Internet Appendix for
"The Cross-Section of Household Preferences"

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Section I discusses data sources and the empirical inputs of the life-cycle model. Section II describes the estimation and procedures used in the main text. Section III reports additional empirical results. Section IV presents additional numerical results that confirm the validity of our estimation procedure. Section V provides a detailed description of the Swedish pension system.

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I. Data, Income, and Pensions

A. Data Sources and Household Variables

Statistics Sweden (SCB) has a parliamentary mandate to collect detailed data on demographics, income and wealth of all Swedish residents using tax returns and information from third parties, such as financial intermediaries, employers and welfare agencies. The demographic characteristics include age, gender, education, and marital status. SCB provides a household identifier since 1991 that allows us to define a household as a family living together with the same adults over time. We define the household head as the adult with the highest average non-financial disposable income; or, if the average income is the same, the oldest; or, if the other criteria fail, the man in the household.

The income registry is available since 1983 and contains information on employment sector and the income components necessary to construct the non-financial disposable income of each Swedish resident. We use income tax and its base to estimate the average income tax rate by education level. Information on retirement income and student allowances allows us to identify retirees and students in the population. The data do not distinguish between DC and DB pension payouts, but they provide pension-qualifying income and thus enable us to impute DC pension contributions and DB pension rights as explained in the main text and in Sections I.B and V of this appendix.

The wealth registry reports debt and disaggregated worldwide financial and real estate holdings at year-end from 1999 to 2007. Bank account balances, stock and mutual fund investments, and real estate holdings are observed at the level of each account, security, or property.[1] Even though the wealth registry does not provide the value of security holdings, it reports ISIN identification codes, which allows us to value and classify each financial asset by using FINBAS, a financial database maintained by the Swedish House of Finance. For

[1]Bank account balances are reported if the account yields more than 100 Swedish kronor during the year (1999 to 2005 period), or if the year-end bank account balance exceeds 10,000 Swedish kronor (2006 and 2007). We impute unreported cash balances by following closely the method reported in Bach, Calvet and Sodini (2020).

securities not covered by FINBAS, we use data from Citygate, Morningstar, Datastream, and the stock exchanges NGM and OMX. Positions in fixed income securities, capital guaranteed products and capital insurance are instead reported at market values. The data do not contain information on defined contribution retirement wealth, which we impute as explained in the main text and in Sections I.B and V of this appendix.

Real estate wealth includes residential real estate properties (i.e., primary and secondary residences), and commercial properties (i.e., rental, industrial, and agricultural properties) serving as business or investment vehicles. Real estate prices are compiled by Statistics Sweden from two main sources. Every 3 to 7 years, tax authorities assess the tax value of housing properties using detailed property characteristics and hedonic pricing. In addition, Statistics Sweden continuously collects data on every real estate transaction in the country, which permits the construction of sales-to-tax-value multipliers for different geographic locations and property types. The transaction data are also used to value apartments at the level of each residential building.

Debt is the sum of mortgages and all other liabilities to financial institutions.[2] Because we do not observe durable goods (such as appliances, cars and boats), the value of household debt can exceed the value of the assets we observe for some households. To avoid this problem, the debt variable is defined as the minimum of the total debt and real estate wealth reported in the registry. This approach is consistent with the fact that we proxy the borrowing rate by the average mortgage rate offered by Swedish institutions.

We explain in the main text the definition of the composite asset, which is a weighted portfolio of liquid financial wealth, real estate wealth, DC retirement wealth, and debt. In aggregate Swedish data in 1999, the shares of these four components in total net wealth are 36%, 76%, 13%, and −25%, respectively.

[2]We exclude student debt because it is exclusively provided by the state and heavily subsidized during our sample period.

B. Pensions

We explain the organization of the Swedish pension system and the imputation methodology we use in the main text. We refer the reader to Section V of this appendix for a full discussion of these topics.

B.1. The Swedish Pension System

The Swedish pension system consists of three pillars: state pensions, occupational pensions, and private pensions.

The *state pension* system requires each worker in Sweden to contribute 18.5% of their pension qualifying income: 16% to the pay-as-you-go defined benefit (DB) system and the remaining 2.5% to a defined contribution (DC) system called premiepension system. DC contributions are invested in a default fund, that mirrors the world index during our sample period, unless the worker opts out and chooses a portfolio of funds among those offered on the state DC platform. State DB payouts are a function of the pension qualifying income earned during the entire working life.

Occupational pensions were introduced to Sweden in 1991. They are regulated for the vast majority of Swedish residents by four collective agreements applying to blue-collar private-sector workers, white-collar private-sector workers, central government employees, and local government employees. Since these agreements specify workers' monthly pension contributions, the fraction directed to DB and DC pension plans, and the DC choices available to workers, we are able to impute both DC contributions and DB entitlements at the household level.

The collective agreements specify DC contributions as a percentage of pension qualifying income. These contributions are invested through insurance companies in either variable annuity products (called TradLiv in Sweden), or in portfolios of mutual funds, chosen by workers from a selection provided by the insurance company.

Defined contribution *private pensions* have existed in Sweden for a long time but our dataset provides us with individual private pension contributions from 1991. We assume that they are invested like occupational and state DC contributions. Section V of this appendix provides a comprehensive description of the three pillars of the Swedish pension system.

B.2. Summary of Imputation Methodology

We now explain how pension qualifying income can be used to impute DB pension payouts and DC contribution. We also describe how we capitalize private pension savings and DC retirement wealth.

State DB payouts are a function of the pension qualifying income earned during the entire working life. Since our individual income data begin in 1983, we cannot observe the full income history for older individuals in our dataset. To handle this, we back-cast their income back to age 25 by using real per-capita GDP growth and inflation before 1983. We then use the state DB payout rules to impute state DB pension payments for each individual retiring during our sample period.

Occupational DB pension payouts can be accurately imputed because in all the collective agreements they are a function of at most the last 7 years of pension qualifying income during working life, and data on pension qualifying income are available from 1991 (the year occupational pensions were introduced).

Defined contribution *private pensions* have existed in Sweden for a long time but our dataset provides us with individual private pension contributions from 1991. More specifically, exact information on private pension contributions is available from 1994, whereas from 1991 the data reports only a capped version. We impute full contributions from 1991 to 1993 taking into account both age effects and individual savings propensities in subsequent years. We assume that these contributions are invested in the same way as occupational and state DC contributions. We follow Bach, Calvet and Sodini (2020) and allocate 58% of the aggregate stock of private pension wealth in 1991 to workers.[3] Across workers, we allocate pension wealth proportionately to their private pension contributions in 1991.

To calculate DC retirement wealth at each point in time, we accumulate contributions from all three pillars of the Swedish pension system. To do this for the state DC system, we follow the investment policy and cost of the system's default fund and assume that equity contributions are invested in the MSCI equity world index, without currency hedging, and are subject to a fee of 15 basis points. The equity share in the state DC system mirrors the allocation rules of the system's default fund: a 130% levered position in the world index up to the age of 55, which is then gradually rebalanced with age to an increasingly conservative portfolio. For occupational and private DC pensions, we assume that equity contributions are invested in the unhedged MSCI equity world index, subject to the 70 basis point fee that prevailed during our sample period. This assumption reflects the high degree of international diversification observed in Swedish equity investments (Calvet, Campbell, and Sodini 2007). The equity share in each household's occupational and private DC retirement portfolio is rebalanced with age following the representative age pattern of life-cycle funds available in Sweden during our sample period. We assume that all DC wealth not invested in equities is invested in cash. We refer the reader to Section V for a full description of the imputation methodology.

C. Labor Income

Methodology. We estimate the income process from consecutive observations of household yearly income data over the period 1992 to 2007,[4] excluding the first and last year of labor income to avoid measuring annual income earned over less than 12 months.[5] We consider

[3]This allocation is chosen to satisfy the condition that imputed pension wealth should be roughly the same just before and just after retirement.

[4]Since our individual income data begin in 1983, we cannot observe the full income history for older individuals in our dataset. To handle this, we back-cast their income back to the age of 25 by using real per-capita GDP growth and inflation before 1983. We then use the state DB payout rules to impute state DB pension payments for each individual retiring during our sample period.

[5]In each year, we winsorize non-financial real disposable income to a minimum level of 1000 kronor or about $150. We also winsorize the pooled data from above at the 0.01% level to take care of extreme outliers

the total income received by all members of the household, but classify households by the head's education level and age. Since the vast majority of Swedish residents retire at 65, we consider two age groups: (i) non-retired households older than 19 and less than 65, and (ii) retired households that are at least 65.

For active households younger than 65, we estimate the coefficients a_c and b of the labor income process, $\log(Y_{h,t}) = a_c + b'x_{h,t} + \nu_{h,t} + \varepsilon_{h,t}$, by running pooled regressions for each of the three education levels. The vector of explanatory variables $x_{h,t}$ includes age dummies, which we then regress on a third-degree polynomial in age and use the fitted third-degree polynomial in our life-cycle model. By construction, the residual, $y_{h,t} = \log(Y_{h,t}) - a_c - b'x_{h,t}$, satisfies

$$y_{h,t} = \nu_{h,t} + \varepsilon_{h,t} = \xi_t + z_{h,t} + \varepsilon_{h,t}.$$

We use the sample mean, $\bar{y}_t = \sum_h y_{h,t}/N$, as an estimate of the permanent aggregate component ξ_t. We estimate the variance of permanent aggregate shocks, σ_u^2, by the sample variance of $\bar{y}_t - \bar{y}_{t-1}$ $(t = 2, \ldots, T)$. Let $y^*_{h,t} = y_{h,t} - \bar{y}_t$ denote the idiosyncratic component of income. We estimate the variance of the permanent and transitory idiosyncratic labor income shocks, σ_w^2 and σ_e^2, as in Carroll and Samwick (1997).

For retired households, we impute the state and occupational after-tax pension benefit of each individual from 1999 to 2007, as explained in section I.B of this appendix. We fill forward the imputed pension benefit in real terms until 2007 at individual level, and aggregate income at the household level in each year. The replacement ratio is estimated for each education group as the fraction of the average income of non-retired 64-year-old households to the average income of retired 65-year-old households across the 1999 to 2007 period.

Results. Table IA.1 presents the size of education and income risk categories and Table IA.2 reports the employment sectors of education and income risk categories. The pat-

at the top of the income distribution.

terns are intuitive, with relatively little income risk in the public sector and in mining and quarrying, electricity, gas, and water supply, and relatively high income risk in hotels and restaurants, real estate activities, construction for less educated workers, and the financial sector for more educated workers.

Table IA.3 shows that educated households, particularly those with higher education, face higher transitory income risk and lower idiosyncratic permanent income risk than less educated households. This result is a likely consequence of the following features of the Swedish labor market. First, it is straightforward for companies to downsize divisions, but extremely difficult for them to lay off single individuals unless they have a high managerial position. Second, companies that need to downsize typically restructure their organizations by bargaining with unions. Third, unions are nationwide organizations that span large areas of employment and pay generous unemployment benefits. Fourth, the pay cut due to unemployment is larger for better paid jobs. After an initial grace period, an unemployed person will be required to enter a retraining program or will be assigned a low-paying job by a state agency. All these features imply that unemployment is slightly more likely and entails a more severe proportional income loss for workers with higher levels of education.

The correlation between permanent income shocks and wealth shocks is a key ingredient of our model. In the main text, we report that the average value of this correlation across the nine education-sector categories. Table IA.4 reports these correlations for each of the nine categories and for three types of wealth. Risky liquid financial wealth is only weakly correlated to permanent income shocks, except for educated households employed in high-risk sectors. Real estate wealth and non-cash net wealth exhibit substantially stronger correlations, which are even more pronounced for educated households in high-risk employment sectors. As we noted in the main text, the correlation between the numeraire risky asset and individual income growth is much smaller because most individual income risk is idiosyncratic.

II. Estimation and Testing Procedure: Theory

A. Estimation Procedure

The indirect inference procedure explained in the main text is implemented as follows. For each group g, we simulate the lifecycle model on a grid of preference parameters, as Section II.B of this appendix explains.

The grid is defined by 12 values of the RRA ranging from 3 to 12, 11 values of the TPR ranging from -0.05 to 0.22, and 14 values of the EIS ranging from 0.1 to 2.5. The grid values of the RRA are 3, 4, 5, 5.5, 6, 6.5, 7, 7.5, 8, 9, 10 and 12. To construct the grid values of the TPR, we assume that the patience parameter $\delta \in \{0.80, 0.85, 0.90, 0.92, 0.94, 0.96, 0.97, 0.98, 0.99, 1.00, 1.05\}$, so that TPR$= -\ln(\delta)$ is contained in $\{-0.05, 0, 0.01, 0.02, 0.03, 0.04, 0.06, 0.08, 0.11, 0.16, 0.22\}$. The grid values of the EIS are 0.1, 0.2, 0.3, 0.4, 0.5, 0.6, 0.7, 0.8, 1.2, 1.4, 1.6, 1.8, 2, and 2.5. Overall, the grid contains 1,848 ($= 12 \times 11 \times 14$) parameter vectors θ.

For each vector θ on the grid, we calculate the value of the objective function:

$$[\tilde{\mu}_S^g(\theta) - \hat{\mu}^g]' \ \Omega \ [\tilde{\mu}_S^g(\theta) - \hat{\mu}^g] \,,$$

where Ω is the weighting matrix defined in the main text and S is the number of Monte Carlo simulations used to compute $\tilde{\mu}_S^g(\theta)$.[6]

We next evaluate the objective function on a finer grid defined as follows. The RRA grid has a grid step of 0.01 and contains 81 equally-spaced grid points ranging from 2 to 10. The EIS grid has grid step of 0.01 and contains 241 values of the EIS ranging from 0.1 to 2.5. We consider an evenly spaced grid of the patience parameter δ ranging from 0.8 to 1.05 with a grid step of 0.001, which generates a TPR grid containing 251 values of the TPR ranging from -0.05 to 0.22. The finer grid therefore contain 4,899,771 ($= 81 \times 241 \times 251$) preference parameters. We evaluate the objective function on the finer grid by interpolating the values

[6]See Section II.B of this appendix for further details.

computed on the original grid. We use modified Akima cubic Hermite interpolation, which is known to reduce interpolation overshoots and oscillations compared to standard spline methods.

For each group, we obtain the indirect inference estimator by determining the parameter vector on the finer grid that minimizes the objective function. This value may occasionally be slightly negative due to interpolation error.

For inference purposes, such as calculations of the root mean squared error or the Jacobian matrix, we compute each of the 16 auxiliary statistics by a separate interpolation. These interpolations are based on a cubic spline using not-a-knot end conditions.

B. *Simulation Details*

For a given group g and a given preference parameter θ, we compute the estimator of the binding function, $\tilde{\mu}_S^g(\theta)$, by Monte Carlo simulations. The period-t simulations take as given the group's wealth-income ratio at the end of year $t-1$, the realized group-level income shock, and the numeraire asset return during the year, which are all set equal to their respective empirical values. The simulations then proceed in four steps.

i. We simulate $S = 10,000$ households/paths in the group over year t. For each simulated unit $i \in \{1..., S\}$, we simulate idiosyncratic (permanent and transitory) labor income shocks and thereby obtain labor income, $\tilde{Y}_{i,t}$, and permanent income, $\tilde{Y}_{i,t}^P$, in period t. We set wealth at the beginning of period t, $\tilde{W}_{i,t-1}$, equal to $\tilde{Y}_{i,t}^P$ times the group's average wealth-income ratio at the end of period $t-1$. Using the lifecycle model's policy functions $\alpha_t^*(\,\cdot\,)$ and $C_t^*(\,\cdot\,)$, we compute the risky share, $\tilde{\alpha}_{i,t-1} = \alpha_t^*(\tilde{Y}_{i,t}, \tilde{W}_{i,t-1}, \tilde{Y}_{i,t}^P; \theta)$, and consumption, $\tilde{C}_{i,t} = C_t^*(\tilde{Y}_{i,t}, \tilde{W}_{i,t-1}, \tilde{Y}_{i,t}^P; \theta)$, of each simulated unit during year t. Consistent with the model, the simulated unit sets these quantities at the end of year $t-1$ and keeps them constant during year t.[7]

[7] This methodology exploits the homogeneity of $\alpha_t^*(\,\cdot\,)$ and $C_t^*(\,\cdot\,)$, with respect to $(\tilde{Y}_{i,t}, \tilde{W}_{i,t-1}, \tilde{Y}_{i,t}^P)$.

ii. We compute the predicted wealth of each simulated unit at the end of year t :

$$\hat{W}_{i,t} = (R_f + \tilde{\alpha}_{i,t-1} R^e_{N,t})(\tilde{W}_{i,t-1} + \tilde{Y}_{i,t} - \tilde{C}_{i,t}). \tag{IA.1}$$

The prediction incorporates the empirical return on the numeraire in year t.

iii. We obtain the group's predicted wealth-income ratio at the end of year t:

$$\widetilde{\mu}^g_{1,t}(\theta) = \frac{\sum_{i=1}^S \hat{W}_{i,t}}{\sum_{i=1}^S \tilde{Y}_{i,t}}. \tag{IA.2}$$

iv. We observe the information set available at the end of year t, we sample S households, and we compute the group's predicted risky share at the end of year t:

$$\widetilde{\mu}^g_{2,t}(\theta) = \frac{\sum_{i=1}^S \tilde{\alpha}_{i,t} \tilde{W}_{i,t}}{\sum_{i=1}^S \tilde{W}_{i,t}}. \tag{IA.3}$$

We stack the resulting values into the column vector $\tilde{\mu}^g_S(\theta)$.

C. *Asymptotic Properties of Our Estimator*

If our model is correctly specified, the indirect inference estimator $\hat{\theta}^g$ converges to the true preference parameter vector as the number of households in each group increases, as we now explain.

The empirical auxiliary estimator $\hat{\mu}^g$ is asymptotically normal:

$$\sqrt{N^g}\,[\hat{\mu}^g - \mu^g(\theta)] \xrightarrow{d} \mathcal{N}(0, W_g) \tag{IA.4}$$

as the group size N^g goes to infinity. This result follows from the delta method and the fact that the auxiliary statistics (defined in equations (15) and (16) of the main text) can be interpreted as ratios of sample moments. We estimate the asymptotic variance covariance matrix of $\hat{\mu}^g$ by the jackknife estimator

$$\frac{\hat{W}_g}{N^g} = \frac{N^g - 1}{N^g} \sum_{i=1}^{N^g} (\hat{\mu}^J_{[i]} - \overline{\mu^J})(\hat{\mu}^J_{[i]} - \overline{\mu^J})', \tag{IA.5}$$

where $\hat{\mu}^J_{[i]}$ is the auxiliary estimator obtained by excluding the i^{th} household, and $\overline{\mu^J} = (N^g)^{-1} \sum \hat{\mu}^J_{[i]}$.

The indirect inference estimator is asymptotically normal:

$$\sqrt{N^g} \left(\hat{\theta}^g - \theta^g \right) \xrightarrow{d} \mathcal{N}(0, V^g), \tag{IA.6}$$

as Gouriéroux, Monfort, and Renault (1993) show. Furthermore, the asymptotic variance-covariance matrix is given by

$$V^g = (1 + s_g^{-1})\,(D_g\,\Omega\,D_g')^{-1}\,D_g\,\Omega\,W_g\,\Omega\,D_g'(D_g\,\Omega\,D_g')^{-1}, \tag{IA.7}$$

where the ratio $s_g = S/N_g$ accounts for simulation noise and $(D_g)' = \partial \mu^g(\theta^g)/\partial \theta'$ is the Jacobian matrix of the binding function $\mu^g(\ \cdot\ )$ evaluated at the true parameter θ^g. In practice, we estimate the asymptotic variance-covariance matrix of V_g by its sample equivalent $\hat{V}^g = (1 + s_g^{-1})\,(\hat{D}_g\,\Omega\,\hat{D}_g')^{-1}\,\hat{D}_g\,\Omega\,\hat{W}_g\,\Omega\,\hat{D}_g'(\hat{D}_g\,\Omega\,\hat{D}_g')^{-1}$, where $\hat{D}_g$ is a finite-difference approximation of D_g.

When the size of each group g is large, we could achieve efficient estimation by setting the second-stage weighting matrix equal to the inverse of the jackknife estimator: $\Omega^{(2)} = \hat{W}_g^{-1}$, and then solving the optimization problem defined in equation (17). of the main text Efficient estimation, however, is not feasible in our sample because most groups are too small to obtain a reliable estimator of W_g^{-1}. The median group size is 63, while the symmetric matrix W_g contains 136 ($= 16 \times 17/2$) distinct elements. A related problem is that in many groups, the weighting matrix $\Omega^{(2)} = \hat{W}_g^{-1}$ assigns almost all the weight to the risky share, while the wealth-income ratio plays essentially no role in estimation. Efficient estimation is therefore unsatisfactory in our sample on statistical and economic grounds.[8] For these reasons, we henceforth focus on one-step estimation based on the diagonal weighting matrix Ω defined in the main text. Since this approach does not provide global specification tests based on the value of the objective function (17), we focus on measures of fit based on root mean squared error or economic significance.

D. *Hypothesis Testing*

We now explain the methodology used to test hypotheses about preference parameters in Table 7 of the main text.

D.1. Group-Specific Tests

We consider a null hypotheses of the form $\mathbf{H_0} : R(\theta^g) = 0$, where $R(\theta^g)$ is a column vector function of dimension r. Since $\sqrt{N^g}(\hat{\theta}^g - \theta^g) \xrightarrow{d} \mathcal{N}(0, V^g)$, the delta method implies that

$$\sqrt{N^g} R(\hat{\theta}^g) \xrightarrow{d} \mathcal{N}\left[0, \frac{\partial R}{\partial \theta'}(\theta^g)\ V^g\ \frac{\partial R'}{\partial \theta}(\theta^g)\right]$$

under the null. In Table 7 of the main text, we report the results of χ^2 tests based on

$$N^g\ R(\hat{\theta}^g)'\ \left[\frac{\partial R}{\partial \theta'}(\hat{\theta}^g)\ \hat{V}^g\ \frac{\partial R'}{\partial \theta}(\hat{\theta}^g)\right]^{-1} R(\hat{\theta}^g) \xrightarrow{d} \chi^2(r), \tag{IA.8}$$

where $\hat{V}^g$ is defined in Section II.C of this appendix.

Expected utility. Following this methodology, we test in each group g the null hypothesis that households in the group exhibit expect utility:

$$\mathbf{H_0} : \psi^g = 1/\gamma^g.$$

[8]These difficulties are consistent with the finite-sample inaccuracy of two-step generalized method of moments studied in Hwang and Sun (2018).

The restriction function is $R(\theta^g) = \gamma^g \psi^g - 1$, where $\theta^g = (\delta^g, \gamma^g, \psi^g)'$. The Jacobian matrix of R is the 1×3 vector

$$\frac{\partial R}{\partial \theta'}(\theta^g) = (0, \psi^g, \gamma^g)$$

under the null hypothesis. The corresponding χ^2 tests are reported in Table 7 of the main text.

Time Preference Rate. Since $TPR^g = -\ln(\delta^g)$, the delta method implies that the variance of TPR is the variance of beta divided by the squared of beta: $Var(\widehat{TPR^g}) \approx Var(\hat{\delta}^g)/(\hat{\delta}^g)^2$. This result allows us to test hypotheses on the time preference rate reported in Table 7 of the main text.

D.2. Tests Involving All Groups

In Table 7 of the main text, we also test restrictions involving the preference parameters of all groups. For instance, we assess if the vector of preference parameters, θ^g, or each of its components, are homogeneous across groups. These tests are conducted as follows.

Let $N = N^1 + ... + N^G$ denote the total number of observations, and let

$$k^g = N^g/N$$

denote the fraction of observations in group g. We stack the group-level parameters into $\theta = [(\theta^1)', \ldots, (\theta^G)']'$ and $\hat{\theta} = [(\hat{\theta}^1)', \ldots, (\hat{\theta}^G)']'$. Since $\sqrt{N^g}(\hat{\theta}^g - \theta^g) \xrightarrow{d} \mathcal{N}(0, V^g)$, we infer that $\sqrt{N}(\hat{\theta}^g - \theta^g) \xrightarrow{d} \mathcal{N}(0, V^g/k^g)$ and therefore

$$\sqrt{N}(\hat{\theta} - \theta) \xrightarrow{d} \mathcal{N}(0, V),$$

where $V = \text{diag}(V^1/k^1, \ldots, V^G/k^G)$.[9] A finite-sample estimator of the asymptotic variance-

[9] We denote by $\text{diag}(A^1, \ldots, A^n)$ the block diagonal matrix with diagonal blocks $A^1, \ldots, A^n$.

covariance matrix V is given by

$$\hat{V} = \mathrm{diag}(\hat{V}^1/k^1, \ldots, \hat{V}^G/k^G),$$

where $\hat{V}^1, \ldots, \hat{V}^G$ are defined in Section II.C of this appendix.

We consider the null hypothesis

$$\mathbf{H_0} : R(\theta) = 0, \tag{IA.9}$$

where $R(\cdot)$ is a column vector function of dimension r. Under the null, the delta method implies that

$$\sqrt{N} R(\hat{\theta}) \xrightarrow{d} \mathcal{N}\left[0, \frac{\partial R}{\partial \theta'}(\theta)\ V\ \frac{\partial R'}{\partial \theta}(\theta)\right] \tag{IA.10}$$

and

$$N\ R(\hat{\theta})'\ \left[\frac{\partial R}{\partial \theta'}(\hat{\theta})\ \hat{V}\ \frac{\partial R'}{\partial \theta}(\hat{\theta})\right]^{-1}\ R(\hat{\theta}) \xrightarrow{d} \chi^2(r). \tag{IA.11}$$

The result holds for fixed proportions $k^1, \ldots, k^G$, and for a total number of observations N going to infinity.

Equality of Group-Level Parameters to the Cross-Sectional Mean. Following this methodology, we test the null hypothesis:

$$\mathbf{H_0} : \theta^g = \sum_{i=1}^{G} k^i\ \theta^i,$$

where θ^i and k^i denote the parameter vector and relative size of each group i. The restriction function is

$$R^g(\theta) = \theta^g - \sum_{i=1}^{G} k^i\ \theta^i.$$

Let I_3 denote the identity matrix of size 3, let e_g denote the row vector of dimension G with g^{th} element equal to unity and other elements equal to 0, and let $k = (k^1, \ldots, k^G)$. We note that $\theta^g = (e_g \otimes I_3)\theta$, and $\sum_{i=1}^{G} k^i \theta^i = (k \otimes I_3)\theta$. The restriction function can therefore be rewritten in matrix form as

$$R^g(\theta) = A^g \ \theta,$$

where $A^g = (e_g - k) \otimes I_3$ for every g. The corresponding tests, based on equations (IA.10) and (IA.11) of this appendix, are reported in Table 7 of the main text.

III. Additional Empirical Results

A. Distribution of Preference Parameters

Figure IA.1 plots the univariate distribution of relative risk aversion, the time preference rate and the elasticity of intertemporal substitution across the Swedish households.

Figure IA.2 plots heats maps for estimates of RRA and its standard error (top panel), the TPR and its standard error (middle panel), and the EIS and its standard error (bottom panel) across Swedish households. The figure reveals that the asymptotic standard error of the EIS is positively correlated with the level of the estimated EIS.

B. Relation Between Preference Parameters and Characteristics

The lower portion of Table 5 in the main text explores correlation patterns among preference parameters and observables. Tables IA.5 of this internet appendix report multiple regressions rather than univariate correlations. Most patterns are similar, but controlling for the initial wealth-income ratio, the growth of wealth-income predicts the EIS positively rather than negatively.

The results in the main text are weighted by the number of households in each group. While this is the natural weighting scheme in household finance applications, asset pricing economists may be interested in wealth-weighted average preference parameters of households. In appendix Table IA.6, we weight groups by their average wealth during the sample period rather than by their size. Compared to equally-weighted averages, we find a similar mean risk aversion of 7.14, much lower mean time preference rate of 2.63%, and a somewhat higher mean EIS of 1.19. The cross-sectional standard deviations of these parameters are similar to the equally weighted case.

C. *Dispersion of Parameters within Wealth Quintiles*

In Table IA.7, we present the standard deviation of preference parameters within wealth quintiles. Our calculation of the standard deviation assigns equal weights to households. The table reveals that heteogeneity in preference parameters is more pronounced in bottom quintiles than in top quintiles. This empirical regularity appears most strongly for the TPR, whose standard deviation declines from 7.6% in the two lowest quintiles to 2.8% in the top quintile. This relationship is also apparent for other parameters. The standard deviation of the RRA declines from 1.21 in the two lowest quintiles to 0.89 in the top quintile, while the standard deviation of the EIS decreases from 0.965 to 0.90 between the bottom and the top quintiles.

D. *Fitted and Observed Life Cycle Profiles*

Figure IA.3 illustrates the life-cycle profiles observed in the data and in the model as a function of age. Panel A reports the average risky share and Panel B reports the wealth-income ratio. The plots are computed by averaging across all 4276 groups of Swedish households, where each group is weighted by its wealth share. The figure shows that on average, the model fits well the patterns of portfolio age and wealth accumulation over the life-cycle.

E. *Alternative Treatment of Real Estate*

In the main text, we assume that real estate earns the FASTPI index return net of a 22% real estate capital gain tax. We now consider how our results are modified when real estate is treated as riskless. This change affects both the calculation of the risky share and the formation of household groups, as defined in Section1.3 of the main text.

In Table IA.8, we report the resulting wealth-income ratio and risky share for groups of households sorted by education and income risk, as defined in the main text. The cross-sectional mean of the risky share is lower than in Table 1 of the main text, as one expects.

The cross-sectional standard deviation of the risk-share and the wealth-income statistics are close to the values obtained in the main text.

In Table IA.9, we report panel regressions of the wealth-income ratio (column 1) and the risky share (columns 2 and 3) on group characteristics. Compared to Table 2 in the main text, the coefficient of the risky share on income volatility gains significance in column 2 and loses significance in column 3, while keeping the same signs. Post high school education loses significance in column 2 and gains significance in column 3. The results are otherwise remarkably similar to the ones obtained in Table 2 of the main text.

F. *Model Fit under Homogeneous Preferences*

Table IA.10 shows how model fit deteriorates under homogeneous preferences. The mean RMSE-scaled objective function more than doubles to 16.0% if we fix RRA at its cross-sectional mean. Fixing TPR at its cross-sectional mean produces a mean RMSE-scaled objective function of 8.6%, and similarly restricting the EIS delivers a mean RMSE-scaled objective function of 7.7%. Fixing all parameters at their cross-sectional means is disastrous in the sense that it increases the mean RMSE-scaled objective function to 24.8%. A life-cycle model with homogeneous preferences, under our maintained assumption of homogeneous rational beliefs, delivers an extremely poor fit to the cross-section of household behavior.

G. *Heterogeneous Beliefs*

In Tables IA.11-IA.13, we consider a simple form of heterogeneity in beliefs by considering three alternative assumptions about the Sharpe ratio: the base value of 0.26, a high value of 0.40, and a low value of 0.15. Then, for each group we pick the Sharpe ratio and preference parameters that minimize the objective function. The base case Sharpe ratio is selected for groups representing 54% of households, while the low Sharpe ratio and the high Sharpe ratio are each selected for 23% of households.

In Table IA.11, we report the resulting size-weighted preference parameter estimates.

Allowing for heterogeneity in household beliefs has only a modest impact on the average preference parameters we estimate. Mean RRA is now 7.80, the mean TPR is 4.72%, and the mean EIS is 1.01. The cross-sectional standard deviations of the TPR and the EIS are similar to those we estimate in the homogeneous-beliefs case, but the cross-sectional standard deviation of risk aversion is over twice as large at 2.74. The explanation is that the model uses heterogeneous beliefs to better fit wealth accumulation, and offsets belief heterogeneity with RRA heterogeneity to avoid counterfactual heterogeneity in the risky share.

Table IA.12 reports the cross-sectional correlations of preference parameters. and observabke characteristics. As in Table 5 of the main text, preference parameters exhibit only weak cross-sectional correlations. The EIS and the TPR have very similar correlations to observable characteristics as in Table 5, while the RRA coefficient is less correlated to wealth variables under heterogeneous beliefs.

Heterogeneous beliefs necessarily improve the fit of our model by adding free parameters, but the degree of improvement is modest. Table IA.13 shows that the mean RMSE-scaled objective function declines only from 7.03 in the homogeneous-beliefs case to 6.52 in the heterogeneous-beliefs case. Importantly, the mean RMSE-scaled objective function is a disastrous 21.44 when we combine heterogeneous beliefs with homogeneous preferences.

H. Time Series of Wealth-Weighted Average Preference Parameters

In Table IA.15, we report the wealth-weighted average preference parameters in each year of our panel, as well as their time-series mean and standard deviation. The preference parameters are based on the baseline estimation method. The wealth of each household is measured at the end of each year, so that time variation in mean preference parameters stems only from time variation in the distribution of household wealth. The mean risk aversion, mean TPR and mean EIS are remarkably stable over time. The time-series standard deviation of the wealth-weighted mean is 0.03 for risk aversion (time series mean = 7.14), 0.13% for the TPR (time series mean = 2.63%), and 0.01 for the EIS (time series mean =

1.19).

IV. Estimation Procedure: Additional Results

A. *Monte Carlo Simulations*

We evaluate the finite-sample performance of our procedure by a simple Monte Carlo exercise. For each group in our sample, we simulate our model under the group's initial conditions and the preference parameters we estimated for the group. We combine simulated households into hypothetical groups each containing N_g^* households, where N_g^* is a measure of the effective empirical group size. We repeat this procedure to obtain 1,000 hypothetical groups and calculate the mean parameter estimate. A comparison of this mean with the preference parameters under which the model was simulated allows us to assess finite-sample bias in our estimation method.

This Monte Carlo analysis does not fully capture the heterogeneity in household-level data, even under the assumption that our model holds without error at the household level and that all households in each group have identical preferences. This is because we simulate each household in the group assuming that the household has the group average wealth-income ratio at the start of the period. In the data, by contrast, and in the ergodic distribution of wealth-income ratios implied by the model, different households have different income and wealth levels at each point of time reflecting the influence of past idiosyncratic income shocks. Hence, the group average wealth-income ratio is more strongly influenced by those households with higher wealth. To partially capture this effect, we adjust our simulations to set the effective group size N_g^* equal to the reciprocal of the sum of squared wealth shares of individual households in the group, rather than the number of households in the group N_g. We find that N_g^* is on average about 3/4 of N_g, with relatively little variation in this ratio across groups.

In Table IA.14 we report regression coefficients of Monte Carlo mean parameter estimates on the parameter estimates that were used to generate the simulated data ("true" parameters for the purpose of this exercise). The results are very good for RRA, which has a slope coefficient of 1.005, insignificantly different from one, and an R^2 statistic of 94%. The regression for TPR has a slope coefficient of 0.917 and an R^2 statistic of 90%. Results are not quite as good for EIS, which has a slope coefficient of 0.651 and an R^2 statistic of 64%. This regression places most of its weight on the high EIS estimates, which are noisy; but results are similar for the log of the EIS. An important lesson of these results is that small-sample bias cannot explain the substantial cross-sectional heterogeneity in our preference parameter estimates. There is almost no small-sample bias for RRA, and minimal bias for the TPR; and while there is some bias in our EIS estimates, a bias correction would have little effect on the cross-sectional dispersion of the EIS.

B. *Measuring Preference Parameter Heterogeneity*

B.1. *Theory*

We denote by $\mathbb{E}$ the expectation operator computed across random realization of income shocks and by E^* the cross-sectional expectation operator. The mean preference parameter vector in the population is $\mu_\theta = E^*(\theta) = \sum_{g=1}^{G} k^g \ \theta^g$, where k^g is the population share of group g and θ^g is the preference vector of households in the group. The cross-sectional variance-covariance matrix of the preference vector is $V_\theta = E^*[(\theta - \mu_\theta)(\theta - \mu_\theta)']$. Our objective is to estimate V_θ from the group-level indirect inference estimators $\hat{\theta}^g$ $(g = 1, \ldots, G)$ defined in the main text, controlling for estimation error.

The estimation of V_θ requires us to take account of estimation noise. We let

$$\begin{aligned}(\hat{\theta}^g - \mu_\theta)(\hat{\theta}^g - \mu_\theta)' &= (\hat{\theta}^g - \theta^g)(\hat{\theta}^g - \theta^g)' + (\theta^g - \mu_\theta)(\theta^g - \mu_\theta)' \\ &\quad + (\hat{\theta}^g - \theta^g)(\theta^g - \mu_\theta)' + (\theta^g - \mu_\theta)(\hat{\theta}^g - \theta^g)'.\end{aligned}$$

We take the expectations across realization of income shocks:

$$\begin{aligned}\mathbb{E}\left[(\hat{\theta}^g - \mu_\theta)(\hat{\theta}^g - \mu_\theta)'\right] &= \mathbb{E}\left[(\hat{\theta}^g - \theta^g)(\hat{\theta}^g - \theta^g)'\right] + (\theta^g - \mu_\theta)(\theta^g - \mu_\theta)' \\ &\quad + \mathbb{E}(\hat{\theta}^g - \theta^g)\ \ (\theta^g - \mu_\theta)' + (\theta^g - \mu_\theta)\ \ \mathbb{E}(\hat{\theta}^g - \theta^g)'.\end{aligned}$$

We apply the cross-sectional expectation operator and obtain:

$$\Psi^{(1)} = \Psi^{(2)} + V_\theta + \Psi^{(3)} + [\Psi^{(3)}]', \qquad \text{(IA.12)}$$

where $\Psi^{(1)} = E^*\mathbb{E}\left[(\hat{\theta}^g - \mu_\theta)(\hat{\theta}^g - \mu_\theta)'\right]$, $\Psi^{(2)} = E^*\mathbb{E}\left[(\hat{\theta}^g - \theta^g)(\hat{\theta}^g - \theta^g)'\right]$, and $\Psi^{(3)} = E^*\left[\mathbb{E}(\hat{\theta}^g - \theta^g)\ \ (\theta^g - \mu_\theta)'\right]$. When the estimators are unbiased, equation (IA.12) is equivalent to

$$\Psi^{(1)} = E^*\left[Var(\hat{\theta}^g)\right] + V_\theta,$$

as the law of total variance implies.

In finite samples, we estimate V_θ by as follows. First, we estimate μ_θ by the size-weighted mean of group estimates:

$$\bar{\theta} = \sum_g k^g\ \hat{\theta}^g,$$

where $k^g = N^g/(\sum_k N^k)$ is the share of group g in the population. The estimator $\bar{\theta}$ is a consistent estimator of μ_θ as the group sizes $N^1, \ldots, N^G$ go to infinity.

Second, we estimate $\Psi^{(1)}$ by the size-weighted variance of group estimates:

$$\hat{\Psi}^{(1)} = \sum_{g=1}^{G} k^g\ (\hat{\theta}^g - \bar{\theta})\ (\hat{\theta}^g - \bar{\theta})'.$$

We estimate $E^*\left[Var(\hat{\theta}^g)\right]$ by the average variance-covariance matrix of $\hat{\theta}^g$:

$$\hat{\Psi}^{(2)} = \sum_{g=1}^{G} k^g\ \frac{\hat{V}^g}{N^g},$$

where $\hat{V}^g$ is the asymptotic variance-covariance matrix defined in section II.C of this appendix. We therefore estimate the variance-covariance of θ by

$$\hat{V}_\theta = \hat{\Psi}^{(1)} - \hat{\Psi}^{(2)}.$$

B.2. Results

Our asymptotic standard errors can be used to adjust our estimates of the heterogeneity in true preference parameters. Table 4 and Figure 3 of the main text describe the cross-sectional distribution of our parameter estimates, but this is increased by noise in the estimation procedure. Since our asymptotic standard errors estimate the noise for each group, in principle we can correct for the effect of noise on the estimated cross-sectional variance of parameters by subtracting the cross-sectional average squared standard error from the cross-sectional variance of our estimates.

A practical difficulty in doing this is that some groups have extremely high standard errors. Although these high standard errors are not pervasive enough to undermine our ability to reject homogeneous preferences for most households in the group-specific tests reported in Table 7, they do have a strong influence on the cross-sectional average of squared standard errors. In fact, if we do not limit the influence of outliers the average squared standard error is higher than the cross-sectional variance of estimates for TPR and EIS, implying a negative cross-sectional variance for true TPR and EIS.

We obtain more reasonable results if we winsorize the group-specific standard errors at the 90th percentile of the cross-sectional distribution. Table IA.16 shows that this procedure implies a cross-sectional standard deviation of 1.05 for RRA, 5.82% for the TPR, and 0.55 for the EIS, as compared with the cross-sectional standard deviations of estimates reported in Table 4 of the main text which are 1.06, 6.96%, and 0.90 respectively.

V. Detailed Description of Swedish Pension System

This Section explains the rules of the Swedish pension system applying to households in the 1999 to 2007 sample. Section V.A defines key indexes used in the calculation of pensions. Section V.B discusses the organization of public pensions, and Section V.C the organization of occupational pensions. Section V.D explains the allocation rule of the default fund in the public DC pension system. Section V.E describes the imputation of private pension contributions and wealth. Section V.F analyzes the impact of taxation.

A. Key Indexes

The calculations of public and occupational pensions rely on the following three income indexes defined in the Swedish Social Insurance Code.

- PBB_t denotes the "price-related base amount" (*prisbasbelopp*). It is set by the government based on calculations produced by Statistic Sweden.
- $HBPA_t$ is the "higher price-related base amount" (*förhöjt prisbasbelopp*). It is set by the government based on calculations produced by Statistic Sweden.
- IBB_t denotes the "income-related base amount" (*inkomstbasbelopp*). It is set by the government based on calculations produced by the Swedish Pensions Agency.

The price-related base amount, PBB_t, and the higher price-related base amount, $HBPA_t$, are available from Statistics Sweden's website. The income-related base amount, IBB_t, is available from the Swedish Parliament's website. For the year 2003, the price-related base amount is 38,600 SEK, the higher-price related base amount is 39,400 SEK, and the income-related base amount is 40,900 SEK.

B. *Public Pension System*

The Swedish Pensions Agency administers both the defined benefit (DB) and defined contribution (DC) components of public pensions.[10] An individual's public pension has four components:

(i) the DB income pension (*inkomstpension*), $IP_{i,t}$, which is based on the individual's lifetime income;

(ii) the DC premium pension (*premiepension*), $PPM_{i,t}$, which is contingent on the returns on the compulsory DC premium pension contributions invested by the individual;

(iii) the DB guaranteed pension (*garantipension*), $GP_{i,t}$, if the individual has a low income pension $IP_{i,t}$ or no earnings-related pensions;

(iv) the DB supplementary pension (*tilläggspension*), $ATP_{i,t}$, if the individual is born in 1953 or earlier.

Our terminology closely matches the terminology used in the annual Orange Reports of the Swedish Pensions Authority (see, e.g., 2005, 2015). Sections V.B.1 to V.B.4 of this appendix provide detailed definitions of the four components of public pensions.

An individual born before 1938 receives a total public pension equal to

$$\text{Public Pension}_{i,t} = ATP_{i,t} + GP_{i,t}. \tag{IA.13}$$

[10]See the Swedish Pensions Agency's website (`https://www.pensionsmyndigheten.se`) and Röstberg et al. (2004) for detailed explanations.

An individual born between 1938 and 1953 receives

$$\text{Public Pension}_{i,t} = \begin{cases} 4/20\ (IP_{i,t} + PPM_{i,t}) + 16/20\ ATP_{i,t} + GP_{i,t} & \text{if born in 1938,} \\ 5/20\ (IP_{i,t} + PPM_{i,t}) + 15/20\ ATP_{i,t} + GP_{i,t} & \text{if born in 1939,} \\ \dots \\ 19/20\ (IP_{i,t} + PPM_{i,t}) + 1/20\ ATP_{i,t} + GP_{i,t} & \text{if born in 1953.} \end{cases} \tag{IA.14}$$

An individual born after 1953 receives a total public pension equal to

$$\text{Public Pension}_{i,t} = IP_{i,t} + PPM_{i,t} + GP_{i,t}. \tag{IA.15}$$

All variables in (IA.13)-(IA.15) are pre-tax. The implications of taxes on public and occupational pensions are analyzed in Section V.F of this appendix.

B.1. DB Income Pension

The DB income pension is a pay-as-you-go system that applies to households in all cohorts, as defined by the 1998 Earnings Related Old Age Pension Act. The rules of the income pension system mimic the organization of individual funded accounts. During an individual's working life, contributions are credited to a notional individual account. Each year a notional interest payment is credited to the account. At retirement, the notional account becomes an annuity, whose amount is determined by the individual's notional balance and his/her cohort's life expectancy.

The precise calculation of the DB income pension proceeds in three steps.

Step 1 (Notional accumulation). The notional accumulation is computed from annual income and credits as follows.

- The relevant income of individual i in year t, $RI_{i,t}$, consists of the year's annual earnings, including sickness cash benefits, parental cash benefits, and unemployment cash benefits.

- The pension qualifying income, $PQI_{i,t}$, is defined as

$$PQI_{i,t} = RI_{i,t} - 7\% \times \min(RI_{i,t}, 8.07\ IBB_t),$$

  where IBB_t is the income-related base amount defined in Section V.A of this appendix.

- The pension qualifying amount, $PQA_{i,t}$, is a fictitious income computed for years of child care, compulsory national service, studies, and sickness or activity compensation.[11]

- The pension base (*pensionunderlag*) is the sum of the pension-qualifying income and pension-qualifying amounts, capped at 7.5 times the income-related base amount:

$$\text{Pension Base}_{i,t} = \max[PQI_{i,t} + PQA_{i,t}\ ;\ 7.5\ IBB_t]. \tag{IA.16}$$

  Beginning in 2003, the pension base is zero if $PQI_{i,t} + PQA_{i,t}$ is lower than $42.3\% \times PBB_t$.

- The pension right accumulated by individual i in year t is

$$\text{Pension Right}_{i,t} = 16.0\% \times \text{Pension Base}_{i,t}.$$

[11] Pension-qualifying amounts for *child care* apply for a period of up to 4 years after a child's birth; the pension-qualifying amount in year t is the maximum of the following three quantities: (i) the parent's pension-qualifying income in the year prior to the child's birthyear, (ii) 75% of year t's average pension-qualifying income of all insured persons aged between 16 and 64, and (iii) the year's income-related base amount, IBB_t. The pension-qualifying amount for *compulsory national service* is set to 50% of the average pensionable qualifying-income for all insured persons aged between 16 and 64. Pension-qualifying amounts for *study* are calculated as 138% of disbursed study grants. Pension-qualifying amounts for *sickness and activity compensation* are set to 93% of the income that the insured person would have likely received if he/she had worked according to the Swedish Social Insurance Agency (*Försäkringskassan*). See Swedish Ministry of Health and Social Affairs (2009) for further details.

Step 2 (Notional balance). The notional *pension balance at t* is

$$\begin{aligned}\text{Pension Balance}_{i,t} &= 99.95\% \times \text{Inheritance Gain Factor}_{i,t} \\ &\times \left[\sum_{t=S_i}^{T_i} \left(\text{Pension Right}_{i,t} \times \frac{\text{IncIndex}_{T_i}}{\text{IncIndex}_t}\right)\right],\end{aligned} \tag{IA.17}$$

where Inheritance Gain $\text{Factor}_{i,t}$ is specific to the individual's cohort in year t, S_i is the individual's first year with a strictly positive pension-qualifying income, T_i is the year of retirement which is typically the 65th birthday, and is at the earliest the 61st birthday. IncIndex_t is the income index for year t. The 99.95% factor corresponds to administrative costs of 0.05%.

The inheritance gain factor (*arvvinstfaktorer*) is computed by the National Social Insurance Board (*Riksförsäkringsverket*). It is designed to allocate the pension balances of deceased persons to survivors in the same age group. For cohorts up to the age of 60, the inheritance gain factor is equal to unity plus the ratio of the pension balances of the deceased to the pension balances of survivors. For cohorts aged more than 60, the inheritance gain factor is a survival rate, which permits the homogeneous treatment of economically active and retired individuals in the same cohort.[12]

The Income Index is computed by the National Social Insurance Board and is available from its website. By definition, the growth rate of the income index, $\text{IncIndex}_t/\text{IncIndex}_{t-1}$, is the product of 1-year inflation with the 3-year moving average of real earnings.[13]

Step 3 (Payout). The annual DB income pension is calculated by dividing the individual's pension account balance (IA.17) by his/her cohort's annuitization divisor (*delningstal*):

$$IP_{i,t} = \frac{\text{Pension Balance}_{i,t}}{\text{Annuitization Divisor}_i}.$$

[12]The exact definitions of the Inheritance Gain Factor is provided in Appendix A of Swedish Pensions Agency (e.g., 2005).

[13]See Appendix A of Swedish Pensions Agency (e.g., 2005).

The annuitization divisor is computed by the National Social Insurance Board. It is common to individuals born in the same year and does not vary after age 65.

B.2. DC Premium Pension

The DC component of public pensions, which was introduced in 1995, is designed as a funded system. During the working life of individual i, the DC premium contribution in every year t is

$$\text{Premierate}_i \times \text{Pension Base}_{i,t},$$

where Pension Base$_{i,t}$ is given by (IA.16). The rate of contribution, Premierate$_i$, is 2.5% for years 1999 and later, and 2% in years 1995-1998. Premium pension contributions are invested by default in the AP7 fund (see section V.D of this appendix for details). Individual can opt out and choose funds among a large menu of mutual funds available in the system.[14]

At retirement, the insured person can either convert the funds into an annuity or keep them invested in mutual funds.

- If the individual chooses to receive an annuity, the pension is calculated as a guaranteed life-long annuity payable in nominal monthly installments.
- Alternatively, the pension savings remain in the account and are invested in mutual funds chosen by the insured. The premium pension is recalculated once a year based on the value of fund shares in December.

In both cases, the premium pension, $PPM_{i,t}$, is the value of the premium pension account divided by an annuity divisor, which is based on forecasts of future life expectancy.[15]

B.3. DB Guaranteed Pension

The guaranteed pension provides basic retirement income for individuals who have had little or no previous earnings. Unlike other components of the public pension system, the

[14]In June 2003, 655 funds were available for active choice on the Premium pension platform

[15]See Appendix A of Swedish Pensions Agency (e.g., 2005) for further details.

guaranteed pension is funded by the state budget. Residents of Sweden are eligible beginning at age 65. In order to obtain a maximum guaranteed pension, one needs to have resided in Sweden for 40 years after age 25.[16] In our simulations, we assume for simplicity that all households meet this criterion.

The level of the guaranteed pension is determined by the DB income pension, $IP_{i,t}$, defined in section V.B.1 of this appendix. The guaranteed pension, $GP_{i,t}$, is given by

$$GP_{i,t} = \begin{cases} 2.13\ PBB_t - IP_{i,t} & \text{when } IP_{i,t} < 1.26\ PBB_t \\ 0.87\ PBB_t - 0.48(IP_{i,t} - 1.26\ PBB_t) & \text{when } 1.26\ PBB_t \leq IP_{i,t} \leq 3.07\ PBB_t \\ 0 & \text{when } IP_{i,t} > 3.07\ PBB_t \end{cases}$$

for unmarried individuals, and by

$$GP_{i,t} = \begin{cases} 1.90\ PBB_t - IP_{i,t} & \text{when } IP_{i,t} < 1.14\ PBB_t \\ 0.76\ PBB_t - 0.48(IP_{i,t} - 1.14\ PBB_t) & \text{when } 1.14\ PBB_t < IP_{i,t} < 2.72\ PBB_t \\ 0 & \text{when } IP_{i,t} > 2.72\ PBB_t \end{cases}$$

for married pensioners, where PBB_t is the price-related base amount defined in section V.A of this appendix.

B.4. DB Supplementary Pension

Individuals born in 1953 or earlier are entitled to a public DB supplementary pension, which can be computed in three steps.

Step 1. We compute the individual's 15 best years of earnings, out of 30. For each of these 15 years we calculate:

$$\text{Pension Points}_{i,t} = (PQI_{i,t} - HBPA_t)/HBPA_t, \tag{IA.18}$$

[16] Years of residence in another EU or EEA country also provide guaranteed pension credits.

where $HBPA_t$ is the higher price-related base amount defined in section V.A of this appendix. The average pension points for the best 15 years is denoted by AvPoints_i.

Step 2. The annual supplementary DB pension is

$$SuppDB_{i,t} = 60\% \times \text{AvPoints}_i \times PBB_t, \tag{IA.19}$$

where PBB_t is the price-related base amount. An individual needs to work 30 years in order to get the full supplementary DB pension. The benefit (IA.19) is otherwise multiplied by a factor of $N/30$, where N is the number of years for which the individual received pension rights. In our simulations, we set N equal to 30 for all individuals in the panel.

Step 3. We compute the *folkspension*:

$$Folkspension_{i,t} = \begin{cases} 96\% \times PBB_t & \text{for an unmarried person,} \\ 78.5\% \times PBB_t & \text{for a married person,} \end{cases}$$

where PBB_t is the price-related base amount.

Step 4. The total supplementary pension is

$$ATP_{i,t} = SuppDB_{i,t} + Folkspension_{i,t}.$$

C. Occupational Pension System

We now provide a detailed description of some of the main collective labor agreements in Sweden:

- STP and SAF-LO, which cover private-sector blue collar workers (1.3 million persons in 2002);

- ITP2, which covers private-sector white collar workers (610,000 persons in 2002);

- PA 91 and PA 03, which cover central government employees (250,000 persons in 2002);
- PFA 98 and KAP-KL, which cover local government employees (940,000 persons in 2002).

Overall, these agreements cover about 90% of the Swedish workforce. We refer the reader to Sjögren and Wadensjö (2005, ch. 6.4.1) for an excellent detailed description of these occupational plans.

C.1. Private-Sector Blue Collar Workers

In 1973, the Swedish Employers Association (SAF) and the central confederation of blue-collar workers (LO) agreed to provide a special defined benefit pension plan, the *särskild tilläggspension* (STP). The STP introduced a defined contribution component in 1991, which came into effect in 1992. Since 1996, the main pension scheme for blue-collar workers is the SAF-LO defined contribution agreement. However, workers previously enrolled in the STP defined benefit scheme still receive pension payments between 1999 and 2007.

Defined Benefits. Yearly DB pension payouts are given by

$$10\% \times (\text{Relevant Pension Points}_i + 1) \times PBB_t,$$

where Relevant Pension Points$_i$ is the average of Pension Points$_{i,t}$ (defined in (IA.18)) over the best 3 years between the ages of 55 and 59. Following the introduction of SAF-LO, the DB pension payouts are adjusted for birthyear using the same ratios as the ones applied to the public DC premium pension explained in section V.B.2 of this appendix.

Defined Contributions. The average yearly premium is

$$
\begin{aligned}
&0.53\% \times PQI_{i,t} && \text{in 1992,} \\
&0.54\% \times PQI_{i,t} && \text{in 1993 and 1994,} \\
&3.90\% \times PQI_{i,t} && \text{in 1996,} \\
&4.08\% \times PQI_{i,t} && \text{in 1997,} \\
&3.50\% \times PQI_{i,t} && \text{from 1998 to 2007.}
\end{aligned}
\tag{IA.20}
$$

There was no contribution in 1995 as the pension system transitioned to a pure defined contribution scheme.

The premium in (IA.20) is the sum of the contributions paid by both the worker and the employer. Specifically, the worker and the employer respectively contributed $2\% \times PQI_{i,t}$ and $1.9\% \times PQI_{i,t}$ to the pension plan in 1996, $2\% \times PQI_{i,t}$ and $2.08\% \times PQI_{i,t}$ in 1997, and $2\% \times PQI_{i,t}$ and $1.5\% \times PQI_{i,t}$ in 1998 and 1999. Starting in 2000, employees pay a contribution of $3.5\% \times PQI_{i,t}$ (no ceiling), and there are no contributions from employers.

The age at which defined contributions start accruing is 28 until 1999, 22 in 2000 and 2001, and 21 in 2002 onward. DC premia continue to be paid if the worker is on sick leave, pregnancy leave, or parental leave.[17] Blue collar workers covered by SAF-LO can allocate their contributions to traditional insurance (characterized by conservative investments and guaranteed payments) and unit-linked insurance (riskier investments).

C.2. Private-Sector White Collar Workers

The *industrins och handelns tilläggspension* (ITP) covers white-collar workers from the private sector. It consists of two schemes: ITP2 for members born before 1979 and ITP1 for members born therafter. ITP1 is not relevant for our paper and is therefore not discussed here. In addition, the ITPK agreement was introduced in 1977 as a compulsory DC scheme

[17] In these cases, the premia are paid by an insurance tool, called *premiebefrielseförsäkring*, instead of being deducted from the salary. The rationale is that long-term illness, for instance, should not result in lower occupational pension on retirement.

for private-sector white collar workers (Hagen 2013).

Defined Benefits. ITP2 provides a specific share of the final salary after thirty years of service. Let $SIYBR_i$ denote the salary in the year before retirement. The ITP2 DB payout is given by the following formula:

$$
\begin{aligned}
&(10\% \times \text{portion of } SIYBR_i \text{ below } 7.5\ IBB_t) \\
&+(65\% \times \text{portion of } SIYBR_i \text{ between } 7.5\ IBB_t \text{ and } 20\ IBB_t) \\
&+(32.5\% \times \text{portion of } SIYBR_i \text{ between } 20\ IBB \text{ and } 30\ IBB_t).
\end{aligned}
$$

Levels of $SIYBR_i$ above 30 IBB_t do not provide an ITP2 payout. Before 2003, the formula was based on the increased price base amount, $HBPA_t$, instead of the income base amount, IBB_t.

Defined Contributions. The yearly contribution to the ITPK defined-contribution plan is

$$2\% \times \min(PQI_{i,t}; 30\ IBB_t)$$

over the entire period, where $PQI_{i,t}$ is the pension qualifying income that applies to enrolled workers. Workers start contributing to the ITPK plan at age 28. DC premia continue to be paid if the worker is on sick leave, pregnancy leave, or parental leave.

Before 1990, the ITPK plan was handled as if each payment gave rise to a benefit and ITPK payouts were part of the collective refund (i.e. with base year and pension allowance). Pre-1990 ITPK can thererefore be seen as a hybrid between a defined contribution and a defined benefit plan. In 1990 this changed and it became possible to invest ITPK premia in garanteed or unit-linked products provided by companies such as Collectum or Valcentralen. For these reasons, we include pre-1990 ITPK in the DB payout defined for regular ITP2.

High Earners. Workers earning more than 7.5 IBB_t in a given year, who are known as *tiotaggare* (high earners), are entitled to leave their standard ITP arrangement and join a defined contribution scheme, in agreement with their employer. According to Collectum (2015), the contributions to these alternative schemes, are calculated using either the *frilagd premie* (free premium) or the *premietrappa* (premium ladder) for the portion of income between 7.5 IBB_t and 30 IBB_t.

Our dataset does not provide the type of premium chosen by high earners and their employers. However, a feature of the system allow us to make a simplifying assumption. The "free premium" matches the costs the employers would incur within the DB component of ITP 2; it tends to vary substantially, which makes it difficult for employers to budget and control their pension costs. By contrast, the "premium ladder" is stable and easy to budget, and the variation of pension premia with age and salary levels can easily be explained to employees. The premium ladder is therefore the more popular option, so we assume that high earners who leave the ITP 2 system always choose the "premium ladder."

The dataset does not allow us to determine if a high earner is enrolled in the ITP 2 system or in an alternative scheme. As a compromise, we assume that a high earner invests half of the premium in the ITP 2 DB scheme and the other half in the alternative DC scheme.

C.3. Central Government Employees

Central government employees are covered by two agreements: PA 91 from 1992 to 2002, and PA 03 since 2003.

Defined Benefits. The pension basis is the average of the 5 last salaries before retirement. The PA 91 defined benefit payout is

$$
\begin{aligned}
&(10\% \times \text{portion of pension basis below } 7.5\ HBPA_t) \\
&+(65\% \times \text{portion of pension basis between } 7.5\ HBPA_t \text{ and } 20\ HBPA_t) \\
&+(32.5\% \times \text{portion of pension basis between } 20\ HBPA_t \text{ and } 30\ HBPA_t).
\end{aligned}
$$

The PA 03 defined benefit payout is

$$(0\% \times \text{portion of pension basis below } 7.5\ HBPA_t)$$
$$+(60\% \times \text{portion of pension basis between } 7.5\ HBPA_t \text{ and } 20\ HBPA_t)$$
$$+(30\% \times \text{portion of pension basis between } 20\ HBPA_t \text{ and } 30\ HBPA_t).$$

Payments of PA 91 continued after the introduction of PA 03 for persons born before 1943.

Defined Contributions. Yearly PA 91 pension contributions are made under the Kåpan scheme and are defined as follows:

$$1.5\% \times PQI_{i,t} \text{ in years 1992 to 1994,}$$
$$1.7\% \times PQI_{i,t} \text{ in years 1995 to 2002,}$$

for workers who are at least 28 years of age. DC premia continue to be paid if the worker is on sick leave, pregnancy leave, or parental leave.

Yearly PA 03 pension contributions are made up of two parts: the *individuell ålderspension* premium and the Kåpan premium.[18] From the age of 23, central government employees pay the *individuell ålderspension* premium:

$$2.3\% \times PQI_{i,t} \text{ in 2003,}$$
$$2\% \times PQI_{i,t} \text{ thereafter.}$$

From the age of 28, employees also pay the Kåpan premium:

$$1.9\% \times PQI_{i,t} \text{ in 2003,}$$
$$2.3\% \times PQI_{i,t} \text{ thereafter.}$$

[18]Source: Sjögren and Wadensjö (2005, p. 159-162).

Central government employees younger than 23 pay no DC contributions.

C.4. Local Government Employees

Local government workers are covered by two occupational pensions schemes during the period: PFA 98 between 1998 and 2006, and KAP-KL thereafter.

Defined Benefits. The pension payout is

$$
\begin{aligned}
&(0\% \times \text{portion of pension basis below } 7.5\ IBB_t) \\
&+(PP1 \times \text{portion of pension basis between } 7.5\ IBB_t \text{ and } 20\ IBB_t) \\
&+(PP2 \times \text{portion of pension basis between } 20\ IBB_t \text{ and } 30\ IBB_t),
\end{aligned}
$$

where the pension payout basis is the annual average salary in the 5 best years out of the last 7 years before retirement. For the 1999 to 2006 period, PP1 = 62.5% and PP2 = 31.25%. For the year 2007, the percentages PP1 and PP2 remain the same only for the cohort born in 1946 and decrease with birth-year for younger cohorts according the KAP-KL agreement.[19]

Defined Contributions. The DC premia are not the same whether the employes is a union member of (i) OFR-förbunden and Akademikeralliansen or (ii) Kommunal. Unfortunately, the data set does not provide the union membership of individual employees. Since the OFR-förbunden and Akademikeralliansen have more members overall, we apply their pension rules to all local government employees.

For the period 1998-2006, the PFA 98 premium is

$$
\begin{aligned}
&3.5\% \times (\text{portion of } PQI_{i,t} \text{ up to } 7.5\ HBPA_t) \\
&+1.1\% \times (\text{portion of } PQI_{i,t} \text{ above } 7.5\ HBPA_t),
\end{aligned}
$$

[19]Source: KAP-KL Avtal (annex 1, section 21) available at `www.ocr.se`.

where $PQI_{i,t}$ is the pension-qualifying income. For 2007, the KAP-KL yearly premium is[20]:

$$4\% \times (\text{portion of } PQI_{i,t} \text{ up to } 30\ IBB).$$

Between 1998 and 2004, DC premia are paid by local government employees who are at least 28 years old. In 2004, that age was reduced to 21 for many local government employees. DC premia continue to be paid if the worker is on sick leave, pregnancy leave, or parental leave.

D. *Allocation Rule of the PPM Default Fund*

To calculate retirement wealth held in the PPM system, we assume that it is invested fully in the AP7 fund, which is the default option in the DC public pension system (PPM). According to the AP7 fund policy during our sample period, each worker's PPM wealth is invested in only two funds: the AP7 Fixed Income Fund (*AP7 Räntefond*) and the AP7 Equity Fund (*AP7 Aktiefond*). The fixed income fund is effectively invested in the risk free rate. The equity fund is 130% levered in the unhedged MSCI World Index.

The allocation depends on the age of the worker. Until age 55, the AP7 fund is fully invested in the equity fund. At 56, the weight of equity fund is 97%. Between 57 and 60, the share of the equity fund decreases by 3.5 percentage points per year to 83%. Between 61 and 65 years of age, the share of the equity fund decreases by 3.2 percentage points per year to 67%. Between 66 and 70 years of age, it decreases by 3.4 percentage points per year until it is 50% at 70 years old.

To convert this proportion into the equity share invested in the unlevered MSCI World Index, we multiply the allocation in the equity fund by 130%.

[20]Source: KAP-KL Avtal (annex 1, section 11) available at `www.ocr.se`.

E. *Private Pension Contributions and Wealth*

This section discusses the imputation of individual private pension contributions and wealth. The Swedish wealth registry contains exact information on private pension contributions from 1994, and only reports a capped version of the variable from 1991. We impute full contributions from 1991 to 1993 using information in the years during which we observe both variables (i.e. 1994 onwards). More specifically, for each of the three education groups, we separately estimate the following individual fixed effect regression model over the years 1994 to 2007:

$$\begin{aligned}\log(T_{i,t}) = a_i + b_1 RI^1_{i,t} + \cdots + b_{20} RI^{20}_{i,t} + c\log(C_{i,t}) + c_1 C^1_{i,t} + \cdots + c_{20} C^{20}_{i,t} \\ + c_{21} age^{21}_{i.t} + \cdots + c_{64} age^{64}_{i.t} + e_{i,t},\end{aligned}$$

where $T_{i,t}$ is the true contribution and $C_{i,t}$ is the capped contribution. Variables with superscripts are dummy variables. That is, $RI^1_{i,t}, \ldots, RI^{20}_{i,t}$ are dummies for equally spaced percentiles (9-5%, 5%-10%, ..., 90%-95%, 95%-100%) of the relevant income variable defined in section V.B.1 of this appendix. Similarly, $C^1_{i,t}, \ldots, C^{20}_{i,t}$ are dummies for equally spaced percentiles of the capped contribution. We use the estimated model to predict individual contributions $T_{i,t}$ over the years 1991-1993.[21]

Since we do not have information on private pension contribution before 1991, we redistribute the aggregate stock of private pension wealth held by the working population in proportion to the individual share of private pension contributions in 1991. To calculate the stock of private pension wealth held by the working population we proceed in two steps. First, in 1991, the national accounts do not distinguish between private pensions and capital insurance (*kapitalförsäkring*), and only report the aggregate stock of both. We take the closest year national accounts have information on the split, and use the 80/130 fraction reported in 1989. Second, as noted in the main text, we follow Bach, Calvet and Sodini

[21] We exclude observations for which the true variable or the capped variable are equal to zero.

(2020) and further assume that 58% of private pension wealth in 1991 belongs to workers.

F. *Taxation of DC Contributions and DB Payouts*

The tax rate for withdrawing your pension before the age of 65 is higher than the tax rate for withdrawing the pension after the age of 65.

First, we calculate the tax rate on the pension income of already retired people, splitting them into three education groups (basic or missing education, high school, post-high school).

Second, we apply the obtained tax rate (specific for each education group) from the retired people and tax with it the DB payouts of the working/retiring people having the same education level. On average, the tax rate is almost invariably about 32% irrespective of education level. Additionally, the standard deviations of the tax rate by education level are negligibly small. We then tax the DC accumulations with a flat rate of 32%.

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VI. Tables

Table IA.1
Sizes of Education and Income Risk Categories

	No High School	High School	Post-High School	All
Low	17,929	44,224	47,988	110,141
Medium	16,036	53,920	49,001	118,957
High	11,362	29,403	28,677	69,442
All	45,327	127,547	125,666	298,540

This table reports the number of households in groups with 3 levels of education and working in sectors with 3 levels of income volatility given in Table IA.2, and for aggregates of these categories.

Table IA.2
Employment Sectors of Education and Income Risk Categories

	No High School	High School	Post-High School
Low	Mining and utilities	Mining and utilities	Mining and utilities
	Manufacturing	Public sector	Public sector
	Public sector	Manufacturing	Manufacturing
	Healthcare	Healthcare	Education
Medium	Education	Finance	Healthcare
	Finance	Education	Construction
	Transportation	Transportation	Other services
	Other services	Construction	Transportation
High	Construction	Wholesale	Finance
	Wholesale	Real estate	Wholesale
	Real estate	Other services	Real estate
	Hotels	Hotels	Hotels

This table reports the classification of employment sectors into 3 levels of income volatility, separately for households with 3 levels of education. Sectors are listed in ascending order of income volatility within each education level.

Table IA.3
Percentage Volatilities of Income Shocks

	Total			Systematic		
	No High School	High School	Post-High School	No High School	High School	Post-High School
Low	13.86	13.69	15.65	2.68	2.80	3.27
Medium	18.06	16.54	17.06	2.95	3.16	3.33
High	21.36	20.75	22.29	2.89	3.18	3.68

	Idiosyncratic permanent			Idiosyncratic transitory		
	No High School	High School	Post-High School	No High School	High School	Post-High School
Low	7.22	6.60	5.15	11.52	11.67	14.41
Medium	8.08	7.45	4.16	15.88	14.43	16.21
High	6.03	6.45	3.91	20.29	19.46	21.63

This table reports the standard deviations of income shocks, in percentage points, for Swedish household groups with 3 levels of education and working in sectors with 3 levels of income volatility. The top left panel reports the total standard deviation of income shocks, the top right panel reports the standard deviation of systematic (group-level) permanent income shocks, the bottom left panel reports the standard deviation of idiosyncratic (household-level) permanent income shocks, and the bottom right panel reports the standard deviation of idiosyncratic transitory income shocks.

Table IA.4
Percentage Correlations of Income Shocks with Wealth Shocks

	Risky Liquid Wealth			Real Estate Wealth		
	No High School	High School	Post-High School	No High School	High School	Post-High School
Low	4.29	6.05	10.46	29.87	32.90	42.61
Medium	3.07	6.21	9.09	30.18	34.10	50.94
High	2.58	5.80	23.33	32.57	35.07	47.93

	Aggregate Wealth		
	No High School	High School	Post-High School
Low	19.11	21.98	30.30
Medium	18.42	22.74	33.83
High	19.38	22.97	42.17

This table reports the correlations of income shocks with wealth shocks, in percentage points, for Swedish household groups with 3 levels of education and working in sectors with 3 levels of income volatility. The top left panel reports the correlation between the excess return on risky liquid wealth net of taxes and group average total permanent income shocks, the top right panel reports the correlation between the excess return on real estate wealth net of taxes and group average total permanent income shocks, the bottom panel reports the average yearly correlation between the excess return on non-cash aggregate net wealth and group average total permanent income shocks.

Table IA.5
Size-Weighted Cross-Sectional Regressions of
Estimated Preference Parameters on Group Financial Characteristics

	(1) RRA	(2) RRA	(3) TPR	(4) TPR	(5) EIS	(6) EIS
Average RS	-3.457***	-3.742***	0.153***	0.138***	0.525***	0.161
	(0.116)	(0.082)	(0.007)	(0.008)	(0.109)	(0.113)
Initial WY	-0.165***	-0.111***	-0.001***	-0.000	0.150***	0.153***
	(0.006)	(0.005)	(0.000)	(0.000)	(0.005)	(0.005)
Growth of WY	5.392***	9.472***	0.389***	0.509***	5.607***	6.540***
	(0.674)	(0.671)	(0.035)	(0.036)	(0.596)	(0.653)
Convexity of WY	0.172	-0.288	-0.051***	-0.052***	-0.570***	-0.393*
	(0.190)	(0.153)	(0.011)	(0.012)	(0.158)	(0.159)
Constant	4.687***	1.731*	-0.449***	-0.553***	-5.932***	-6.752***
	(0.716)	(0.709)	(0.037)	(0.039)	(0.635)	(0.699)
Control variables	No	Yes	No	Yes	No	Yes
R^2	0.436	0.832	0.388	0.438	0.208	0.234

This table reports the cross-sectional regression coefficients across estimated preference parameters and group financial characteristics. All regressions weight groups by their size, to recover the underlying cross-sectional relationships at the household level. Growth of WY is defined in equation (12) and convexity of WY is defined in equation (13). Standard errors are reported in parentheses and statistical significance levels are indicated with stars: * denotes 1-5%, ** 0.1-1%, *** less than 0.1% significance. There are 4,276 groups containing 298,540 households. Control variables are 9 income risk/education categories and cohort fixed effects.

Table IA.6
Wealth-Income Ratio and Risky Share under Riskless Real Estate

	Mean	Median	Std. Dev.	10%	25%	75%	90%
RRA	7.14	7.10	1.03	5.90	6.40	7.90	8.50
TPR (%)	2.63	2.12	5.12	-4.31	0.80	3.25	6.40
EIS	1.19	1.19	0.92	0.10	0.21	2.05	2.50
Average RS	0.59	0.56	0.14	0.42	0.49	0.67	0.77
Initial WY	6.76	5.17	4.71	1.73	3.09	9.09	14.70
Growth of WY	1.06	1.05	0.04	1.02	1.03	1.07	1.10
Convexity of WY	0.25	0.23	0.10	0.15	0.19	0.29	0.36

This table reports the mean, median, standard deviation, and 10th, 25th, 75th, and 90th percentiles of estimated preference parameters and group financial characteristics. All statistics weight groups by their average net wealth. Growth of WY is defined in equation (12) and convexity of WY is defined in equation (13). There are 4,276 groups containing 298,540 households.

Table IA.7
Standard Deviation of Preference Parameters within Wealth Quintiles

Wealth Quintile	(1) RRA	(2) TPR (%)	(3) EIS
Bottom	1.207	7.616	0.965
2	1.205	7.645	0.911
3	1.120	7.077	0.930
4	1.022	4.714	0.946
Top	0.890	2.806	0.902

This table reports the standard deviation of preference parameter estimates within wealth quintiles. Households are equally-weighted within each quintile.

Table IA.8
Wealth-Income Ratio and Risky Share under Riskless Real Estate

Panel A. Cross-Sectional Means

	WY				RS			
	No High School	High School	Post-High School	All	No High School	High School	Post-High School	All
Low	3.66	4.12	5.08	4.46	0.459	0.462	0.454	0.458
Medium	4.46	4.50	4.92	4.67	0.405	0.429	0.423	0.423
High	4.69	5.09	6.16	5.47	0.417	0.435	0.461	0.443
All	4.20	4.51	5.27	4.78	0.429	0.442	0.443	0.441

Panel B. Cross-Sectional Standard Deviations

	WY				RS			
	No High School	High School	Post-High School	All	No High School	High School	Post-High School	All
Low	3.03	3.28	3.61	3.44	0.224	0.216	0.206	0.213
Medium	3.68	3.58	3.77	3.68	0.214	0.215	0.208	0.212
High	3.83	3.87	3.93	3.93	0.214	0.210	0.198	0.207
All	3.51	3.57	3.78	3.67	0.219	0.215	0.206	0.212

Panel A reports cross-sectional means of the wealth-income ratio (WY) and risky share (RS) for Swedish household groups with 3 levels of education and working in sectors with 3 levels of income volatility given in Table 2, and for aggregates of these groups. Panel B reports cross-sectional standard deviations of WY and RS across the groups in each of these categories and their aggregates. In contrast to Table 1 of the main text, we treat real estate as riskless in the calculation of the risky share and the formation of household groups (defined in Section1.3 of the main text). All statistics weight groups by their size, that is by the number of households they contain, to recover the underlying household-level statistics assuming homogeneity of WY and RS within groups. Summary statistics on group size are reported in the online appendix.

Table IA.9
Panel Regressions of Wealth-Income Ratio and Risky Share on Group Characteristics under Riskless Real Estate

	(1) WY	(2) RS	(3) RS
Age	0.158***	-0.008***	-0.005***
	(0.015)	(0.001)	(0.001)
Total income volatility	15.861***	-0.210*	0.088
	(1.794)	(0.100)	(0.093)
High school	0.523***	0.005	0.015*
	(0.131)	(0.008)	(0.007)
Post-high school	1.087***	0.010	0.031***
	(0.132)	(0.007)	(0.007)
WY			-0.019***
			(0.001)
Constant	-6.493***	0.984***	0.862***
	(0.750)	(0.045)	(0.041)
Year fixed effects	Yes	Yes	Yes
R^2	0.104	0.167	0.263

This table reports panel regressions of the wealth-income ratio (WY) and risky share (RS) on group characteristics including the age of households in the group, total income volatility (in natural units), and dummies for high-school and post-high-school education. In contrast to Table 2 of the main text, we treat real estate as riskless in the calculation of the risky share and the formation of household groups (defined in Section1.3 of the main text). All regressions weight groups by their size, to recover underlying relationships at the household level, and include year fixed effects. Standard errors are reported in parentheses and statistical significance levels are indicated with stars: * denotes 1-5%, ** 0.1-1%, *** less than 0.1% significance. There are 41,265 observations on groups, corresponding to 2,694,456 observations on underlying households.

Table IA.10
Size-Weighted Cross-Sectional Distributions of RMSE-Scaled Objective Functions for Alternative Model Specifications

	Mean	Median	Std. Dev.	10%	25%	75%	90%
Unrestricted	7.03	5.95	4.48	3.40	4.40	8.48	12.23
Fixed RRA	16.03	12.60	11.61	5.39	8.26	20.06	31.25
Fixed TPR	8.55	7.06	5.11	4.07	5.20	10.35	15.39
Fixed EIS	7.71	6.52	4.50	3.81	4.80	9.37	13.02
Fixed TPR and EIS	9.80	8.30	5.47	4.81	6.10	11.81	17.38
All Parameters Fixed	24.78	18.61	21.23	7.53	11.27	30.92	47.93

This table reports the mean, median, standard deviation, and 10th, 25th, 75th, and 90th percentiles of the RMSE-scaled objective function for several alternative model specifications. All statistics weight groups by their size to recover the underlying cross-sectional distributions at the household level. The RMSE-scaled objective function is the square root of the objective function divided by 4 and multiplied by 100 to express it in RMSE-equivalent percentage units. The results in the first row are for the unrestricted model estimated in Table 8 of the main text. The results in subsequent rows are for models that fix selected parameters at their size-weighted cross-sectional means estimated in the unrestricted model. There are 4,276 groups containing 298,540 households.

Table IA.11
Size-Weighted Cross-Sectional Distributions
of Estimated Preference Parameters and Group Financial Characteristics, Assuming Heterogeneous Beliefs

	Mean	Median	Std. Dev.	10%	25%	75%	90%
RRA	7.80	7.50	2.74	4.40	5.90	9.00	12.00
TPR (%)	4.72	3.15	6.80	-4.40	1.01	6.19	18.39
EIS	1.01	0.58	0.90	0.10	0.20	1.83	2.50
Average RS	0.65	0.63	0.17	0.45	0.53	0.75	0.90
Initial WY	4.28	3.04	3.90	0.87	1.64	5.22	9.25
Growth of WY	1.08	1.07	0.05	1.03	1.05	1.10	1.14
Convexity of WY	0.24	0.23	0.09	0.15	0.19	0.28	0.35

This table is an equivalent of Table 5, allowing three possible household beliefs about the Sharpe ratio, that it equals 0.15, 0.26, or 0.40. For each group we pick the Sharpe ratio and preference parameters that minimize the objective function. This table reports the mean, median, standard deviation, and 10th, 25th, 75th, and 90th percentiles of estimated preference parameters and group financial characteristics. All statistics weight groups by their size to recover the underlying cross-sectional distributions at the household level. Growth of WY is defined in equation (12) and convexity of WY is defined in equation (13). There are 4,276 groups containing 298,540 households.

Table IA.12
Size-Weighted Cross-Sectional Correlations
of Estimated Preference Parameters and Group Financial Characteristics, Assuming Heterogeneous Beliefs

	RRA	TPR	EIS	Average RS	Initial WY	Growth of WY
RRA	1.000					
TPR	0.087***	1.000				
EIS	0.042***	-0.210***	1.000			
Average RS	-0.176***	0.550***	-0.115***	1.000		
Initial WY	-0.018***	-0.508***	0.365***	-0.501***	1.000	
Growth of WY	0.052***	0.606***	-0.161***	0.600***	-0.709***	1.000

This table is an equivalent of Table 5, allowing three possible household beliefs about the Sharpe ratio, that it equals 0.15, 0.26, or 0.40. For each group we pick the Sharpe ratio and preference parameters that minimize the objective function. This table reports the cross-sectional correlations across estimated preference parameters and group financial characteristics. Correlations weight groups by their size to recover the underlying cross-sectional correlations at the household level. Growth of WY is defined in equation (12) and convexity of WY is defined in equation (13). Statistical significance levels of correlation coefficients are indicated with stars: * denotes 1-5%, ** 0.1-1%, *** less than 0.1% significance. There are 4,276 groups containing 298,540 households.

Table IA.13
Size-Weighted Cross-Sectional Distributions of RMSE-Scaled Objective Functions for Alternative Model Specifications

	Mean	Median	Std. Dev.	10%	25%	75%	90%
Homogeneous beliefs	7.03	5.95	4.48	3.40	4.40	8.48	12.23
Heterogeneous beliefs							
Unrestricted preferences	6.52	5.48	4.46	2.91	3.95	7.98	11.81
Fixed TPR and EIS	8.84	7.43	5.30	4.15	5.32	10.66	15.95
All preferences fixed	21.44	18.00	14.44	7.51	11.26	26.16	40.26
Pref vary only with beliefs	20.02	14.70	16.53	7.13	9.43	25.28	38.81

This table reports the mean, median, standard deviation, and 10th, 25th, 75th, and 90th percentiles of the RMSE-scaled objective function for several alternative model specifications. All statistics weight groups by their size to recover the underlying cross-sectional distributions at the household level. The RMSE-scaled objective function is the square root of the objective function divided by 4 and multiplied by 100 to express it in RMSE-equivalent percentage units. The results in the first row are for the unrestricted model estimated in Table 8. The results in subsequent rows are for models that fix selected parameters at their size-weighted cross-sectional means estimated in the unrestricted model. In the last row, the preference parameters are restricted to only vary with beliefs. There are 4,276 groups containing 298,540 households.

Table IA.14
Size-Weighted Cross-Sectional Regressions of Mean Monte Carlo Preference Parameter Estimates on Indirect Inference Estimates

	(1) RRA	(2) TPR	(3) EIS	(4) Log EIS
Slope coefficient	1.005***	0.917***	0.643***	0.651***
	(0.003)	(0.007)	(0.010)	(0.011)
Constant	-0.030	0.006***	0.374***	0.031***
	(0.023)	(0.001)	(0.013)	(0.009)
R^2	0.936	0.897	0.630	0.636

This table reports cross-sectional regressions of average Monte Carlo estimates of preference parameters on the preference parameters used to generate simulated data, which are set equal to indirect inference parameter estimates for each group. Monte Carlo simulations use the effective group size, the reciprocal of the sum of squared wealth shares of individual households in each group. Average estimates are calculated from 1,000 simulations of each group. All regressions weight groups by their size, to recover the underlying cross-sectional relationships at the household level. Standard errors are reported in parentheses and statistical significance levels are indicated with stars: * denotes 1-5%, ** 0.1-1%, *** less than 0.1% significance. There are 4,276 groups containing 298,540 households.

Table IA.15
Wealth-Weighted Average Preference Parameters

	(1) RRA	(2) TPR (%)	(3) EIS
2000	7.091	2.406	1.207
2001	7.122	2.510	1.196
2002	7.165	2.624	1.188
2003	7.155	2.590	1.188
2004	7.157	2.663	1.180
2005	7.153	2.737	1.178
2006	7.155	2.757	1.176
2007	7.166	2.792	1.172
Mean	7.145	2.635	1.185
Standard deviation	0.026	0.132	0.011

This table reports the wealth-weighted average preference parameters in each year of our panel, as well as their time-series mean and standard deviation. The preference parameters are based on the baseline estimation method. The wealth of each household is measured at the end of each year.

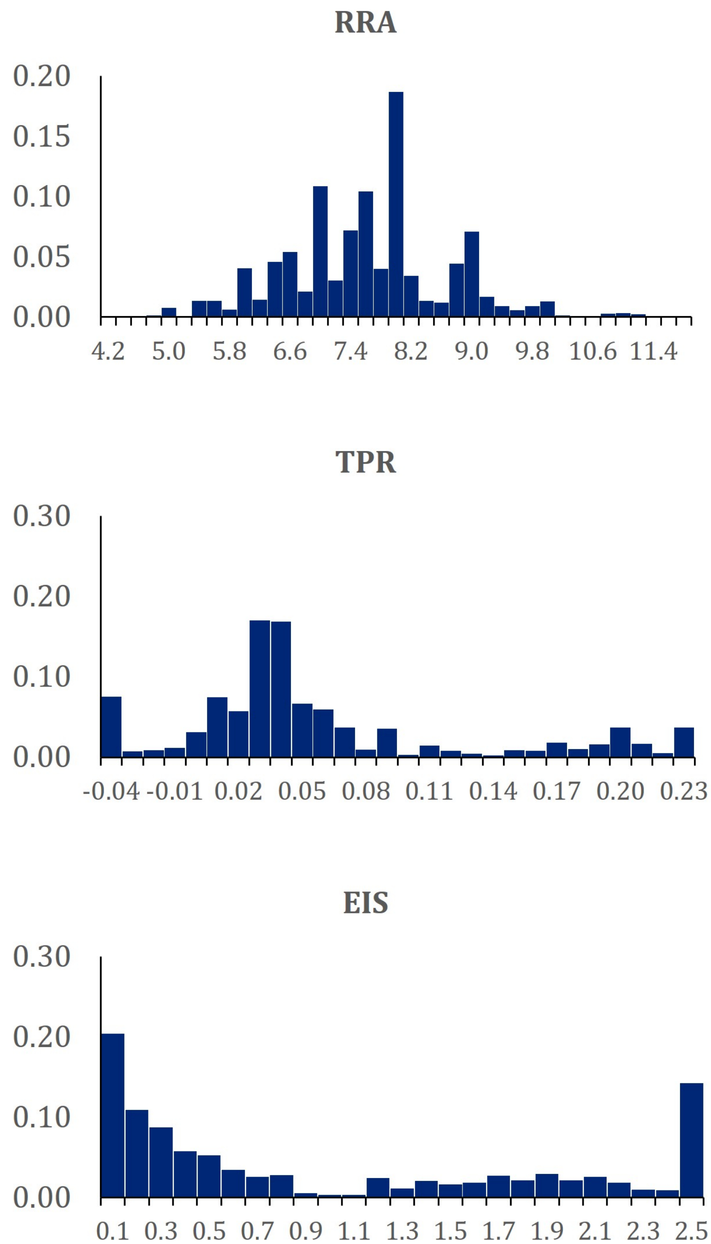
Table IA.16
Adjusting Estimated Parameter Heterogeneity for Estimation Noise

	(1) RRA	(2) TPR (%)	(3) EIS
Unadjusted	1.057	6.96	0.900
Adjusted by winsorized asymptotic standard errors	1.051	5.82	0.545
Adjusted by Monte Carlo standard errors	1.056	6.77	0.782

This table reports the cross-sectional standard deviations of parameter estimates with two alternative adjustments for estimation noise. Unadjusted standard deviations, as in Table 4 of the paper, are reported in the first row. The second row reports standard deviations subtracting the cross-sectional average squared asymptotic standard error, after winsorizing asymptotic standard errors at the 90th percentile of the cross-sectional distribution. The third row reports standard deviations subtracting the cross-sectional average squared Monte Carlo standard error. As discussed in the text, Monte Carlo standard errors are much smaller than asymptotic standard errors but this reflects the limited extent of within-group household-level heterogeneity considered in the Monte Carlo analysis.

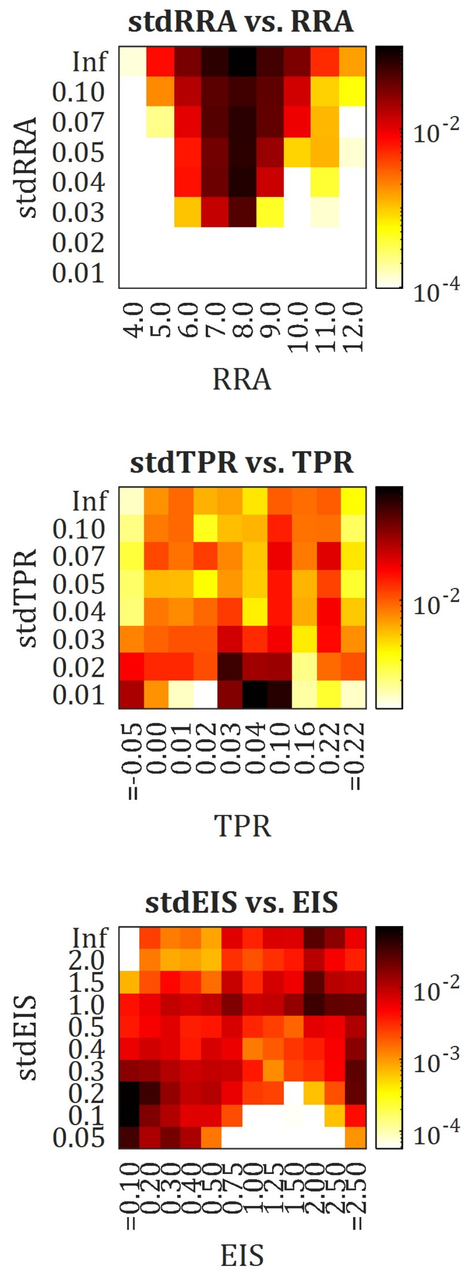
VII. Figures

Figure IA.1. Distribution of Estimated Preference Parameters



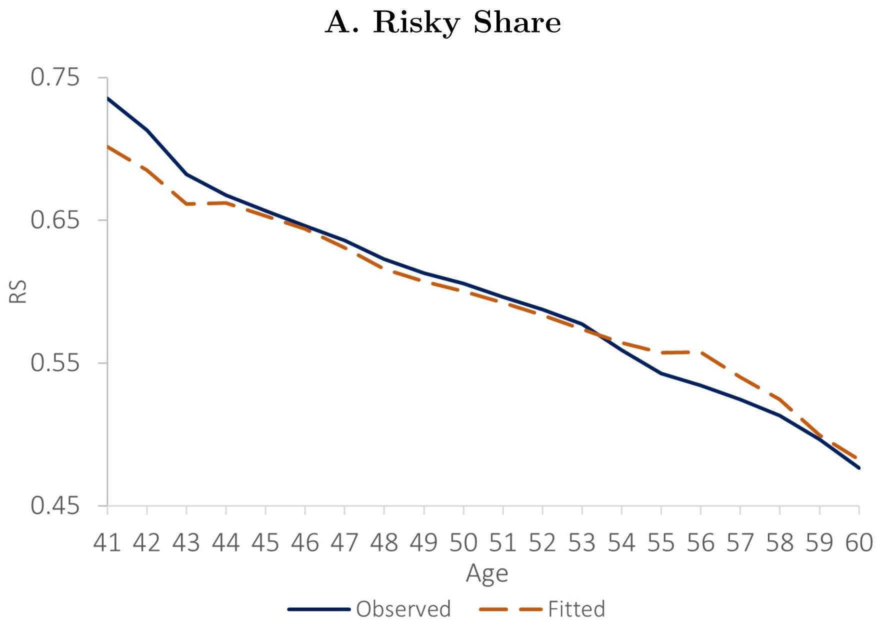
This figure presents histograms for estimates of relative risk aversion (RRA), the time preference rate (TPR), and the elasticity of intertemporal substitution (EIS) across 4,276 groups of Swedish households, size-weighted to recover the underlying distribution across households under the assumption that preferences are homogeneous within groups. Each horizontal axis label shows the upper cutoff value at the right edge of the bin above the label. The vertical axis shows the size-weighted fraction of the sample in each bin.

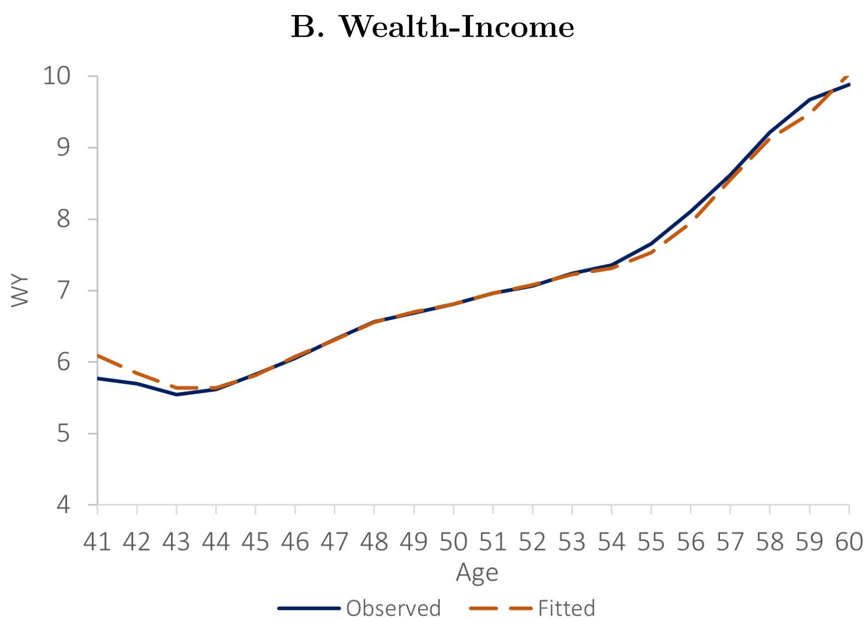
Figure IA.2. Joint Distribution of Estimated Preference Parameters and Respective Standard Errors



This figure presents bivariate heat maps for estimates of RRA and its standard error (top panel), the TPR and its standard error (middle panel), and the EIS and its standard error (bottom panel) across 4,276 groups of Swedish households, size-weighted to recover the underlying distribution across households under the assumption that preferences are homogeneous within groups. Each axis label shows the upper cutoff value of the corresponding bin, except for labels beginning with = which indicate that the bin contains only estimates of the exact value indicated by the label, and the label Inf which indicates that the bin has no upper cutoff but contains all values above the previous bin's cutoff. The logarithmic color scheme indicates the fraction of the sample in each bin. This fraction is equal to 11.0%, 9.7% and 7.0% for the darkest color in the top, middle and bottom panels respectively and 0.0% for the brightest color in all three panels.

Figure IA.3. Empirical and Model-Implied Life-Cycle Profiles





This figure plots the average risky share (Panel A) and wealth-income ratio (Panel B), in the data and in the model, as a function of age. These moments are computed by averaging across all 4276 groups of Swedish households, where each group is weighted by its wealth share.