (i) CGPs are written on the same index as the equity fund, (ii) they are strictly illiquid before maturity, and (iii) the agent can hold at most one type of CGP at given point in time. These assumptions ensure that the demand for CGPs is not driven by an artificially strong diversification motive, or early redemption or rollover strategies that bypass the illiquidity of CGPs.[40] These choices allow us to provide a disciplined assessment of household demand for CGPs and its impact on risk-taking.

3. Budget Constraint. At the beginning of period t, cash on hand X_t is the sum of the period's labor income, the value of holdings in the riskless asset and equity fund, and the value of holdings in the CGP if the contract reaches maturity at t. Capital previously invested in a CGP and still illiquid at date t is denoted by K_t, and time to maturity by τ_t.

The household selects the following variables at t: (i) consumption, C_t, (ii) investment in the illiquid product issued in the period, I_t, and (iii) the share of liquid wealth invested in the equity fund, α_t. We impose the constraint $I_t = 0$ whenever $\tau_t > 0$, so that the agent only invests in one type of CGP. Therefore, cash on hand at the beginning of period $t+1$ is

$$X_{t+1} \;=\; Y_{t+1} \;+\; (X_t - I_t - C_t)\left[1 + R_f + \alpha_t(R_{eq,t+1} - R_f)\right] \;+\; (1 + R_{g,t+1})K_t \mathbb{1}_{\{\tau_t = 1\}}. \quad (14)$$

The last term in (14) expresses that the capital K_t becomes liquid at $t+1$ if $\tau_t = 1$.

4. Information Structure. The household observes every period the returns on the equity index, the equity fund, and the held CGP if it reaches maturity. The observation of index returns helps the agent produce increasingly accurate forecasts of the CGP's return as time goes by. At date t, a sufficient statistic for the information available on the held CGP, issued at date $t-s$, is the cumulative return $CR_t = e^{-qs}(1 + R_{m,t-s+1}) \dots (1 + R_{m,t})$. The agent's position at the beginning of period t is summarized by the state vector (X_t, K_t, CR_t, τ_t). We now close the model by considering the specification of preferences and beliefs.

[40] The investor could diversify by investing directly in the underlying of the capital guarantee product.

B. *The Role of Preferences*

This section investigates the preferences that can explain the empirical results of Section IV. We assume that the household has rational expectations and recursive utility:

$$V_t(X_t, K_t, CR_t, \tau_t) = \max_{(C_t, I_t, \alpha_t)} \left[(1-\delta)\, C_t^{1-1/\psi} + \delta p_t \left(\mu_{t+1}\right)^{1-1/\psi}\right]^{\frac{1}{1-1/\psi}}, \tag{15}$$

where $t \in \{1, ..., T-1\}$, p_t is the probability that the agent is alive at $t+1$ conditional on being alive at date t, and μ_{t+1} is the certainty equivalent of future consumption. We let $V_T = (1-\delta)^{1/(1-1/\psi)}\, C_T$ at the terminal date, which does not include a bequest motive.

For each given specification of the certainty equivalent, μ_{t+1}, we solve the model numerically before and after financial innovation. Capital guarantee products have the median representative design: a maturity of 4 years, the full guarantee of the contract's face value ($g = 0$), a participation rate p of 112%, an underlying index with a risk premium of 6%, a volatility of 20%, a dividend yield of 2%, and an issue price equal to 111% of face value. The return on the CGP is based on the values of the index in the last 13 months of the contract. These parameters imply a markup of 1.5% in annual units. We refer the reader to the Internet Appendix for a full description of the model and solution methodology.

Under the Epstein and Zin (1989) utility: $\mu_{t+1} = [\mathbb{E}_t^{\mathbb{P}}(V_{t+1}^{1-\gamma})]^{1/(1-\gamma)}$, financial innovation does not generate an increase in the risk-taking index, as we show in the Internet Appendix for the baseline specification and a battery of alternative parameter values. Since Epstein-Zin preferences imply second-order relative risk aversion, the equity fund provides an attractive risk premium to the household, which generates strong demand for the equity fund before financial innovation. CGPs offer the partial protection of invested capital and diversification opportunities. The guarantee offers only weak welfare benefits to an investor with second-order risk aversion. The benefits from diversification are also limited since the CGP and the equity fund are both linked to the same equity index. As a result, the life-cycle model with rational expectations and Epstein-Zin utility does not explain the strong increase in risk-taking triggered by financial innovation observed in large segments of the Swedish population.

The natural next step is to consider preferences with first-order risk aversion. As Barberis et al. (2006) explain, the choice of such preferences requires some care in multi-period environments. The presence of other preexisting risks, such as labor income risk, makes the agent act in a second-order risk-averse manner toward independent, delayed gambles. Therefore, first-order risk aversion alone may be insufficient to explain our empirical results. The Internet Appendix confirms this intuition. We report that financial innovation does not substantially increase risk-taking when the household exhibits generalized disappointment aversion, a classic class of loss-averse preferences (Gul, 1991; Routledge and Zin, 2010).[41]

In order to explain the empirical results of Section IV, we next consider preferences combining loss aversion with narrow framing, in which investors separately evaluate changes in financial wealth. We consider the recursive specification of Barberis and Huang (2009):

$$\mu_{t+1} = \left[\mathbb{E}_t^{\mathbb{P}}(V_{t+1}^{1-\gamma})\right]^{\frac{1}{1-\gamma}} + b_0 \mathbb{E}_t^{\mathbb{P}}\left[v(W_{t+1} - W_{t+1}^R)\right], \tag{16}$$

where $b_0 \geq 0$ is a constant, W_{t+1} is the value of liquid financial wealth at the beginning of period $t+1$, $v(\cdot)$ is the piecewise linear function:

$$v(x) = \begin{cases} x & \text{if } x \geq 0, \\ \lambda x & \text{if } x \leq 0, \end{cases}$$

and $\lambda \geq 1$ is a kink parameter. The reference level, W_{t+1}^R, is set equal to the current value of past investments if the agent only invests in the riskless asset: $W_{t+1}^R = (X_t - C_t - I_t)(1+R_f) + K_t\,(1+R_f)^M\,\mathbb{1}_{\{\tau_t=1\}}$. This reference level offers the benefits of not altering the

[41] The certainty equivalent μ_{t+1} is implicitly defined by:

$$(\mu_{t+1})^{1-\gamma} = \mathbb{E}_t^{\mathbb{P}}(V_{t+1}^{1-\gamma}) + (\lambda - 1)\mathbb{E}_t^{\mathbb{P}}\left\{\left[V_{t+1}^{1-\gamma} - (\kappa\mu_{t+1})^{1-\gamma}\right]\mathbb{1}_{\{V_{t+1}<\kappa\mu_{t+1}\}}\right\},$$

where $\lambda \geq 1$ is a kink parameter and κ controls the disappointment threshold. This specification coincides with disappointment aversion (Gul, 1991) if $\kappa = 1$.

consumption-saving path when the household does not invest in risky assets.[42]

In Figure 5, we plot the life-cycle profile of an agent with loss aversion and narrow framing, as defined in equation (16). We set $\gamma = 4$, $\delta = 0.98$, and $\psi = 0.5$. The agent accumulates substantial amounts of CGPs (Panel A), which induces a considerable increase in the risk-taking index until retirement (Panel B). The higher average returns on savings allow the agent to increase her consumption during most of her working life and retirement (Panel C). The CGP therefore fosters risk-taking and consumption during most of the life-cycle. We examine the implications for household welfare in Section VI.

In Panel A of Figure 6, we plot the proportional change in the risk-taking index triggered by innovation as a function of the household's initial level of the index. The solid line illustrates the predictions from the life-cycle model and the dashed line the empirical values. In the model plot, we capture heterogeneity in initial risk appetite by varying the kink parameter λ controlling first-order risk aversion, while other preference parameters are set to the constants used in Figure 5. In practice, we let λ vary between 2 and 5 to span the empirical range of the index before innovation. The model seems reasonably consistent with the data. The proportional increase in the risk-taking index is high for households with low initial risk-taking, and decreases sharply with the initial risk-taking index. The model with narrow framing and loss aversion explains why the innovation has a higher impact on households that are less willing to take risk.

One may ask if the same results would hold under preferences combining *second-order* risk aversion and narrow framing. Such preferences can be obtained by letting $\lambda = 1$ in the Barberis and Huang (2009) specification, or more generally by letting $\mu_{t+1} = \left[\mathbb{E}_t^{\mathbb{P}}(V_{t+1}^{1-\gamma})\right]^{\frac{1}{1-\gamma}} + b_0 \left\{[\mathbb{E}_t^{\mathbb{P}}(W_{t+1}^{1-\gamma})]^{\frac{1}{1-\gamma}} - W_{t+1}^R\right\}$, where W_{t+1}^R is the reference level defined earlier in the section. The Internet Appendix verifies that such specifications do not explain the data. While these tests are not exhaustive, they strongly suggest that the combination of narrow framing and loss aversion is important to explain our empirical results under rational expectations.

[42]This specification of the reference level is consistent with earlier life-cycle applications of Barberis and Huang (2009) preferences available in the literature (Chai and Maurer, 2012).

C. The Role of Subjective Beliefs

Another possible explanation for the portfolio impact of financial innovation is that households hold pessimistic subjective beliefs about the equity index. Pessimistic beliefs assign a higher likelihood to negative outcomes than the physical measure $\mathbb{P}$, which discourages investment in the equity fund. By contrast, CGPs provide a protection against negative realizations of equity markets, which pessimistic households view as quite likely, while also providing an upside potential. Financial innovation can then increase risk-taking, an effect that should be especially strong for households with more pessimistic beliefs.

An extensive literature motivates the use of pessimistic beliefs in our model. Prospect theory points to the importance of pessimistic beliefs in decision-making, and one of its components, probability weighting, has emerged as a key building block of behavioral economics (Barberis, 2013). Complementary survey evidence documents that a substantial fraction of households assign a high probability to the occurrence of a large crash (Goetzmann, Kim, and Shiller, 2017). Pessimism is therefore a plausible driver of the demand for CGPs.[43]

We incorporate pessimism into the life-cycle model by adopting Prelec (1998)'s probability weighting methodology. Let $F_{\mathbb{P}}(r)$ denote the cumulative distribution function of the yearly log return on the equity index, $r_{m,t}$, under the physical probability measure $\mathbb{P}$. The household's subjective belief about $r_{m,t}$ is specified by the cumulative distribution function:

$$F(r; a, b) = \exp\left\{-b\left[-\ln F_{\mathbb{P}}(r)\right]^a\right\},$$

where a and b are strictly positive constants. The parameter a controls the curvature of $F(\,\cdot\,; a, b)$. The Prelec transform $F(r; a, b)$ decreases with b, so a higher value of b implies stronger pessimism.

In Figure 7, we plot the life-cycle profile of an agent with Epstein-Zin utility and Prelec

[43] Of course other households may be irrationally exuberant about stock market investing. However, optimistic households likely have a high risk-taking index before financial innovation and are less likely to drive the demand for guaranteed products.

probability weighting. We set $a = 0.5$ and let the pessimism parameter b vary from 0.6 to 1.3. The results are qualitatively similar to the ones obtained in Figure 5 under rational expectations and Barberis-Huang preferences. The household has a strong demand for CGPs, which is hump-shaped over the life-cycle. This strong demand is associated with an increase in the risk-taking index. The higher average returns on savings triggered by innovation encourage households to slightly reduce consumption in their early years, and then enjoy higher average consumption after 40. Panel B of Figure 6 also illustrates that the proportional increase in the risk-taking index is stronger for households with more pessimistic beliefs and a lower initial risk-taking index, consistent with the data.

The Internet Appendix shows that the results of Figures 6 and 7 are strongly robust to alternative specifications of pessimism. We define the subjective probability distribution as a mixture of a Gaussian and a crash event. Alternatively, we consider that the household believes that the volatility of the index exceeds its volatility level under $\mathbb{P}$, while the mean return remains unchanged. Variation in the crash probability or in volatility misperception induces variation in the risk-taking index analogous to the results reported in Figure 6.

Overall, the portfolio impact of financial innovation documented in Section IV is consistent with a life-cycle model, provided that one departs from the canonical combination of Epstein-Zin preferences and rational expectations. Loss aversion and narrow framing (Barberis and Huang, 2009), or pessimistic subjective beliefs specified by probability weighting, subjective disaster risk, or volatility misperception deliver a model that explains the demand for CGPs and its cross-sectional variation with initial risk-taking.

VI. Implications for Household Welfare

This section measures the welfare implications of financial innovation under a set of assumptions on decision and experienced utilities. Section VI.A measures the total surplus generated by innovation and its allocation to households and institutions. We show that households with pronounced behavioral biases and low initial levels of risk-taking enjoy large welfare gains over the life-cycle, which are equivalent to 6 to 12 months of yearly income and represent a substantial share of the surplus generated by the innovation. In Section VI.B, we reexamine these results when experienced utility, which households use to assess economic well-being (Kahneman et al., 1997), differs from the decision utility used to make consumption-portfolio choices. While the welfare gains of households with pronounced biases remain substantial, households with weaker biases enjoy modest welfare gains and even incur experienced utility losses in some cases.

A. Total Surplus and Its Allocation

The life-cycle model allows us to measure the total surplus per household generated by CGPs. In this subsection, we conduct the analysis under the following assumptions. First, household decision and experienced utilities coincide. Second, we use actual CGP prices. Third, equity returns and the riskless rate have identical properties before and after financial innovation in Sweden, consistent with the global pricing of asset markets. Under these assumptions, a consumption-portfolio strategy that is feasible before the introduction of the new product remains feasible afterward, so that the innovation cannot reduce welfare.

We define the *household benefit from financial innovation* as the wealth transfer that allows the household to attain in the pre-innovation economy the same lifetime utility as the one it achieves in the post-innovation economy without the transfer. For simplicity, the transfer takes place in the first period of the life-cycle. Our measure takes into account the optimization of financial resources via asset markets.

In Table VIII, we report innovation benefits under several specifications of preferences and beliefs. In all cases, the parameters are chosen so that the risk-taking index before innovation is 8%, its 25th percentile in the Swedish population. The introduction of CGPs generates a benefit of about $15,000 for households exhibiting loss aversion and narrow framing or pessimistic beliefs. Alternative specifications of pessimism produce even higher estimates. The gains from innovation represent a substantial fraction of average yearly income. Therefore, financial innovation is highly beneficial to households with strong behavioral biases and low initial risk-taking.

The *bank benefit from financial innovation* is defined as the no-arbitrage value in the first year of the life cycle of the change in profit per household:

$$\text{Bank benefit} = \mathbb{E}^{\mathbb{Q}}\left[\sum_{t=1}^{T} \frac{p_1 \ldots p_{t-1} \Delta(\text{Profit}_t)}{(1+R_b)^{t-1}}\right], \tag{17}$$

where R_b is the funding cost of the bank and p_t denotes the survival probability defined in Section V. Given the limited information at our disposal, we proxy the change in bank profits by the sum of (i) the change in the fees earned on equity funds sold to the household and (ii) the gross profit margin earned on CGP sales. This approach is conservative because our measures of the bank's benefit and surplus share are upper bounds of actual values. We measure the funding cost R_b by the swap rate, which we take as constant, and we assume that the stochastic variation in profit is not priced, so that we take expectations under $\mathbb{P}$. The analysis therefore incorporates the reduction of profit from mutual funds that can be caused by financial innovation, commonly referred to as crowding out effects.

The *total surplus* is the sum of the household and bank benefits. In Table VIII, we report that the bank receives about 50-60% of the surplus and the household correspondingly receives 40-50% across specifications of preferences and beliefs. Thus, pricing by the bank does not appear to be predatory, consistent with the results of Section III.[44]

B. Sensitivity to Decision and Experienced Utilities

We now assess how behavioral biases impact the measured benefits from financial innovation. In particular, we allow that households may assess well-being by way of an experienced utility that differs from the decision utility used to select consumption and investments.

[44] In the Internet Appendix, we show that markups are not strongly tied to IQ, which further confirms that predatory pricing is not a dominant concern for this asset class.

While one can impute the decision utility from observed choices, the experienced utility is considerably more challenging to estimate in the present context. For this reason, we focus on two polar cases. In one scenario, the experienced utility coincides with the decision utility. In a second scenario, households are prone to behavioral biases in decision-making (as explained in previous sections) but not in the assessment of economic well-being.[45] We then assume that the experienced utility exhibits constant relative risk aversion (CRRA) and is evaluated under the physical measure $\mathbb{P}$.

In this expanded framework, we evaluate the welfare implications of financial innovation as follows. For a given decision utility and probability belief, we solve numerically the policy function $(C_t^*, I^*, t, \alpha_t^*)$ and then compute by simulation the experienced utility:

$$V^{\text{EXP}} = \mathbb{E}_0^{\mathbb{P}} \left[\sum_{t=1}^{T} \delta^{t-1} p_1 \dots p_{t-1} \, u(C_t^*) \right],$$

where $u(C) = C^{1-1/\psi}/(1-\psi^{-1})$. To map the experienced utility V^{EXP} into yearly units, we define its *constant consumption equivalent* as the time- and state-invariant yearly consumption level C^{EXP} that achieves the same experienced utility: $\sum_{t=1}^{T} \delta^{t-1} p_1 \dots p_{t-1} \, u(C^{\text{EXP}}) = V^{\text{EXP}}$.[46]

In the left graph of Figure 8, Panel A, we consider households with identical decision and experienced utilities, which are of the Barberis-Huang type. Variation in initial risk-taking is obtained by letting the loss aversion parameter λ vary, while the other preference parameters are set as in Section V.B. The figure plots the constant consumption equivalent before and after innovation as a function of the initial risk-taking index. The innovation increases the constant consumption equivalent by $1,500 per year for households with low

[45] In both cases, the experienced utility exhibits less behavioral traits in preferences or beliefs than the decision utility. If instead the experienced utility has stronger behavioral traits, financial innovation could trigger an increase in risk-taking that would make the household worse off.

[46] The constant consumption equivalent is given by:

$$C^{\text{EXP}} = \left[(1-\psi^{-1}) \left(\sum_{t=1}^{T} \delta^{t-1} p_1 \dots p_{t-1} \right)^{-1} V^{\text{EXP}} \right]^{1-1/\psi}.$$

initial risk-taking, and about \$1,000 for households with high initial risk-taking. The welfare gains are substantial for all households and are most pronounced for households that have a higher loss aversion parameter λ and therefore a lower risk-taking index *ex ante*.

In the right graph of Figure 8, Panel A, we consider households with (i) Barberis-Huang decision utilities with heterogeneous loss aversion parameters λ, and (ii) a common CRRA experienced utility with parameters $\psi = 0.5$ and $\delta = 0.98$. The figure plots the constant consumption equivalent before and after innovation as a function of the initial risk-taking index (corresponding to different levels of λ). While financial innovation increases the experienced utility of households with low initial risk-taking, it now *decreases* the experienced utility of households with high initial risk-taking. Under our chosen specification, the difference in utility breaks even when the risk-taking index is about 20% *ex ante*. Households with high initial risky shares cater to their behavioral biases (loss aversion and narrow framing) by purchasing CGPs, which lowers their risk-taking index and reduces average consumption and experienced utility over the life-cycle.

Panel B of Figure 8 reports similar findings for households with Prelec subjective utilities. The most pessimistic households strongly benefit from financial innovation, while less biased households incur losses in experienced utility.

Overall, households with low initial risk-taking are the prime beneficiaries of the introduction of CGPs across preference and belief specifications. The new products address these households' concerns about very adverse outcomes and allows them to increase their levels of risk-taking, which produces an increase in average consumption. Since the experienced utility is not particularly sensitive to consumption volatility, household welfare improves. By contrast, for households initially more willing to take risk, the introduction of CGPs crowds out equity fund investments, thereby reducing average consumption and experienced utility. Our results suggests that in order to maximize household welfare, financial advisers and institutions should target the sale of CGPs to households with low risk exposures, while continuing to market diversified equity funds to customers with stronger risk appetites.

VII. Conclusion

This study provides empirical evidence that security design can help to alleviate low financial risk-taking by a sizable segment of the household population. The growing class of capital guarantee products provide investors with a substantial share of the equity premium, along with a guarantee typically representing about 90% of invested capital. Using a large administrative data set, we show that the introduction of retail capital guarantee products significantly increases the expected returns of household financial portfolios, especially if the initial willingness to take risk is low.

The present paper illustrates that financial innovation can be used as a laboratory to test theories of portfolio choice. For instance, we show that pessimistic beliefs or preferences combining loss aversion with narrow framing can explain low levels of household risk-taking and the impact of financial innovation, while the combination of second-order risk aversion and rational expectations cannot explain these facts in a standard life-cycle model.

Our work also contributes to the literature that assesses the welfare implications of financial innovation. When experiential utility coincides with decision utility, the introduction of capital guarantee products generates large welfare gains for households with a low initial willingness to take risk, and more modest gains for other households. If instead behavioral biases impact decision utility but not experiential utility, the innovation is only beneficial for households with the strongest biases and the lowest initial equity shares. This analysis suggests that capital guarantee products should be primarily marketed to low risk-takers, while low-fee traditional equity products are better suited for other households.

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Table I
Household Risk-Taking Across Countries in 2015

	National Accounts	Surveys of Households Above 50	
	Share of Equity in Aggregate Financial Wealth in %	**Fraction of Equity Participants in %**	**Median Share of Equity in Financial Wealth of Participants in %**
	(1)	(2)	(3)
Sweden	17.41	68.16	36.64
United States	31.91	32.03	31.61
China	n/a	10.08	19.88
European Union	8.74	25.26	31.25
Selected European countries			
Austria	7.80	18.69	33.33
Belgium	14.24	40.46	33.11
Croatia	n/a	6.93	33.33
Czech Republic	1.18	37.92	21.98
Denmark	28.27	58.31	32.38
Estonia	5.09	8.40	33.26
Finland	15.71	n/a	n/a
France	9.12	30.57	23.37
Germany	6.29	32.91	26.19
Greece	3.14	2.58	27.36
Hungary	3.86	n/a	n/a
Italy	7.13	8.03	30.00
Latvia	3.48	n/a	n/a
Lithuania	4.20	n/a	n/a
Luxembourg	10.98	26.68	36.80
The Netherlands	26.00	n/a	n/a
Norway	15.80	n/a	n/a
Poland	10.12	2.38	36.36
Portugal	3.21	16.36	28.00
Slovakia	0.38	n/a	n/a
Slovenia	8.19	10.97	30.11
Spain	8.90	7.81	31.09
United Kingdom	9.05	25.70	7.06

Notes: This table reports (1) the percentage of aggregate household financial wealth invested in equity, (2) the fraction of households participating in equity markets, and (3) the median share of equity in the financial wealth of participants. The data in column 1 are retrieved from the OECD National Accounts and the US Federal Reserve's Financial Accounts. The statistics in columns 2 and 3 are based on surveys of households representative of the population of people aged 50 years and older, except for China, where the sample is representative of the total population. The surveys are the following: the 2016 wave of the University of Michigan Health and Retirement Study (HRS) for the US, the 6^{th} wave of the Survey of Health, Ageing and Retirement in Europe (SHARE) for European countries including Sweden, the 7^{th} wave of the English Longitudinal Study of Ageing (ELSA) for the United Kingdom, and the 2015 China Household Finance Survey (CHFS) for China. Section II in the Internet Appendix describes the precise methodology.

Table II
Capital Guarantee Products Around the World

Country or Region	Product Type	Outstanding Volume (Billion U.S. $)
North America		1,764
USA	Guaranteed Life Annuity*	1,720
	Retail Structured Products	22
Canada	Retail Structured Products	20
Mexico	Retail Structured Products	2
Europe		1,794
France	Euro Contracts*	1,540
	Retail Structured Products	31
Germany	Retail Structured Products	47
Belgium	Retail Structured Products	45
UK	Retail Structured Products	12
Asia		936
China	Guaranteed Wealth Management Products*	854
	Retail Structured Products	13
South Korea	Retail Structured Products	31
Japan	Retail Structured Products	17
Other	Retail Structured Products	18

Notes: This table reports the types and outstanding volumes of capital guarantee products around the world in 2015. The outstanding volume is obtained from Ellul et al. (2020) for guaranteed life annuities in the United States, Hombert and Lyonnet (2020) for Euro contracts in France, the 2015 Annual Report of China Banking Wealth Management Product by "China Central Depository & Clearing Co. Ltd" for wealth management products in China, and from the same data provider as in Célérier and Vallée (2017) for retail structured products. Retail structured products volume only include issuances offering a capital protection of at least 90% of the capital invested. * *For these products the guarantee is obtained using reserves, possibly complemented by hedging.*

Table III
Design, Markup, and Expected Return of Retail Equity Products

	Mean	p1	p10	p50	p90	p99
Panel A. Full sample of capital guarantee products (1,511 contracts)						
Issuance year	2006	2002	2004	2006	2007	2007
Volume (2000 $ million)	5.2	0.1	0.5	2.6	13.0	29.1
Design parameters:						
- Maturity (months)	40.1	12.0	17.9	37.6	60.5	72.5
- Guarantee (% of face value)	100.2	100.0	100.0	100.0	100.0	108.0
- Issue price (% of face value)	107.0	100.0	101.0	106.0	112.0	122.0
Panel B. Baseline capital guarantee products (809 contracts)						
Issuance year	2006	2002	2004	2006	2007	2007
Volume (2000 $ million)	4.8	0.1	0.5	2.7	11.9	25.9
Design parameters:						
- Maturity (months)	44.4	12.6	24.5	48.0	60.5	72.5
- Guarantee (% of face value)	100.2	100.0	100.0	100.0	100.0	108.0
- Issue price (% of face value)	108.7	100.0	101.5	111.5	112.0	122.0
- Participation rate (%)	112.9	30.0	60.0	110.0	160.0	210.0
- Asian option length (months)	13.6	0.0	4.0	13.0	24.0	60.0
Asset pricing inputs:						
- Historical volatility	0.1	0.0	0.0	0.1	0.1	0.1
- Dividend yield (%)	2.0	0.0	0.5	2.1	3.0	4.5
- CDS premium (%)	18.8	8.0	11.2	15.4	31.5	47.5
- Beta of underlying to world index	1.1	0.5	0.9	1.1	1.3	1.4
Asset pricing outputs:						
- Yearly markup (%)	1.6	-0.7	0.3	1.6	2.7	3.9
- Risk-taking index η	0.44	-0.17	0.02	0.45	0.83	1.06
Panel C. Equity mutual funds (1,376 funds)						
Volume in 2007 ($ million)	21.7	0.0	0.0	0.4	27.2	448
Beta to world index (%)	0.9	0.0	0.5	0.9	1.2	1.5
Yearly fees (%)	2.1	0.6	1.6	1.9	2.8	4.1
Asset pricing outputs:						
- Risk-taking index η	0.55	0.0	0.0	0.58	0.89	1.16

Notes: Panel A reports the average characteristics of retail CGPs issued in Sweden between 2002 and 2007. The capital guarantee, g, is the minimum fraction of face value the household receives at maturity. The issue price, P_0, is expressed as a percentage of face value. Panel B displays summary statistics on the subsample of baseline CGPs with total returns of the form $1 + R_g = [1 + \max(p\, R^*; g)]\, \xi\, F/P_0$, where p is the participation rate, R^* is the average performance of the underlying, and ξ is the fraction of pledged cash flows paid at maturity. The Asian option length is the length of the period over which the underlying's performance is averaged to define R^*. The risk-taking index η is the ratio of the product's risk premium to the world index's risk premium, as defined in equation (6). The yearly markup is computed as defined in Section III.B. Panel C reports summary statistics on all equity mutual funds available in Sweden between 2002 and 2007. Yearly fees include the management and entry fees paid by retail investors.

Table IV
Household Characteristics and Portfolio Allocation

	Full Sample (1)				Traditional Equity Product Participants (2)				Capital Guarantee Product Participants (3)			
	Number of households: N=3,107,893				Number of households: N=2,128,612 (68.5% of total)				Number of households: N=428,337 (13.9% of total)			
	Mean	Median	p10	p90	Mean	Median	p10	p90	Mean	Median	p10	p90
Panel A: 2002												
Financial wealth (in 2000 $, thousands)												
Total	33.7	11.2	2.5	72.8	44.9	17.7	4.6	92.3	72.9	38.0	8.0	149.6
Traditional equity products	15.3	1.2	0.0	29.6	22.4	4.4	0.2	42.9	36.1	11.8	0.2	79.1
Stocks	7.1	0.0	0.0	6.5	10.4	0.3	0.0	11.3	13.5	0.9	0.0	22.4
Equity mutual funds	8.0	0.5	0.0	19.7	11.7	2.6	0.0	28.5	22.1	7.8	0.0	54.5
Bank deposits	13.1	6.3	2.0	27.4	1.6	7.7	2.6	32.4	22.4	10.5	3.0	48.0
Fixed income securities	4.2	0.0	0.0	11.6	5.5	0.0	0.0	14.9	11.4	2.1	0.0	30.5
Demographics												
Family size	2.1	2.0	1.0	4.0	2.3	2.0	1.0	4.0	2.2	2.0	1.0	4.0
Number of children	0.2	0.0	0.0	1.0	0.2	0.0	0.0	1.0	0.2	0.0	0.0	1.0
Disposable income (in 2000 $)	27.5	22.8	10.5	48.0	31.5	28.0	12.3	52.0	35.3	30.9	14.2	57.0
Years of schooling (head)	11.4	11.0	8.0	15.0	11.8	11.0	8.0	15.0	11.9	12.0	8.0	16.0
Male, in % (head)	60.0	100.0	0.0	100.0	63.5	100.0	0.0	100.0	62.9	100.0	0.0	100.0
Age (head)	53.1	52.0	33.0	76.0	52.0	52.0	32.0	73.0	55.2	56.0	37.0	72.0
Equity participants, in %	68.5	-	-	-	100	-	-	-	92.9	-	-	-
Allocation of financial wealth (%, 2002 participants only)												
Traditional equity products	22.5	11.5	0.0	65.0	32.9	27.5	2.5	72.6	37.9	35.9	1.3	76.5
Stocks	6.4	0.0	0.0	20.9	9.3	1.4	0.0	30.2	9.8	2.5	0.0	30.7
Equity mutual funds	15.7	4.3	0.0	48.8	22.9	16.9	0.0	56.3	27.3	23.5	0.0	61.1
Bank deposits	66.5	74.0	17.1	1	54.4	55.0	12.9	95.1	42.9	37.7	8.6	88.1

Table IV (continued)
Household Characteristics and Portfolio Allocation

Panel B: 2007												
Allocation of financial wealth (%)												
Capital guarantee products	1.6	0.0	0.0	2.5	2.1	0.0	0.0	6.2	11.9	7.3	0.0	30.0
Traditional equity products	24.4	11.8	0.0	70.1	34.5	30.5	0.0	76.5	34.4	31.9	0.7	71.0
Stocks	6.4	0.0	0.0	21.8	9.2	0.3	0.0	31.7	9.1	1.6	0.0	30.1
Equity mutual funds	16.5	1.4	0.0	54.5	23.6	15.9	0.0	61.9	23.2	18.7	0.0	55.0
ETFs	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0
Bank deposits	63.8	69.0	14.3	100.0	52.1	49.2	11.2	99.2	36.9	31.6	7.7	75.4
Panel C: Household risk-taking index (η_h)												
Year 2002	0.15	0.06	0.0	0.46	0.22	0.17	0.0	0.53	0.26	0.23	0.01	0.56
Year 2007	0.17	0.06	0.0	0.51	0.25	0.20	0.0	0.58	0.28	0.26	0.01	0.56
2002-2007 % Change					0.7	13.5	-195.1	122.2	17.6	13.6	-78.8	123.8

Notes: This table reports summary statistics on the characteristics, portfolio allocation, and risk-taking behavior of Swedish households. In Panel A, we tabulate the financial characteristics, demographics, equity participation, and share of financial wealth allocated to equity in 2002. Panel B displays equity participation and the equity share of households in 2007. Panel C shows the risk-taking index in 2002 and 2007. The statistics are computed on all households in Sweden in the first set of columns ($N = 3,107,893$ households), participants in traditional equity products in the second set of columns ($N = 2,128,612$ households or 68.5% of the total household population), and participants in capital guarantee products ($N = 428,337$ households or 13.8% of the total population).

Table V
Participation in Capital Guarantee Products and Financial Risk-Taking: Cross Section Analysis

	2002 -2007 Percentage Change in Risk-Taking Index ($\Delta\eta_h$)			
	Total Change		Active Change	
	(1)	(2)	(3)	(4)
$\mathbb{1}_{CGP_h}$	0.24***	0.44***	0.27***	0.51***
	(0.01)	(0.01)	(0.01)	(0.01)
$\mathbb{1}_{CGP_h} \times$ 2002 risk-taking index		-0.85***		-0.99***
		(0.02)		(0.02)
Fixed effects (2002 value)				
Risk-taking index quartiles	Yes	Yes	Yes	Yes
Financial wealth deciles	Yes	Yes	Yes	Yes
Income deciles	Yes	Yes	Yes	Yes
Age deciles	Yes	Yes	Yes	Yes
Gender	Yes	Yes	Yes	Yes
Years of education	Yes	Yes	Yes	Yes
Family size	Yes	Yes	Yes	Yes
Number of children	Yes	Yes	Yes	Yes
Province	Yes	Yes	Yes	Yes
Control				
2002-2007 change in income	Yes	Yes	Yes	Yes
Observations	2,128,612	2,128,612	2,128,612	2,128,612
R^2	0.106	0.061	0.05	0.05

Notes: This table displays OLS regression coefficients of the change in the risk-taking on an indicator variable for participation in capital guarantee products and control variables. In Columns 1 and 2, the dependent variable is the Davis and Haltiwanger (1992) measure of growth in the risk-taking index from 2002 to 2007. In Columns 3 and 4, the dependent variable is the *active* change in risk-taking index from 2002 to 2007. We compute the active change in the risk-taking index as the Davis and Haltiwanger (1992)'s growth rate between the risk-taking index in 2002 and the 2007 "market-neutral" risk-taking index, as described in Section IV.B. The indicator variable $\mathbb{1}_{CGP_h}$ is equal to unity if the household invests at least once in capital guarantee products over the 2002 to 2007 period. In Columns 2 and 4, we interact $\mathbb{1}_{CGP_h}$ with the household 2002 risk-taking index, filtered with household observable characteristics as described in Section IV.C. The sample is restricted to households participating in stock markets in 2002. Standard errors are clustered at the district level and displayed below their coefficient of interest. *, **, and *** represent statistical significance at the 10%, 5%, and 1% confidence levels, respectively.

Table VI
Participation in Capital Guarantee Products and Financial Risk-Taking: Panel Analysis

		Quartiles of 2002 Risk-Taking Index			
	All	Q1	Q2	Q3	Q4
	(1)	(2)	(3)	(4)	(5)
Panel A. Dependent variable: Risk-taking index $\eta_{h,t}$					
CGP Share$_{h,t}$	0.21***	0.40***	0.31***	0.15***	-0.06***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Controls and Observations: see Panel C					
R^2	0.832	0.723	0.731	0.680	0.708
Panel B. Dependent variable: Market-neutral risk-taking index $\eta_{h,t}^{MN}$					
CGP Share$_{h,t}$	0.26***	0.40***	0.33***	0.22***	0.08***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)
Controls and Observations: see Panel C					
R^2	0.629	0.542	0.515	0.566	0.623
Panel C. Control variables and Number of Observations					
Fixed Effects					
Household	Yes	Yes	Yes	Yes	Yes
Year	Yes	Yes	Yes	Yes	Yes
Fixed effects interacted with year fixed effects					
2002 risk-taking index quartiles	Yes	Yes	Yes	Yes	Yes
Financial wealth deciles	Yes	Yes	Yes	Yes	Yes
Income deciles	Yes	Yes	Yes	Yes	Yes
Age deciles	Yes	Yes	Yes	Yes	Yes
Gender	Yes	Yes	Yes	Yes	Yes
Years of education	Yes	Yes	Yes	Yes	Yes
Family size	Yes	Yes	Yes	Yes	Yes
Number of children	Yes	Yes	Yes	Yes	Yes
Province	Yes	Yes	Yes	Yes	Yes
Observations	12,771,671	3,192,917	3,192,908	3,192,912	3,192,916

Notes: This table reports panel regressions of household risk-taking on the share of financial wealth invested in capital guaranteed products, *CGP Share*$_{h,t}$. In Panel A, the dependent variable is the risk-taking index. In Panel B, the dependent variable is the "market-neutral" risk-taking index, as described in Section IV.B. Panel C lists the control variables used in the regressions reported in Panels A and B. The sample is restricted to households participating in stock markets in 2002. Standard errors are clustered at the bank times year level and displayed below their coefficient of interest. *, **, and *** represent statistical significance at the 10%, 5%, and 1% confidence levels, respectively.

Table VII
Instrumental Variable Panel Analysis

Instruments	**Volumes per Depositor**		**Idiosyncratic Supply Shocks**					
	First Stage	Second Stage	First Stage	Second Stage				
	$CGPShare_{h,t}$	Risk-Taking Index	$CGPShare_{h,t}$	Risk-Taking Index				
	Full	Full	Full	Full	Quartiles of Risk-Taking Index			
	Sample (1)	Sample (2)	Sample (3)	Sample (4)	Q1 (5)	Q2 (6)	Q3 (7)	Q4 (8)
$\widehat{\text{CGP Share}}_{h,t}$		0.9***		0.69**	0.79***	0.60***	0.61*	0.33
		(0.9)		(0.30)	(0.20)	(0.31)	(0.36)	(0.43)
Volume issued by main bank	2.84***							
	(0.6)							
Idiosyncratic supply shocks								
Main bank			1.15***					
			(0.03)					
Second main bank			0.56***					
			(0.04)					
Third main bank			0.65***					
			(0.07)					
Fixed effects								
Household	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Fixed effects interacted with year fixed effects								
Risk-taking index quartiles	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Financial wealth deciles	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Income deciles	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age deciles	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Gender	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Years of education	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Family size	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Number of children	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Province	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	8,131,784	8,131,784	4,164,828	4,164,828	1,013,793	1,013,793	1,013,793	1,013,793
R^2	0.49		0.51					
F-statistic	76.2		568.5					

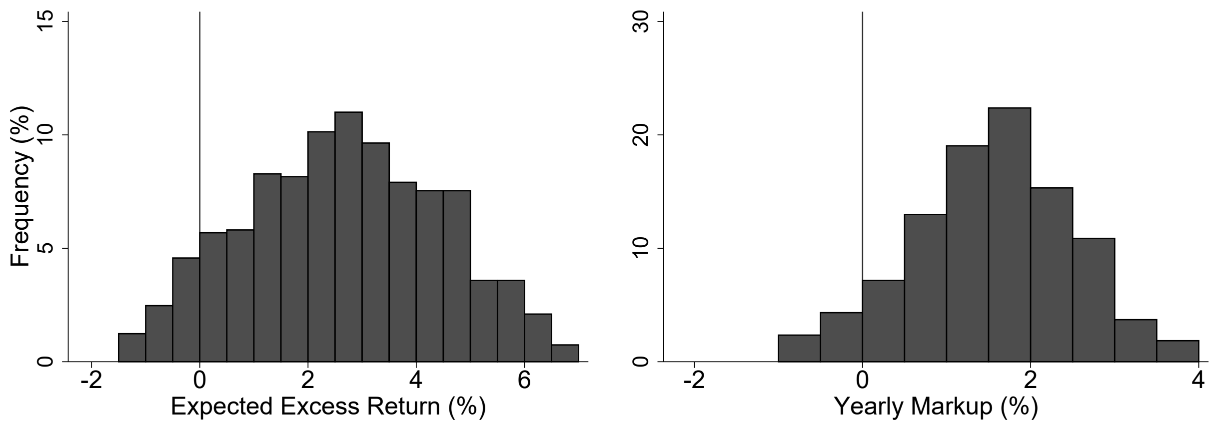
Notes: This table displays the results of the instrumental variable analysis, in which the share of CGPs in the financial wealth of household h in year t, *CGP Share*$_{h,t}$, is instrumented by a measure of *CGP* supply by the main bank(s) with which the household has a relationship in 2002. In columns 1 and 2, we instrument *CGP Share*$_{h,t}$ by the contemporaneous outstanding volume of CGPs per depositor issued by household h's main bank. In Columns 3 to 7, we filter out demand effects from this measure by partitioning the household population into two random sub-samples of equal size. We use the first sub-sample to estimate idiosyncratic bank-level supply shocks, and the second sub-sample to estimate the structural equation. More specifically, in the first sub-sample, we regress *CGP Share*$_{h,t}$ on (i) a vector of indicator variables for every bank b, where the b^{th} indicator is equal to unity if b is one of the household's three main banks at the beginning of the sample period, and (ii) a set of household characteristics. The resulting linear coefficients of bank indicators provide measures of bank-levels idiosyncratic supply shocks. We then use the second random sub-sample to implement two stage least squares (2SLS). In the first stage of 2SLS, we regress *CGP Share*$_{h,t}$ on the supply shocks of the household's three main banks. In the second stage of 2SLS, we regress the household's risk-taking index on the predicted $\widehat{\text{CGP Share}}_{h,t}$ from the first stage. Both stages of 2SLS are panel models with household and year fixed effects. The sample is restricted to household participating in stock markets in 2002. Standard errors are clustered at the bank times year levels and are displayed below their coefficient of interest. *, **, and *** represent statistical significance at the 10%, 5%, and 1% confidence levels, respectively.

Table VIII
Household Welfare Gains Predicted by the Models

Models	**Loss Aversion with Narrow Framing** (1)	**Probability Weighting** (2)
Key parameter value	Utility kink parameter $\lambda = 3.5$	Pessimism parameter b=0.71
Change in risk-taking (%)	95.4	121.6
Household utility gain, in U.S. $	15,737	14,088
Bank revenue gain, in U.S. $	14,741	19,419
Household share of surplus (%)	51.6	42.0

Notes: This table reports the changes in the household risk-taking index, welfare gains, bank revenue gains, and the household share of the surplus generated by the introduction of capital guarantee products under various specifications of preferences and beliefs. Decision and experienced utilities are assumed to be identical. Under all specifications, the starting value is household with an ex-ante risk-taking index of 8%, which corresponds to the 25^{th} percentile in the Swedish population. In column 1, we consider an investor with Barberis and Huang (2009) preferences, which combines loss aversion with narrow framing, and rational expectations. In column 2, we consider an investor with Prelec (1998) probability weighting. The subjective cumulative distribution function of the investor is given by $F(r; a, b) = \exp\{-b[-\ln F_{\mathbb{P}}(r)]\}$, where $F_{\mathbb{P}}(r)$ denotes the cumulative distribution function of the yearly log return on the underlying under the physical measure $\mathbb{P}$.

Panel A. Baseline Capital Guarantee Products (809 Products)



Panel B. Equity Mutual Funds (1,376 Products)

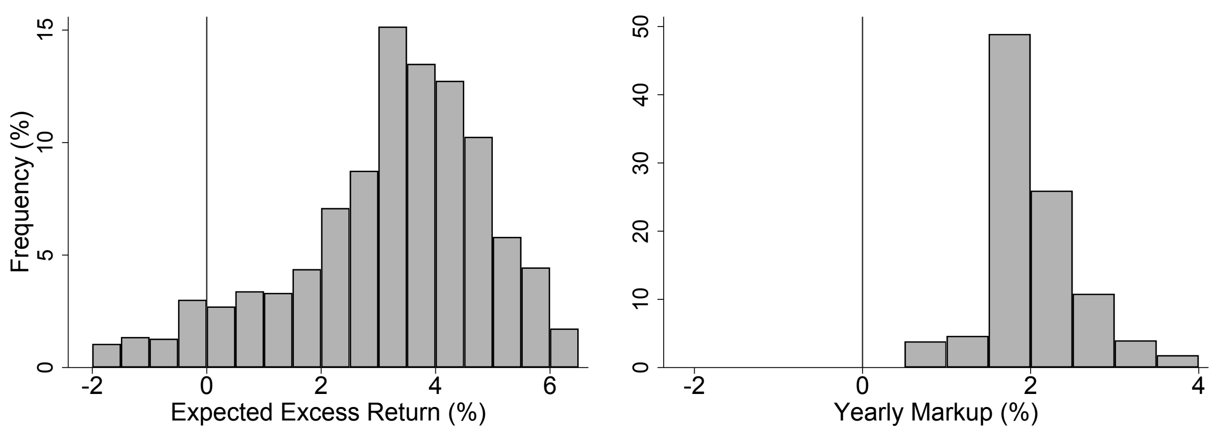


Figure 1. Expected Excess Returns and Yearly Markups of Capital Guarantee Products and Equity Mutual Funds. Panel A shows the histogram of the expected excess return offered by the 809 baseline capital guarantee products issued in Sweden over the 2002-2007 period (left graph) and the histogram of the gross markup of the banks distributing them (right graph). Both measures result are computed by following the asset pricing methodology outlined in Section III. Panel B shows the histograms of the expected excess return (left graph) and gross markup (right graph) of the 1,376 equity mutual funds under management in Sweden over the 2002-2007 period.

Panel A. Full Sample

Panel B. First Quartile of 2002 Risk-Taking Index

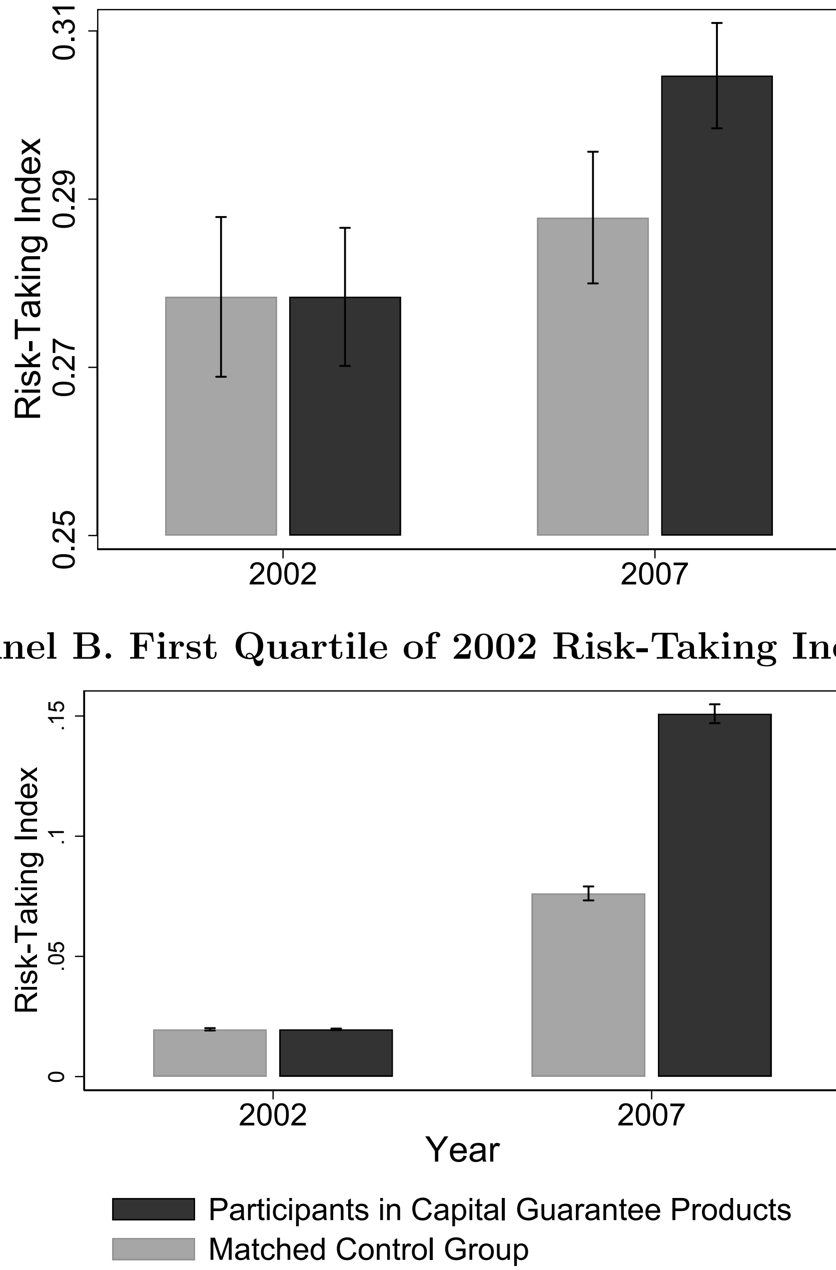


Figure 2. Household Risk-Taking Index in 2002 and 2007. Panel A plots the risk-taking index in 2002 and in 2007 for: (i) capital guarantee product participants, and (ii) a control group of equal size made of stock market participants matched based on their 2002 risk-taking index. Panel B reproduces the same graph when restricting the sample to households in the first quartile of risk-taking index in 2002. The whiskers represent the confidence band at the 95% level.

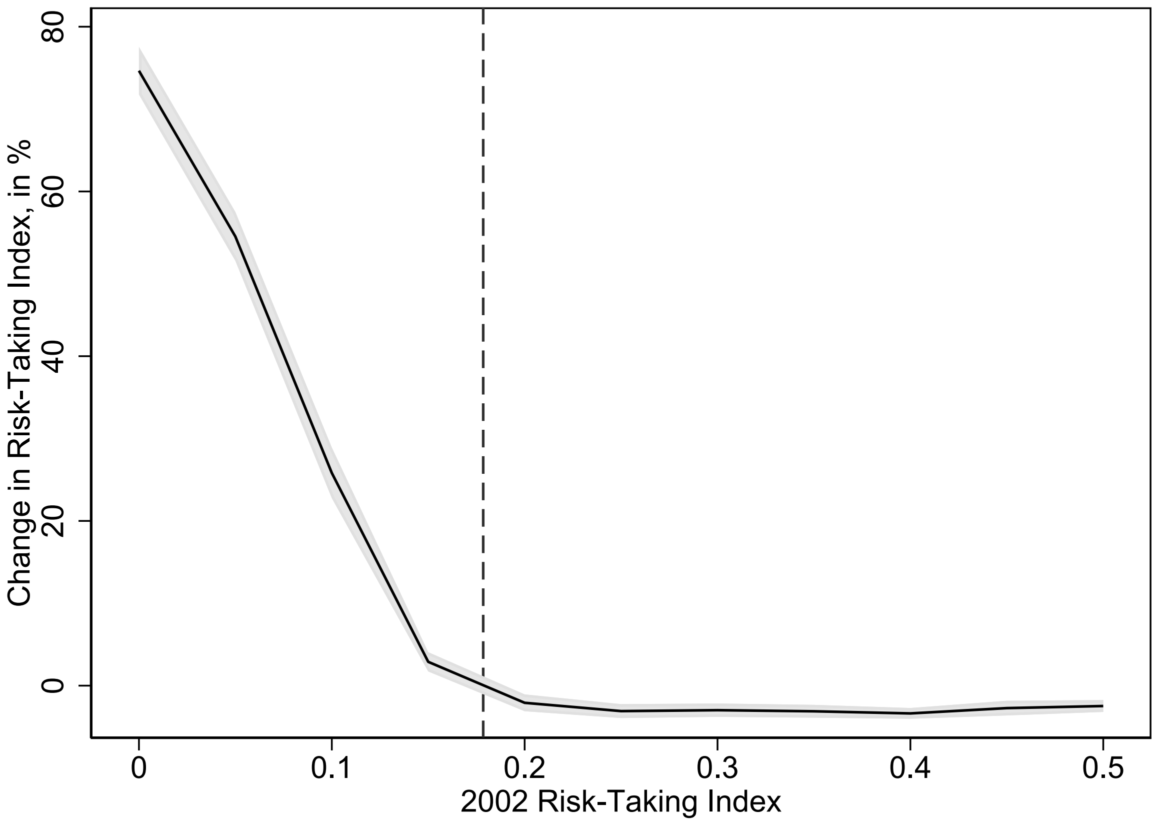


Figure 3. Proportional Change in Risk-Taking Index as a Function of Initial Risk-Taking for Capital Guarantee Product Participants. This figure shows the proportional change in the risk-taking index for CGP participants as a function of their filtered risk-taking index in 2002. The proportional change in risk-taking for CGP participants is the ratio of the predicted incremental change in the risk-taking index for CGP participants (vs. non participants) to their period-average of risk-taking index $\frac{\eta_{h,2002}+\eta_{h,2007}}{2}$. The 2002 risk-taking index is filtered from household observable characteristics, as described in Section IV.C. The vertical dotted line plots the median 2002 risk-taking index. The shaded area represents the confidence band at the 95% level.

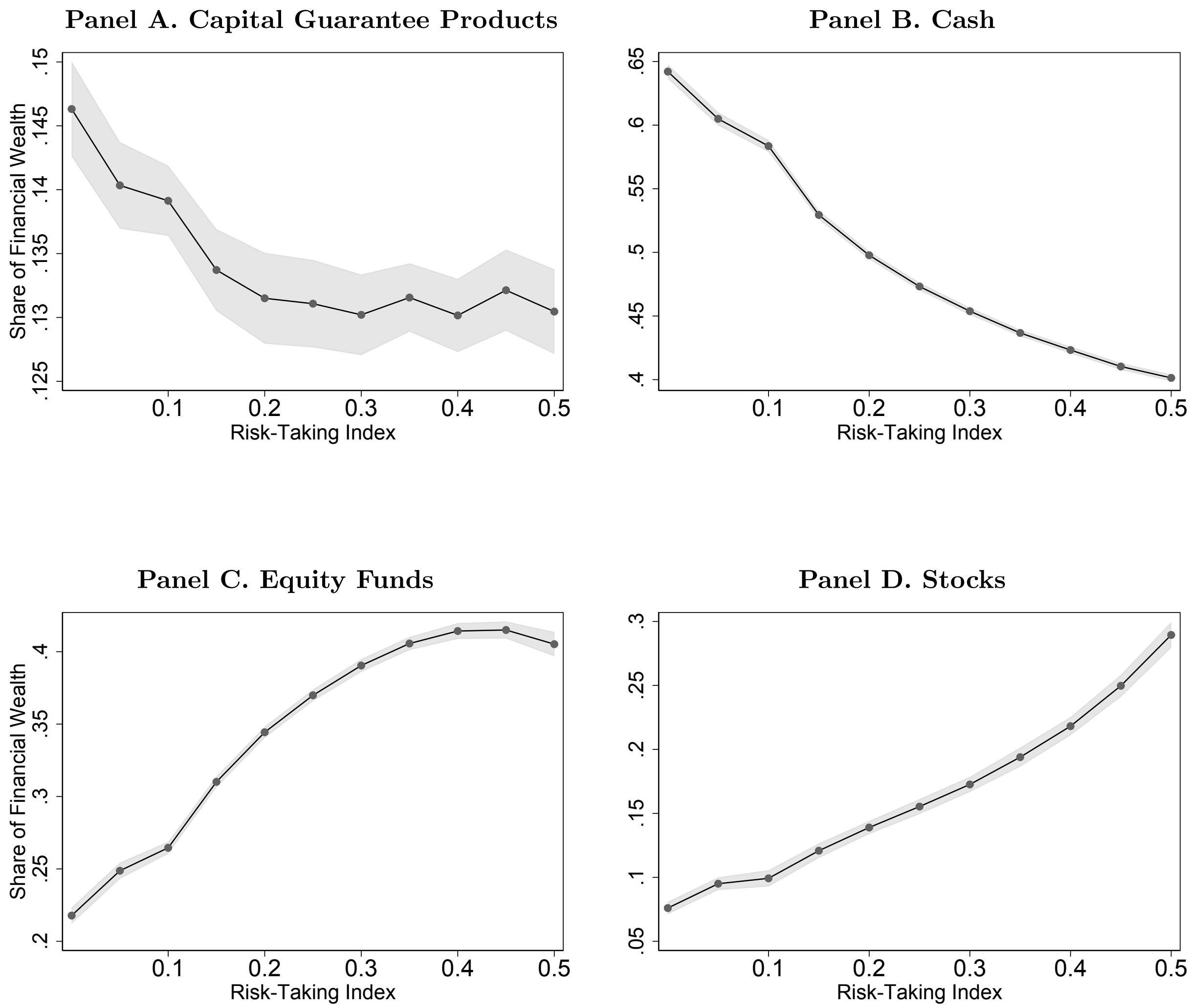


Figure 4. Allocation of 2007 Financial Portfolio as a Function of Initial Risk-Taking. This figure displays the predicted share of household financial wealth invested in capital guarantee products, cash, funds and stocks in 2007 as a function of the 2002 risk-taking index. The 2002 risk-taking index is filtered from household observable characteristics as described in Section IV.C. The sample is restricted to participants in each asset class. The shaded area represents the confidence band at the 95% level.

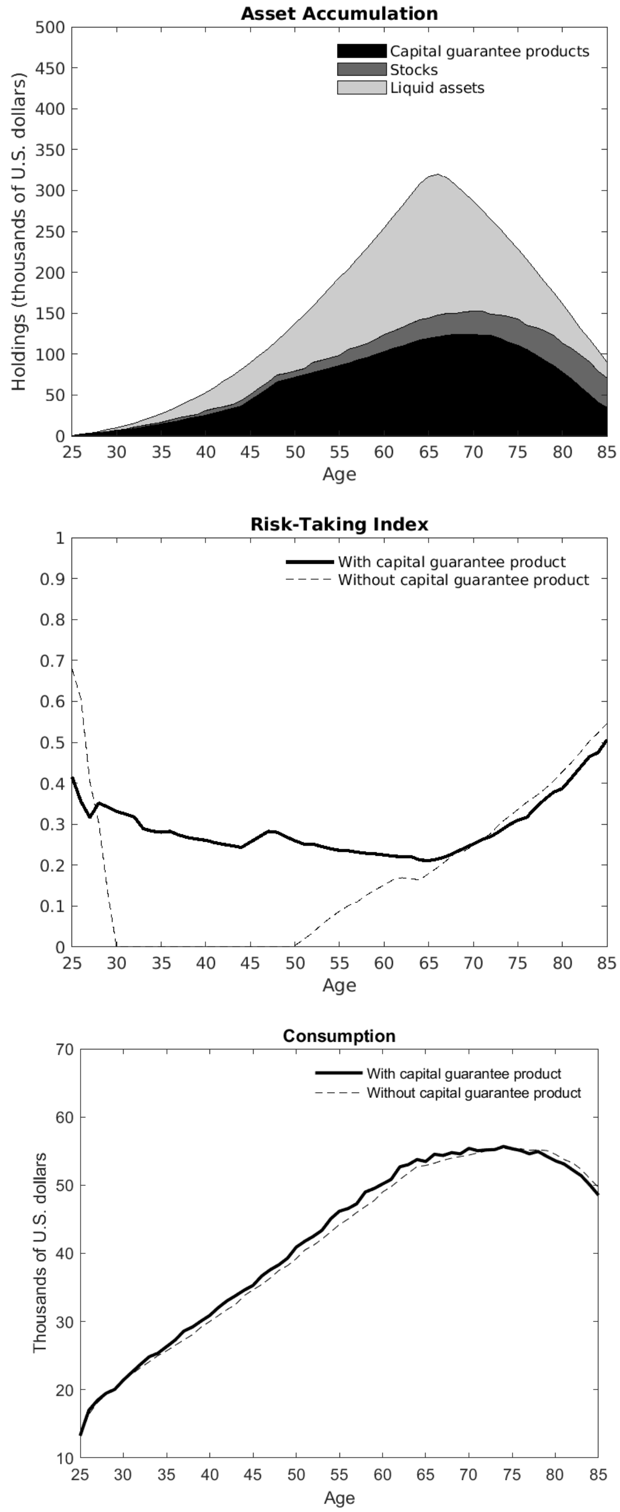
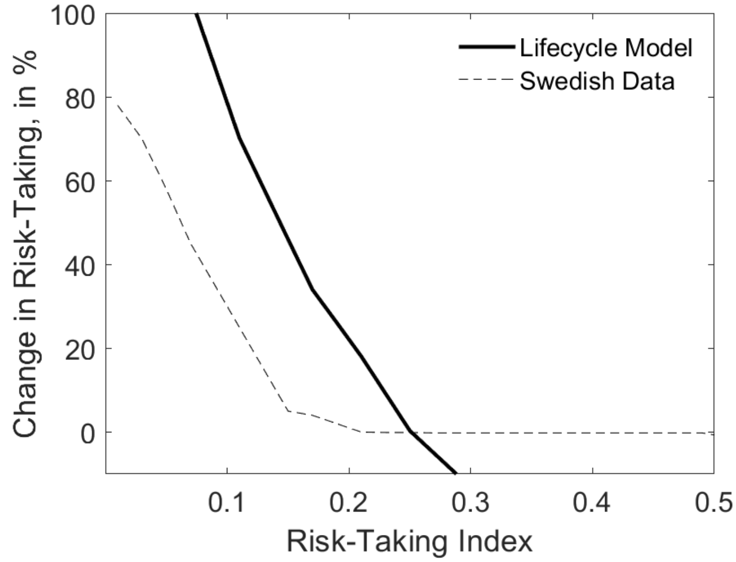


Figure 5. Life-Cycle Model with Loss Aversion and Narrow Framing. This figure displays the average portfolio allocation (Panel A), risk-taking index (Panel B), and consumption (Panel C) in a life-cycle model with equity funds, bonds, and capital guarantee products. The investor has Barberis-Huang utility with parameters $b_0 = 0.05$, $\lambda = 3.3$, $\gamma = 4$, and $\psi = 0.5$.

Panel A. Loss Aversion and Narrow Framing



Panel B. Probability Weighting

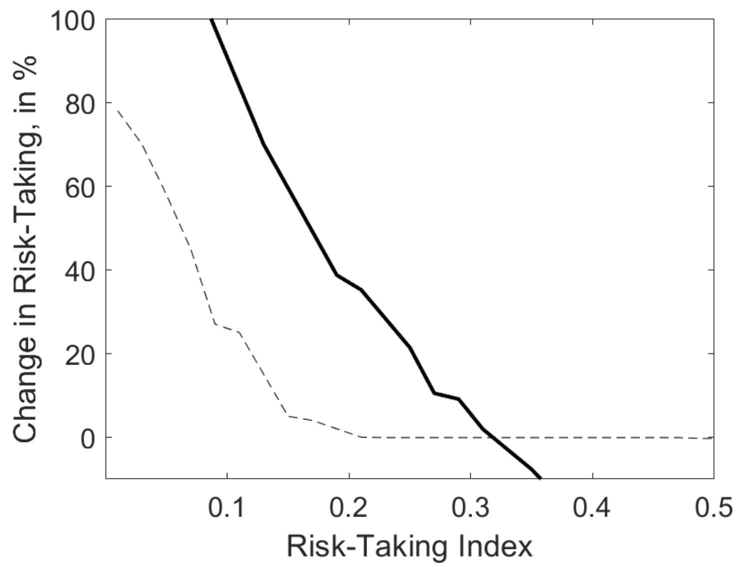


Figure 6. Change in Risk-Taking: Life-Cycle Model versus Data. This figure illustrates the relationship between initial risk-taking and the change in the risk taking index that follows the introduction of capital guarantee products. In each panel, the dashed line corresponds to empirical data, while the solid line plots the value implied by the life-cycle model with Barberis and Huang (2009) utility (Panel A) or Prelec (1998) probability weighting (Panel B). Each point is an average over households with a head between 50 and 60. The solid line is obtained by varying the kink parameter λ (Panel A) or the probability weighting parameter b (Panel B), while all other model parameters are kept constant.

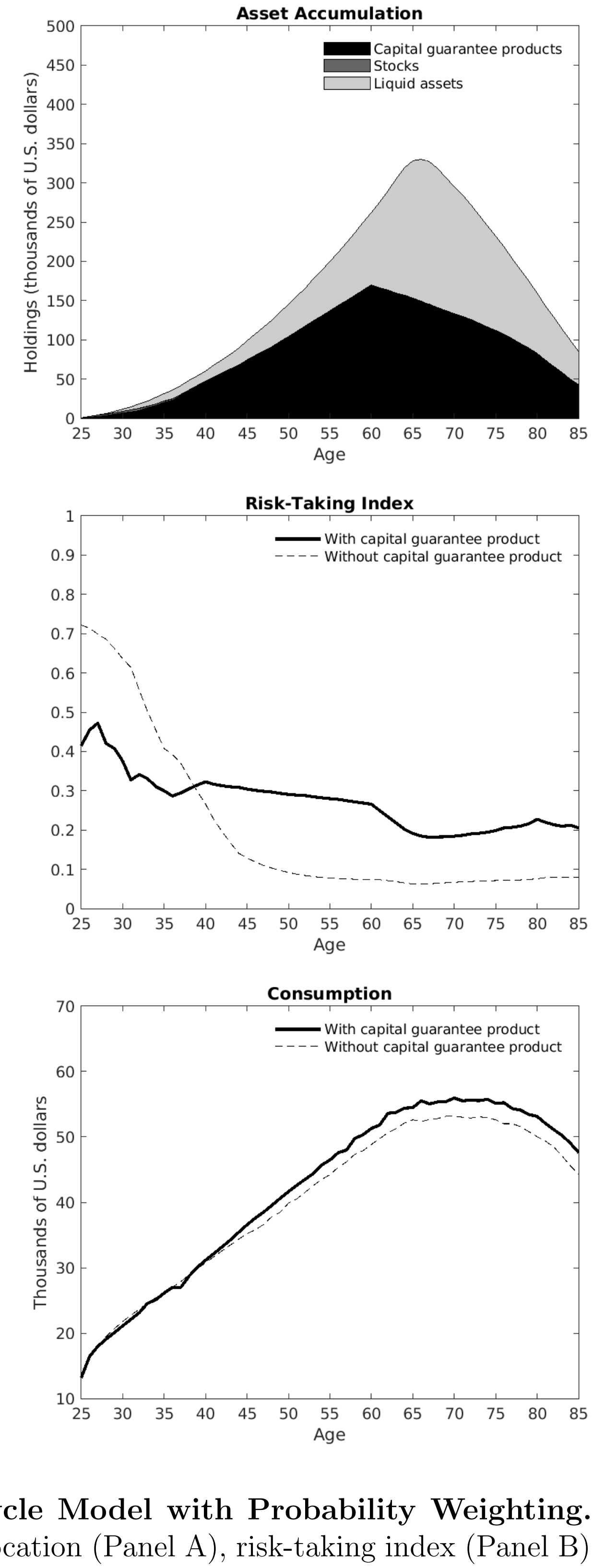


Figure 7. Life Cycle Model with Probability Weighting. This figure displays the average portfolio allocation (Panel A), risk-taking index (Panel B), and consumption (Panel C) in a life-cycle model with equity funds, bonds, and capital guarantee products. The investor has Prelec (1998) utility function with the following parameter: $a = 0.5$, $b = 0.73$, $\gamma = 4$, and $\psi = 0.5$.

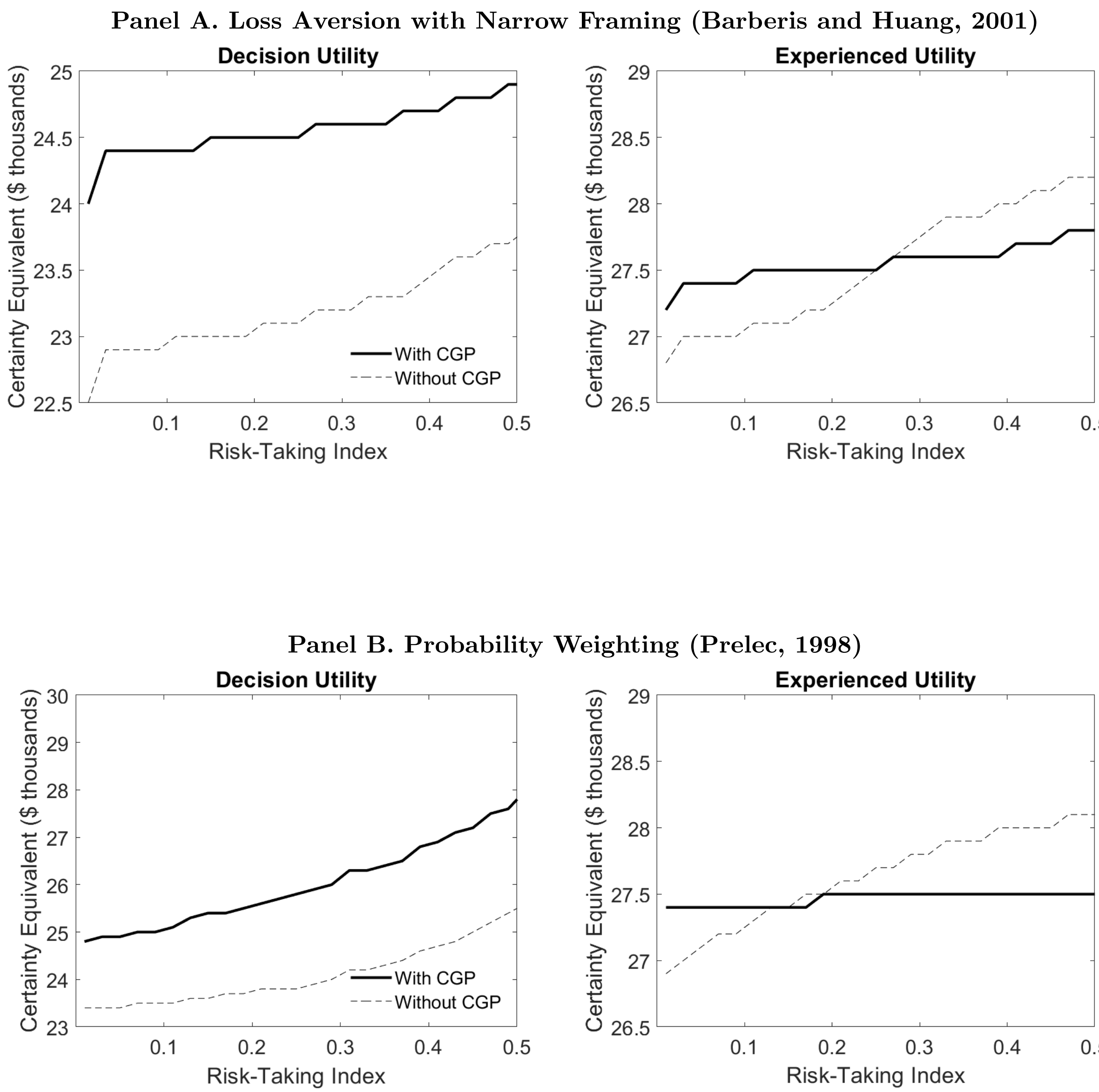


Figure 8. Welfare Implications of Capital Guarantee Products. This figure plots the welfare implications of introducing capital guarantee products under the life-cycle model with Barberis and Huang (2009) utility (Panel A) and Prelec (1998) probability weighting. For each specification, we compute the certainty equivalent before and after the introduction of the product under the decision utility (left subpanel) and the experienced utility (right subpanel). The certainty equivalent is the deterministic level of yearly consumption, assumed for simplicity to be constant over the life-cycle, that provides the same lifetime utility as the lifetime utility predicted by a model.

Internet Appendix to "Can Security Design Foster Household Risk-Taking?"

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Contents

I. The Design of CGPs: A Product Example

The *Aktieobligation Europa Trygg 98, ISIN: SE0001940107,* issued by Nordea Bank in 2004 with a volume of USD19.3m, provides a representative example of a capital guarantee product in Sweden. The payoff of this product is defined as follows.

> *The product has a maturity of 3 years and a fee of 2% is charged at issuance. The product return is linked to the performance of the Eurostoxx 50 index as follows: at maturity the product offers a minimum capital return of 100% plus 80% of the positive performance of the index over the investment period. The performance of the index is calculated as the average of the index return since inception over the last 7 months, and does not include dividends.*

Using the notation of the paper, the product is therefore specified by a maturity $M = 3$ years, a guaranteed net rate of return $g = 0$, and a participation rate $p = 1$. The issue price equal to 102% of the face value of the contract: $P_0/F = 1.02$. The benchmark return is the average performance of the Eurostoxx 50 between the issue date, $t_0 = 0$, and the $n = 7$ observation dates $t_1 = 2.5$ years, $t_2 = 2.5 + 1/12$ years, $\ldots$, $t_7 = 2.5 + 6/12 = 3$ years. Performance measures are ex dividends, consistent with the pricing approach developed in the paper.

II. Household Risk-Taking Across Countries: Design of Table I

Table I of the main text provides three measures of household risk-taking in 2015 across countries: (1) the percentage of aggregate household financial wealth invested in equity, (2) the fraction of households participating in equity markets, and (3) the median share of equity in the financial wealth of equity participants. This section provides details of the procedure used to build the table.

A. *Share of Equity in Aggregate Household Financial Wealth*

For each country or area, the percentage of aggregate household financial wealth invested in equity is the ratio of the aggregate equity holdings of households to their aggregate financial wealth.

Europe

We compute the percentage of aggregate household financial wealth invested in equity for European countries, including Sweden, as follows. We retrieve household aggregate financial wealth and detailed information on equity investments from the OECD's Consolidated Financial Balance Sheets and the Households' Financial Assets and Liabilities data, respectively.[1] Households invest in equity through direct stock ownership, equity mutual funds, mixed funds, pension funds, and life insurance. Stock ownership corresponds to the variable "Listed shares." Equity mutual funds correspond to the variable "Equity fund shares" in the "Non-money market fund shares" category and the variable for mixed funds is "Mixed fund shares." We assume that 50% of mixed funds is allocated to equity.

To obtain equity invested through pension funds, we multiply the variable "Pension entitlements" by the proportion of equity in pension funds in 2015.[2] These 2015 statistics are available from the 2019 OECD pensions report.[3]

The equity in life insurance products is imputed by multiplying the "Life insurance and annuity entitlements" variable with the 2015 equity allocation of life insurers from the OECD's 2016 Global

[1] The Consolidated Financial Balance sheets data set is available at `https://stats.oecd.org/Index.aspx?DataSetCode=SNA_TABLE710`. The Households' Financial Assets and Liabilities data set is available at `https://stats.oecd.org/Index.aspx?DataSetCode=QASA_7HH#`. Due to data availability constraints in these two data sets, we use the combined Household and Non-profit Institutions Serving Households (NPISHs) sector data for Portugal and the UK.

[2] Note that for Finland, France, and Greece, the equity allocation of pensions is unavailable in 2015, so we use the closest available year's proportions in these cases.

[3] The underlying statistical tables for the OECD's 2019 Pension Markets in Focus report is available here: `http://www.oecd.org/pensions/private-pensions/pensionmarketsinfocus.htm`. We use data from Table A.B.8 covering the allocation of assets in pension plans in equities.

Insurance Market Trends.[4]

For the European Union (EU) as a whole, we measure the aggregate equity share by the ratio of the aggregate equity to aggregate financial wealth in the 21 EU countries for which financial data are available. Bulgaria, Croatia, Cyprus, Ireland, Malta, and Romania are excluded from the analysis due to a lack of data.

United States

For the United States, we exploit data from the Financial Accounts database of the Federal Reserve because the OECD data sets are incomplete for this country.[5] More precisely, we use 2015 values from the Balance Sheet of Households and Nonprofit Organizations (Table B101).[6] Stock ownership corresponds to the variable "Corporate Equity" and equity mutual funds to the variable "Mutual fund shares." We assume that 80% of mutual fund assets are allocated to equity.

To obtain equity invested through pension funds and life insurance, we follow the same methodology as the one used for European countries. We multiply *pension entitlements* and *life insurance reserves* by the 2015 proportion of equity in pension funds (2019 OECD pensions report) and life insurance reserves (OECD's 2016 Global Insurance Market Trends), respectively.

B. Participation Rate and Median Share of Financial Wealth Invested in Equity

European Union (excluding the United Kingdom)

For European countries other than the United Kingdom, we exploit data from the 6th wave of the Survey of Health, Ageing, and Retirement in Europe (SHARE).[7] SHARE is a cross-national panel of households comprising data on health, socio-economics status, and social and family networks. The 6th wave took place in 2015 and covers approximately 140,000 households in 14 EU countries, Switzerland, and Israel. The sample is representative of the population aged 50 and above.

In our calculations, a household is classified as an equity participant if it reports that it holds either stocks or equity mutual funds. The participation rate is the number of participating households divided by the total number of respondents.

[4] We use Table 1 in the underlying statistical tables of the 2016 report. It is available at: https://www.oecd.org/pensions/globalinsurancemarkettrends.htm.

[5] This database is available here: https://www.federalreserve.gov/releases/z1/default.htm

[6] This table is available here: https://www.federalreserve.gov/releases/z1/20200611/html/b101.htm

[7] More details on the survey are available at http://www.share-project.org/.

We calculate the median share of financial wealth invested in equity as follows. Our definition of financial wealth includes equity (stocks and mutual funds), bonds, bank account balances, and long-term savings. Long-term savings consist of household retirement savings, housing savings plans, and the face value of life policies.

We assume that households invest in equity through direct stock ownership, mutual funds, and pension funds. We assume that a mutual fund or retirement account has an equity share equal to 80% if the respondent answered "mostly equity" to the asset allocation question, 50% if the respondent answered "half stocks and half bonds," and 20% if the respondent answered "mostly bonds." If the respondent did not answer the asset allocation question, we use the average equity share of participating households. We ignore equity held through life insurance contracts due to a lack of data on the composition of life insurance portfolios.

United Kingdom

For the United Kingdom, we obtain the participation rate and the median share of financial wealth invested in equity from the 7^{th} wave of the English Longitudinal Study of Ageing (ELSA).[8] The survey took place between June 2014 and May 2015 and covers 6,341 households. Since ELSA is harmonized with SHARE, we follow the exact same methodology as the one used for other European countries. The sample covers households aged 50 years and older. We assume that the estimates obtained for England are illustrative of the United Kingdom as a whole.

United States

We retrieve equity participation and the median share of equity in financial wealth from the 2016 wave of the University of Michigan Health and Retirement Study (HRS).[9] In 2016, the HRS covers 14,531 households that are representative of households with a head aged 50 and above. The survey is harmonized with SHARE and ELSA. We primarily use Section Q, which provides information on household assets and income. We supplement it with Section J2 and Section T, which cover pensions and life insurance, respectively.

In our calculations, a household is classified as an equity participant if it reports that it holds either stocks or equity mutual funds, directly or through their pension saving accounts. The partic-

[8]The data is available at: `https://beta.ukdataservice.ac.uk/datacatalogue/series/series?id=200011`.
[9]The HRS is available at: `hrs.isr.umich.edu`.

ipation rate is the number of participating households divided by the total number of respondents.

We calculate the median share of financial wealth invested in equity as follows. Our definition of financial wealth includes equity (stocks and equity mutual funds), checking accounts, interest-earning accounts (savings accounts, money market funds, certificates of deposit, governments savings bonds, and T-bills), and long-term savings. Long-term savings consist of household retirement savings (IRA and Keogh accounts) and the face value of life policies.

We assume that households invest in equity through direct stock ownership, equity mutual funds, and retirement saving accounts. We obtain the amount invested in equity through retirement saving accounts using a variable that asks respondents to declare the equity percentage of these accounts. In instances where this variable is missing, we use another variable asking if the account has an equity share less than 50%, about 50%, or more than 50%, and assume the equity share is 20%, 50%, and 80%, respectively. We ignore equity held through life insurance contracts due to lack of data on the composition of life insurance portfolios.[10]

China

We rely on the 2015 wave of the China Household Finance Survey (CHFS). The 2015 CHFS covers 29 provinces, 351 counties, and 1396 villages. It is a nationally, provincially, and sub-provincially representative sample of 37,289 households. We estimate the participation rate by the number of respondents that self-report as equity participants divided by the total number of respondents.

The financial wealth of a household in the CHFS is the total value of the cash, checking account balances, time deposit balances, financial products, funds, equity, bonds, and other financial assets that it owns. The household's equity share is the ratio of its total equity holdings divided by financial wealth.

[10]This number should be small. The OECD Life Insurance Asset Allocation Report reports that in the U.S., only 3.6% of the life insurance portfolios is allocated to equity in 2015.

III. Asset Pricing Methodology

As is explained in the main text, we assume that under the physical measure $\mathbb{P}$, the underlying S_t follows a geometric Brownian motion with constant drift and volatility:

$$\frac{dS_t}{S_t} = (\mu - q)\, dt + \sigma\, dZ_t, \tag{IA-1}$$

where q denotes the dividend yield. The capital guarantee product has a random payoff ratio $\xi \in [0, 1]$, which is taken to be independent of S_t.

The risk-adjusted measure $\mathbb{Q}$ is defined by the Radon-Nikodym derivative:

$$\frac{\mathrm{d}\,\mathbb{Q}}{\mathrm{d}\,\mathbb{P}} = \exp\left[-\left(\frac{\mu - r_f}{\sigma}\right) Z_M - \frac{M}{2}\left(\frac{\mu - r_f}{\sigma}\right)^2\right].$$

Under $\mathbb{Q}$, the underlying follows the geometric Brownian motion:

$$\frac{dS_t}{S_t} = (r_f - q)dt + \sigma dZ_t^{\mathbb{Q}}, \tag{IA-2}$$

where r_f is the continuous-time interest rate, q is the continuous-time dividend yield, σ denotes volatility, and $Z_t^{\mathbb{Q}} = Z_t + \sigma^{-1}(\mu - r_f)\, t$ is a standard Brownian motion under $\mathbb{Q}$. The payoff ratio ξ and underlying price S_t remain independent under $\mathbb{Q}$. Let $\mathbb{E}_0^{\mathbb{P}}(\,\cdot\,)$ and $\mathbb{E}_0^{\mathbb{Q}}(\,\cdot\,)$ denote the expectation operators under $\mathbb{P}$ and $\mathbb{Q}$, respectively, conditional on the information available at date 0. We denote by $\kappa = 1 - \mathbb{E}_0^{\mathbb{P}}(\xi) = 1 - \mathbb{E}_0^{\mathbb{Q}}(\xi)$ the expected loss due to default on a \$1 promise.

A. *Expected Return of Capital Guarantee Product under $\mathbb{P}$*

The return on the capital guarantee product is given by:

$$1 + R_g = \frac{F}{P_0}\left[1 + \max(p\, R^*; g)\right] \xi. \tag{IA-3}$$

We now derives its properties.

PROPOSITION 1 (Expected return under $\mathbb{P}$): *The expected return on the guaranteed product under*

$\mathbb{P}$ *over the life of the contract is given by*

$$\mathbb{E}_0^{\mathbb{P}}(1+R_g) = (1-\kappa)\,\frac{F}{P_0}\left[1+\mathbb{E}_0^{\mathbb{P}}\max(p\,R^*; g)\right]. \tag{IA-4}$$

The expected return decreases with the issue price, P_0, and increases with the guarantee rate, g. It also increases with the participation rate p if $g \geq 0$ or $\mu \geq q$.

Proof. Equation (IA-4) immediately implies that the expected return on the capital guarantee product, $ER(P_0, g, p) \equiv \mathbb{E}_0^{\mathbb{P}}(1+R_g)$, decreases with the issue price P_0.

Consider two contracts with guarantee rates g_1 and g_2, with $g_1 \geq g_2$, while all the other parameters are unchanged. We note that

$$\max(p\,R^*; g_1) \geq \max(p\,R^*; g_2) \tag{IA-5}$$

for every realization of R^*. Hence $\mathbb{E}_0^{\mathbb{P}}(\max(p\,R^*; g_1)) \geq \mathbb{E}_0^{\mathbb{P}}[\max(p\,R^*; g_2)]$, and the expected return on the first contract is higher than the expected return on the second contract: $ER(P_0, g_1, p) \geq ER(P_0, g_2, p)$.

A similar argument implies that the expected return (IA-4) increases with the participation rate p if $g \geq 0$. By contrast, if $g < 0$, we cannot apply the same simple argument because a contract with a high participation rate p_1 does not dominate a contract with a low participation rate p_2 for every realization of R^*. Indeed, a high participation rate p generates higher losses when $R^* \in [g/p; 0]$. However, we easily verify that:

$$\frac{\partial ER}{\partial p} = (1-\kappa)\,\frac{F}{P_0}\,\mathbb{E}_0^{\mathbb{P}}\left(R^*\,\mathbb{1}_{\{p\,R^* \geq g\}}\right).$$

If $g < 0$ and $\mu \geq q$, we infer that

$$\frac{\partial ER}{\partial p} \geq (1-\kappa)\,\frac{F}{P_0}\,\mathbb{E}_0^{\mathbb{P}}\left(R^*\right) = (1-\kappa)\,\frac{F}{P_0}\left[\frac{1}{n}\sum_{i=1}^{n} e^{(\mu-q)(t_i-t_0)} - 1\right] \geq 0,$$

and we conclude that the expected return increases with the participation rate. ■

The expected return (IA-4) is easily computed by Monte Carlo simulations of the underlying S_t and the resulting benchmark return R^*. A simulation $b \in \{1, \ldots, B\}$ is obtained as follows. We draw $\varepsilon_1^{(b)}, \ldots, \varepsilon_n^{(b)}$ independently from a standard Gaussian, we let

$$S_{t_i}^{(b)} = S_{t_{i-1}}^{(b)} \exp\left[\left(\mu - q - \frac{\sigma^2}{2}\right)(t_i - t_{i-1}) + \sigma\,(t_i - t_{i-1})^{1/2}\varepsilon_i^{(b)}\right],$$

and we compute the simulated benchmark return[11]

$$R^{*(b)} = \frac{S_{t_1}^{(b)} + \cdots + S_{t_n}^{(b)}}{n\,S_0}.$$

The average of $\max(p\,R^{*(b)}; g)$ across simulations provides an unbiased and consistent estimate of $\mathbb{E}_0^{\mathbb{P}}[\max(p\,R^*; g)]$.

B. *Fair Issue Price of the Capital Guarantee Product*

Under the risk-adjusted measure $\mathbb{Q}$, the expected return on the capital guarantee product is equal to the risk-free rate:

$$\mathbb{E}_0^{\mathbb{Q}}(1 + R_g) = e^{r_f\,M}. \quad \text{(IA-6)}$$

We therefore obtain the following pricing result.

PROPOSITION 2 (Fair issue price): *The fair issue price of the capital guarantee product is given by:*

$$P_0 = (1-\kappa)\,F\,e^{-r_f\,M}\,\mathbb{E}_0^{\mathbb{Q}}\left[1 + \max(p\,R^*; g)\right]. \quad \text{(IA-7)}$$

The fair issue price increases with the guaranteed return g. It also increases with the participation rate $p \in (0, +\infty)$ if $g \geq 0$ or if $r_f \geq q$.

In practice, we calculate the fair issue price by Monte Carlo simulations of the stock price under $\mathbb{Q}$.

[11] We draw an additional standard Gaussian $\varepsilon_0^{(b)}$ and stock price S_{t_0} if $t_0 > 0$.

C. Fair Price of Stream of Mutual Fund Fees

The fees earned by a mutual fund company are stochastic and earned over multiple years. To compare them to the markup of a capital guarantee product, we compute the fair value earned on an initial capital of \$1 that is fully invested in a fund over the period $t = 0, \ldots, M$. There are no intermediate withholdings or contributions. The mutual fund charges an annual fee φ and is invested in a fairly priced asset or portfolio A with annual return $1 + R_{A,t}$.

The \$1 investment in the fund generates a fee of φ and an investment of $1 - \varphi$ in A at date $t = 0$. At $t = 1$, the investment is worth $(1 - \varphi)(1 + R_{A,1})$, the mutual fund company earns a fee of $\varphi(1 - \varphi)(1 + R_{A,1})$, and the capital $(1 - \varphi)^2(1 + R_{A,1})$ remains in the fund. The fair value of the flow of fees is therefore:

$$\varphi + \sum_{t=1}^{M-1} \varphi (1 - \varphi)^t \, e^{-r_f t} \, \mathbb{E}_0^{\mathbb{Q}}[(1 + R_{A,1}) \ldots (1 + R_{A,t})] = 1 - (1 - \varphi)^M. \tag{IA-8}$$

Since the portfolio A is fairly priced, we know that $e^{-r_f t} \, \mathbb{E}_0^{\mathbb{Q}}[(1 + R_{A,1}) \ldots (1 + R_{A,t})] = 1$. The fair value of the stream of fund fees is therefore

$$\sum_{t=0}^{M-1} \varphi (1 - \varphi)^t = 1 - (1 - \varphi)^M.$$

This formula controls for the time value of money and risk premia, as equation (IA-8) shows.

The gross markup on the capital guarantee product has the same value as a stream of yearly fees over M years with the annual rate φ if $(P_0 - P_0^{\text{fair}})/P_0 = 1 - (1 - \varphi)^M$. The formula

$$\varphi_{CGP} = 1 - (P_0^{\text{fair}}/P_0)^{1/M}$$

allows us to convert the markup of a capital guaranteed product with a maturity of M years into its yearly mutual fund fee equivalent. We note that $\varphi_{CGP} \approx (P_0 - P_0^{\text{fair}})/(M \, P_0)$ for small markups.

D. Capital Guarantee Products with a Cap

A subset of contracts include a cap on the return that can be earned. The return on the capital guarantee product is then

$$1 + R_g = \frac{F}{P_0} \{1 + \min[\max(p\, R^*; g); cap]\}\ \xi, \tag{IA-9}$$

where the cap rate, *cap*, is higher than the guaranteed rate g.

The expected return and fair issue price of the capped capital guarantee product are given by:

$$\mathbb{E}_0^{\mathbb{P}}(1 + R_g) = (1 - \kappa)\, \frac{F}{P_0}\, \mathbb{E}_0^{\mathbb{P}} \{1 + \min[\max(p\, R^*; g); cap]\}\,.$$

$$P_0 = (1 - \kappa)\, F\, e^{-r_f\, M}\, \mathbb{E}_0^{\mathbb{Q}} \{1 + \min[\max(p\, R^*; g); cap]\}\,.$$

They are easily computed by Monte Carlo.

IV. Closed-Form Pricing Approximations

This section computes the moments of the benchmark return and develops approximate closed-form pricing formulas in the style of Black and Scholes, which provide further intuition on the properties of capital guarantee products. Section IV.A computes the exact first and second moments of the benchmark return. Section IV.B develops approximate closed-form formulas for the expected return and issue price of a capital guarantee product. Section IV.C provides the joint distribution of the logarithmic returns on the underlying and the capital guarantee product, which we use in the numerical implementation of the life-cycle model.

A. Moments of the Benchmark Return

We derive the moments of the benchmark return,

$$1 + R^* = \frac{S_{t_1} + \cdots + S_{t_n}}{n\, S_{t_0}}, \tag{IA-10}$$

used in the definition of the final payoff of the capital guarantee product.

PROPOSITION 3 (Moments of benchmark return): *The benchmark return satisfies*

$$m_1^{\mathbb{P}} = \mathbb{E}_0^{\mathbb{P}}(1 + R^*) = \frac{1}{n} \sum_{i=1}^{n} e^{(\mu - q)(t_i - t_0)}, \tag{IA-11}$$

$$m_2^{\mathbb{P}} = \mathbb{E}_0^{\mathbb{P}}\left[(1 + R^*)^2\right] = \frac{1}{n^2} \sum_{i=1}^{n} \sum_{j=1}^{n} e^{[2(\mu - q) + \sigma^2][\min(t_i, t_j) - t_0] + (\mu - q)|t_j - t_i|}, \tag{IA-12}$$

under the physical measure $\mathbb{P}$. *Similarly, the moments of benchmark return are given by*

$$m_1^{\mathbb{Q}} = \mathbb{E}_0^{\mathbb{Q}}(1 + R^*) = \frac{1}{n} \sum_{i=1}^{n} e^{(r_f - q)(t_i - t_0)}, \tag{IA-13}$$

$$m_2^{\mathbb{Q}} = \mathbb{E}_0^{\mathbb{Q}}\left[(1 + R^*)^2\right] = \frac{1}{n^2} \sum_{i=1}^{n} \sum_{j=1}^{n} e^{[2(r_f - q) + \sigma^2][\min(t_i, t_j) - t_0] + (r_f - q)|t_j - t_i|}, \tag{IA-14}$$

under the risk-adjusted probability measure $\mathbb{Q}$.

Proof. We infer from (IA-1) and Ito's lemma that

$$\ln(S_t/S_{t_0}) \sim \mathcal{N}\left[(\mu - q - \sigma^2/2)(t - t_0)\,,\, \sigma^2(t - t_0)\right] \tag{IA-15}$$

for every $t > t_0$. Hence $\mathbb{E}_0^{\mathbb{P}}(S_{t_i}/S_{t_0}) = e^{(\mu-q)(t_i-t_0)}$ for every $i \in \{1, \ldots, n\}$.

The first moment of the benchmark return (IA-10) is therefore

$$m_1^{\mathbb{P}} = \frac{1}{n}\sum_{i=1}^{n} \mathbb{E}_0^{\mathbb{P}}\left(\frac{S_{t_i}}{S_{t_0}}\right) = \frac{1}{n}\sum_{i=1}^{n} e^{(\mu-q)(t_i-t_0)}.$$

so that equation (IA-11) holds.

Since $(1 + R^*)^2 = n^{-2}\sum_{i=1}^{n}\sum_{j=1}^{n}(S_{t_i}/S_{t_0})(S_{t_j}/S_{t_0})$, the second moment of the benchmark return satisfies

$$m_2^{\mathbb{P}} = \frac{1}{n^2}\sum_{i=1}^{n}\sum_{j=1}^{n} \mathbb{E}_0^{\mathbb{P}}\left(\frac{S_{t_i}}{S_{t_0}}\frac{S_{t_j}}{S_{t_0}}\right).$$

If $i \leq j$, the product

$$\frac{S_{t_i}}{S_{t_0}}\frac{S_{t_j}}{S_{t_0}} = \left(\frac{S_{t_i}}{S_{t_0}}\right)^2 \frac{S_{t_j}}{S_{t_i}} \tag{IA-16}$$

has mean

$$\mathbb{E}_0^{\mathbb{P}}\left(\frac{S_{t_i}}{S_{t_0}}\frac{S_{t_j}}{S_{t_0}}\right) = e^{(\mu-q)(t_j-t_i)}\,\mathbb{E}_0^{\mathbb{P}}\left[\left(\frac{S_{t_i}}{S_{t_0}}\right)^2\right].$$

We infer from (IA-15) that

$$\begin{aligned} \mathbb{E}_0^{\mathbb{P}}\left[\left(\frac{S_{t_i}}{S_{t_0}}\right)^2\right] &= e^{2(\mu-q-\sigma^2/2)(t_i-t_0)+2\sigma^2(t_i-t_0)} \\ &= e^{[2(\mu-q)+\sigma^2](t_i-t_0)}. \end{aligned}$$

Hence

$$\mathbb{E}_0^{\mathbb{P}}\left(\frac{S_{t_i}}{S_{t_0}}\frac{S_{t_j}}{S_{t_0}}\right) = e^{[2(\mu-q)+\sigma^2](t_i-t_0)+(\mu-q)(t_j-t_i)}$$

for all $i \leq j$. Thus

$$\mathbb{E}_0^{\mathbb{P}}\left(\frac{S_{t_i}}{S_{t_0}}\frac{S_{t_j}}{S_{t_0}}\right) = e^{[2(\mu-q)+\sigma^2][\min(t_i,t_j)-t_0]+(\mu-q)|t_j-t_i|}.$$

for all i and j, and equation (IA-12) holds.

By a similar derivation, the first and second moments of the benchmark return under the risk-adjusted measure $\mathbb{Q}$ satisfy (IA-13) and (IA-14). ■

Specialized Example. Assume that the benchmark is computed f times a month over the last L months of the product. For instance, the frequency f is 1 when the index is recorded once month, 2 if it is recorded twice month, or 0.5 if it is recorded every other month. In our notation, the number of observations is $n = L\,f$ and the time interval between two consecutive observations is $t_i - t_{i-1} = 1/(12\,f)$ in annual units. The index is therefore recorded at dates

$$t_i = M - \frac{n-i}{12\ f}, \tag{IA-17}$$

where $i = 1, ..., n$. We also assume that $t_0 = 0$.

Let

$$\begin{aligned} a^{\mathbb{P}} &= e^{\frac{\mu-q}{12\,f}}, \qquad b^{\mathbb{P}} = e^{\frac{2(\mu-q)+\sigma^2}{12f}}, \\ a^{\mathbb{Q}} &= e^{\frac{r_f-q}{12\,f}}, \qquad b^{\mathbb{Q}} = e^{\frac{2(r_f-q)+\sigma^2}{12f}}. \end{aligned}$$

We easily infer from Proposition 3 the following results, which facilitate the implementation of the closed-form pricing approximations.

PROPOSITION 4: *When the benchmark return is computed from values of the index recorded on the evenly spaced time grid* (IA-17), *the first two moments of R^* satisfy*

$$m_1^{\mathbb{P}} = \frac{(a^{\mathbb{P}})^{12fM}}{n} \frac{1-(a^{\mathbb{P}})^{-n}}{1-(a^{\mathbb{P}})^{-1}}, \tag{IA-18}$$

$$m_2^{\mathbb{P}} = \frac{(b^{\mathbb{P}})^{12fM}}{n^2\,(a^{\mathbb{P}}-1)} \left[2a^{\mathbb{P}} \frac{1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-n}}{1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-1}} - (a^{\mathbb{P}}+1)\frac{1-(b^{\mathbb{P}})^{-n}}{1-(b^{\mathbb{P}})^{-1}}\right], \tag{IA-19}$$

$$m_1^{\mathbb{Q}} = \frac{(a^{\mathbb{Q}})^{12fM}}{n} \frac{1-(a^{\mathbb{Q}})^{-n}}{1-(a^{\mathbb{Q}})^{-1}}, \tag{IA-20}$$

$$m_2^{\mathbb{Q}} = \frac{(b^{\mathbb{Q}})^{12fM}}{n^2\,(a^{\mathbb{Q}}-1)} \left[2a^{\mathbb{Q}} \frac{1-(b^{\mathbb{Q}}/a^{\mathbb{Q}})^{-n}}{1-(b^{\mathbb{Q}}/a^{\mathbb{Q}})^{-1}} - (a^{\mathbb{Q}}+1)\frac{1-(b^{\mathbb{Q}})^{-n}}{1-(b^{\mathbb{Q}})^{-1}}\right]. \tag{IA-21}$$

Proof. By (IA-11), the first moment of the benchmark return under $\mathbb{P}$ satisfies:

$$m_1^{\mathbb{P}} = \frac{1}{n}\sum_{i=1}^{n} e^{(\mu-q)\left(M-\frac{n-i}{12\,f}\right)} = \frac{e^{(\mu-q)\,M}}{n}\sum_{i=0}^{n-1}(a^{\mathbb{P}})^{-i},$$

and equation (IA-18) holds.

We note that

$$\sum_{j=1}^{n} e^{[2(\mu-q)+\sigma^2]\min(t_i,t_j)+(\mu-q)|t_j-t_i|} = \sum_{j=1}^{i} e^{[2(\mu-q)+\sigma^2]\,t_j+(\mu-q)(t_i-t_j)} + \sum_{j=i+1}^{n} e^{[2(\mu-q)+\sigma^2]\,t_i+(\mu-q)(t_j-t_i)}$$

can be rewritten as

$$\begin{aligned}
& e^{[2(\mu-q)+\sigma^2]\,t_i}\left[\sum_{j=1}^{i}\left(\frac{b^{\mathbb{P}}}{a^{\mathbb{P}}}\right)^{-(i-j)} + \sum_{j=i+1}^{n}(a^{\mathbb{P}})^{j-i}\right] \\
&= (b^{\mathbb{P}})^{12fM-n+i}\left\{\frac{1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-i}}{1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-1}} + \frac{a^{\mathbb{P}}}{a^{\mathbb{P}}-1}[(a^{\mathbb{P}})^{n-i}-1]\right\} \\
&= \frac{(b^{\mathbb{P}})^{12fM}}{1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-1}}\left[(b^{\mathbb{P}})^{-n+i}-(b^{\mathbb{P}})^{-n}\,(a^{\mathbb{P}})^{i}\right] + \frac{a^{\mathbb{P}}\,(b^{\mathbb{P}})^{12fM}}{a^{\mathbb{P}}-1}\left[\left(\frac{b^{\mathbb{P}}}{a^{\mathbb{P}}}\right)^{-n+i}-(b^{\mathbb{P}})^{-n+i}\right].
\end{aligned}$$

This expression, combined with (IA-12), implies that the second moment of the benchmark return satisfies:

$$m_2^{\mathbb{P}} = X_1 + X_2, \tag{IA-22}$$

where

$$\begin{aligned}
X_1 &= \frac{(b^{\mathbb{P}})^{12fM}}{n^2\,[1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-1}]}\left\{\frac{1-(b^{\mathbb{P}})^{-n}}{1-(b^{\mathbb{P}})^{-1}} - \frac{a^{\mathbb{P}}}{a^{\mathbb{P}}-1}\,(b^{\mathbb{P}})^{-n}\,[(a^{\mathbb{P}})^{n}-1]\right\}, \\
X_2 &= \frac{a^{\mathbb{P}}\,b^{12fM}}{n^2\,(a^{\mathbb{P}}-1)}\left[\frac{1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-n}}{1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-1}} - \frac{1-(b^{\mathbb{P}})^{-n}}{1-(b^{\mathbb{P}})^{-1}}\right].
\end{aligned}$$

We note that

$$
\begin{aligned}
X_1 &= \frac{(b^{\mathbb{P}})^{12fM}}{n^2\left[1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-1}\right]}\left\{\frac{1-(b^{\mathbb{P}})^{-n}}{1-(b^{\mathbb{P}})^{-1}}+\frac{a^{\mathbb{P}}}{a^{\mathbb{P}}-1}\left[1-\left(\frac{b^{\mathbb{P}}}{a^{\mathbb{P}}}\right)^{-n}\right]-\frac{a^{\mathbb{P}}}{a^{\mathbb{P}}-1}\left[1-(b^{\mathbb{P}})^{-n}\right]\right\} \\
&= \frac{(b^{\mathbb{P}})^{12fM}}{n^2\left[1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-1}\right](a^{\mathbb{P}}-1)}\frac{1-(b^{\mathbb{P}})^{-n}}{1-(b^{\mathbb{P}})^{-1}}\left\{a^{\mathbb{P}}-1-a^{\mathbb{P}}\left[1-(b^{\mathbb{P}})^{-1}\right]\right\} \\
&\quad +a^{\mathbb{P}}\,\frac{(b^{\mathbb{P}})^{12fM}}{n^2(a^{\mathbb{P}}-1)}\frac{1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-n}}{1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-1}} \\
&= -\frac{(b^{\mathbb{P}})^{12fM}}{n^2(a^{\mathbb{P}}-1)}\frac{1-(b^{\mathbb{P}})^{-n}}{1-(b^{\mathbb{P}})^{-1}}+a^{\mathbb{P}}\,\frac{(b^{\mathbb{P}})^{12fM}}{n^2(a^{\mathbb{P}}-1)}\frac{1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-n}}{1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-1}},
\end{aligned}
$$

and therefore

$$
X_1 = \frac{(b^{\mathbb{P}})^{12fM}}{n^2(a^{\mathbb{P}}-1)}\left[a^{\mathbb{P}}\,\frac{1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-n}}{1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-1}}-\frac{1-(b^{\mathbb{P}})^{-n}}{1-(b^{\mathbb{P}})^{-1}}\right]. \tag{IA-23}
$$

We infer from equations (IA-22) and (IA-23) that equation (IA-19) holds. This derivation, applied to the special case $\mu = r_f$, implies that equations (IA-20) and (IA-21) also hold. ■

B. Approximate Pricing of Capital Guarantee Product

The expected return on the capital guarantee product satisfies an approximate Black-Scholes type formula.

PROPOSITION 5 (Closed-form approximation of the expected return under $\mathbb{P}$): *The expected return on the capital guarantee product* (IA-4) *is approximately given by*

$$
\mathbb{E}^{\mathbb{P}}(1+R_g) \approx (1-\kappa)\,\frac{F}{P_0}\left[1+g+p\,m_1^{\mathbb{P}}N(d_1^{\mathbb{P}})-(p+g)N(d_2^{\mathbb{P}})\right], \tag{IA-24}
$$

where $m_1^{\mathbb{P}} = \mathbb{E}_0^{\mathbb{P}}(1+R^*)$ *and* $m_2^{\mathbb{P}} = \mathbb{E}_0^{\mathbb{P}}\left[(1+R^*)^2\right]$ *denote the first two moments of the benchmark return computed in Proposition 3,*

$$
(w^{\mathbb{P}})^2 = \ln\left[m_2^{\mathbb{P}}/(m_1^{\mathbb{P}})^2\right] \tag{IA-25}
$$

$$
d_1^{\mathbb{P}} = \frac{1}{w^{\mathbb{P}}}\left[\ln\left(\frac{p}{p+g}\right)+\ln(m_1^{\mathbb{P}})+\frac{(w^{\mathbb{P}})^2}{2}\right], \tag{IA-26}
$$

and $d_2^{\mathbb{P}} = d_1^{\mathbb{P}} - w^{\mathbb{P}}$*. Furthermore, the volatility of the underlying,* σ*, increases the approximate expected return* (IA-24)*.*

The approximation is based on the Edgeworth expansion of the distribution of the benchmark index. It is exact if the option is European ($n = 1$) and is increasingly accurate when $n \to n$.

Proof. By the Edgeworth expansion (Turnbull and Wakeman, 1991), we approximate the distribution of the benchmark return $1 + R^*$ by a lognormal with mean $m_1^{\mathbb{P}}$ and second moment $m_2^{\mathbb{P}}$. The lognormal approximation is exact if the capital guarantee contract is a European option, that is if $n = 1$. It also applies with increasing accuracy as the number of observations n goes to infinity. The variance of the log benchmark return, $Var^{\mathbb{P}}[\ln(1 + R_M^*)]$, is then approximately equal to (IA-25), which follows directly from the properties of the lognormal distribution.

We use the following standard result.[12] If V is lognormally distributed and the standard deviation of $\ln(V)$ is s, then

$$\mathbb{E}\left[\max(V - K, 0)\right] = \mathbb{E}(V)\, N(d_1) - K\, N(d_2) \qquad \text{(IA-27)}$$

for every $K > 0$, where $d_1 = \left\{\ln[\mathbb{E}(V)/K] + s^2/2\right\}/s$ and $d_2 = d_1 - s$.

The return on the capital guarantee product can be rewritten as

$$1 + R_g = \frac{F}{P_0}\left[1 + g + p \max(1 + R^* - 1 - g/p; 0)\right] \xi.$$

We infer from (IA-27) that

$$\mathbb{E}_0^{\mathbb{P}}\left[\max(1 + R^* - 1 - g/p; 0)\right] \quad \approx \quad m_1^{\mathbb{P}}\, N(d_1^{\mathbb{P}}) - (1 + g/p) N(d_2^{\mathbb{P}}).$$

We conclude that the Proposition holds. ■

We similarly develop a closed-form approximation for the fair issue price.

PROPOSITION 6 (Approximation of fair issue price): *The fair issue price of the capital guarantee product is approximately given by*

$$P_0 \approx (1 - \kappa)\, F\, e^{-r_f M}\left[1 + g + p\, m_1^{\mathbb{Q}}\, N(d_1^{\mathbb{Q}}) - (p + g)\, N(d_2^{\mathbb{Q}})\right], \qquad \text{(IA-28)}$$

[12] See, e.g., Hull (2017) for a textbook derivation.

where $m_1^{\mathbb{Q}}$ and $m_2^{\mathbb{Q}}$ are defined in Proposition 3,

$$\begin{aligned}
(w^{\mathbb{Q}})^2 &= \ln\left[m_2^{\mathbb{Q}}\,(m_1^{\mathbb{Q}})^{-2}\right], && \text{(IA-29)}\\
d_1^{\mathbb{Q}} &= \frac{1}{w^{\mathbb{Q}}}\left[\ln\left(\frac{p}{p+g}\right) + \ln(m_1^{\mathbb{Q}}) + \frac{(w^{\mathbb{Q}})^2}{2}\right] && \text{(IA-30)}\\
d_2^{\mathbb{Q}} &= d_1^{\mathbb{Q}} - w^{\mathbb{Q}}. && \text{(IA-31)}
\end{aligned}$$

The fair issue price IA-28 increases with the volatility of the underlying.

C. *Joint Distribution of Underlying and Benchmark Returns in Last Year of Contract*

We now derive the joint distribution of the return on the underlying and the benchmark return in the last year of a contract. This distribution is an important input for the numerical implementation of the lifecycle model defined in the main text.

The contract is issued at date $t-(M-1)$ and reaches maturity at date $t+1$. The cumulative return at date t is

$$CR_t = \frac{S_t}{S_{t-M+1}} = e^{-(M-1)\,q}\,(1+R_{m,t-M+2})\ldots(1+R_{m,t}).$$

In the last year of the contract, the log return on the underlying is:

$$r_{m,t+1} = \ln(1+R_{m,t+1}) = \ln\left(\frac{S_{t+1}}{S_t}\,e^{q}\right).$$

The log benchmark return is

$$r^*_{t+1} = \ln(1+R^*_{t+1}) = \ln\left(\frac{S_{t_1}+\ldots+S_{t_n}}{n\,S_{t-M+1}}\right).$$

We assume that the observation used to compute the benchmark return are all contained in the last year of the contract: $t \le t_1 < \cdots < t_n \le t+1$, as is the case of most capital guarantee products in our data set.

PROPOSITION 7 (Conditional distribution of underlying and benchmark returns): *Conditional on the cumulative return CR_t, the log return on the underlying and the log benchmark return are* *approximately jointly Gaussian under $\mathbb{P}$:*

$$\begin{pmatrix} r_{m,t+1} \\ r^*_{t+1} \end{pmatrix} \sim \mathcal{N}\left[\begin{pmatrix} \mu - \frac{\sigma^2}{2} \\ \mu^* \end{pmatrix}, \begin{pmatrix} \sigma^2 & \sigma_{m,*} \\ \sigma_{m,*} & \sigma^2_{**} \end{pmatrix}\right], \tag{IA-32}$$

where

$$\begin{aligned}
m_1^{**} &= \frac{1}{n}\sum_{i=1}^{n} e^{(\mu-q)(t_i-t)}, \\
m_2^{**} &= \frac{1}{n^2}\sum_{i=1}^{n}\sum_{j=1}^{n} e^{[2(\mu-q)+\sigma^2][\min(t_i,t_j)-t]+(\mu-q)|t_j-t_i|}, \\
\mu^* &= \ln\left[CR_t\,(m_1^{**})^2\,(m_2^{**})^{-0.5}\right], \\
\sigma^2_{**} &= \ln[m_2^{**}/(m_1^{**})^2] \qquad \text{(IA-33)}\\
\sigma_{m,*} &= \ln\left[\frac{\sum_{i=1}^{n} e^{(\mu-q+\sigma^2)(t_i-t)}}{\sum_{i=1}^{n} e^{(\mu-q)(t_i-t)}}\right]. \qquad \text{(IA-34)}
\end{aligned}$$

Furthermore, under the regular time grid (IA-17), *the moments simplify to*

$$\begin{aligned}
m_1^{**} &= \frac{(a^{\mathbb{P}})^{12f}}{n}\frac{1-(a^{\mathbb{P}})^{-n}}{1-(a^{\mathbb{P}})^{-1}}, \\
m_2^{**} &= \frac{(b^{\mathbb{P}})^{12f}}{n^2(a^{\mathbb{P}}-1)}\left[2a^{\mathbb{P}}\frac{1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-n}}{1-(b^{\mathbb{P}}/a^{\mathbb{P}})^{-1}} - (a^{\mathbb{P}}+1)\frac{1-(b^{\mathbb{P}})^{-n}}{1-(b^{\mathbb{P}})^{-1}}\right], \\
\sigma_{m,*} &= \ln\left[e^{\sigma^2}\frac{1-(c^{\mathbb{P}})^{-n}}{1-(c^{\mathbb{P}})^{-1}}\frac{1-(a^{\mathbb{P}})^{-1}}{1-(a^{\mathbb{P}})^{-n}}\right],
\end{aligned}$$

where $c^{\mathbb{P}} = e^{\frac{\mu-q+\sigma^2}{12f}}$.

Proof. The benchmark return can be decomposed as

$$R^*_{t+1} = \frac{S_{t_1}+...+S_{t_n}}{n\,S_{t-M+1}} = \frac{S_{t_1}+...+S_{t_n}}{n\,S_t}\;\frac{S_t}{S_{t-M+1}} = (1+R^{**}_{t+1})\;CR_t, \tag{IA-35}$$

where

$$1+R^{**}_{t+1} = \frac{S_{t_1}+...+S_{t_n}}{n\,S_t}.$$

By an Edgeworth expansion, $1+R^{**}_{t+1}$ is approximately lognormal, with mean m_1^{**} and second moment m_2^{**}.

Equation (IA-35) implies that the log benchmark return satisfies:

$$r^*_{t+1} = \ln(CR_t) + r^{**}_{t+1}.$$

where $r^{**}_{t+1} = \ln(1 + R^{**}_{t+1})$. Since $1 + R^{**}_{t+1}$ is approximately lognormal, we infer that $\mathbb{E}_t(r^{**}_{t+1}) = \ln\left[(m_1^{**})^2 \, (m_2^{**})^{-0.5}\right]$ and variance $Var_t(r^{**}_{t+1}) = \ln[m_2^{**}/(m_1^{**})^2]$. Conditional on CR_t, the log benchmark return r^*_t is therefore approximately normal with mean $\mathbb{E}_t(r^*_{t+1}) = \ln[CR_t \, (m_1^{**})^2 \, (m_2^{**})^{-0.5}]$ and variance $Var_t(r^*_{t+1}) = \ln[m_2^{**}/(m_1^{**})^2]$.

The conditional covariance of the log return on the underlying and the log benchmark return satisfies:

$$\mathrm{Cov}_t(r_{m,t+1}, r^*_{t+1}) = \mathrm{Cov}_t(r_{m,t+1}, r^{**}_{t+1}).$$

Recall that if $X = (X_1, X_2)$ is bivariate normal with mean $(\mu_1, \mu_2)'$ and variance-covariance matrix $\Sigma = (\sigma_{i,j})_{1\le i,j\le 2}$, then

$$\mathbb{E}^{\mathbb{P}}(e^{X_1+X_2}) = \exp\left(\mu_1 + \mu_2 + \frac{\sigma_{1,1} + \sigma_{2,2} + 2\sigma_{1,2}}{2}\right) = \mathbb{E}(e^{X_1})\,\mathbb{E}(e^{X_2})\,\exp(\sigma_{1,2}),$$

or equivalently

$$\sigma_{1,2} = \ln\left[\frac{\mathbb{E}(e^{X_1+X_2})}{\mathbb{E}(e^{X_1})\mathbb{E}(e^{X_2})}\right].$$

The conditional covariance of $r_{m,t+1}$ and r^{**}_{t+1} is therefore

$$\sigma_{m,*} = \ln\left\{\frac{\mathbb{E}^{\mathbb{P}}_t\left[(1 + R_{m,t+1})(1 + R^{**}_{t+1})\right]}{\mathbb{E}^{\mathbb{P}}_t(1 + R_{m,t+1})\;\mathbb{E}^{\mathbb{P}}_t(1 + R^{**}_{t+1})}\right\}.$$

We note that

$$\begin{aligned}
\mathbb{E}^{\mathbb{P}}_t\left[(1 + R_{m,t+1})(1 + R^{**}_{t+1})\right] &= \mathbb{E}^{\mathbb{P}}_t\left[\frac{S_{t+1}e^q}{S_t}\frac{S_{t_1} + S_{t_2} + \cdots + S_{t_n}}{n\;S_t}\right] \\
&= \frac{e^q}{n}\sum_{i=1}^n \mathbb{E}^{\mathbb{P}}_t\left[\left(\frac{S_{t_i}}{S_t}\right)^2\frac{S_{t+1}}{S_{t_i}}\right], \\
&= \frac{e^q}{n}\sum_{i=1}^n e^{[2(\mu-q)+\sigma^2](t_i-t)}e^{(\mu-q)(t+1-t_i)},
\end{aligned}$$

and therefore

$$\mathbb{E}_t^{\mathbb{P}}\left[(1+R_{m,t+1})(1+R_{t+1}^{**})\right] = \frac{e^{\mu}}{n}\sum_{i=1}^{n} e^{(\mu-q+\sigma^2)(t_i-t)}.$$

We conclude that equation (IA-34) holds. ■

V. Empirical Analysis: Robustness Tables and Figures

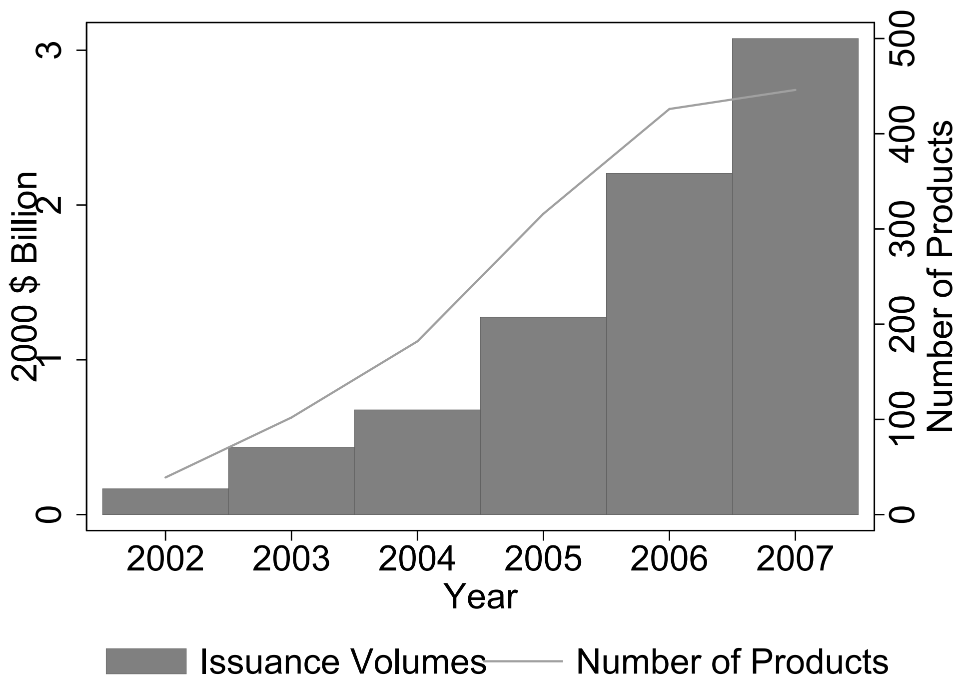


Figure IA.1. Yearly Issuance of Capital Guarantee Products in Sweden (2002-2007). This figure shows the volumes and numbers of yearly issuance of CGPs over the 2002 to 2007 period in the Swedish market. Volumes are in 2000 $ billion.

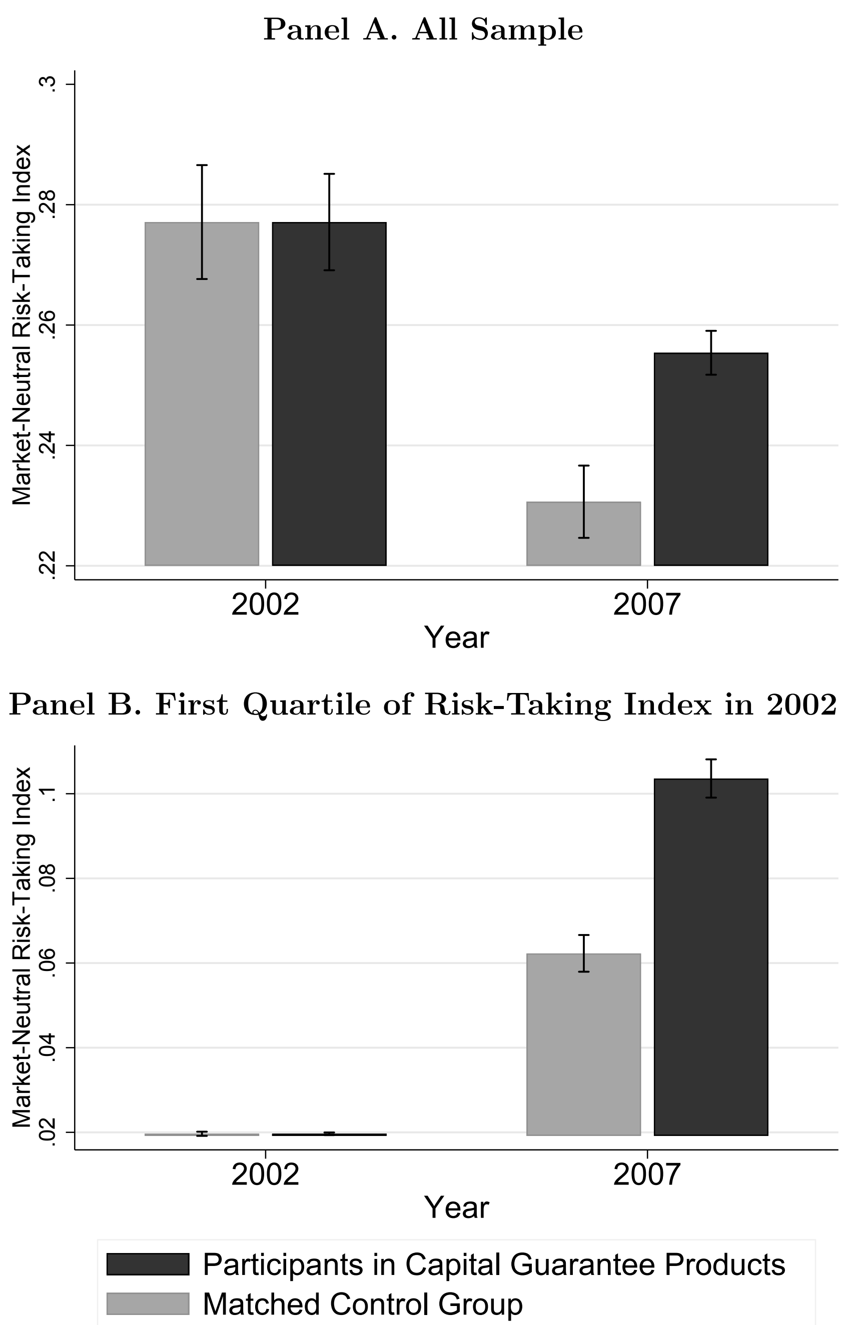


Figure IA.2. Household Market-Neutral Risk-Taking Index in 2002 and 2007. Panel A plots the "market neutral" risk-taking index in 2002 and in 2007 for: (i) capital guarantee product participants, and (ii) a control group of equal size containing stock market participants matched based on their 2002 risk-taking index. Panel B reproduces the same graph when restricting the sample to households in the first quartile of risk-taking index in 2002.

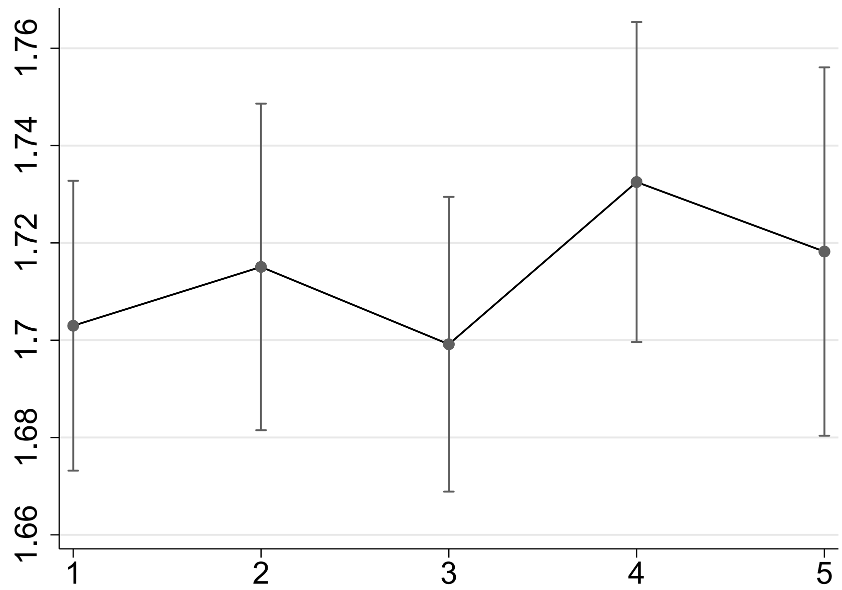


Figure IA.3. CGP Markup and IQ. This figure displays the predicted markup of CGPs as a function of household IQ setting the other household characteristics at the mean. The IQ data, resulting from military tests, is available for men born after 1945. Our analysis being conducted at the household level, we use the man's IQ when available as a proxy of the average IQ in the household. Section III describe the asset pricing exercise to obtain the CGP markups.

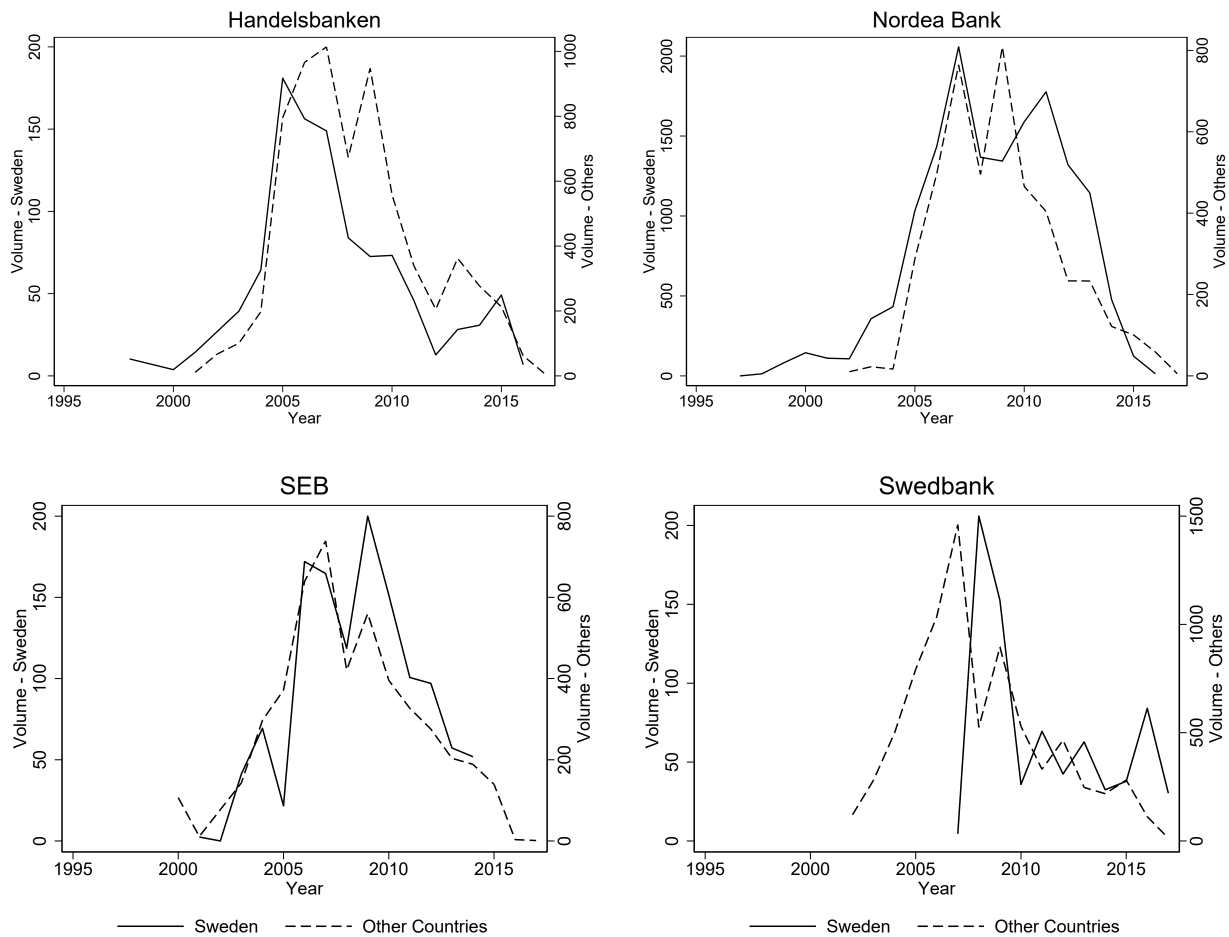


Figure IA.4. Volumes of Capital Guarantee Products Issued by the four Major Swedish Banks in Sweden versus other Markets. This figure compares the market volumes of CGPs issued by each of the four major Swedish banks in Sweden and in other countries over the 2000-2017 period. Market conditions are likely driving the downward trend in issuance volume from 2009, as it would require a significantly higher issuance price to offer the same design when interest rates are low. Volumes are in Million USD.

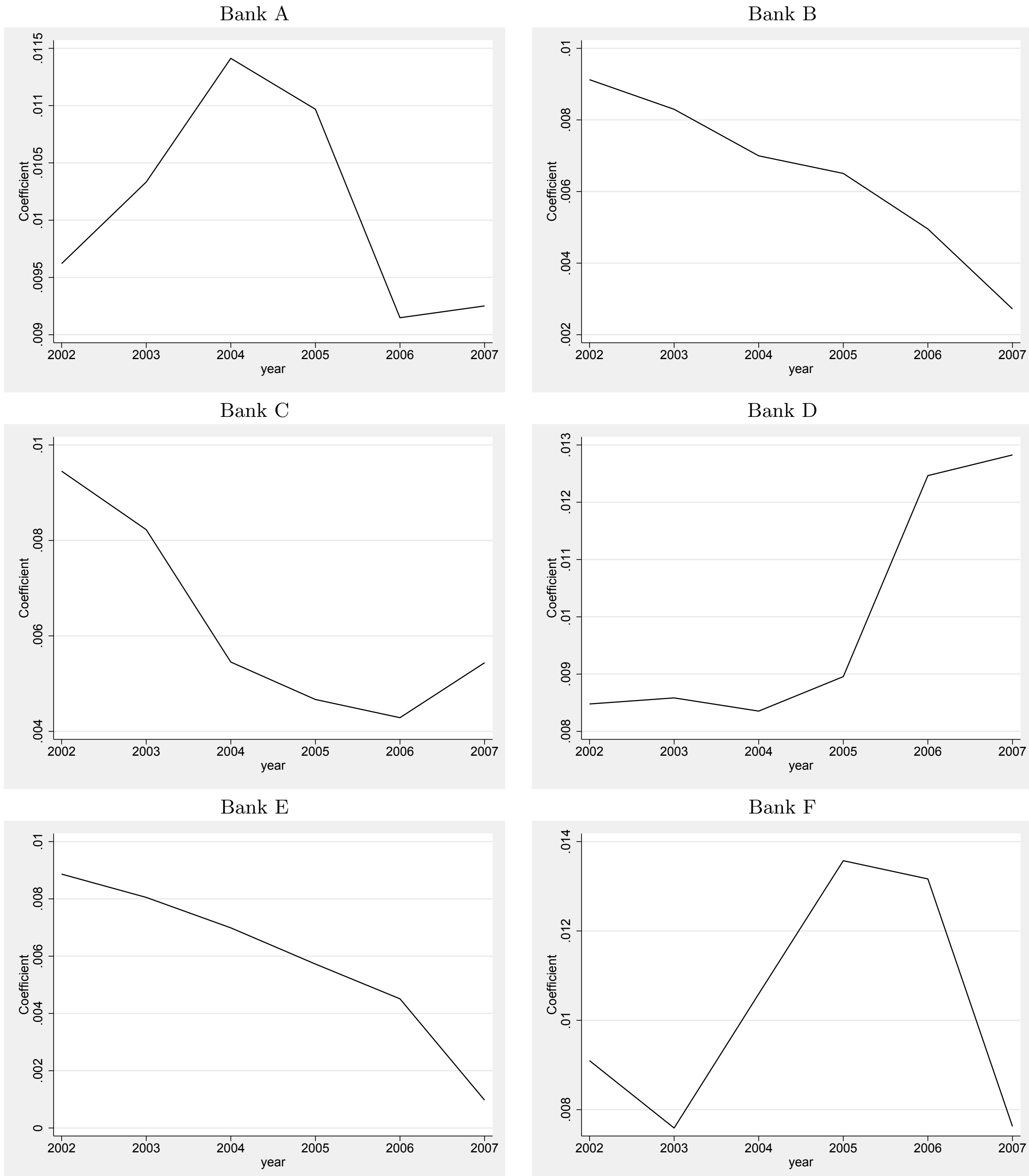


Figure IA.5. Bank Idiosyncratic Supply Shocks. This figure plots the time-varying idiosyncratic shocks to the supply of capital guarantee products. Section IV describes the methodology we use to measure the shocks.

Table IA.1. Sensitivity Analysis of the Asset Pricing Exercise

	Market Volatility (Yearly SD)				
	18%	**19%**	**20%**	**21%**	**22%**
	Yearly Mark-up (in %)				
	1.94	1.76	1.58	1.40	1.23
	Yearly Expected Excess Return (in %)				
Market Risk Premium					
4%	0.83	1.00	1.16	1.33	1.49
5%	1.60	1.75	1.90	2.06	2.21
6%	2.39	2.53	2.67	2.81	2.96
7%	3.20	3.33	3.46	3.60	3.74
8%	4.03	4.15	4.27	4.40	4.53
	Risk-Taking Index η (in %)				
Market Risk Premium					
4%	21.1	25.1	29.1	33.2	37.2
5%	32.1	35.0	38.1	41.1	44.3
6%	39.9	42.1	44.5	46.9	49.4
7%	45.7	47.5	49.4	51.4	53.4
8%	50.4	51.9	53.4	55.0	56.6

Notes: This table reports the expected excess return, yearly markup, and risk-taking index of CGPs across different parameter values for the market risk premium and the market volatility. The expected excess return, yearly markup, and risk-taking index result from the asset pricing exercise described in Section III. In the paper, the volatility parameter is the historical volatility over the 1990-2007 period and amounts to 20%, while the market risk premium amounts to 6%.

Table IA.2. Retail Structured Product Issuance across Countries and Banks

	Retail Strutured Product Volume (Million USD)					
Sample	All 2001-2015		All 2001-2008		CGPs 2001-2008	
	(1)	(2)	(3)	(4)	(5)	(6)
Adj. R^2	0.039	0.113	0.031	0.141	0.031	0.190
Bank Group × Year FE	N	Y	N	Y	N	Y
Country × Year FE	Y	Y	Y	Y	Y	Y
Observations	3,739	3,338	1,502	1,400	1,319	1,195

This table displays the results of an OLS regression where the dependent variable is the issuance volume of structured products at the bank-country-year level. Columns 1, 3 and 5 only include country-year fixed effects, while columns 2, 4 and 6 display the results of the same OLS regressions with additional bank-year fixed effects. The sample is the full sample in columns 1 and 2, the sample of products issued between 2001 and 2008 in columns 3 and 4, and CGPs issued between 2001 and 2008 in columns 5 and 6.

VI. Life-Cycle Model

This section provides the details of the lifecycle model defined in the main text. We derive the Bellman equations that define the policy functions before and after retirement. These results provide the basis of the numerical implementation.

A. Laws of Motion

A household's position in financial assets at the beginning of a period t consists of (i) cash on hand, X_t, (ii) the illiquid capital previously invested in a capital guarantee product, K_t, (iii) the cumulative return on the product's underlying index since origination, CR_t, and (iv) the time to maturity of the capital guarantee investment held by the household, or more compactly:

$$Z_t = (X_t, K_t, CR_t, \tau_t). \qquad \text{(IA-36)}$$

The vector Z_t fully characterizes the household's position in financial assets at the beginning of period t. If the household holds no capital guarantee product ($K_t = 0$), we set $CR_t = 1$ and $\tau_t = 0$. In the next subsections, we will define the household's state vector by complementing Z_t with a state variable characterizing the income process.

During period t, the household chooses the following control variables: (i) consumption, C_t, (ii) investment in the illiquid product issued at t, I_t, and (iii) the share of liquid wealth invested in the equity fund, α_t. We impose that the agent can only hold one type of capital guarantee product in any given period, so that $I_t = 0$ whenever $\tau_t > 0$.

At the beginning of period $t+1$, the household's position in financial assets, Z_{t+1}, is determined by the initial position, Z_t, the control variables (C_t, I_t, α_t) chosen at date t, the exogenous labor income Y_{t+1}, and the vector of asset returns, $R_{t+1} = (R_f, R_{eq,t+1}, R_{g,t+1})$. Consistent with the main text, R_f and $R_{eq,t+1}$ denote the returns on the riskless asset and the equity fund between t and $t+1$, while $R_{g,t+1}$ is the return on a capital guarantee product between dates $t+1-M$ and $t+1$. The components of Z_{t+1} are computed as follows.

First, cash on hand, X_{t+1}, is the sum of labor income, the value of holdings in the riskless asset and the equity fund, and the value of investments in the capital guarantee product reaching maturity at $t+1$:

$$X_{t+1} = Y_{t+1} + (X_t - I_t - C_t)\left[1 + R_f + \alpha_t(R_{eq,t+1} - R_f)\right] + (1 + R_{g,t+1})K_t\mathbb{1}_{\{\tau_t=1\}}. \tag{IA-37}$$

The term $(1 + R_{g,t+1})K_t\mathbb{1}_{\{\tau_t=1\}}$ in equation (IA-37) expresses that past investments in the capital guarantee product become liquid at $t+1$ if $\tau_t = 1$.

Second, the capital stock, K_{t+1}, is given by

$$K_{t+1} = \begin{cases} I_t \ \text{ if } \tau_t = 0, \\ 0 \quad \text{if } \tau_t = 1, \\ K_t \ \text{if } \tau_t > 1. \end{cases} \tag{IA-38}$$

Without loss of generality, the illiquid wealth invested in the capital guarantee product is not marked to market at intermediate dates. This convention has no impact on the agent's cash on hand and consumption in later periods.

Third, the cumulative return, CR_{t+1}, is given by

$$CR_{t+1} = \begin{cases} e^{-q}(1 + R_{m,t+1}) \text{ if } \tau_t = 0 \text{ and } I_t > 0, \\ 1 \quad \text{if } \tau_t = 0 \text{ and } I_t = 0, \\ e^{-q}\left(1 + R_{m,t+1}\right) CR_t \text{ if } \tau_t > 1. \end{cases} \tag{IA-39}$$

The cumulative return is reset in the year following an investment in a capital guaranteed product and is then updated each period by the ex dividend return on the underlying asset.

Fourth, time to maturity, τ_{t+1}, is

$$\tau_{t+1} = \begin{cases} \tau_t - 1 \text{ if } \tau_t > 0 \\ 0 \quad \text{if } \tau_t = 0 \text{ and } I_t = 0, \\ M - 1 \text{ if } \tau_t = 0 \text{ and } I_t > 0. \end{cases} \tag{IA-40}$$

Equations (IA-37) to (IA-40) specify the laws of motion of the system for given choices of the control variables. We summarize these relationships by the mapping:

$$Z_{t+1} = \zeta(Z_t, C_t, I_t, \alpha_t, Y_{t+1}, R_{t+1}).$$

We now investigate how an agent selects the optimal policy function (C_t, I_t, α_t) at each node.

B. *Optimal Policy Function During Retirement*

The household's financial position at the beginning of period t is fully specified by the state vector:

$$(Z_t, Y_{RA}^P),$$

where the vector Z_t defined in (IA-36) characterizes asset holdings and Y_{RA}^P denotes the permanent component of income in the last year of working life.

The value function after retirement,

$$V_t(Z_t, Y_{RA}^P) = V_t(X_t, K_t, CR_t, \tau_t, Y_{RA}^P),$$

satisfies the Bellman equation

$$V_t(Z_t, Y_{RA}^P) = \max_{\{C_t, I_t, \alpha_t\}} \left\{(1-\delta)C_t^{1-1/\psi} + \delta\, p_t \left[\mu_{t+1}(Z_t, C_t, I_t, \alpha_t, Y_{RA}^P, R_{t+1})\right]^{1-1/\psi}\right\}^{\frac{1}{1-1/\psi}} \quad \text{(IA-41)}$$

for every $t \in \{RA, \ldots, T-1\}$. We also assume that the household consumes all its cash on hand at the terminal date, so that $V_T \equiv (1-\delta)^{\frac{1}{1-1/\psi}} X_T$.

The certainty equivalent is specified as in Barberis and Huang (2009):

$$\mu_{t+1}(Z_t, C_t, I_t, \alpha_t, Y_{RA}^P, R_{t+1}) = \left\{\mathbb{E}_t[V_{t+1}(Z_{t+1}, Y_{RA}^P)^{1-\gamma}]\right\}^{\frac{1}{1-\gamma}} + b_0 \mathbb{E}_t[v(W_{t+1} - W_{t+1}^R)], \quad \text{(IA-42)}$$

where $Z_{t+1} = \zeta(Z_t, C_t, I_t, \alpha_t, Y_{RA}^P, R_{t+1})$ characterizes the household's asset position next period, $b_0 \geq 0$ is a constant, $v(\cdot)$ is a piecewise linear function specified by the kink parameter $\lambda \geq 1$:

$$v(x) = \begin{cases} x & \text{if } x \geq 0, \\ \lambda x & \text{if } x \leq 0, \end{cases}$$

W_{t+1} is the value of liquid financial wealth at the beginning of period $t+1$:

$$W_{t+1} = (X_t - I_t - C_t)\ [1 + R_f + \alpha_t(R_{eq,t+1} - R_f)] + K_t\,(1 + R_{g,t+1})\mathbb{1}_{\{\tau_t=1\}},$$

and W^R_{t+1} is a reference level. The reference level is set equal to the value of the investment if the agent invests in the riskless asset:

$$W^R_{t+1} = (X_t - C_t - I_t)(1 + R_f) + K_t\,(1 + R_f)^M\,\mathbb{1}_{\{\tau_t=1\}}. \tag{IA-43}$$

This specification implies loss aversion and ambiguity aversion if $b_0 > 0$. It reduces to Epstein-Zin preferences if $b_0 = 0$, and to constant relative risk aversion (CRRA) utility if $b_0 = 0$ and $\gamma = 1/\psi$.

The utility specification implies that the value function is homogeneous of degree 1 with respect to monetary state variables.

PROPOSITION 8 (Homogeneity): *The value function after retirement is homogenous of degree* 1 *with respect to monetary state variables:*

$$V_t(a\,X_t, a\,K_t, CR_t, \tau_t, a\,Y^P_{RA}) = a\,V_t(X_t, K_t, CR_t, \tau_t, Y^P_{RA},)$$

for every $a > 0$.

Proof. We show this property by a backward recursion. Since $V_T \equiv (1-\delta)^{\frac{1}{1-1/\psi}}\,X_T$, the proposition holds for $t = T$. We now assume that the property holds at $t+1$. Since $v(\,\cdot\,)$ is homogeneous of degree 1, we infer from (IA-42) that

$$\mu_{t+1}(a\,X_t, a\,K_t, CR_t, \tau_t, a\,C_t, a\,I_t, \alpha_t, a\,Y^P_{RA}, R_{t+1}) = a\,\mu_{t+1}(X_t, K_t, CR_t, \tau_t, C_t, I_t, \alpha_t, Y^P_{RA}, R_{t+1}),$$

for every $a \in \mathbb{R}_{++}$ and admissible value of $(X_t, K_t, CR_t, \tau_t, C_t, I_t, \alpha_t, Y^P_{RA}, R_{t+1})$.[13] The function

$$G(Z_t, C_t, I_t, \alpha_t, Y^P_{RA}, R_{t+1}) = \left\{(1-\delta)C_t^{1-1/\psi} + \delta\,p_t\,[\mu_{t+1}(Z_t, C_t, I_t, \alpha_t, Y^P_{RA}, R_{t+1})]^{1-1/\psi}\right\}^{\frac{1}{1-1/\psi}}$$

[13] That is, $(X_t, K_t, CR_t, \tau_t, C_t, I_t, \alpha_t, Y^P_{RA}, R_{t+1}) \in \mathbb{R}^3_+ \times \{0, \ldots, M-1\} \times \mathbb{R}^2_+ \times [0, 1] \times \mathbb{R}_{++} \times [-1, \infty)$.

is therefore homogeneous of degree 0 with respect to the monetary variables $(X_t, K_t, C_t, I_t, Y^P_{RA})$. We infer that the proposition holds. ■

The homogeneity property allows us to reduce the dimensionality of the optimization problem. We normalize all flow and stock variables by Y^P_{RA}. That is, we denote normalized cash on hand by $X^*_t = X_t/Y^P_{RA}$, normalized holdings of the structured product by $K^*_t = K_t/Y^P_{RA}$, the normalized state vector by

$$Z^*_t = (X^*_t, K^*_t, CR_t, \tau_t), \tag{IA-44}$$

and the normalized value function by

$$V^*_t(Z^*_t) = V_t(1, X^*_t, K^*_t; CR_t, \tau_t)$$

We now derive the optimization problem satisfied by V^*_t.

PROPOSITION 9 (Policy Function After Retirement): *At every $t \geq RA$, the normalized value function, $V^*_t(Z^*_t)$, satisfies the Bellman equation*

$$V^*_t(Z^*_t) = \max_{\{C^*_t, I^*_t, \alpha_t\}} \left[(1-\delta)\,(C^*_t)^{1-1/\psi} + \delta\, p_t\,(\mu^*_{t+1})^{1-1/\psi}\right]^{\frac{1}{1-1/\psi}}, \tag{IA-45}$$

where

$$\begin{aligned}
\mu^*_{t+1} &= \left\{\mathbb{E}_t[V_{t+1}(Z^*_{t+1})^{1-\gamma}]\right\}^{\frac{1}{1-\gamma}} + b_0\,\mathbb{E}_t[v(W^*_{t+1} - W^{*R}_{t+1})],\\
Z^*_{t+1} &= \zeta(Z^*_t, C^*_t, I^*_t, \alpha_t, \lambda, R_{t+1}),\\
W^*_{t+1} &= (X^*_t - I^*_t - C^*_t)\;[1 + R_f + \alpha_t(R_{eq,t+1} - R_f)] + K^*_t\,(1 + R_{g,t+1})\mathbb{1}_{\{\tau_t=1\}},\\
W^R_{t+1} &= (X^*_t - C^*_t - I^*_t)(1 + R_f) + K^*_t\,(1 + R_f)^M\,\mathbb{1}_{\{\tau_t=1\}},
\end{aligned}$$

for every t.

Proof. We divide the Bellman equation (IA-41) by Y^P_{RA} and obtain:

$$V^*_t(Z^*_t) = \max_{\{C^*_t, I^*_t, \alpha^*_t\}} \left\{(1-\delta)(C^*_t)^{1-1/\psi} + \delta\, p_t\,[\mu_{t+1}(Z^*_t, C^*_t, I^*_t, \alpha_t, 1, R_{t+1})]^{1-1/\psi}\right\}^{\frac{1}{1-1/\psi}}. \tag{IA-46}$$

The function $v(\cdot)$ is homogenous of degree 1. By (IA-42), the certainty equivalent in equation (IA-52) is therefore

$$\mu_{t+1}(Z_t^*, C_t^*, I_t^*, \alpha_t, 1, R_{t+1}) = \left\{\mathbb{E}_t[V_{t+1}^*(Z_{t+1}^*)^{1-\gamma}]\right\}^{\frac{1}{1-\gamma}} + b_0 \mathbb{E}_t\left[v\left(\frac{W_{t+1} - W_{t+1}^R}{Y_{RA}^P}\right)\right], \quad \text{(IA-47)}$$

and we conclude that the Proposition holds. ■

C. *Working Years*

Before retirement ($t < RA$), the financial position of the household is fully summarized by the state vector:

$$(Z_t, Y_t^P),$$

where Z_t is the vector defined by equation (IA-36) that quantifies asset holdings and Y_t^P is the permanent component of labor income. As is explained in the main text, $Y_t^P = e^{f(t;Z_t)+\nu_t}$. Let $\eta_t = \nu_{t+1} - \nu_t$. Hence $Y_{t+1}^P / Y_{t,P} = \exp(f(t+1; Z_t) - f(t; Z_t) + \eta_{t+1})$. For simplicyt, we let $g_{t+1} = \exp(f(t+1; Z_t) - f(t; Z_t)$.

The value function before retirement, $V_t(Z_t, Y_t^P) = V_t(X_t, K_t, CR_t, \tau_t, Y_t^P)$, satisfies the Bellman equation

$$V_t(Z_t, Y_t^P) = \max_{\{C_t, I_t, \alpha_t\}} \left\{(1-\delta)C_t^{1-1/\psi} + \delta\, p_t\, [\mu_{t+1}(Z_t, C_t, I_t, \alpha_t, Y_t^P, R_{t+1})]^{1-1/\psi}\right\}^{\frac{1}{1-1/\psi}} \quad \text{(IA-48)}$$

for every $t \in \{RA, \ldots, T-1\}$, where

$$\mu_{t+1}(Z_t, C_t, I_t, \alpha_t, Y_t^P, R_{t+1}) = \left\{\mathbb{E}_t[V_{t+1}(Z_{t+1}, Y_{t+1}^P)^{1-\gamma}]\right\}^{\frac{1}{1-\gamma}} + b_0 \mathbb{E}_t[v(W_{t+1} - W_{t+1}^R)]. \quad \text{(IA-49)}$$

The certainty equivalent can be rewritten as

$$\mu_{t+1}(Z_t, C_t, I_t, \alpha_t, Y_t^P, R_{t+1}) = \left\{\mathbb{E}_t[V_{t+1}(Z_{t+1}, Y_t^P\, f_t\, e^{\eta_{t+1}})^{1-\gamma}]\right\}^{\frac{1}{1-\gamma}} + b_0 \mathbb{E}_t[v(W_{t+1} - W_{t+1}^R)]. \quad \text{(IA-50)}$$

PROPOSITION 10 (Homogeneity): *The value function before retirement is homogenous of degree*

1 *with respect to monetary state variables:*

$$V_t(a\,X_t, a\,K_t, CR_t, \tau_t, a\,Y^P_{RA}) = a\,V_t(X_t, K_t, CR_t, \tau_t, Y^P_{RA},)$$

for every $a > 0$.

Proof. This result follows from the fact that when one multiplies the permanent income Y_t^P by a fixed factor $a \in \mathbb{R}_{++}$, the entire labor income process between t and retirement RA is multiplied by a. The rest of the proof is the same as the proof of Proposition 8. ■

Let $X_t^* = X_t/Y_t^P$ denote normalized cash on hand, and let $K_t^* = K_t/Y_t^P$ denote the normalized holdings of the structured product. Consider the normalized state vector

$$Z_t^* = (X_t^*, K_t^*, CR_t, \tau_t),$$

and the normalized value function $V_t^*(Z_t^*) = V_t(X_t^*, K_t^*, CR_t, \tau_t, 1)$.

PROPOSITION 11 (Policy Function During Working Years): *At every* $t \leq RA$, *the normalized value function satisfies the Bellman equation:*

$$V_t^*(Z_t^*) = \max_{\{C_t^*, I_t^*, \alpha_t\}} \left[(1-\delta)\,(C_t^*)^{1-1/\psi} + \delta\, p_t\,(\mu_{t+1}^*)^{1-1/\psi}\right]^{\frac{1}{1-1/\psi}}, \qquad \text{(IA-51)}$$

where

$$\begin{aligned}
\mu_{t+1}^* &= \left\{\mathbb{E}_t\left[\left(\frac{Y_{t+1}^P}{Y_t^P}\right)^{1-\gamma} V_t(Z_{t+1}^*)^{1-\gamma}\right]\right\}^{\frac{1}{1-\gamma}} + b_0\,\mathbb{E}_t[v(W_{t+1}^* - W_{t+1}^{*R})],\\
X_{t+1}^* &= Y_{t+1}^H + \frac{Y_t^P}{Y_{t+1}^P}(X_t^* - I_t^* - C_t^*)[1 + R_f + \alpha_t^*(R_{eq,t+1} - R_f)] + (1 + R_{g,t+1})\,K_t^*\,\mathbb{1}_{\{\tau_t=1\}},\\
K_{t+1}^* &= \frac{Y_t^P}{Y_{t+1}^P}(I_t^*\,\mathbb{1}_{\{\tau_t=0\}} + K_t^*\,\mathbb{1}_{\{\tau_t>1\}}),\\
CR_{t+1} &= \mathbb{1}_{\{\tau_t=0 \text{ and } I_t=0\}} + e^{-q}\,(1 + R_{m,t+1})\,\mathbb{1}_{\{\tau_t=0 \text{ and } I_t>0\}} + e^{-q}\,(1 + R_{m,t+1})\,CR_t\,\mathbb{1}_{\{\tau_t>1\}},\\
\tau_{t+1} &= (\tau_t - 1)\,\mathbb{1}_{\{\tau_t>0\}} + (M-1)\,\mathbb{1}_{\{\tau_t=0 \text{ and } I_t>0\}},
\end{aligned}$$

$$
\begin{aligned}
W_{t+1}^{*} &= (X_t^* - I_t^* - C_t^*)\,[1 + R_f + \alpha_t(R_{eq,t+1} - R_f)] + K_t^*\,(1 + R_{g,t+1})\mathbb{1}_{\{\tau_t = 1\}}, \\
W_{t+1}^{R} &= (X_t^* - C_t^* - I_t^*)(1 + R_f) + K_t^*\,(1 + R_f)^M\,\mathbb{1}_{\{\tau_t = 1\}},
\end{aligned}
$$

at the beginning of period $t + 1$.

Proof. We divide the Bellman equation (IA-41) by Y_{RA}^P and obtain:

$$
V_t^*(Z_t^*) = \max_{\{C_t^*, I_t^*, \alpha_t^*\}} \left\{(1-\delta)(C_t^*)^{1-1/\psi} + \delta\, p_t\,[\mu_{t+1}(Z_t^*, C_t^*, I_t^*, \alpha_t, 1, R_{t+1})]^{1-1/\psi}\right\}^{\frac{1}{1-1/\psi}}. \quad \text{(IA-52)}
$$

The function $v(\,\cdot\,)$ is homogenous of degree 1. By (IA-42), the certainty equivalent in equation (IA-52) is therefore

$$
\mu_{t+1}(Z_t^*, C_t^*, I_t^*, \alpha_t, 1, R_{t+1}) = \left\{\mathbb{E}_t[V_{t+1}^*(Z_{t+1}^*)^{1-\gamma}]\right\}^{\frac{1}{1-\gamma}} + b_0 \mathbb{E}_t\left[v\left(\frac{W_{t+1} - W_{t+1}^R}{Y_{RA}^P}\right)\right], \quad \text{(IA-53)}
$$

and we conclude that the Proposition holds. ■

VII. Life-Cycle Model under Alternative Preferences and Beliefs

A. Preferences

Consistent with Section VI, we consider a household with recursive preferences of the form

$$V_t(X_t, K_t, CR_t, \tau_t) = \max_{(C_t, I_t, \alpha_t)} \left[(1-\delta)\, C_t^{1-1/\psi} + \delta p_t\, (\mu_{t+1})^{1-1/\psi}\right]^{\frac{1}{1-1/\psi}} \tag{IA-54}$$

where $t \in \{1, ..., T-1\}$, p_t is the probability that the agent is alive at $t+1$ conditional on being alive at date t, and $V_T = (1-\delta)^{1/(1-1/\psi)}\, C_T$ at the terminal date. We solve the life-cycle model under several alternative specifications of the certainty equivalent μ_{t+1}.

Epstein-Zin Utility

In Figure IA.6, we consider an investor with the Epstein-Zin certainty equivalent

$$\mu_{t+1} = \left[\mathbb{E}_t(V_{t+1}^{1-\gamma})\right]^{\frac{1}{1-\gamma}}.$$

We set $\gamma = 20$, $\psi = 0.5$, and $\delta = 0.98$, which produces an average risk-taking index of 7% between the agents of 50 and 60 before the capital guarantee product is introduced. The figure illustrates that Epstein-Zin investors purchase the CGP, but the risk-taking index does not increase after the innovation is introduced.

Panel A of Figure IA.11 further confirms this result. The increase in risk-taking is close to zero or negative for wide range of the risk aversion parameter γ, and the model does not match the correlation between initial risk-taking and the change in the index triggered by financial innovation.

Narrow Framing with Second Order Risk Aversion

In Figure IA.7, we simulate the model when the investor exhibits narrow framing but not loss aversion. The certainty equivalent is given by:

$$\mu_{t+1} = \left[\mathbb{E}_t(V_{t+1}^{1-\gamma})\right]^{\frac{1}{1-\gamma}} + b_0\left[CE_t(W_{t+1}) - CE_t(W_{t+1}^R)\right],$$

where $CE_t(W_{t+1}) = [\mathbb{E}_t(W_{t+1}^{1-\gamma})]^{\frac{1}{1-\gamma}}$ and the benchmark W_{t+1}^R is given by (IA-43). Consistent with the approach to narrow framing in the main text, this specification does not impact the saving behavior of an agent who is only investing in bonds. We set $\gamma = 20$, $\psi = 0.5$ and $\delta = 0.98$ in the simulations reported in Figure IA.7. As is the case with Epstein-Zin investors, the household with narrow framing and second-order risk aversion purchase the CGP, but the risk-taking index does not increase as a result for households between 45 and 75. Panel B of Figure IA.11 further confirms this result.

Loss Aversion without Narrow Framing

In Figure IA.8, we consider the life-cycle model under generalized disappointment aversion (Gul, 1991; Routledge and Zin, 2010), a form of preferences that includes loss aversion but not narrow framing. The certainty equivalent is implicitly defined by:

$$(\mu_{t+1})^{1-\gamma} = \mathbb{E}_t(V_{t+1}^{1-\gamma}) + (\lambda - 1)\mathbb{E}\left\{\left[V_{t+1}^{1-\gamma} - (\kappa\,\mu_{t+1})^{1-\gamma}\right]\mathbb{1}_{\{V_{t+1}<\kappa\,\mu_{t+1}\}}\right\},$$

where $\lambda \geq 1$ is a kink parameter and κ controls the disappointment threshold. This specification reduces to disappointment aversion (Gul, 1991) if $\kappa = 1$. The investor has kink parameter $\lambda = 9.5$ in the simulations reported in Figure IA.8. General disappointment aversion does not increase the risk-taking index throughout the life-cycle. Panel C of Figure IA.11 provides further confirmation of this result.

B. *Alternative Specifications of Beliefs*

Crash Risk

In Figure IA.9, we assume that the investor's decision utility is based on a subjective probability distribution of the log return of the underlying, defined as the mixture of a crash event and a Gaussian.

- A crash occurs with probability $p = 51\%$ and corresponds to an ex dividend log return on the underlying of $r_{crash} = -\ln(4)$.
- The Gaussian distribution can be written as

$$\mathcal{N}(c + \mu - q - \sigma^2/2, \sigma^2),$$

where c denotes is a constant that acts as a compensator for crash risk. The standard deviation σ coincides with the standard deviation of the underlying under the physical measure $\mathbb{P}$ (so that $\sigma = 20\%$ per year). The compensator c is chosen so that the expected return of the underlying (cum or ex dividends) is the same as under $\mathbb{P}$. Specifically, it satisfies $p\, e^{r_{crash}} + (1-p)\, e^{\mu-q+c} = e^{\mu-q}$, or equivalently

$$c = \ln\left(\frac{1 - p\, e^{r_{crash}-\mu+q}}{1-p}\right).$$

This approach is motivated by the U.S. Crash Confidence Index conducted by Yale's International Center for Finance and discussed in Goetzmann, Kim, and Shiller (2017).[14] In any given month since the survey started in 1990, more than 60% U.S. households have expected that the probability of a crash over the following six months exceeded 10%, an estimate that largely exceeds the historical probability of a crash.

Figure IA.9 shows that financial innovation triggers a very substantial increase in the risk-taking index under this specification of pessimistic beliefs, consistent with the probability weighting specification used in the main text. Furthermore, Panel A of Figure shows that subjective crash risk also explains the relationship between the initial risk-taking index and the increase in risk-taking caused by innovation.

Volatility Misperception

In Figure IA.10, we assume that the household's subjective volatility of the underlying, σ_{mis}, is higher than the volatility under the physical measure, σ. Under the subjective probability, the log return of the underlying has Gaussian distribution

$$\mathcal{N}(\mu - q - \sigma_{\text{mis}}^2/2, \sigma_{\text{mis}}^2),$$

We can view σ_{mis} as the sum of the volatility under the physical measure, σ, and a misperception term. In IA.10, we set $\sigma = 20\%$ and the miperception term to 55%, so that the subjective probability of the underlying is 75% in annual units. The figure shows a strong increase in the risk-taking

[14]This survey is available at https://som.yale.edu/faculty-research-centers/centers-initiatives/international-center-for-finance/data/stock-market-confidence-indices/united-states-stock-market-confidence-indices.

index under a subjective volatility equal to 3.75 times historical volatility. In Panel B of Figure , we let the size of misperception vary. Heterogeneity in volatility misperception also accounts for the relationship between initial risk-taking and the increase in the risk-taking index caused by innovation.

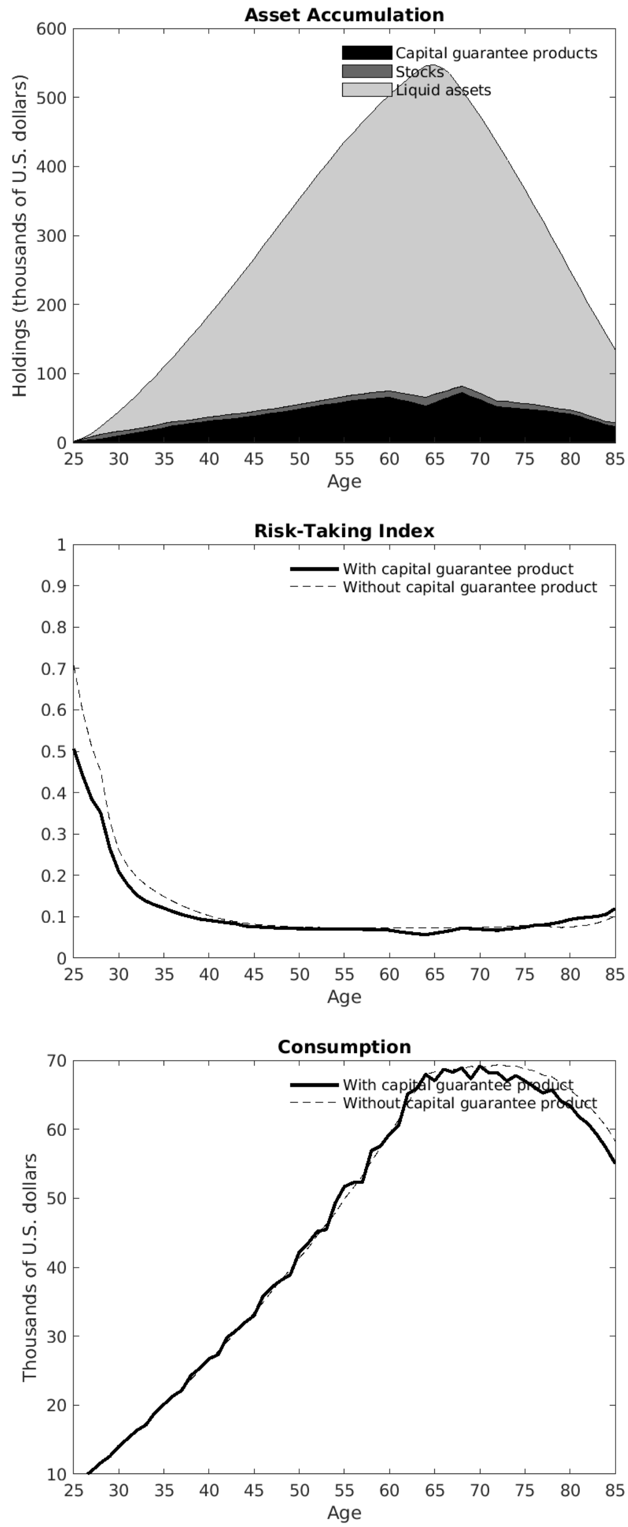


Figure IA.6. Life-Cycle Profile of an Epstein-Zin Investor. This figure illustrates the average path of asset accumulation, the risk-taking index, and consumption of an Epstein-Zin investor with parameters $\gamma = 20$, $\psi = 0.5$, and $\delta = 0.98$. This specification generates an average risk-taking index of 7% (between the ages of 50 and 60) before the CGP is introduced.

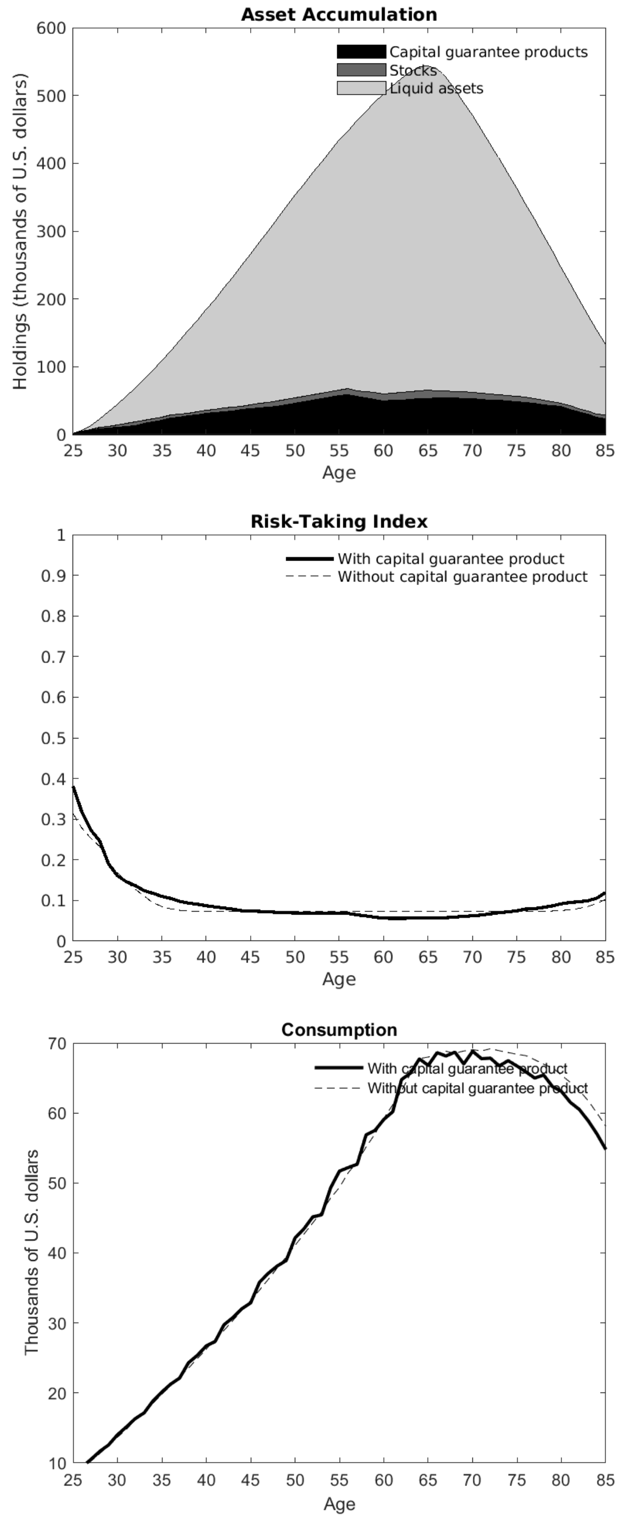


Figure IA.7. Life-Cycle Profile of an Investor with Narrow Framing and Second-Order Risk Aversion. This figure illustrates the average path of asset accumulation, the risk-taking index, and consumption of an investor with narrow-framing and second-order risk aversion. The utility parameters are $\gamma = 20$, $\psi = 0.5$, and $\delta = 0.98$. This specification generates an average risk-taking index of 7% (between the ages of 50 and 60) before the CGP is introduced.

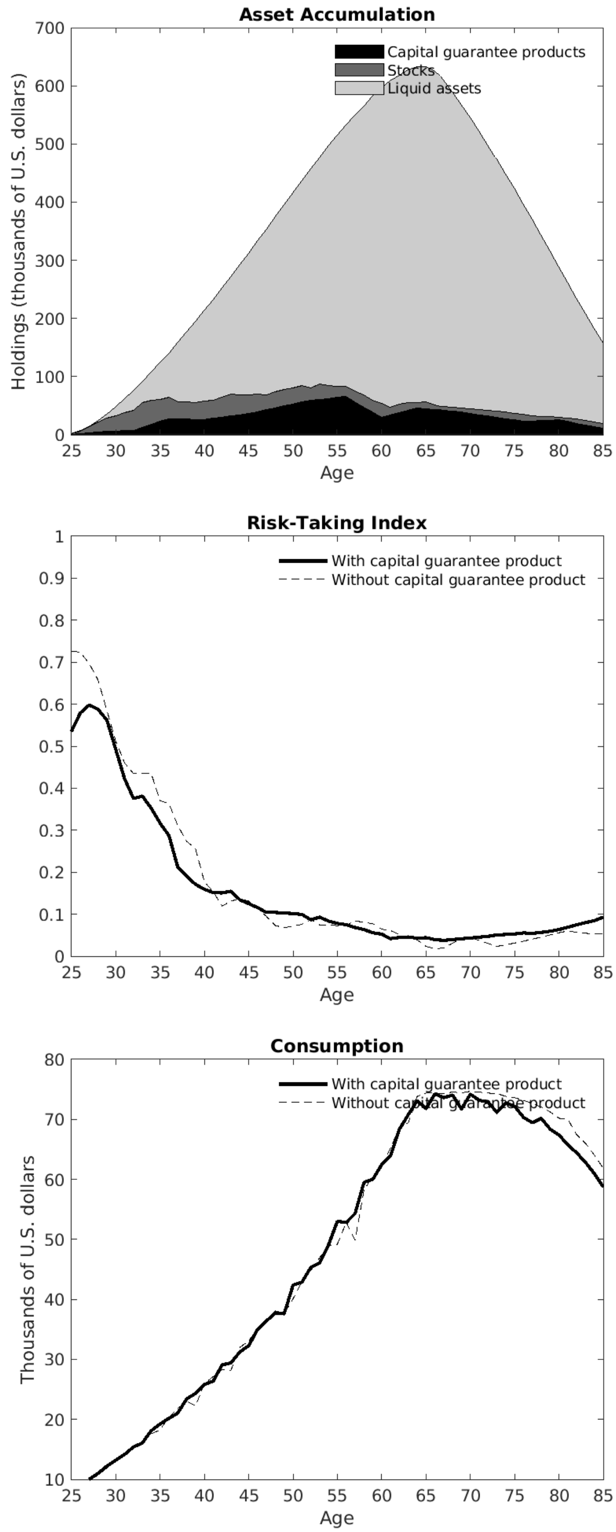


Figure IA.8. Life-Cycle Profile of an Investor with Generalized Disappointment Aversion. This figure illustrates the average path of asset accumulation, the risk-taking index, and consumption of an investor with generalized disappointment aversion. The kink parameter is $\lambda = 9.5$. This specification generates an average risk-taking index of 7% (between the ages of 50 and 60) before the CGP is introduced.

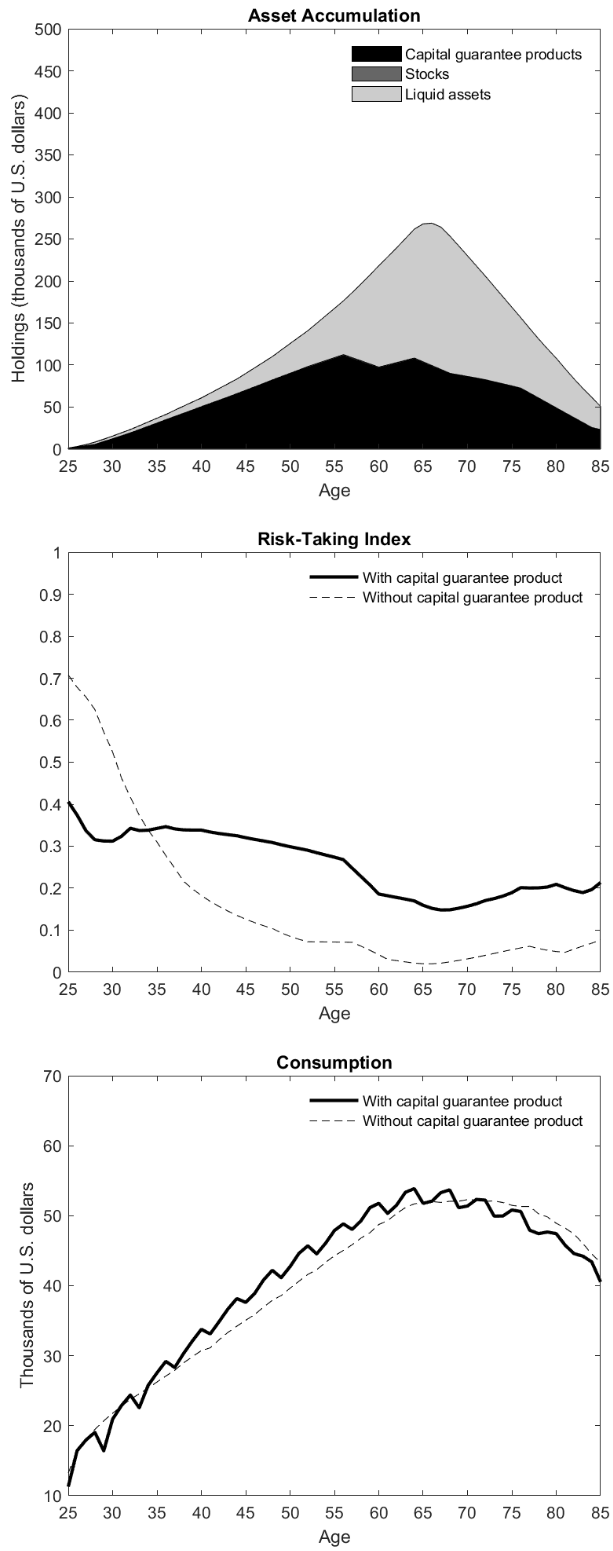


Figure IA.9. Life-Cycle Profile of an Investor with Subjective Crash Probability. This figure illustrates the average path of asset accumulation, the risk-taking index, and consumption of an investor with a subjective probability that the ex dividend log return drops by $-\ln(4)$ with probability of 51%. This specification generates an average risk-taking index of 7% (between the ages of 50 and 60) before the CGP is introduced.

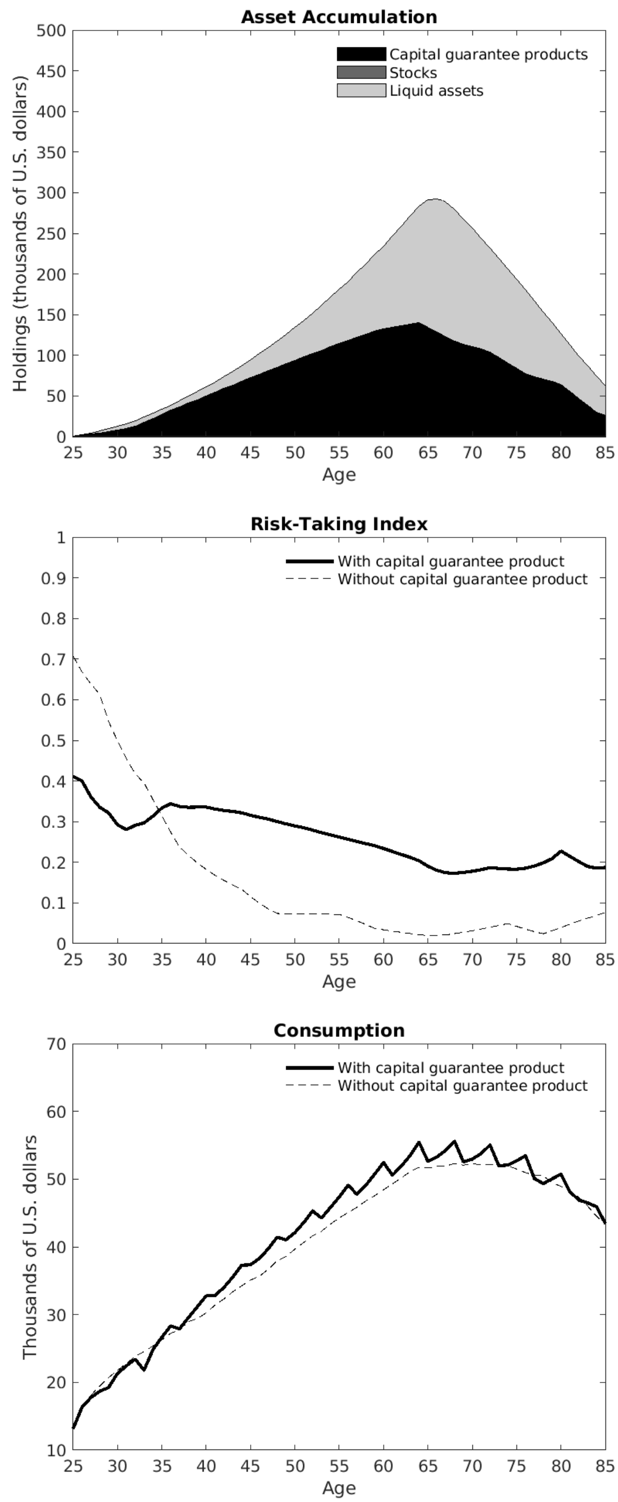


Figure IA.10. Life-Cycle Profile of an Investor with Volatility Misperception. This figure illustrates the average path of asset accumulation, the risk-taking index, and consumption of an investor with a volatility misperception parameter of 55%. This specification generates an average risk-taking index of 7% (between the ages of 50 and 60) before the CGP is introduced.

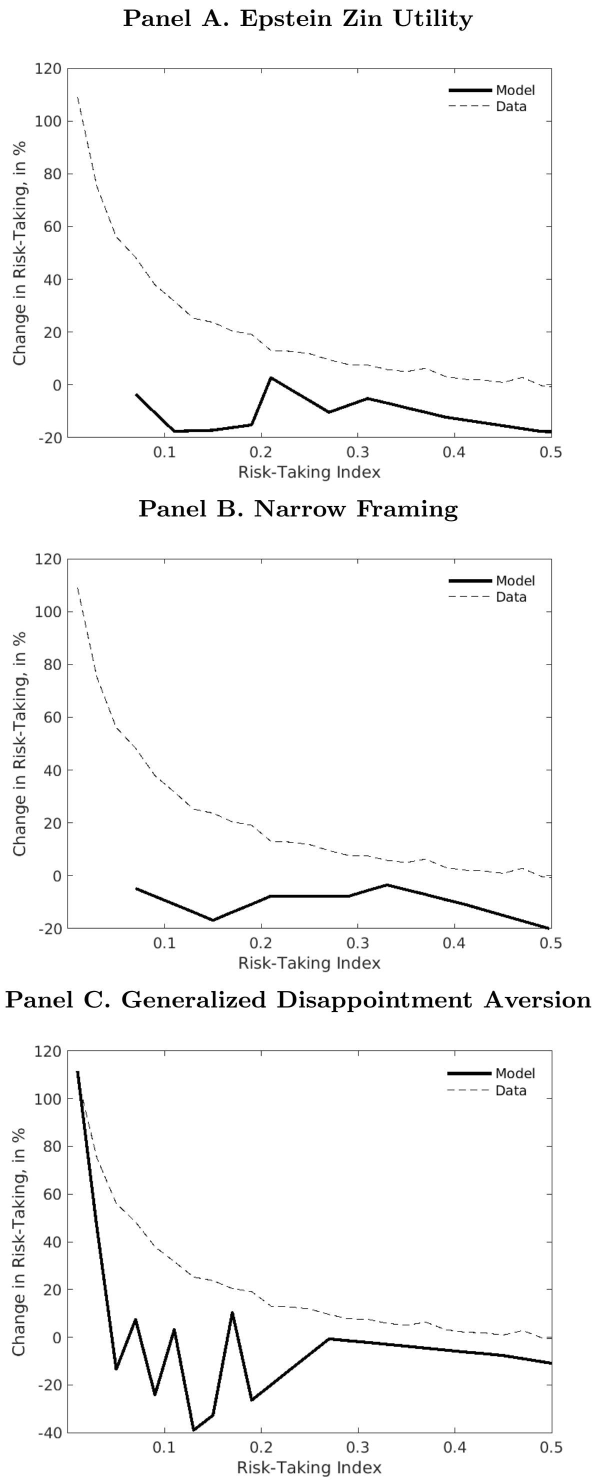
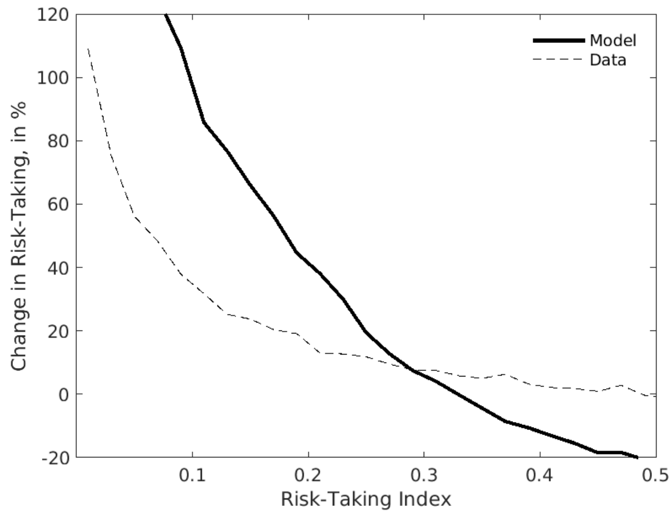


Figure IA.11. Change in Risk-Taking Index under Alternative Specifications of Preferences. This figure illustrates the relationship between initial risk-taking and the change in the risk-taking index triggered by CGPs under the life-cycle model with Epstein-Zin utility (Panel A), narrow framing (Panel B), and generalized disappointment aversion (Panel C). Each point is an average over households with a head between 50 and 60. The solid line is obtained by varying the coefficient of relative risk-aversion γ (Panel A) or the kink parameter λ (Panels B and C), while all other model parameters are kept constant.

Panel A. Volatility Misperception



Panel B. Probability of Crash

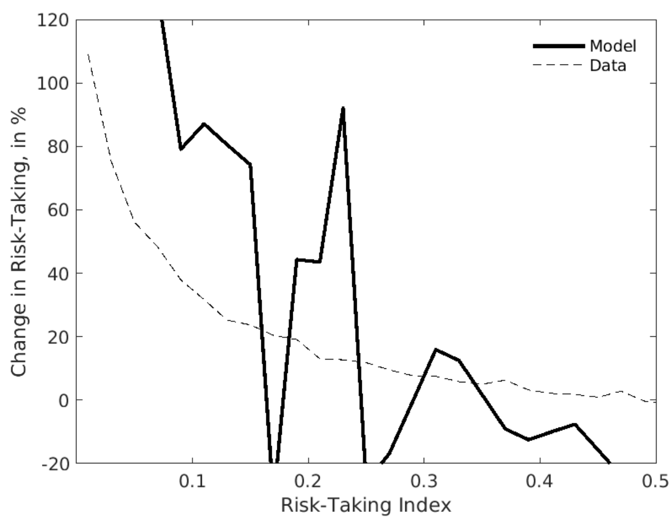


Figure IA.12. Change in Risk-Taking Index under Alternative Specifications of Beliefs. This figure illustrates the relationship between initial risk-taking and the change in the risk-taking index that follows the introduction of capital guarantee products. In each panel, the dashed line corresponds to empirical data, while the solid line plots the value implied by the life-cycle model with volatility misperception (Panel A), and probability of crash (Panel B). Each point is an average over households with a head between 50 and 60. The solid line is obtained by varying the level of volatility misperception (Panel A) and the probability of crash (Panel B), while all other model parameters are kept constant.

VIII. Welfare Analysis

Consider the utility, $V_1(X_1, K_1, CR_1, \tau_1, Y_1^P) = V_1(X_1, 0, 1, 0, Y_1^P)$ at date $t = 1$, and let $J(X_1; Y_1^P) = [V_1(X_1, 0, 1, 0, Y_1^P)]^{1-1/\psi}$. We note that

$$J_X(X_1; Y_1^P) = (1 - \psi^{-1})\,[V_1(X_1, 0, 1, 0, Y_1^P)]^{-1/\psi}\,V_{1,X}(X_1, 0, 1, 0, Y_1^P).$$

The function $J(X_1, Y_1^P)$ satisfies the Bellman equation:

$$J(X_1; Y_1^P) = \max_{\{C_1, I_1, \alpha_1\}} \left[(1-\delta)\,C_1^{1-1/\psi} + \delta\,p_1\,(\mu_2)^{1-1/\psi}\right].$$

The envelope theorem implies that

$$J_X(X_1; Y_1^P) = (1-\delta)\,(1-\psi^{-1})\,C_1^{-1/\psi}.$$

The utility V_1 therefore satisfies

$$V_{1,X} = (1-\delta)\left(\frac{C_1}{V_1}\right)^{-1/\psi}$$

A wealth transfer θ at date $t = 1$ increases utility by

$$\Delta J \approx (1-\delta)\,(1-\psi^{-1})\,C_1^{-1/\psi}\,\theta.$$

In practice, X_1 is stochastic because the labor income received in period 1 is random. A non-random transfer θ_1 increases average utility by $\Delta\mathbb{E}(J)$ if

$$\theta \approx \frac{\Delta\mathbb{E}(J)}{(1-\delta)\,(1-\psi^{-1})\,\mathbb{E}(C_1^{-1/\psi})}. \qquad \text{(IA-55)}$$

We compute the change in utility before and after innovation, $\Delta\mathbb{E}(J)$, and use (IA-55) to obtain the benefit from financial innovation, θ.

Table IA.3
Household Welfare Gains Predicted by the Models

Models	**Volatility Misperception** (1)	**Crash Risk** (2)
Key parameter value	Misperception gap=0.55	Crash probability 0.51
Change in risk-taking (%)	130.8	109.0
Household utility gain, in U.S. $	25,747	21,306
Bank revenue gain, in U.S. $	17,345	13,403
Household share of surplus (%)	59.7	61.4

Notes: This table reports the changes in the household risk-taking index, welfare gains, bank revenue gains, and the household share of the surplus generated by the introduction of CGPs under various specifications of preferences and beliefs. Under all specifications, the starting value is household with an ex-ante risk-taking index of 8%, which corresponds to the 25^{th} percentile in the Swedish population. In column 1, we consider that the investor believes that the volatility of the underlying exceeds its volatility level under $\mathbb{P}$ by 55 basis points. In column 4, we assume that the subjective probability measure is a mixture of the physical measure $\mathbb{P}$ and a crash event - the ex dividend log return drops by $-\ln(4)$ - with probability of 51%.

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