Macro-Financial Trends in a Model with Concentrated Ownership of Capital

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Abstract

This paper investigates the macroeconomic and asset pricing consequences of the upward trend in financial market participation observed in the U.S. since the late 1980s. In a limited participation two-agent Real Business Cycle model where stockholders feature external habit preferences, higher participation produces higher equity premium and stock market volatility, while reducing the risk-free rate and the standard deviation of aggregate consumption. When coupled with a lower volatility of aggregate shocks, this mechanism helps rationalizing a period characterized by milder aggregate fluctuations but increased perceived risk in asset markets, such as the Great Moderation. I show that these results stem from a novel mechanism whereby an increase in the participation rate improves risk-sharing but raises the representative investor's average risk-aversion, with the latter channel dominating the former. Using household-level data on consumption from the U.S. Consumer Expenditure Survey for the sample 1984-2017, I show that the *risk-aversion channel* is consistent with the empirical evidence.

JEL Codes: D31, E13, E21, E25, E32, E44, G12, G51.

Keywords: Asset Pricing, Limited Participation, Stock Market, Risk-free Rate, Macro-Finance Trends, Production Economies, Great Moderation.

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1 Introduction

Asset pricing models in production economies aim to identify the mechanisms underlying the observed behavior of real and financial variables, and the importance of the macroeconomic environment in shaping risk premia. A large number of studies (Danthine and Donaldson, 2002; Guvenen, 2009; De Graeve et al., 2010; Lansing, 2015, among the others) have explored the potential of the "participation puzzle", i.e. the empirical regularity that only a small fraction of population participates in the stock market (Guiso and Sodini, 2013), to explain asset pricing facts within empirically plausible macroeconomic models. As shown in recent works (Lettau et al., 2018; Greenwald et al., 2019), since the 1980s a great part of the variations in stock prices and risk-premia was driven by redistributive production-factors shares shocks. This evidence points toward the crucial importance of the heterogeneity between workers and shareholders. Remarkably, the percentage of U.S. households accessing financial markets rose dramatically since the late 1980s. While in 1962 only around 19% of families held stocks directly or indirectly (Poterba et al., 1995), since 1989 the degree of stock market participation jumped from 31.6% to around 53% in 2007, and then stabilized at similar levels. Furthermore, from the mid-1980s the U.S. experienced a more stable macroeconomic environment, a phenomenon dubbed "Great Moderation", together with a persistent increase in the equity premium and fall in the risk-free rate (Farhi and Gourio, 2018).

Motivated by this evidence, this work explores the macroeconomic and asset pricing implications of the upward trend in the degree of financial market participation. I employ a Real Business Cycle model with concentrated ownership of capital featuring both technology and distribution shocks where only a fraction of population has full access to financial markets. I assume that the capitalists' habit stock depends on aggregate per capita consumption. Under this assumption, the gap between capitalists' and aggregate per capita consumption plays a crucial role for asset prices, by determining the endogenous effective risk-aversion in the economy. I calibrate the model to match key macroeconomic moments in the 1950-1983 sample, with capitalists representing 20% of total population. Similarly to Lansing (2015), the model produces a sizeable equity premium with a low average risk-aversion, a dividend growth process that is calibrated to match the data and an empirically plausible relative volatility between capitalists' and workers' consumption growth.

In representative-agent consumption-based asset pricing models, macroeconomic risk plays a direct role in the determination of asset prices. A drop in aggregate consumption volatility entails a lower equity premium and a lower Sharpe ratio (Lettau et al., 2007). Furthermore, canonical limited participation models (Guvenen, 2009; Favilukis, 2013;

Lansing, 2015; Morelli, 2021, among the others) predict a negative (positive) relationship between the degree of participation and the average equity premium (risk-free rate). According to the existing literature, the joint fall in aggregate volatility and rise in financial participation should have implied a decline in the risk compensation required by investors on the stock market, which is at odds with the data. In my model economy, an increase in stock market participation generates a sizeable rise in both the equity premium and the volatility of stock returns while reducing the average risk-free rate and the volatility of aggregate consumption. Moreover, higher participation is associated with a moderate increase in the price-dividend ratio. The increase in equity premium and stock volatility produced by the higher participation substantially counteracts the shrinking effect due to the lower aggregate uncertainty associated with the Great Moderation, providing a novel explanation for the U.S. macro-financial phenomena observed over the last three decades.

I show that these results stem from a mechanism whereby higher participation reduces the average investor's risk-tolerance. As participation rises, the representative capitalist's consumption converges to the aggregate per capita consumption. As a consequence, the surplus-consumption ratio, which depends on the gap between the capitalist's consumption and the habit level, shrinks. Since stockholders exhibit external habit utility, this results in an increase in the economy average risk-aversion. In existing limited participation models, the degree of financial market participation affects only the quantity of risk borne by the average stockholder, hence the distribution of consumption is irrelevant for asset prices. Higher participation only improves risk-sharing, reducing the covariance between capitalists' consumption growth and equity returns, which tends to depress the equity premium. This mechanism is also present in the model employed here, but is dominated by the novel *risk-aversion channel*.

To test the model predictions, I employ household-level data on consumption from the U.S. Consumption Expenditure Survey. Following the procedure proposed by Malloy et al. (2009), I construct an annual consumption series for stockholders from 1984 to 2017. I show that the model-implied average risk-aversion for the representative stockholder trended upward over time closely tracking the rate of participation, while the opposite holds for the quantity of risk and the stockholder-to-aggregate consumption ratio. Moreover, I exploit the state-level data available in the survey to document a negative relationship between the degree of participation and the stockholder-to-aggregate consumption ratio in the cross-section of U.S. states. Hence, the empirical analysis supports the main model implications.

This paper builds on the large literature exploring the potential of limited asset market participation to explain several asset pricing puzzles in production economies (Danthine and Donaldson, 2002; Guvenen, 2009; Lansing, 2015, among the others). Unlike most of these studies, I investigate the macroeconomic and asset pricing consequences of the upward trend in asset market participation observed in the U.S. since the 1980s.

In this respect, particularly related to this work are the contributions by Favilukis (2013) and Morelli (2021). My work differs in that the degree of participation affects not only the quantity of risk but also the economy's average risk-aversion. Furthermore, while Favilukis (2013) does not look at the implications of the trend in participation on macroeconomic variables, Morelli (2021) is silent about the implications of higher participation for the risk-free rate. This work also relates to the few studies that look at the asset pricing implications of the Great Moderation (Lettau et al., 2007; Pancrazi, 2014). However, these papers employ asset pricing models where the macroeconomic processes are exogenously specified, and are therefore unable to provide a joint description of the macroeconomy and the stock market, which is instead the main object of study in this work.

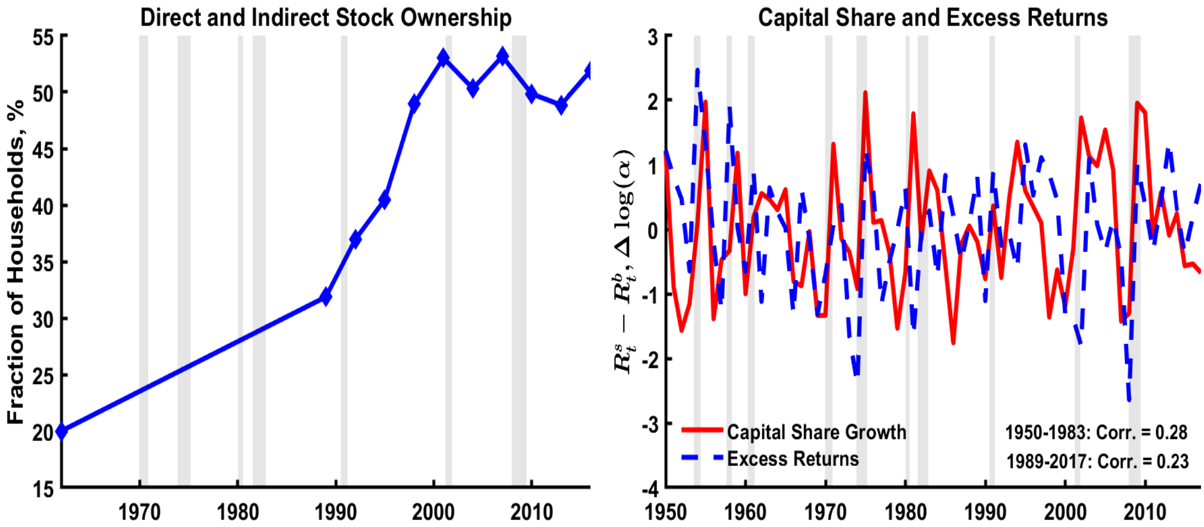
Finally, a growing literature is recently focusing on the role of several structural changes, such as ageing population, rising savings supply and market concentration in shaping the recent U.S. macro-financial phenomena (Caballero et al., 2017; Farhi and Gourio, 2018; Corhay et al., 2020, among the others). For example, Farhi and Gourio (2018) document how over the last thirty years the U.S. experienced a moderate increase in the price-dividend ratio and the equity premium, together with stable average stock returns and a decreasing average risk-free rate. According to the accounting framework proposed, rising market power and effective risk-aversion played a crucial role in determining such trends. My work complements these studies by showing how the increase in financial participation represents another important determinant of the same phenomena, and individuates in the developments of the representative stockholders' surplus-consumption ratio a key driver of the increase in investors' effective risk-aversion.

The rest of the article is organized as follows. In Section 2, I report the empirical stylized facts that motivate this work. Section 3 presents the model setup, while Section 4 discusses the calibration adopted and the quantitative results of the model. Section 5 analyzes the main mechanisms underlying the results, which are then empirically validated and discussed in Section 6. Finally, Section 7 concludes.

2 Some macro-financial facts

From the late 1980s, the U.S. economy experienced a sustained increase in the percentage of households accessing financial markets. Moreover, an extensive literature has documented a remarkable decline in the volatility of main U.S. macroeconomic indicators

Figure 1: Macro-Financial Facts



Notes: Left panel: degree of stock market participation in the U.S. from Poterba et al. (1995) and SCF (2016). Right panel: correlation between (realized) excess returns and capital share growth. In the right panel, the series are normalized and the sample period is 1950-2017. Shaded bands indicate NBER recessions.

over the same decades, together with a falling risk-free rate and a rising equity premium.

Stock market participation The fraction of U.S. households holding stock directly or indirectly rose sharply. As shown in the left panel of Figure 1, in 1962 only around 19% of U.S. families held stocks directly or indirectly (Poterba et al., 1995). According to the data reported in the triennial Survey of Consumer Finances (SCF), since 1989 the degree of stock market participation jumped from 31.6% to around 53% in 2007, and then stabilized at similar levels. In Appendix A, I show that a very similar upward trend can be observed for more refined measures of participation in the stock market.[1]

Macro-financial moments Table 1 displays, in the top panel, the standard deviation of annual growth in real per capita GDP Δy, consumption Δc, investment Δi, total factor productivity Δa, relative price of investment to consumption goods Δp^I, capital share of income $\Delta \alpha$ and macroeconomic dividends[2] Δd over the three sub-samples 1950-

[1]Specifically, I also consider holdings above specified dollar amount thresholds and wealth-weighted participation rates, in the same spirit as Lettau et al. (2018).

[2]The inclusion of macroeconomic dividends, defined as capital income (capital share multiplied by GDP) net of investment, and of the variations in the capital share is motivated by recent works (Lansing, 2015; Lettau et al., 2018; Greenwald et al., 2019) proving both theoretically and empirically how these two variables are crucial determinants of stock returns. Effectively, both variables display a strong correlation (23%) with realized excess returns over the post-war sub-sample. These correlations are remarkably higher than the one exhibited by the S&P500 real dividend growth (10%).

Table 1: Effects of the Great Moderation

	1950-1983	1989-2017	2001-2017
Macroeconomic Variables			
$\sigma(\Delta y)$	3.1	1.54 [−50%***]	1.57 [−50%***]
$\sigma(\Delta c)$	1.65	1.07 [−35%***]	1.06 [−36%**]
$\sigma(\Delta i)$	6.17	4.65 [−25%*]	4.8 [−22%]
$\sigma(\Delta a)$	2.11	1.31 [−38%***]	1.32 [−37%**]
$\sigma(\Delta p^I)$	2.55	1.3 [−49%***]	1.16 [−55%***]
$\sigma(\Delta \alpha)$	2.49	2.42 [−3%]	2.6 [4%]
$\sigma(\Delta d)$	6.69	7.33 [9%]	7.93 [18%]
Financial Variables			
$\sigma(R^b)$	2.38	2.02 [−15%]	1.4 [−41%***]
$\sigma(R^s)$	18.45	17.3 [−6%]	17.77 [−4%]
$E(R^b)$	1.23	0.74 [−39%]	−0.56 [−145%*]
$E(R^{e,s})$	6.17	6.22 [0.9%]	6.09 [−1%]
$E(R^{e,s} - R^b)$	4.52	5.48 [21%***]	6.65 [47%***]

Notes: All values are reported in percent. Lower case letters denote the logarithm of the variables. The numbers in square brackets are the relative variation from the first sample. * = p-value < 0.1, ** = p-value < 0.05, *** = p-value < 0.01. The p-values refer to the F-test for the standard deviations, and to the Chow test for the means. In both cases the null-hypothesis is of no structural break in the moment of interest. The data for the expected return on equity ($R^{e,s}$) is available only from 1961, while the data on the relative price of investment (p^I) is available up to 2016.

1983, 1989-2017 and 2001-2017.[3] The choice of 1983 as a break date for the post-WWII period follows the large literature on the Great Moderation (Stock and Watson, 2002). The choice of the sample 2001-2017 follows instead recent works (Favilukis, 2013; Farhi and Gourio, 2018) studying the macro-financial trends also considered here. The bottom panel reports summary statistics for financial data over the same sub-samples. Namely, the mean and standard deviation of the risk-free rate R^b, the average expected return on equity $R^{e,s}$ together with the volatility of historical stock returns R^s and the (expected) excess returns $E(R^{e,s} - R^b)$.[4]

[3] A detailed description of the data used in this section can be found in Appendix A.

[4] As in Farhi and Gourio (2018), the equity premium is not estimated using historical excess returns, as they are extremely noisy. Thus, using them would be essentially pointless over short periods and detecting changes in the average equity premium over the two sub-samples would be impossible. The series for the expected return on equity is calculated from the version updated through 2019 of Damodaran (2013), and is available only from 1961.

From the late 1980s the volatility of output, consumption, investment, TFP and relative price of investment significantly dropped, with a relative variation around $40\% - 50\%$ (output, TFP and relative price of investment) and $25 - 35\%$ (investment and consumption). To measure the statistical significance of this change, I follow Nason and Smith (2008). The statistic $F_x = \sigma^2(x)_{1950-83}/\sigma^2(x)_{1989-2017}$, i.e. the ratio of the variance of variable x between the pre-Great Moderation and in the Great Moderation period, is distributed under Normality as a $F(T_{1950-83}, T_{1989-2017})$, with T denoting the number of observations in the sample. The p-value, reported in brackets for the second sample, is obtained by locating F_x in this density. The p-values confirm, at standard levels, a significant drop in the standard deviation of all the six variables. Moreover, when repeating the same exercise with the latter sub-period (2001-2017), we can see that very similar conclusions still hold. However, the test does not support a significant change in the standard deviation of capital share growth across sub-samples. Both Table 1 and the right panel of Figure 1 show that capital share growth remained volatile and strongly correlated with equity excess returns throughout the period. Such evidence bears important asset pricing implications. As shown in the model, the steady volatility of capital share growth is part of the explanation for the behavior of asset prices over the last few decades. Finally, the volatility of macroeconomic dividend growth has, if anything, increased during the Great Moderation and in particular over the last 17 years of the sample.

Turning to the second panel, the risk-free rate substantially dropped in terms of both mean and standard deviation especially during the more recent period, with both the Chow-test (for the mean) and the F-test (for the standard deviation) supporting a statistically significant shift. In contrast, the expected return on equity remained essentially unchanged, while the variation in the standard deviation of realized stock returns was moderate and not statistically significant. As a consequence of the significant drop in the risk-free rate and the concurrent stability of expected returns on stocks, the average equity premium increased by almost one percentage point over the Great Moderation period, and by above two percentage points over the last two decades. These results are in line with Delle Monache et al. (2020), who document a recent increase in the average equity premium mostly due to falling interest rates rather than rising stock returns.

It is worth noting that these results are somewhat puzzling. Representative-agent consumption-based asset pricing models predict a positive relationship between the volatility of consumption growth and the equity premium. Under constant relative risk-aversion utility, the Sharpe Ratio can be approximated as

$$SR_t \equiv \frac{E_t\left(R^{e,s}_{t+1} - R^{b}_{t+1}\right)}{\sigma_t\left(R^{r,s}_{t+1}\right)} \approx RRA \times \sigma_t(\Delta c_t)$$

where RRA is the parameter of relative risk-aversion. Therefore, given the essentially stable stock returns volatility, a decline in macroeconomic risk should imply a lower equity premium and a lower Sharpe ratio, as shown by Lettau et al. (2007). The evidence in Table 1 indicates that both the equity premium and the Sharpe Ratio[5] increased, if anything, since the advent of the Great Moderation.[6] Therefore, the behaviour of a traditional macroeconomic fundamental such as consumption volatility is difficult to reconcile with the recent macro-financial empirical evidence. The relevance of the fluctuations in the capital share of income points toward the crucial role of the heterogeneity between workers and shareholders, which is at the core of the limited stock market participation literature. In the next sections I investigate the role played by the upward trend in financial market participation in shaping the recent behavior of the U.S. economy.

3 The model

I employ a RBC model featuring workers, capitalists and competitive firms. Workers, who constitute a fraction γ of the total population (normalized to 1), are assumed to be excluded from the bond and stock markets because of the existence of (un-modeled) transaction costs.[7] Capitalists exhibit external habit preferences, own firms through equity shares and can smooth consumption intertemporally by trading one-period bonds. Both workers and capitalists are assumed to inelastically supply their entire time-endowment to firms and earn the same wage. Moreover, since workers do not price securities, they can in principle have exactly the same preferences as capitalists without affecting the equilibrium conditions. The two agents differ only for their ability to access financial markets, which helps keep the model as simple as possible to clearly identify the effects of changes in the level of stock market participation.

[5] Using the above formula, the Sharpe Ratio increased from 0.245 during the period 1950-1983 to 0.374 during the period 2001-2017.

[6] This result is robust to considering the sample 1989-2007, i.e. excluding the financial crisis, or to using realized in place of expected excess returns.

[7] In a working paper version of his work, Bilbiie (2008) shows that any level of non-participation in financial markets can be supported by the existence of a sufficiently high proportional cost. On the same line, in Appendix B I perform a simple exercise that shows how a reduction in aggregate volatility is not able to explain the increase in asset market participation observed in the data, supporting the assumption that the trend was caused by regulatory reforms that facilitated financial investments for most households and that can be seen as exogenous to business cycle dynamics.

3.1 Workers

Workers are assumed to incur a transaction cost that prevents them from accessing financial markets. Being unable to smooth consumption intertemporally, workers consume their labor income hand-to-mouth, entailing that

$$C_t^w = W_t N_t^w \tag{1}$$

where W_t is the wage and I assume $N_t^w = 1$, i.e. workers do not value leisure and supply their entire time-endowment to firms.

3.2 Capitalists

Capitalists have full access to financial markets. The maximization problem of the individual capitalist reads:

$$\max_{C_t^c, Q_{t+1}^s, Q_{t+1}^b} E_0 \sum_{t=0}^{\infty} \beta^t \frac{(C_t^c - H_t)^{1-\sigma} - 1}{1-\sigma} \tag{2}$$

subject to the constraint

$$C_t^c + P_t^s Q_{t+1}^s + P_t^b Q_{t+1}^b = (P_t^s + D_t) Q_t^s + Q_t^b + W_t N_t^c \tag{3}$$

Labor supply is inelastic, $N_t^c = 1$. The budget constraint states that consumption and purchase of equity shares in quantity Q_{t+1}^s at price P_t^s and of one period bonds in quantity Q_{t+1}^b at price P_t^b must be financed by labor income $W_t N_t^c$ and the returns on their financial investments. Shares purchased in the previous period yield a dividend D_t, while the one period-bonds yield a single consumption unit per-bond in the following period.

I assume that capitalists exhibit external habit utility, where the habit stock H_t evolves according to the law of motion[8]

$$H_t = m H_{t-1} + (1-m) C_{t-1} \tag{4}$$

where C_{t-1} denotes aggregate per capita consumption at time $t-1$. The parameter m allows to introduce a slow-moving component in habit formation, similarly to Campbell and Cochrane (1999), as for $m > 0$ the habit stock does not fully depreciate within the period. On the other hand, the coefficient $(1-m)$ captures how sensitive the reference level is to changes in aggregate per capita consumption. The assumed utility function for capitalists bears crucial implications for the effects of the stock market participation

[8] The specification of the habit as a moving average of past consumption follows Cochrane (2017).

rate on macroeconomic and asset pricing variables, as discussed more in detail later.

The first order conditions to the maximization problem are:

$$\Lambda_t = (C_t^c - H_t)^{-\sigma} \tag{5}$$

$$P_t^s = E_t M_{t,t+1}(P_{t+1}^s + D_{t+1}) \tag{6}$$

$$P_t^b = E_t M_{t,t+1} \tag{7}$$

where Λ_t denotes the Lagrangean multiplier on the budget constraint and $M_{t,t+1} \equiv \beta E_t(\Lambda_{t+1}/\Lambda_t)$ is the capitalist's stochastic discount factor. The F.O.C.s (6) and (7) govern the asset pricing dynamics of the model. In particular, the risk-free rate will be given by $R_{t+1}^b = 1/P_t^b = 1/E_t M_{t,t+1}$, while the stock return is defined as $R_{t+1}^s = \frac{(P_{t+1}^s + D_{t+1}^s)}{P_t^s}$.

3.3 Firms

Firms operate in perfect competition and produce according to the standard Cobb-Douglas technology

$$Y_t = A_t N_t^{1-\alpha_t} K_t^{\alpha_t},\ \alpha_t \in (0,1) \tag{8}$$

where N_t is aggregate employment and the total factor productivity productivity A_t evolves exogenously according to the stationary process

$$\log(A_t) = \rho_a \log(A_{t-1}) + \epsilon_t^a,\ \rho_a \in (0,1) \tag{9}$$

where $\epsilon_t^a \sim N(0, \sigma_a^2)$. As proposed in Lansing (2015), the capital share of income α_t is allowed to fluctuate over time in response to a distribution shock. Specifically

$$\alpha_t = \alpha exp(\nu_t) \tag{10}$$

where α denotes the steady state capital income share and

$$\nu_t = \rho_\nu \nu_{t-1} + \epsilon_t^\nu,\ \rho_\nu \in (0,1) \tag{11}$$

is the distribution shock, which follows a stationary AR(1) process in logs, denoted by lower-case letters, and $\epsilon_t^\nu \sim N(0, \sigma_\nu^2)$. Following Jermann (1998), capital accumulation follows a law of motion featuring capital adjustment costs

$$K_{t+1} = (1-\delta)K_t + \Phi\left(\frac{I_t}{K_t}\right)K_t \tag{12}$$

where δ is the depreciation rate and

$$\Phi\left(\frac{I_t}{K_t}\right) = \left[\frac{a_1}{1-1/\chi_k}\left(\frac{I_t}{K_t}\right)^{1-1/\chi_k} + a_2\right] \tag{13}$$

is the standard concave adjustment cost function. In particular, $\chi_k \to 0$ (∞) implies higher (lower) adjustment costs.

Define dividends as:

$$D_t = Y_t - W_t N_t - \frac{I_t}{P_t^I} \tag{14}$$

Following Greenwood et al. (1988) and Liu et al. (2013), P_t^I is interpreted as the investment-specific technological change, which is assumed to evolve according to the process

$$\log(P_t^I) = \rho_{p^I} \log(P_{t-1}^I) + \epsilon_t^{p^I},\ \rho_{p^I} \in (0,1) \tag{15}$$

where $\epsilon_t^{p^I} \sim N(0, \sigma_{p^I}^2)$ is the investment-specific technology (IST) shock.

The firm's problem is to choose labor, capital and investment to maximize

$$\underset{I_t, N_t, K_{t+1}}{max} E_0 \sum_{t=0}^{\infty} M_{t,t+1}\left\{D_t - Q_t[K_{t+1} - (1-\delta)K_t - \Phi(I_t/K_t)K_t]\right\} \tag{16}$$

subject to the constraints (8), (12) and (13). Q_t is the shadow price of the capital accumulation constraint, equivalent to marginal q.

The first order conditions are given by[9]

$$W_t = (1-\alpha_t)Y_t/N_t \tag{17}$$

implying that dividends can be rewritten as

$$D_t = \alpha_t Y_t - \frac{I_t}{P_t^I} \tag{18}$$

whereas the F.O.C. with respect to investment is

$$\Phi'\left(\frac{I_t}{K_t}\right) = \frac{1}{P_t^I Q_t} \tag{19}$$

[9]Note that capitalists and workers are assumed to be equally productive and therefore earn the same wage. This assumption is quite standard in both macroeconomic and asset pricing literature, see for example Bilbiie (2008), Guvenen (2009) or Debortoli and Galí (2017).

with

$$\Phi'\left(\frac{I_t}{K_t}\right) = a_1 \left(\frac{I_t}{K_t}\right)^{-1/\chi_k} \tag{20}$$

Finally, the firm's optimal decision regarding capital yields

$$Q_t = E_t\left\{M_{t,t+1}\left[\alpha_{t+1}\frac{Y_{t+1}}{K_{t+1}} + Q_{t+1}\left((1-\delta) + \Phi\left(\frac{I_{t+1}}{K_{t+1}}\right) - \Phi'\left(\frac{I_{t+1}}{K_{t+1}}\right)\frac{I_{t+1}}{K_{t+1}}\right)\right]\right\} \tag{21}$$

3.4 Equilibrium

In the competitive equilibrium, all agents take prices as given. The competitive equilibrium in this economy is defined by a sequence of prices and quantities such that the optimality conditions (1), (5), (6), (7), (17), (19) and (21) hold, all constraints are satisfied and all markets clear. More specifically, labor market clearing requires that

$$N_t = \gamma N_t^w + (1-\gamma)N_t^c = 1 \tag{22}$$

while equilibrium in the good market implies

$$Y_t = C_t + I_t \tag{23}$$

where

$$C_t = \gamma C_t^w + (1-\gamma)C_t^c \tag{24}$$

is aggregate per capita consumption. Assuming that the bond market is in zero net supply entails that in equilibrium $Q_t^b = 0$, $\forall t$. Moreover, assuming that the stock market is in unit supply yields the stock market clearing condition

$$(1-\gamma)Q_t^s = 1 \tag{25}$$

where the left hand side is the aggregate demand of stocks, since only a fraction $(1-\gamma)$ of the population participates in the stock market. Therefore, the budget constraint (3) for the individual capitalist reads

$$C_t^c = W_t N_t^c + \frac{D_t}{1-\gamma} \tag{26}$$

in equilibrium. Plugging (1) and (26) into equation (24) yields

$$C_t = \gamma W_t N_t^w + (1-\gamma)\left[W_t N_t^c + \frac{D_t}{1-\gamma}\right] \tag{27}$$

which, given the assumption that both workers and capitalists supply all their time-endowment to firms ($N_t^w = N_t^c = 1$), becomes

$$C_t = W_t + D_t \tag{28}$$

that is, aggregate consumption consists of labour income plus dividends.

3.5 Risk-sharing and risk-aversion channels

It is worth discussing the two key channels through which the rate of stock market participation affects equilibrium dynamics in this economy. On the one hand, higher participation reduces the exposure of the average participant to fluctuations in the stock market, because of improved risk-sharing (*risk-sharing channel*). On the other hand, the degree of participation affects the cross-sectional distribution of consumption and, as a consequence, investors' average effective risk-aversion (*risk-aversion channel*).

Regarding the first channel, equation (25) shows that the number of stock shares held by the representative capitalist is inversely proportional to the degree of participation $1 - \gamma$. As participation rises, the total supply of stocks is diluted over a wider public of market participants, implying that the risk coming from stockholdings is shared among a larger number of investors. This is reflected in the representative capitalist's equilibrium budget constraint (26), where the weight of dividends declines as the fraction $1 - \gamma$ increases. Therefore, investors' consumption becomes less exposed to volatile dividend income. Note that the *risk-sharing channel* is intrinsic to the limited asset market participation structure, and does not derive from other specific assumptions. Indeed, this channel is the one usually explored in the extant literature (Guvenen, 2009; Favilukis, 2013; Lansing, 2015; Morelli, 2021), whereby an increase in participation naturally delivers a decline in the equity premium and a rise in the risk-free rate.

On the other hand, a novel mechanism identified by this work pertains to the relationship between consumption inequality and average risk-aversion. According to the law of motion in equation (4), the habit stock depends on aggregate rather than capitalists' average consumption. Therefore, as in Chan and Kogan (2002) and Kogan et al. (2020), the habit stock can be interpreted as a standard of living in the economy. Under this assumption, the distribution of consumption between capitalists and workers becomes crucial for asset prices, since the gap between the representative capitalist's consumption and aggregate per capita consumption (i.e. the weighted average between the two agents' consumption) directly affects the endogenous effective risk-aversion in the economy.[10] To

[10] Limited participation to financial markets implies that, in equilibrium, the representative capitalist consumes more than the cross-sectional average thus ensuring that the utility function is always well-

clarify this point, notice that in the case of habit utility the relative risk-aversion is given by

$$RRA_t = \frac{\sigma C_t^c}{C_t^c - H_t} = \frac{\sigma}{S_t} \tag{29}$$

where $S_t \equiv \frac{C_t^c - H_t}{C_t^c}$ is the surplus-consumption ratio. As capitalists constitute a higher fraction of population, their level of consumption converges to aggregate (per capita) consumption, which can be seen by comparing equations (26) and (28). This in turn reduces the denominator of equation (29), generating a higher risk aversion as the individual capitalist's consumption is now closer to the habit (or subsistence) level. In other words, higher participation translates into a more equal distribution of consumption, as captured by the declining ratio between capitalist's and average aggregate consumption. In turn, a more equal distribution generates a lower surplus-consumption ratio for the representative capitalist. In contrast with the *risk-sharing channel*, the *risk-aversion channel* entails a positive (negative) relationship between the degree of participation and the average equity premium (risk-free rate).

In Appendix B, I show that if the habit stock depended on capitalists' average consumption (as in Lansing, 2015), the average effective risk-aversion would be independent of the degree of participation. In other words, the *risk-aversion channel* would be shut down. The model would therefore fail to provide a plausible explanation for the behavior of U.S. macroeconomic and financial variables over the last few decades. As shown in Section 4.5, the standard *risk-sharing channel* alone would reinforce the negative (positive) pressure on the equity premium (risk-free rate) due to lower aggregate volatility. Thus, the joint decline in aggregate risk and rise in financial participation would produce counterfactual variations in the main asset pricing moments for the Great Moderation period. In presence of the *risk-aversion channel*, instead, the model is able to rationalize a period characterized by a decline in aggregate risk but increased perceived risk in financial markets.

4 Quantitative results

In this section I present the main quantitative results. The model is solved by using second-order perturbation methods.[11] I show that the model is able to replicate key macro-financial moments for the U.S. in the pre-Great Moderation period and how, in

defined.

[11] Malkhozov (2014) shows that the macroeconomic and asset pricing moments computed with second-order perturbation methods are essentially identical to those obtained with global solution algorithms for a wide set of models, including RBC models with habit utility and capital adjustments costs (as in Jermann, 1998).

this economy, an increase in financial participation raises the equity premium and stock market volatility while reducing the standard deviation of aggregate consumption growth and the mean risk-free rate. I subsequently show that, based on these results, the increase in financial participation played a quantitatively relevant role in shaping the behavior of U.S. macroeconomic and financial variables during the Great Moderation.

4.1 Calibration

To study the implications of an increase in financial market participation starting from an empirically plausible economy, I calibrate the model to match some key macroeconomic and financial moments for the U.S. in the pre-Great Moderation period (1950-1983). Table 2 summarizes the baseline calibration adopted.

A time period in the model is taken to be one year. The fraction of workers γ is set to 0.8, implying a degree of financial market participation of 20%. This is consistent with the evidence reported in Section 2 and with Guvenen (2009). The parameter β is set to 0.9535 to match an average price to dividend ratio around 26, consistent with the S&P500 stock index mean value for the 1950-1983 sample. Given this parameter value, the steady state capital share of income, the capital depreciation rate and the parameters of the adjustment cost function in equation (13) are set to standard values in the Real Business Cycle literature. In particular, $\alpha = 0.37$ consistent with the average value for the 1950-1983 sample and $\delta = 0.115$ implying a yearly depreciation rate of 11.5%. Such combination of β, α and δ delivers a steady state capital-to-output ratio of 2.26, which is in line with the evidence reported in Rios-Rull and Santaeulalia-Llopis (2010) for the post-WWII. Following Jermann (1998), the parameters a_1 and a_2 are constructed so that $\Phi\left(\frac{I}{K}\right) = \delta$, $\left(\frac{I}{K}\right) = \delta$ and $\Phi'\left(\frac{I}{K}\right) = 1$ in steady state. Thus, $a_1 = \delta^{1/\chi_k}$ and $a_2 = \delta - \frac{\delta}{1-1/\chi_k}$.

The persistence of the technology, distribution shock and investment-specific technology (IST) processes are set to $\rho_a = 0.97$, $\rho_\nu = 0.99$ and $\rho_{p^I} = 0.92$. The latter parameter is taken from Justiniano and Primiceri (2008), who estimate it over the post-WWII subsample. On the other hand, the persistence parameter of the distribution shock follows Greenwald et al. (2019), who document how the capital share of income exhibits a very persistent but still stationary process. The combination of parameters ρ_a, ρ_ν and ρ_{p^I} guarantees a weak and positive autocorrelation in both output and consumption growth, as in the data. The calibrated ρ_ν also helps matching a high autocorrelation coefficient and a low volatility for the risk-free rate.

I calibrate the volatility of the IST shock to exactly match the standard deviation of the growth in the relative price of investment over the sample 1950-1983, as reported in Table 1. This is achieved by setting $\sigma_{p^I} = 2.51\%$. Regarding the distribution shock,

Table 2: Baseline parameter values

Description	Parameter	Value
Fraction of Workers	γ	0.8
Discount Rate	β	0.9535
Capital Share of Income	α	0.37
Depreciation Rate	δ	0.115
Technology Persistence	ρ_a	0.97
Distribution Shock Persistence	ρ_ν	0.99
IST Persistence	ρ_{p^I}	0.92
IST Shock Volatility	σ_{p^I}	0.0251
Distribution Shock Volatility	σ_ν	0.04
Technology Shock Volatility	σ_a	0.017
Local Utility Curvature	σ	3
Habit Stock Persistence	m	0.8
Capital Adjustment Cost	χ_k	0.6
Leverage Factor	χ	0.635

Notes: The model is calibrated at the annual frequency.

the volatility $\sigma_\nu = 4\%$ is set to roughly match the empirical volatility of macroeconomic dividend growth of 6.7%. Given these values, the parameters governing the local curvature of the capitalist's utility function, σ, and the persistence of the habit stock process, m, are set jointly with the relative volatility of the technology and distribution shocks to achieve a standard deviation for annual aggregate consumption growth around 1.65%. Specifically, the curvature parameter σ is set to 3. The volatility of the technology shock is set to $\sigma_a = 1.7\%$, implying $\frac{\sigma_\nu}{\sigma_a} = 2.35$.[12] Moreover, $m = 0.8$, similar to Jaccard (2014) for the case of inelastic labor supply. Finally, the degree of capital adjustment costs, governed by χ_k, is set to a standard value of 0.6, achieving a volatility of investment growth of 6.53%, quite close to the data, and a relative volatility between capitalists' and workers' consumption growth around 1.63. This value lies well above the minimum ratio of 1.5 identified by the literature, as reported in Guvenen (2009).

Regarding the financial moments generated by the model, I follow Favilukis (2013). Specifically, in the tables below I report the moments related to the levered equity return. I exploit the relationship $R^{lev} = R^s + \chi(R^s - R^b)$, where χ is the leverage factor. The levered equity return R^{lev} has the same Sharpe Ratio as the stock return R^s, but higher mean and volatility. The parameter χ is set to 0.635, in order to exactly match the standard deviation of realized stock returns.

[12] This ratio is slightly higher than the 1.855 employed in Lansing (2015). This is because the author considers labor augmenting, permanent technology shocks in a model that features also long-run growth. Differently, I consider TFP, transitory technology shocks in a model that abstracts from long-run growth.

As a check of the plausibility of the calibration adopted, it is worth stressing that the model generates a steady state consumption distribution across the two groups of agents that is consistent with the evidence reported in Guvenen (2006). According to the data from the 1989 Panel Study of Income Dynamics (PSID), the bottom 80% of the wealth distribution, a group identified as consumers, contribute almost 70% of total consumption, while the top 20%, identified as investors, concurred to the the remaining 30%. In per capita terms, the author calculates that the average investor consumes only 1.7 times more than the average consumer. In the calibrated model, at the steady state capitalists make up 32% of aggregate consumption $\left(\frac{(1-\gamma)C^c)}{C}\right)$ while the average capitalist consumes 1.87 times more than the average worker $\left(\frac{C^c}{C^w}\right)$.

4.2 Model predictions

The macroeconomic and asset pricing predictions of the baseline model are reported in Table 3 and are directly compared to the pre-Great Moderation data.

The model is able to match the moments targeted in the calibration. The standard deviation of aggregate consumption growth is 1.69%, while investment growth is about 2.46 times more volatile than output growth. The volatility of investment growth generated by the model is 6.53%, which compares well with the empirical 6.17%. On the other hand, the theoretical output growth volatility of 2.65% falls short of the empirical point estimate, but lies well within the 95% confidence interval. The model also yields a relative volatility of capitalists' and workers' aggregate consumption growth of 1.63, in line the empirical evidence reported in Guvenen (2009) suggesting that stockholders' consumption is about $1.5 - 2$ times as volatile as non-stockholders'.

While the standard deviation of the growth in the relative price of investment is matched exactly, the volatility of TFP growth in the model is lower than in the data. To the contrary, the model over-predicts the volatility of the growth in the capital share of income, which is high at 4.01% compared to the empirical 2.49%. However, the 6.85% volatility of dividends in the model provides a very good match with the empirical counterpart. Notice that the volatility of dividend growth is strictly related to wages growth and hence the relative volatility of capitalists' and workers' consumption growth. As shown later, the model delivers a volatility of workers' consumption growth of about 1.9%. Since workers consume only labor income and supply labor inelastically, this volatility corresponds to the volatility of wages growth. Therefore, wages growth is sufficiently smoother than output growth, with a relative volatility of 0.71.[13] The relative smoothness of wages clearly positively impacts on dividends growth volatility, thus helping also to match the

[13]The relative smoothness of wages is a well-established business cycle fact, see also De Graeve et al. (2010), among the others.

Table 3: Pre-Great Moderation: model versus data

Variable	Empirical		Simulated
		Volatility	
Δy	3.1	[2.37,3.68]	2.65
Δc	1.65	[1.22,1.99]	1.69
Δi	6.17	[4.81,7.29]	6.53
Δd	6.69	[5.11,7.97]	6.85
$\Delta \log(A)$	2.11	[1.4,2.64]	1.71
$\Delta \log(P^I)$	2.55	[1.04,3.45]	2.56
$\Delta \log(\alpha)$	2.49	[1.93,2.95]	4.01
$\Delta c^c, \Delta c^w$	$> 1.5 - 2^*$	-	1.63
R^b	2.38	[1.29,3.11]	2.16
R^s	18.45	[13.36,22.42]	18.45
		Persistence: AC(1)	
Δy	0.07	[-0.24,0.38]	0.13
Δc	0.26	[-0.13,0.65]	0.20
Δi	-0.06	[-0.35,0.22]	0.06
Δd	-0.04	[-0.33,0.25]	-0.01
$\Delta \log(A)$	0.06	[-0.16,0.28]	-0.015
$\Delta \log(P^I)$	-0.02	[-0.46,0.41]	-0.04
$\Delta \log(\alpha)$	-0.07	[-0.37,0.23]	-0.005
R^b	0.52	[0.05,0.98]	0.97
R^s	-0.05	[-0.43,0.34]	-0.001
		Mean	
R^b	1.23	[0.41,2.04]	1.69
$E(R^{e,s} - R^b)$	4.52	[3.88,5.16]	4.5

Notes: Results of the baseline model for the pre-Great Moderation period. Lower case letters denote the logarithm of the variable. Moments are all annual. The third column reports, in brackets, the lower and upper bounds of the 95 percent confidence interval around the moment point estimate (second column). Volatility and mean values are in percent. *The range is taken from Guvenen (2009).

ratio $\sigma(\Delta c^c)/\sigma(\Delta c^w)$.

Regarding asset pricing moments, we can see that the calibration perfectly matches the standard deviation of stock returns despite a very low leverage factor ($\chi = 0.635$). The combination of habit utility and high capital adjustments costs helps match this moment (Chen, 2016, 2017). At the same time, as it is well-known in the literature (Jermann, 1998; Jaccard, 2014; Chen, 2017, among the others) these features usually entail excessive fluctuations in the risk-free rate. Nevertheless, the highly persistent distribution shock process counteracts such effect and helps achieve a very plausible value for the volatility of the risk-free rate. The model also produces a sufficiently low average interest rate (1.69% in the model compared to the empirical 1.23%) and generates a mean equity premium of

4.5%, very close to the estimated value of 4.52%.

Finally, the model produces first-order autocorrelation coefficients, AC(1), with the correct sign and of the same order of magnitude as in the data. It is worth noting that for all the variables these coefficients are small and not statistically significant. Overall, the model is able to provide a good fit with the data for the pre-Great Moderation period.

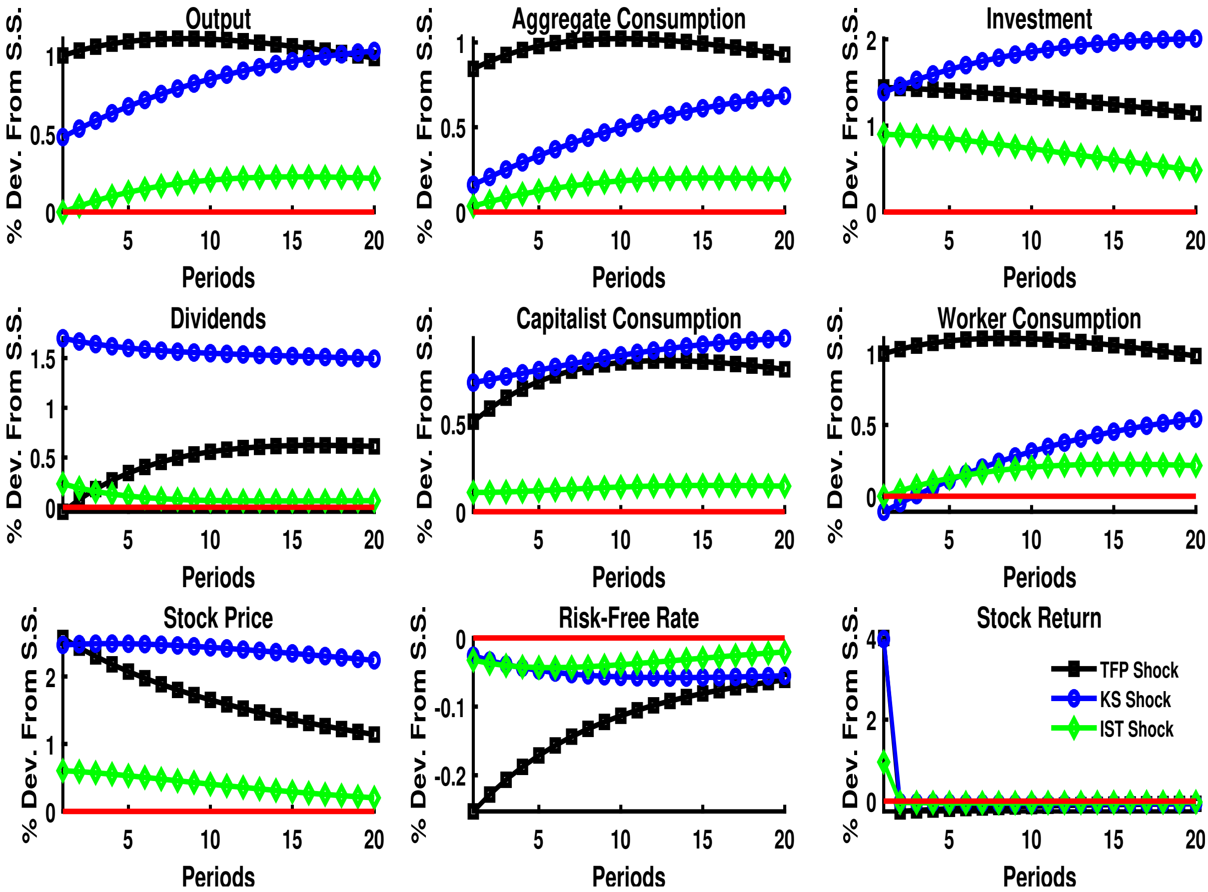
4.3 Impulse response functions

To gain intuition about the model implied dynamics, Figure 2 displays the impulse response function of the main variables to a one-percent technology shock (black-squares line), distribution shock (blue-circles line) and IST shock (green-diamonds line). The IRFs are computed on the logarithm of the variable of interest and the numbers on the y-axis represent relative variations from the non-stochastic steady state (in percent).

A positive neutral and investment-specific technology shock raises all the macroeconomic variables, with investment and aggregate consumption being more and less volatile than output, respectively. Also dividends positively comove with output, although they weakly respond to the shocks. The mild response of dividends, together with capitalists' ability to smooth consumption intertemporally thanks to their access to financial markets, implies that capitalist's consumption increases less than worker's consumption in response to a neutral technology shock. On the other hand, the immediate response of dividends to a IST shock implies that capitalist's consumption reacts more promptly compared to worker's consumption. It is worth noting that the two agents' consumption however comoves positively after a neutral or investment-specific technology shock. Also, notice that the external habit preferences introduce a slight hump-shape in capitalist's consumption IRF. The positive response of dividends in turn raises the price of stocks and their realized return. To the contrary, the risk-free rate is only mildly countercyclical.

Compared to the technology shocks, a distribution shock more strongly affects investment. A distribution shock increases the capital share of income thus making physical capital more productive. Hence, capitalists invest more resources to raise the capital stock which in turn boosts output. Although a strong investment response tends to shrink dividends, the joint increase in the capital share of income and in output still allows a strong procyclical response in the dividends accruing to capital owners (recall equation (18)). This helps matching a high volatility of capitalists' consumption relative to workers', since the latter only rely on labor income. Moreover, unlike the technology shocks, a distribution shock redistributes resources away from workers in favour of capitalists making their consumption comove negatively at impact. This negative comovement, together with the fact that capitalists constitute a small fraction of total population, entails a mild

Figure 2: Impulse response functions



Notes: IRFs to a one-percent technology shock (black-squares line), distribution shock (blue-circles line) and IST shock (green-diamonds line). The IRFs are computed on the logarithm of the variables and are generated with a first-order approximation of the model. Numbers on the y-axis therefore represent relative deviations from the non-stochastic steady state (in percent).

response in aggregate consumption. However, in the medium term the positive effect on output positively influence workers' consumption as well, although only slightly.

From an asset pricing perspective, the distribution shock makes stock prices and realized stock returns strongly procyclical and volatile, while the response of the risk-free rate is more muted but persistent. Taken together, these results are well in line with those obtained by Lansing (2015) and clarify how distribution shocks can help matching high and volatile stock returns while generating a sufficiently smooth aggregate consumption process.

4.4 Stock market participation and macro-financial trends

According to Figure 1, before the late 1980s, only 20% of households held stocks directly or indirectly. This participation rate was employed in the baseline calibration of the model. Since 1989, as reported in the Survey of Consumer Finances, the participation rate trended upward until the early 2000s', when it stabilized around a value of 50%. Therefore, to analyze the effects of increasing asset market participation on macroeconomic and financial moments, I vary the fraction of population represented by capitalists within the interval $[0.2, 0.5]$.

Macroeconomic effects Figure 3 shows heterogeneous effects on the main macroeconomic variables. An increase in the degree of financial participation has negligible effects on output growth and workers' consumption growth volatility, which are almost exclusively determined by the size of the technology and distribution shocks given the assumption of inelastic labor supply. Increasing participation from 20% to 50% raises output (wage) volatility by about 5 basis point(s). On the other hand, the impact on the standard deviation of aggregate consumption and investment growth is remarkable. From a quantitative perspective, increasing market participation from 20% to 50% reduces the volatility of aggregate consumption from 1.69% to about 1.32%, i.e. by about 22%. Qualitatively, the sign of the effect is quite intuitive. It seems reasonable to expect that higher degrees of access to financial instruments should improve the economy's ability to smooth consumption intertemporally.[14] In contrast, higher fractions of capitalists are associated with higher levels of investment growth volatility.

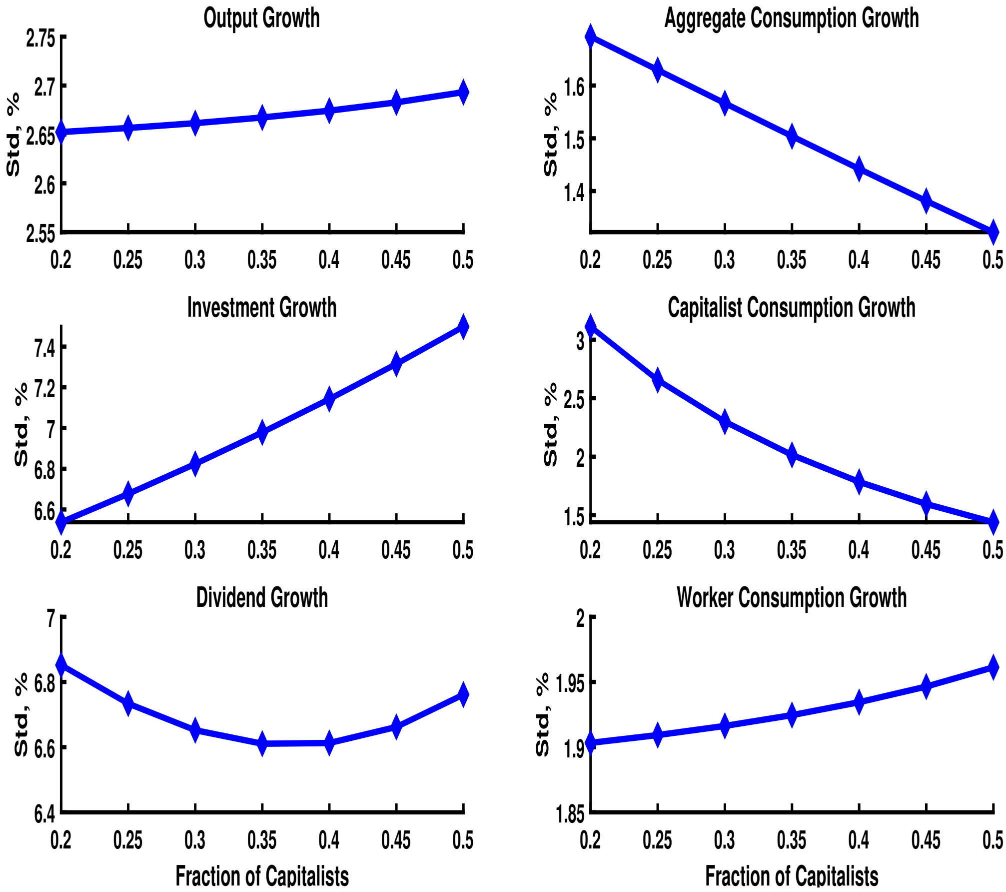
These results stem from the utility specification for capitalists and the *risk-aversion channel* discussed in Section 3.5. Note that the elasticity of intertemporal substitution (EIS) is given by the inverse of the relative risk aversion:

$$EIS_t = \frac{1}{RRA_t} = \frac{S_t}{\sigma} \tag{30}$$

An increase in participation therefore determines a decline in the average EIS, which translates into a stronger consumption smoothing motive for investors. As shown in Figure 3, the decline in the EIS and the lower reliance on volatile dividend income imply that the standard deviation of capitalists' consumption growth drops, and this pattern is inherited by aggregate consumption. On the other hand, investment is used more aggressively to smooth consumption over time, which explains the increase in its volatility.

[14] A similar interpretation has been provided by earlier works on the topic (Blanchard and Simon, 2001; Stock and Watson, 2003; Campbell and Hercowitz, 2006, among the others). For a critical view on the relevance of the financial liberalization process as a driver of the Great Moderation, see Den Haan and Sterk (2011).

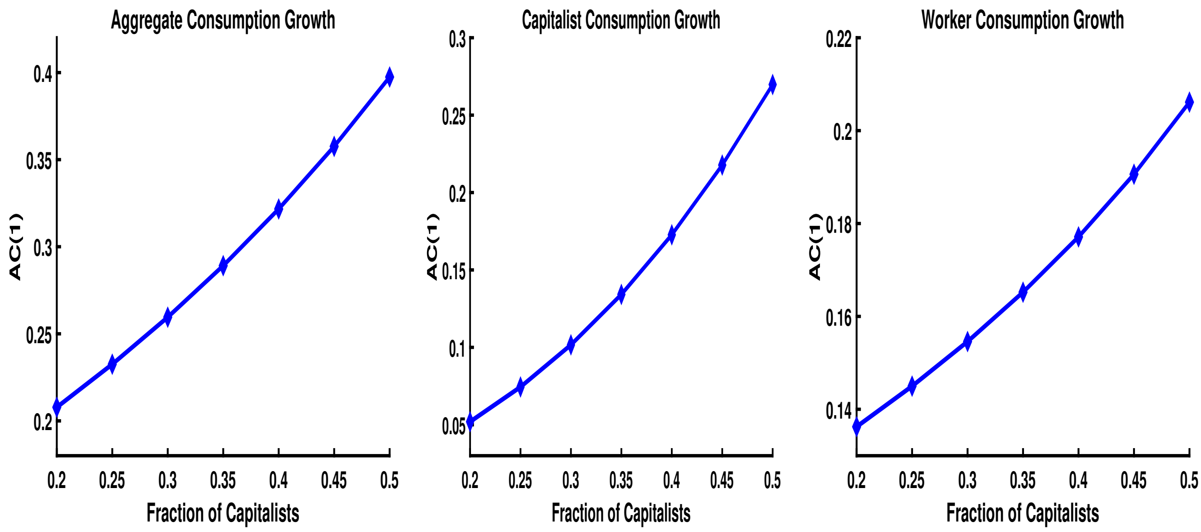
Figure 3: Macroeconomic effects - Standard deviations



Notes: All standard deviations are reported in percent. The degree of asset market participation, captured by the fraction of capitalists, is discretely varied between 20 and 50 percent.

The stronger consumption smoothing motive is also reflected into the degree of persistence of aggregate and agent specific consumption, captured by the first-order autocorrelation coefficient. As displayed in Figure 4, the increase in participation significantly raises not only the persistence of the representative capitalist's consumption growth (from 0.05 to 0.27), but also that of the representative worker (from 0.13 to 0.20). To smooth consumption intertemporally, investors require a more uniform path of capital over time, which influences wages in the same direction. Both forces clearly affect the aggregate consumption growth autocorrelation structure, with the AC(1) coefficient increasing from 0.21 to 0.4. Hence, the *risk-aversion channel* endogenizes the two mechanisms stressed in Pancrazi (2014). Namely, the "variance effect" due to lower consumption growth volatility, which tends to reduce the equity premium and to increase the risk-free rate; and the "autocorrelation effect" related to the larger autocorrelation coefficients, which has the opposite effect on asset pricing variables. As discussed below, the second effect prevails

Figure 4: Macroeconomic effects - Autocorrelations

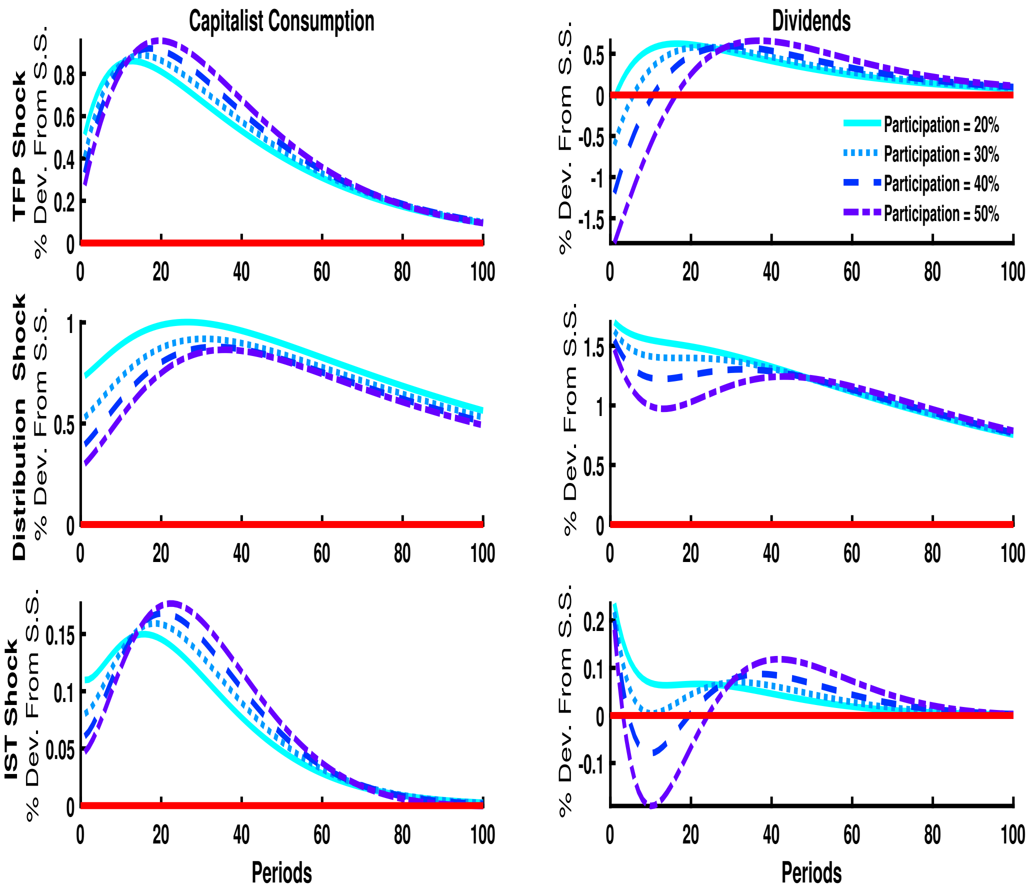


Notes: The degree of asset market participation, captured by the fraction of capitalists, is discretely varied between 20 and 50 percent.

here.

This mechanism can be alternatively gauged by inspecting Figure 5, which reports the IRFs of capitalist's consumption and dividends to both neutral and investement-specific technology shocks (TFP and IST, respectively) and distribution shocks for different levels of market participation. Darker lines denote higher degrees of participation. Regarding capitalist's consumption (left column), as financial participation increases two opposite effects are at work. First, the impact response becomes less sizeable, implying a lower volatility of their realized consumption growth. Second, reductions in the immediate response generates higher volatility of future expected consumption, because the response becomes more hump-shaped. Therefore, future expected marginal utility of consumption becomes more volatile relative to the current realization, which translates into a higher volatility of the stochastic discount factor. As highlighted by equations (19) and (21), investment volatility is thus affected as well. The volatility of investment also affects the behavior of dividends, defined as capital income net of investment according to equation (18). According to Figure 3, the standard deviation of dividend growth (bottom-left panel) is overall only mildly affected. This result is strictly linked to the different impact of technology and distribution shocks on dividends. To see this point, consider again Figure 5. While the KS and IST shocks always increase dividends at impact, the sign of the response to a TFP shock depends on the level of participation. For higher levels of participation investment is used more aggressively in response to TFP shocks in order to smooth consumption, eventually reverting the sign of the immediate dividends' response.

Figure 5: Capitalist's consumption and dividends (IRFs)

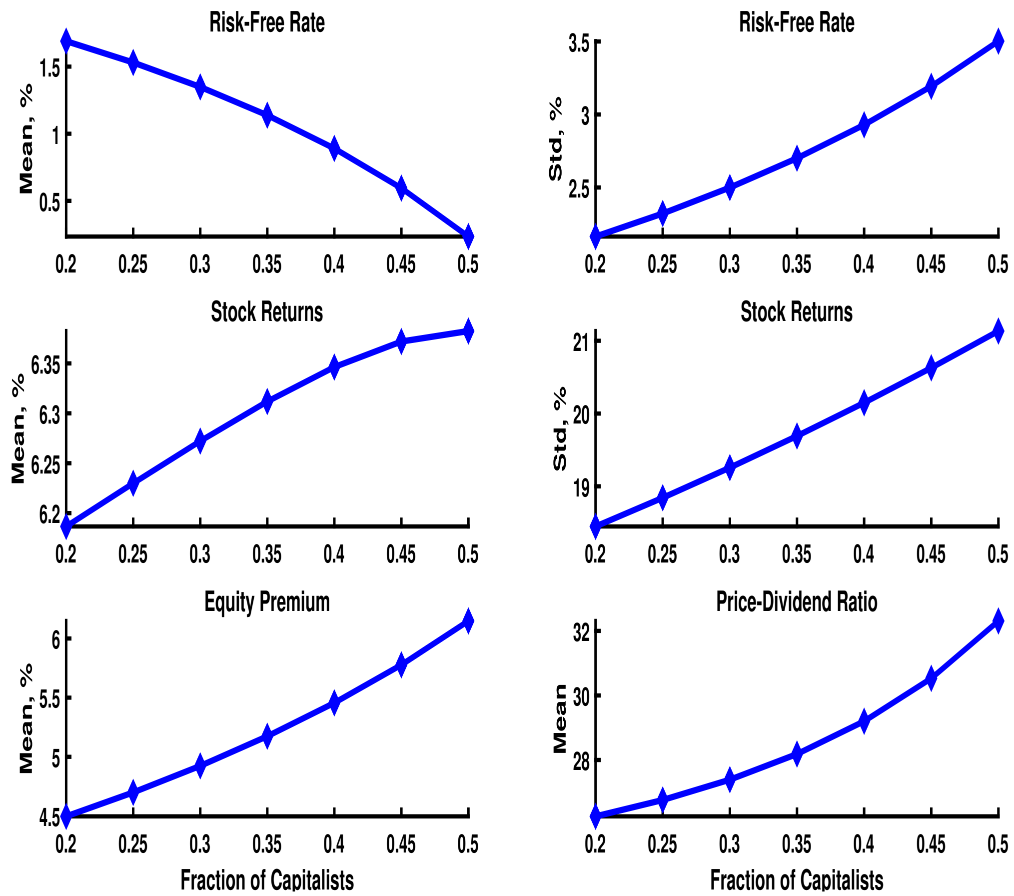


Notes: IRFs of capitalist's consumption and dividends to a one-standard deviation neutral technology shock (TFP, top panels), distribution shock (middle panels) and investment-specific technology shock (IST, bottom panels) for different levels of financial participation. Darker lines denote higher levels of participation.

The opposite sign of the response to the different shocks rationalizes the behavior of dividends volatility.

Asset pricing effects The *risk-aversion channel* drives the asset pricing effects of higher participation too. The increase in the average-risk aversion lowers the risk-free rate and produces an increase in the mean equity premium. Similarly, the associated decline in the EIS tends to raise the volatility of both risk-free rate and stock returns (Chen, 2017). The changes are quantitatively non-negligible. According to Figure 6, an increase in the fraction of capitalists from 20% to 50% raises the equity premium from 4.5% to about 6.15%, while increasing the volatility of stock returns from 18.45% to above 21.13%. Conversely, the risk-free rate drops from 1.69% to 0.23%, i.e. by about 86%. Furthermore, an interesting result pertains to the price-dividend ratio. As we can see in

Figure 6: Asset pricing effects



Notes: All moments are reported in percent, except for the price-dividend ratio. The degree of asset market participation, captured by the fraction of capitalists, is discretely varied between 20 and 50 percent.

the bottom-left panel of Figure 6, an increase in participation raises it from 26 to 32, i.e. by about 23%.

These findings suggest that easier access to financial markets was one of the factors that contributed significantly to the run-up in stock prices relative to fundamentals observed since the 1990s, complementing the increasing importance of distribution shocks supported by Greenwald et al. (2019). Higher participation is also able to account for a strong fall in the average risk-free rate, as documented by Del Negro et al. (2019) and Delle Monache et al. (2020). Interestingly, according to Figure 6 average stock returns only mildly increase as the degree of participation rises. Therefore, consistent with Table 1 and the findings by Farhi and Gourio (2018) and Delle Monache et al. (2020), the increase in the average equity premium is almost entirely driven by a decline in the safe asset returns. The model provides a novel explanation for the asset pricing trends documented in Farhi and Gourio (2018), who also find a moderate increase in the price-dividend ratio together with an increasing equity premium and a secular decrease in the average risk-free rate over the last thirty years.

These results are novel. Standard models with concentrated ownership of capital, which do not feature the *risk-aversion channel*, predict that an increase in financial participation counterfactually lowers the equity premium and stock market volatility while raising the standard deviation of aggregate consumption growth and the mean level of the risk-free rate.[15] As shown in the next section, these models fail to provide a plausible explanation for the behavior of U.S. macroeconomic and financial variables over the last few decades.

4.5 Application to the Great Moderation

The results presented so far highlight the relevance of a secular increase in financial participation in shaping the macroeconomic volatility and the stock market. In particular, a realistic increase in the degree of participation from 20% to 50% is able to account for a decrease in aggregate consumption growth volatility and an increase in the equity premium of the same order of magnitude as in the data reported in Table 1. Moreover, these tendencies are shown to be consistent with a downward trend in the mean risk-free rate and a run-up in stock prices relative to fundamentals. As documented in Section 2, since the mid-1980s the U.S. experienced not only an upward trend in financial market participation but also a structural break in the volatility of main macroeconomic indicators. Arguably the Great Moderation reflected "good luck", insofar the volatility of the shocks hitting the economy substantially dropped after 1984 (Stock and Watson, 2002).[16]

Therefore, to provide a description of the last 30 years together with a quantitative intuition of the historical relevance of the observed trend in financial participation, I simulate the model with a lower volatility of the exogenous shocks and a high level of participation. In other words, similarly to Farhi and Gourio (2018), I perform a steady state analysis which compares the pre-Great Moderation period, interpreted as a regime with low participation and high volatility of exogenous shocks, and the Great Moderation period, characterized by higher participation and lower aggregate risk. More specifically, the level of participation is increased to 45%, as an approximation of the average rate for the second sub-sample. I then reduce the volatility of the technology and distribution shocks in order to exactly match the relative variation in the standard deviation of TFP,

[15] Details about this version of the model, including the calibration and the predictions for the pre-Great Moderation period, can be found in Appendix B.

[16] Other explanations include improvements in the monetary policy conduct and smaller dependence on oil, which are out of the scope of this paper.

capital share and relative price of investment growth from the pre to the Great Moderation period.[17]

I also compare the results with those that would be obtained in a counterfactual regime characterized by low volatility but no increase in financial participation, i.e. keeping the fraction of capitalists at 20%. The results for this regime will isolate the effects due only to the reduced aggregate risk. Moreover, I consider an alternative scenario where the participation rate is raised to 45% but the *risk-aversion channel* discussed in Section 3.5 is shut down. This exercise is therefore apt to provide a measure of the quantitative relevance of higher participation and of the crucial role played by the novel *risk-aversion channel* to generate predictions that are in line with the data. As a caveat, it is worth stressing that this analysis is however not meant to match all the moments considered. Many other factors that are identified by the literature as relevant for the analysis of the period under scrutiny, such as shifts in the monetary policy conduct (Bilbiie, 2008; Justiniano and Primiceri, 2008), easier credit access (Jensen et al., 2018, 2020) or increasing market power and savings supply (Farhi and Gourio, 2018) are neglected here.

A decrease in the volatility of exogenous shocks would reduce the overall aggregate risk in the economy, thus shrinking both the volatility of all the variables and the equity premium, while substantially increasing the risk-free rate. For simplicity, and without loss of generality, consider a canonical representative-agent model with power utility and neutral technology shocks only. Under these assumptions, Lettau (2003) shows that the (log) equity premium can be written as[18]

$$r_t^{e,s} - r_t^b = \sigma \times (\eta_{ca}\eta_{ra})\, \sigma_a^2$$

where η_{ca} and η_{ra} are, respectively, the elasticity of consumption and the stock return to technology shocks, and depend only on the deep parameters of the model. Figure 2 clarified how these elasticities are positive, since both consumption and stock returns react positively to a TFP shock. The impulse response functions suggest that the elasticities to IST and redistributive shocks are positive too. Therefore, a decrease in the variance of the exogenous shocks reduces both aggregate volatility and the equity premium in the model, while implying an increase in the average risk-free rate. Through the *risk-sharing channel*, an increase in financial participation reinforces these effects on asset prices. Better risk-sharing implies that the average investor is less exposed to the stock market and would therefore require a lower compensation for holding stocks. On the other hand, in presence

[17] As Justiniano and Primiceri (2008) point out, changes in the autocorrelation structure of the exogenous shocks would not be able to explain the volatility shift observed during the Great Moderation. Therefore, in this exercise I only focus on the decline in the standard deviation of the shocks.

[18] Lower-case letters denote the logarithm of the variable of interest.

Table 4: Relative variation from pre-GM to GM period

	Empirical	Simulated		
		$1-\gamma = 20\%$ (Baseline)	$1-\gamma = 45\%$ (No RA channel)	$1-\gamma = 45\%$ (Baseline)
Volatility				
Δy	-50% [3.1→1.54]	-16% [2.65→2.23]	-16% [2.68→2.25]	-15% [2.65→2.25]
Δc	-35% [1.65→1.07]	-28% [1.69→1.22]	-24% [1.72→1.30]	-38% [1.69→1.04]
Δi	-25% [6.17→4.65]	-12% [6.53→5.77]	-21% [6.77→5.38]	-4% [6.53→6.3]
Δd	9% [6.69→7.33]	-3% [6.85→6.62]	28% [6.15→7.85]	-10% [6.85→6.13]
R^s	-6% [18.45→17.3]	-8% [18.45→16.93]	-18% [18.45→15.09]	-0.4% [18.45→18.37]
R^b	-15% [2.38→2.02]	-13% [2.16→1.88]	-35% [2.88→1.86]	23% [2.16→2.65]
Mean				
R^b	-39% [1.23→0.74]	20% [1.69→2.03]	100% [1.46→2.93]	-32% [1.69→1.15]
$E(R^{e,s} - R^b)$	21% [4.52→5.48]	-13% [4.5→3.91]	-41% [4.42→2.59]	7% [4.5→4.83]

Notes: The empirical variation from the 1950-1983 (pre-GM) to the 1989-2017 (GM) period is compared to the variation obtained by simulating the model with unchanged (third column) or increased (last column) participation in the baseline model, or increased participation but in the alternative model where the *risk-aversion channel* (RA channel) is shut down (fourth column). In brackets is reported the shift (indicated by the arrow) in the moments from the first to the second sub-sample. The empirical relative variation in the volatility of TFP (-38%), capital share (-3%) and relative price of investment (-49%) growth is exactly matched in all simulated models by construction.

of the *risk-aversion channel* the increase in financial participation brings about a lower risk-tolerance by investors, because of a decline in the gap between their consumption and the subsistence (habit) level of consumption. This channel plays an opposite pressure on asset prices, and helps rationalize the observed financial trends despite the decline in aggregate fluctuations.

Table 4 reports the results of this quantitative exercise. Recall that, by construction, the relative variation in the volatility of TFP, capital share and relative price of investment growth is exactly matched in all simulations.[19] The third column isolates the effects of lower aggregate volatility, as the participation rate is kept at the baseline value. Consistent with the reasoning above, in this scenario the standard deviation of all variable declines. Moreover, lower aggregate risk produces an increase in the average risk-free rate, while the average equity premium drops. The fourth column of the table highlights the effects of the *risk-sharing* channel. When the *risk-aversion channel* is shut

[19] This is achieved by reducing σ_a from 1.7% to 1.054%, σ_ν from 4% to 3.88% and σ_{p^I} from 2.51% to 1.28%.

down, the rise in financial participation reinforces the effects of reduced aggregate risk on asset prices. Under this scenario, the volatility of asset returns decreases remarkably more than in the data, and the equity premium (risk-free rate) declines by a striking 41% (doubles), which is completely at odds with the empirical evidence.

Finally, the last column shows the crucial role played by the novel *risk-aversion channel* in generating model predictions that are in line with the data. The baseline model with high participation generates a remarkable 15% decrease in output growth volatility and an empirically plausible 38% decrease in consumption growth standard deviation, together with a 4% decline for investment growth.[20] Crucially, these results are accompanied by an increase in the average equity premium and decline in the average risk-free rate that are of the same order of magnitude as in the data. Taken together, these results point to a quantitatively relevant impact of the trend in financial participation as one of the factors contributing to the behavior of macroeconomic and financial variables over the last 30 years.

5 Inspecting the mechanism

The results of the counterfactual exercises shed light on the key role played by the novel *risk-aversion channel* to rationalize a period of lower aggregate volatility but increased perceived risk on financial markets, such as the Great Moderation. In this Section, I show that the baseline model predicts, as a result of more widespread participation, a reduction in the quantity of risk borne by stockholders consistent with improved risk-sharing (*risk-sharing channel*), but at the same time a rise in average risk-aversion (*risk-aversion channel*). This is because capitalists' consumption converges to aggregate per-capita consumption, which in turn brings the former closer to their habit (or subsistence) consumption level. Standard analytical decompositions exemplify how these two channels exert opposite forces on both risky and safe asset returns. In the model economy, the *risk-aversion channel* dominates and produces an overall positive (negative) relationship between the degree of participation and the equity premium (risk-free rate). Moreover, I demonstrate how the *risk-aversion channel* helps generating an increase in the volatility of asset returns as participation rises, which explains why the volatility of stock returns and the risk-free rate remained essentially stable during the Great Moderation despite a reduction in aggregate risk.

To obtain closed form solutions and to highlight the main mechanisms at work, in the

[20]The decline in output and investment volatility falls short of the empirical counterpart. This is strictly linked to the effects of increasing participation studied in Figure 3. In presence of the *risk-aversion channel*, higher participation weakens the effect of lower aggregate risk on investment growth volatility. The result for output growth is instead independent of the level of participation.

rest of the analysis I consider the simple case where capitalists have CRRA preferences.[21] Appendix C discusses additional analytical results, with particular focus on how distribution shocks help generate a sufficiently high relative consumption volatility between the capitalists and workers.

5.1 Participation and average asset returns

I first study the impact of asset market participation on the (unconditional) average equity premium and risk-free rate. In the power utility, log-linear framework adopted here the log-the expected excess stock return over the risk-free rate is given by[22]

$$E(r^{e,s} - r^b) = RRA \times Cov(\Delta c^c, r^s) \tag{31}$$

while the log-risk-free rate is

$$E(r^b) = RRA \times E(\Delta c^c) - \frac{1}{2}RRA^2 \times Var(\Delta c^c) - \log\beta \tag{32}$$

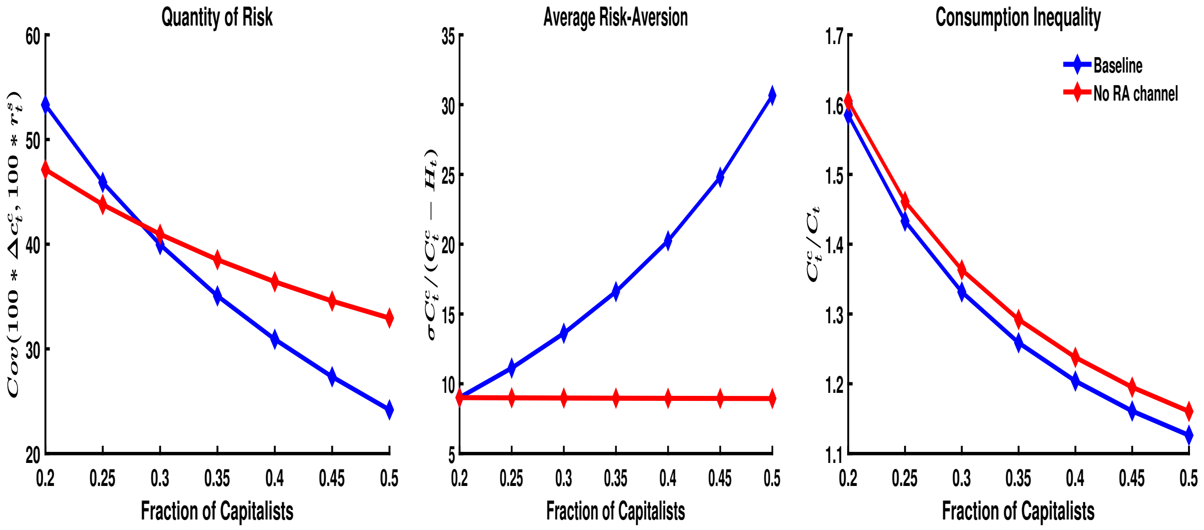
Equation (31) shows that the average equity premium depends on both the relative risk aversion, RRA, and the quantity of risk, $Cov(\Delta c^c, r^s)$. In the baseline model, the degree of participation affects both components of the equity premium. As shown in the left panel of Figure 7, the *risk-sharing channel* implies that higher participation rates are associated with a lower covariance between capitalist's consumption growth and stock returns. Quantitatively, the covariance between capitalist's consumption growth and stock returns is halved when participation increases from 20% to 50%. Notice that this channel is also present in the version of the model where the *risk-aversion channel* is shut down (red line). Therefore, standard models with limited participation predict a negative relationship between participation rates and average equity premium.

However, in presence of the *risk-aversion channel* this effect is overturned, as the reduction in the stockholder-to-aggregate consumption ratio raises the average relative risk-aversion, RRA. According to the middle-panel, in the baseline model the average effective risk-aversion strongly rises, while the same variable is independent of participation when this mechanism is not present. In the model, this channel dominates and produces an overall strong increase in the equity premium. This pattern is strictly linked to the assumed law of motion in equation (4), where the habit stock depends on aggregate rather

[21] Preserving the assumption of habit utility would complicate the derivations without altering the main mechanism under analysis.

[22] To keep notation simple, I omit the Jensen's Inequality terms. The expressions for the average equity premium and risk-free rate in the log-linear framework are standard, see for example Campbell (2003) and Lettau (2003).

Figure 7: Inspecting the mechanism



Notes: This figure shows the model implied: covariance between capitalist's consumption growth and stock returns (left panel); implied average risk-aversion (middle panel); and capitalist-aggregate consumption ratio (right panel), as a function of the fraction of capitalists. These variables are reported for the baseline model (blue lines) and the alternative model where the *risk-aversion channel* is shut down (red lines).

than capitalists' average consumption. Under this assumption, the gap between the representative capitalist's consumption and aggregate per capita consumption is a crucial determinant of asset prices since it directly affects the endogenous effective risk-aversion in the economy. Indeed, the right panel of Figure 7 shows that the ratio between capitalists' and aggregate consumption declines from above 1.5 to below 1.2 when half population participates in asset markets, and this happens also in the alternative standard economy. Nevertheless, only in the baseline model does the drop in the surplus-consumption ratio determine a strong rise in the average risk-aversion, whereas it leaves the same variable unaltered in the alternative economy where the distribution of consumption does not affect equilibrium allocations and prices.

A similar reasoning can be applied to the average risk-free rate, which according to Equation (32) depends positively on the average representative capitalist's consumption growth rate and negatively on its variance. Given the transitory nature of the exogenous shocks, the theoretical economy does not feature trend growth, implying that $E(\Delta c^c) = 0$. Moreover, the discount factor parameter β is fixed and independent of the level of participation. Thus, in the model economy the participation rate only affects the average risk-free rate through the variance term, which enters equation (32) with the negative sign. An increase in participation exerts two opposite forces on the safe asset return. On the one hand, the *risk-sharing channel* tends to increase the risk-free rate. As the fraction of capitalists rises, volatile dividend income weighs less on the representative investor's budget constraint (recall equation (26)). As a consequence, $Var(\Delta c^c)$ declines (as shown in Figure 3) and reduces the precautionary savings motive, exerting a positive pressure on the risk-free rate. As before, the *risk-aversion channel* dominates over the *risk-sharing channel*, thereby reducing the safe asset average returns.

5.2 Participation and asset returns volatility

How does higher participation affect the volatility of stock returns? To show this, I follow Campbell and Shiller (1988). Notice that unexpected log-returns on stocks can be written in terms of revisions in expected future dividends and returns

$$r^s_{t+1} - E_t r^s_{t+1} = (E_{t+1} - E_t) \sum_{j=0}^{\infty} \rho^j \Delta d_{t+1+j} - (E_{t+1} - E_t) \sum_{j=1}^{\infty} \rho^j r^s_{t+1+j} \tag{33}$$

where ρ is a linearization constant smaller but close to 1.

For simplicity, assume that capitalists finance their consumption only with dividend income, i.e.

$$C^c_t = \frac{D_t}{1-\gamma}$$

which in logs, and ignoring constants, becomes

$$c^c_t = d_t \tag{34}$$

It can be shown that the Euler equation, in the case of CRRA preferences, can be log-linearized as

$$RRA \times E_t(\Delta c^c_{t+1}) = E_t(r^s_{t+1}) \tag{35}$$

where RRA is the parameter of relative risk-aversion. Thus, substituting (34) and (35) into (33), I obtain

$$r^s_{t+1} - E_t r^s_{t+1} = (E_{t+1} - E_t) \sum_{j=0}^{\infty} \rho^j \Delta c^c_{t+1+j} - RRA(E_{t+1} - E_t) \sum_{j=1}^{\infty} \rho^j \Delta c^c_{t+1+j} \tag{36}$$

where notice the first summation starts from $j = 0$ while the second one from $j = 1$. Hence

$$r^s_{t+1} - E_t r^s_{t+1} = \underbrace{(\Delta c^c_{t+1} - E_t \Delta c^c_{t+1})}_{\equiv R^{t,t+1}_{\Delta c^c}} + (1 - RRA) \underbrace{(E_{t+1} - E_t) \sum_{j=1}^{\infty} \rho^j \Delta c^c_{t+1+j}}_{\equiv R^{t+1,\infty}_{\Delta c^c}} \tag{37}$$

where $R^{t,t+1}_{\Delta c^c}$ captures the revision in the capitalist's consumption growth rate between t and $t+1$ while $R^{t+1,\infty}_{\Delta c^c}$ captures the revisions between $t+1$ and the infinite future.[23] The decomposition provides a natural framework to analyze asset returns volatility as a function of the stockholder's consumption growth process. As for stock returns, one obtains

$$Var(r^s_{t+1} - E_t r^s_{t+1}) = \\ Var(R^{t,t+1}_{\Delta c^c}) + (1 - RRA)^2 Var(R^{t+1,\infty}_{\Delta c^c}) + 2(1 - RRA) Cov(R^{t,t+1}_{\Delta c^c}, R^{t+1,\infty}_{\Delta c^c}) \quad (38)$$

while, considering that real bonds pay no dividends, the variance of the risk-free rate can be expressed as[24]

$$Var(r^b_{t+1} - E_t r^b_{t+1}) = RRA^2 Var(R^{t+1,\infty}_{\Delta c^c}) \quad (39)$$

Equation (38) shows that the volatility of stock returns depends on the volatility of realized capitalists' consumption growth, $Var(R^{t,t+1}_{\Delta c^c})$, the volatility of the discounted sum of revisions in expected future capitalists' consumption growth, $Var(R^{t+1,\infty}_{\Delta c^c})$, and the covariance between the two. Note that, for a parameter of risk-aversion greater than 1, the coefficient $(1 - RRA)$ is negative and larger than 1 in modulus. Moreover, the covariance term is negative in presence of temporary shocks. As argued by Kaltenbrunner and Lochstoer (2010), a positive shock to realized consumption growth is necessarily followed by negative expected growth rates in the long-run, as consumption reverts to the steady state. Therefore, all the three addends of equation (38) contribute positively to the variance of stock returns. In contrast, according to equation (39) the variance of the risk-free rate exclusively depends on fluctuations in future expected consumption growth rates.

The above expressions highlight the opposite effects produced by the *risk-sharing* and *risk-aversion channels*. On the on hand, an increase in participation reduces the volatility of asset returns through a reduction in the fluctuations of both realized and expected consumption growth, as captured by the declining variance of capitalists' consumption growth (recall Figure 3). As shown in Appendix B, standard models featuring the *risk-sharing channel* alone predict a negative relationship between the participation rate and the standard deviation of the risk-free rate and stock returns. On the other hand, the *risk-aversion channel* amplifies the sensitivity of asset returns to shocks to the consumption growth process. Therefore, even smaller fluctuations in expected consumption produce larger fluctuations in the returns on stocks and bonds. This mechanism helps explaining

[23]The derivation of equation (37) is standard, see Campbell (2003).

[24]This expression is obtained by setting dividends equal to zero in equation (33) and using the Euler equation for bonds. For more details, see Campbell (2003).

why the volatility of stock returns and the risk-free rate remained essentially stable even during the Great Moderation.

6 Validating the mechanism

The previous section highlights the main mechanisms at work in the model as a result of increasing financial market participation. First, a decline in the quantity of risk for the representative stockholder, as measured by the covariance between his consumption growth rate and stock returns (*risk-sharing channel*). Second, an upward trend in the stockholder's average risk-aversion as a result of a shrinking stockholders'-to-aggregate consumption ratio (*risk-aversion channel*). In other words, the limited participation model economy displays higher quantity of risk and lower average risk-aversion compared to the full-participation (representative agent) limit case, where the distribution of consumption cannot affect equilibrium outcomes by construction.

In this section, I document that the two model mechanisms are supported by novel evidence from the U.S. Consumption Expenditure Survey for the sample 1984-2017.[25] Specifically, to test the predictions of Figure 7, I calculate annual consumption expenditures for the representative stockholder[26]. The consumption and stockholding status definitions follow Malloy et al. (2009). Consumption consists of non-durable goods and some services aggregated from the disaggregated expenditure categories reported in the survey. Moreover, the financial information reported both in the CEX and the Survey of Consumer Finances allows to sort the population in stockholders and non-stockholders. All the details about the construction of the dataset are reported in Appendix D.

6.1 Time-series evidence from the CEX

The three panels of Figure 8 contrast the degree of stock market participation estimated from the CEX (in red) with the empirical counterpart of the measures discussed in Figure 7 (in blue).[27] The left panel shows the estimated covariance between stock returns and the representative investor's consumption growth.[28] In line with the model's

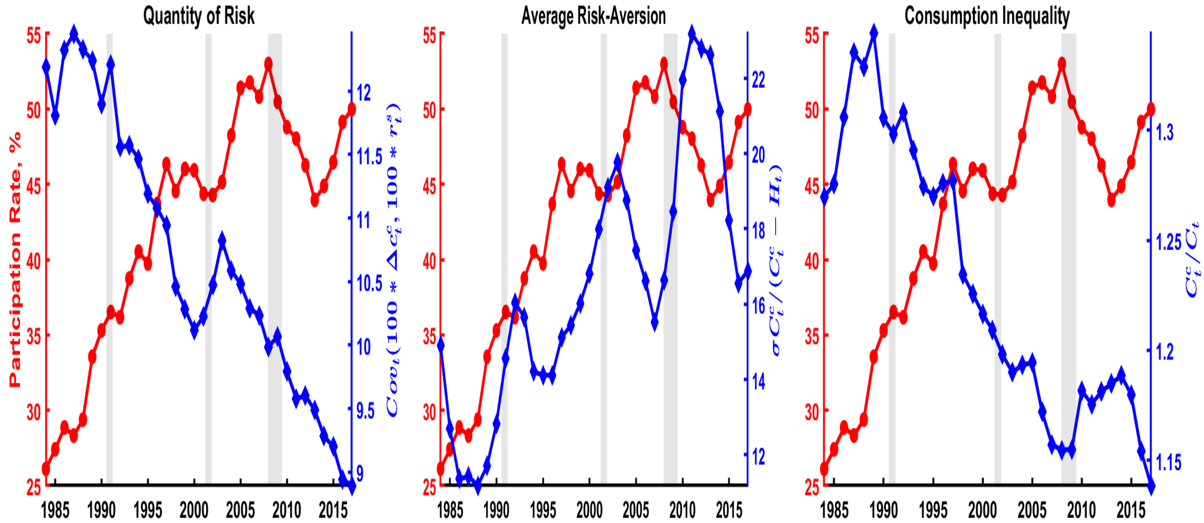
[25] I am grateful to Myroslav Pidkuyko for sharing his STATA dictionaries for the years 1980-1995 with me.

[26] As discussed in Appendix D, the representative agent assumption implies perfect risk-sharing within groups, thus ignoring uninsurable idiosyncratic consumption shocks. Nevertheless, Malloy et al. (2009) show that the comovement between asset returns and within-group cross-household inequality plays a minor role in explaining risk-premia. Moreover, the representative agent assumption is consistent with the two-agent model, where within-group heterogeneity is ruled out by construction.

[27] In Appendix E I verify that the results discussed below are robust along both the stock-holding status and the consumption definition dimensions.

[28] The time-varying covariance is estimated as an exponentially weighted moving-average covariance with smoothing parameter equal to 0.99 (annual data) on the demeaned series. Initial values are set to

Figure 8: Validating the mechanism: time-series evidence



Notes: This figure shows the: time-varying covariance between stockowners' consumption growth and stock returns (left); implied average risk-aversion (middle); and stockholders/aggregate consumption ratio (right), plotted against the participation rate. All quantities are estimated from the CEX, except for the dividend and wage growth series. Shaded bands indicate NBER recessions.

prediction, such covariance trended strongly downward in face of higher participation. In particular, an increase in participation is associated with improved risk-sharing and therefore a decline in the quantity of risk borne by investors. Nevertheless, the estimated covariance is smaller in magnitude than the model-implied one, which can be due to the fact that the model abstracts from many shocks that are likely to affect this measure in reality. In the middle panel a model-based empirical measure of relative risk-aversion is reported. For the sake of comparability with the theoretical model, I compute the relative risk-aversion as in equation (29) and the habit stock as in equation (4). The parameters are set to the baseline model calibration values.[29] The resulting series displays a strong positive trend in tandem with the degree of stock market participation, providing suggesting evidence in favor of the theoretical model mechanism. In quantitative terms the measure of relative risk-aversion oscillates between 11 and 22, a range which is fully compatible with the model. In addition, the countercyclical behavior of the measure is

the unconditional covariance for the sample 1984-1990.

[29]The initial value for H_t is set equal to the average aggregate per capita real consumption, as estimated from the CEX, in the period 1984-1990. It is worth stressing that the value of the local utility curvature parameter, σ, does not affect the correlation between the degree of participation and the estimated average risk-aversion, but only the level of the latter. Indeed, such correlation exclusively depends on the dynamics of the (inverse of the) surplus-consumption ratio, $\frac{C_t^c}{C_t^c - H_t}$, which is fully data driven except for the presence of the persistence parameter for the habit stock, m. Nevertheless, the results are very robust to changing such parameter over a wide range of values.

Table 5: Long-run covariation with participation rate

Variable	ρ_{LR}		β_{LR}	
	$\hat{\rho_{LR}}$	67% CI	$\hat{\beta_{LR}}$	67% CI
$Cov_t\left(\Delta c_t^c, r_t^s\right)$	-0.67	$[-0.92, -0.184]$	-0.117	$[-0.178, -0.02]$
$\sigma C_t^c/(C_t^c - H_t)$	0.757	$[0.25, 0.938]$	0.305	$[0.087, 0.441]$
C_t^c/C_t	-0.67	$[-0.912, -0.184]$	-0.007	$[-0.01, -0.002]$

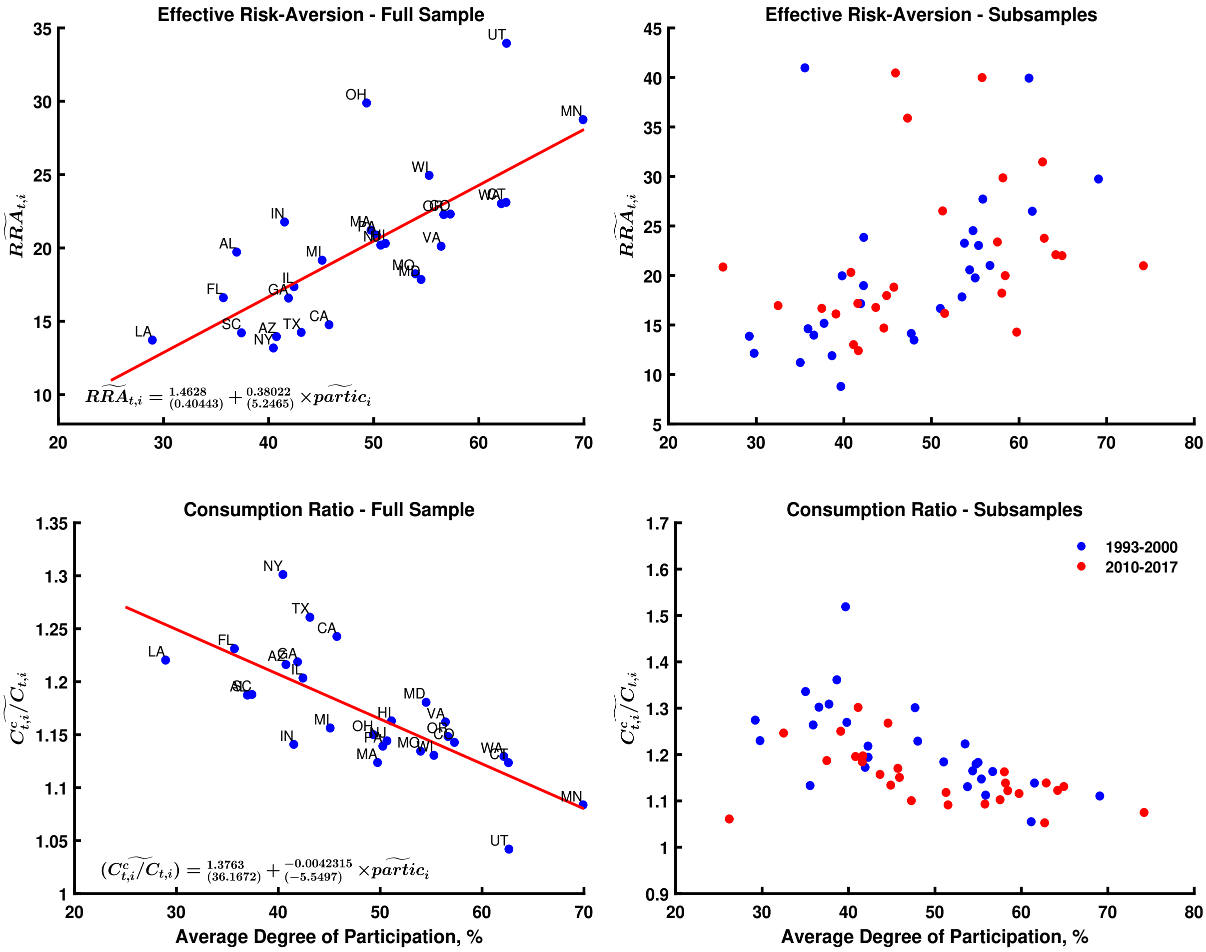
Notes: This table reports the long-run covariation measures proposed by Müller and Watson (2018) between the variables in the first column and the participation rate estimated from the CEX. The covariation is computed for periods longer than 20 years, to capture the covariation in trend components. ρ_{LR} denotes the long-run correlation coefficient while β_{LR} is the long-run slope coefficient of the regression where $Y = variable$ and $X = participation$. Hats denote the point estimate, reported along with the 67% confidence interval.

perfectly in line with the theory.

Finally, in the right panel I report the estimated stockholder-to-aggregate per capita consumption ratio (C_t^c/C_t in the model notation). The measure trends downward over the sample too, suggesting that more widespread access to financial markets implied a convergence of the representative investor consumption toward the aggregate standard of living in the economy. Similarly to the previous measures, the cyclical behavior of the consumption ratio is in line with economic intuition, with the ratio systematically dropping around recessions. The consumption of the stockholders identified in the CEX appears to be more sensitive to stock market crashes. Moreover, a consumption ratio around 1.35 at the beginning of the sample is quantitatively close to the 1.45 implied by the model for a participation rate of 25%, as in Figure 7, supporting the baseline calibration of the model. Also, in the model the ratio drops below 1.2 as the participation rate approaches 50%, which well compares with the empirical 1.15.

Table 5 reports evidence on the long-run comovement between the participation rate and the variables considered in Figure 8. Specifically, I compute the measures proposed by Müller and Watson (2018) for periods longer than 20 years. As noted by the authors, such length allows to abstract from both business cycles and medium-run fluctuations. The table displays the point estimate (denoted by the hat) of the long-run correlation coefficient ρ_{LR} and of the slope coefficient β_{LR} of a long-run regression of the variable of interest (first column) on the participation rate. Focusing on the second and third columns, the estimated long-run correlations are of the expected sign. The evidence in favor of the *risk-sharing channel* and *risk-aversion channel* (third and fourth row, respectively) seems particularly strong, with point estimates being as great as 67% and 76% in absolute terms, respectively. Similarly, the consumption ratio is found to significantly decline in association with and increase in the participation rate. The results are con-

Figure 9: Validating the mechanism: cross-sectional evidence



Notes: This figure displays the median (denoted by the tilde) effective risk-aversion (top panels) and stockholder-to-aggregate per capita consumption ratio (bottom panels) against the median participation rate at the state-level. Left panels: medians computed on the full sample 1993-2017. Right panels: medians computed on the 1993-2000 (blue dots) and 2010-2017 (red dots) subsamples.

firmed when looking at the estimated long-run betas (last two columns). In particular, an increase of one percentage point in the participation rate is associated with a decline of 0.117 points in the covariance between stockholders' consumption growth and stock returns, and a rise of 0.305 units in the measure of risk-aversion.

6.2 Cross-sectional evidence from the CEX

The time-series results provide evidence supporting the main model predictions, and in particular the novel *risk-aversion channel*. The CEX dataset also allows to exploit the cross-sectional variation at the state-level to test the main model mechanism. I estimate the participation rate and consumption series for both the aggregate and the representative stockholder for each of the U.S. states that are considered in the survey.[30] I then analyze the cross-sectional relationship between participation and effective risk-aversion or consumption inequality, measured again as the stockholder-to-aggregate per capita consumption ratio.

Figure 9 displays the results. The left panels depict the relationship between median effective risk-aversion (top) or consumption inequality indicator (bottom) and median participation rate at the state-level over the full sample.[31] The right panels report the same information, but for the two subsamples 1993-2000 and 2010-2017. In the top-left scatterplot, the cloud neatly follows an upward-sloping regression line, with a strongly significant and positive slope coefficient of about 0.38. Therefore, the evidence suggests that states registering higher degrees of participation exhibit a higher effective risk-aversion. It is interesting to notice that the size of the coefficient is very similar to the one reported in Table 5. In the top-right panel, for both clouds again a clearly positive relationship emerges. Moreover, the red cloud appears to be slightly shifted to the right compared to the blue one. This indicates that most states experienced an upward trend in the participation rate, coupled with an increase in the median effective risk-aversion.

Opposite results hold for the consumption inequality indicator. The cloud in the bottom-left scatterplot indicates a clearly negative correlation with the participation rate, with a slope coefficient of about -0.004 that is close to the time series estimate in Table 5. In line with the model predictions, the evidence suggests that states registering higher degrees of participation exhibit lower degrees of consumption inequality between the average investor and the average consumption level. Regarding the subsample analysis, the red cloud appears to be shifted down-right compared to the blue one. This indicates that in most states the participation rate rose over time while the average consumption ratio declined.

7 Conclusion

The empirical regularity that only a fraction of U.S. households invest in the stock market has spurred a number of studies analyzing the asset pricing consequences of limited stock market participation in production economies. Recent works have pointed to the heterogeneity between workers and stockholders as a feature of the U.S. economy that plays a crucial role in the determination of the equity premium and the volatility of

[30]As thoroughly discussed in Appendix D, the state-level data is available only from 1993 and not for all the 50 U.S. states. Given also the sample restrictions adopted, the evidence provided here employs data from 27 states.

[31]Given the small sample size available at the state level, I report the median rather than the mean of each variable, since the former moment is more robust to outliers.

stock returns. I employ a workers-capitalists Real Business Cycle model with technology and distribution shocks to study the macroeconomic and asset pricing implications of the upward trend in stock market participation observed in the U.S. since the late 1980s. The combination of high participation and low volatility of technology shocks is able to produce a substantial decrease in the volatility of output and consumption growth while moving the average equity premium and risk-free rate in the direction observed during the last thirty years. This work contributes to the limited participation literature by shedding light on a new channel through which the heterogeneity between workers and stockholders affects both the macroeconomy and the stock market. The model adds a new dimension of analysis to the existing literature, by identifying the developments in the representative investor's surplus-consumption ratio as a relevant determinant of the macro-financial trends observed in the U.S. over the last decades.

References

Attanasio, O. P., Banks, J., and Tanner, S. (2002). Asset holding and consumption volatility. *Journal of Political Economy*, 110(4):771–792.

Bee, A., Meyer, B. D., and Sullivan, J. X. (2012). The validity of consumption data: are the consumer expenditure interview and diary surveys informative? *NBER Working Paper Series*, 18308.

Bilbiie, F. O. (2008). Limited asset markets participation, monetary policy and (inverted) aggregate demand logic. *Journal of Economic Theory*, 140(1):162–196.

Bilbiie, F. O. and Straub, R. (2013). Asset market participation, monetary policy rules, and the great inflation. *Review of Economics and Statistics*, 95(2):377–392.

Blanchard, O. and Simon, J. (2001). The long and large decline in us output volatility. *Brookings Papers on Economic Activity*, 2001(1):135–174.

Caballero, R. J., Farhi, E., and Gourinchas, P.-O. (2017). Rents, technical change, and risk premia Accounting for secular trends in interest rates, returns on capital, earning yields, and factor shares. *American Economic Review*, 107(5):614–20.

Campbell, J. R. and Hercowitz, Z. (2006). The role of collateralized household debt in macroeconomic stabilization. *Foerder Institute for Economic Research Working Papers*, 275707.

Campbell, J. Y. (1994). Inspecting the mechanism: An analytical approach to the stochastic growth model. *Journal of Monetary Economics*, 33(3):463–506.

Campbell, J. Y. (2003). Consumption-based asset pricing. *Handbook of the Economics of Finance*, 1:803–887.

Campbell, J. Y. and Cochrane, J. H. (1999). By force of habit: A consumption-based explanation of aggregate stock market behavior. *Journal of Political Economy*, 107(2):205–251.

Campbell, J. Y. and Shiller, R. J. (1988). The dividend-price ratio and expectations of future dividends and discount factors. *Review of Financial Studies*, 1(3):195–228.

Chan, Y. L. and Kogan, L. (2002). Catching up with the joneses: Heterogeneous preferences and the dynamics of asset prices. *Journal of Political Economy*, 110(6):1255–1285.

Chen, A. (2016). Semi-parametric restrictions on production-based asset pricing models. Available at SSRN: http://dx.doi.org/10.2139/ssrn.2742323.

Chen, A. Y. (2017). External habit in a production economy: A model of asset prices and consumption volatility risk. *Review of Financial Studies*, 30(8):2890–2932.

Cloyne, J., Ferreira, C., and Surico, P. (2019). Monetary policy when households have debt: New evidence on the transmission mechanism. *Review of Economic Studies*, 0:1–28.

Cochrane, J. H. (2017). Macro-finance. *Review of Finance*, 21(3):945–985.

Corhay, A., Kung, H., and Schmid, L. (2020). Q: Risk, rents or growth? Available at SSRN: http://dx.doi.org/10.2139/ssrn.3546667.

Damodaran, A. (2013). Equity risk premiums (erp): Determinants, estimation and implications–the 2013 edition. *Managing and measuring risk: Emerging global standards and regulations after the financial crisis*, pages 343–455.

Danthine, J.-P. and Donaldson, J. B. (2002). Labour relations and asset returns. *Review of Economic Studies*, 69(1):41–64.

De Graeve, F., Dossche, M., Emiris, M., Sneessens, H., and Wouters, R. (2010). Risk premiums and macroeconomic dynamics in a heterogeneous agent model. *Journal of Economic Dynamics and Control*, 34(9):1680–1699.

Debortoli, D. and Galí, J. (2017). Monetary policy with heterogeneous agents: Insights from TANK models. Technical report, mimeo.

Del Negro, M., Giannone, D., Giannoni, M. P., and Tambalotti, A. (2019). Global trends in interest rates. *Journal of International Economics*, 118:248–262.

Delle Monache, D., Petrella, I., and Venditti, F. (2020). Price dividend ratio and long-run stock returns: a score driven state space model. *Journal of Business & Economic Statistics*, Forthcoming.

Den Haan, W. J. and Sterk, V. (2011). The myth of financial innovation and the Great Moderation. *Economic Journal*, 121(553):707–739.

Farhi, E. and Gourio, F. (2018). Accounting for macro-finance trends: Market power, intangibles, and risk premia. *NBER Working Paper Series*, 25282.

Favilukis, J. (2013). Inequality, stock market participation, and the equity premium. *Journal of Financial Economics*, 107(3):740–759.

Fernald, J. (2014). A quarterly, utilization-adjusted series on total factor productivity. Federal Reserve Bank of San Francisco.

Greenwald, D. L., Lettau, M., and Ludvigson, S. C. (2019). How the wealth was won: Factors shares as market fundamentals. *NBER Working Paper Series*, 25769.

Greenwood, J., Hercowitz, Z., and Huffman, G. W. (1988). Investment, capacity utilization, and the real business cycle. *American Economic Review*, pages 402–417.

Guiso, L. and Sodini, P. (2013). Household finance: An emerging field. In *Handbook of the Economics of Finance*, volume 2, pages 1397–1532. Elsevier.

Guvenen, F. (2006). Reconciling conflicting evidence on the elasticity of intertemporal substitution: A macroeconomic perspective. *Journal of Monetary Economics*, 53(7):1451–1472.

Guvenen, F. (2009). A parsimonious macroeconomic model for asset pricing. *Econometrica*, 77(6):1711–1750.

Israelsen, R. D. (2010). Investment based valuation and managerial expectations. Available at SSRN: http://dx.doi.org/10.2139/ssrn.1563502.

Jaccard, I. (2014). Asset returns and labor supply in a production economy. *Journal of Money, Credit and Banking*, 46(5):889–919.

Jensen, H., Petrella, I., Ravn, S. H., and Santoro, E. (2020). Leverage and deepening business-cycle skewness. *American Economic Journal: Macroeconomics*, 12(1):245–81.

Jensen, H., Ravn, S. H., and Santoro, E. (2018). Changing credit limits, changing business cycles. *European Economic Review*, 102:211–239.

Jermann, U. J. (1998). Asset pricing in production economies. *Journal of Monetary Economics*, 41(2):257–275.

Justiniano, A. and Primiceri, G. E. (2008). The time-varying volatility of macroeconomic fluctuations. *American Economic Review*, 98(3):604–41.

Kaltenbrunner, G. and Lochstoer, L. A. (2010). Long-run risk through consumption smoothing. *Review of Financial Studies*, 23(8):3190–3224.

Kogan, L., Papanikolaou, D., and Stoffman, N. (2020). Left behind: Creative destruction, inequality, and the stock market. *Journal of Political Economy*, 128(3):855–906.

Krueger, D. and Perri, F. (2006). Does income inequality lead to consumption inequality? Evidence and theory. *Review of Economic Studies*, 73(1):163–193.

Lansing, K. J. (2015). Asset pricing with concentrated ownership of capital and distribution shocks. *American Economic Journal: Macroeconomics*, 7(4):67–103.

Lettau, M. (2003). Inspecting the mechanism: Closed-form solutions for asset prices in real business cycle models. *Economic Journal*, 113(489):550–575.

Lettau, M., Ludvigson, S., and Ma, S. (2018). Capital share risk and shareholder heterogeneity in u.s. stock pricing. *Journal of Finance*, Forthcoming.

Lettau, M., Ludvigson, S. C., and Wachter, J. A. (2007). The declining equity premium: What role does macroeconomic risk play? *Review of Financial Studies*, 21(4):1653–1687.

Liu, Z., Wang, P., and Zha, T. (2013). Land-price dynamics and macroeconomic fluctuations. *Econometrica*, 81(3):1147–1184.

Malkhozov, A. (2014). Asset prices in affine real business cycle models. *Journal of Economic Dynamics and Control*, 45:180–193.

Malloy, C. J., Moskowitz, T. J., and Vissing-Jørgensen, A. (2009). Long-run stockholder consumption risk and asset returns. *Journal of Finance*, 64(6):2427–2479.

Morelli, J. M. (2021). Limited participation in equity markets and business cycles. Finance and Economics Discussion Series 2021-026, Board of Governors of the Federal Reserve System U.S.

Müller, U. K. and Watson, M. W. (2018). Long-run covariability. *Econometrica*, 86(3):775–804.

Nason, J. M. and Smith, G. W. (2008). Great moderation(s) and US interest rates: Unconditional evidence. *The BE Journal of Macroeconomics*, 8(1).

Pancrazi, R. (2014). How beneficial was the Great Moderation after all? *Journal of Economic Dynamics and Control*, 46:73–90.

Poterba, J. M., Samwick, A. A., Shleifer, A., and Shiller, R. J. (1995). Stock ownership patterns, stock market fluctuations, and consumption. *Brookings Papers on Economic Activity*, 1995(2):295–372.

Rios-Rull, J.-V. and Santaeulalia-Llopis, R. (2010). Redistributive shocks and productivity shocks. *Journal of Monetary Economics*, 57(8):931–948.

Stock, J. H. and Watson, M. W. (2002). Has the business cycle changed and why? *NBER Macroeconomics Annual*, 17:159–218.

Stock, J. H. and Watson, M. W. (2003). Has the business cycle changed? Evidence and explanations. In *Monetary policy and uncertainty: Adapting to a changing economy*, pages 9–56. Citeseer.

Wong, A. (2019). Refinancing and the transmission of monetary policy to consumption. Technical report, mimeo.

Appendices

A Dataset for macro-financial facts

The data commented in the main text are all at the annual frequency and cover the period 1949-2017, except for real S&P500 dividends and the S&P500 price-to-dividend ratio, which span the sample 1949-2012; the expected stock returns, which are available only over the period 1961-2017; and the relative price of investment goods, available up to 2016. All macroeconomic data are expressed in real per capita terms, where appropriate. Real variables are obtained by deflating the nominal variables by the annual Consumer Price Index (CPI) obtained from the Bureau of Labor Statistics.[32] Per capita measures are obtained by dividing the real variables by the U.S. population, obtained from the National Income and Products Accounts (NIPA, Table 2.1, line 39). Growth rates are constructed as the first-difference in the logarithm of the variables.

GDP is constructed as Nominal GDP (NIPA, Table 1.1.5, Line 1). Consumption is defined as the sum of Nominal Expenditures on Nondurable Goods (NIPA, Table 2.3.5, Col. 8) plus Nominal Expenditures on Services (NIPA, Table 2.3.5, Col. 13). Investment is defined as the sum of Nominal Private Nonresidential Fixed Investment (PNFIA series from the FRED website) plus Nominal Expenditures on Durable Goods (Table 2.3.5, Col. 3). The capital share of income and the macroeconomic dividends are constructed following Lansing (2015). The former is defined as one minus the ratio between the compensation of employees (NIPA, Table 1.14, Line 4) and the gross value added of corporate business (NIPA, Table 1.14, Line 1). Dividends are constructed by subtracting real per capita investment to the product between the capital share of income and real per capita GDP. Consistent with the definition of the capital share of income, wages are constructed as the compensation of employees. The (not utilization adjusted) total factor productivity (TFP) growth series is from Fernald (2014).[33] Finally, the relative price of investment to consumption goods is defined as the Quality-Adjusted Price of New Equipment and Software from Israelsen (2010).[34]

Regarding financial variables, the annual real risk-free rate is defined as the annual gross return on the 90-days T-Bill return net of annual CPI inflation. The former is downloaded from the WRDS (Wharton Research Data Services) website. Annual historical S&P500 stock returns are taken from the version of Damodaran (2013) updated

[32] Available at: https://data.bls.gov/PDQWeb/cu

[33] Available at: https://www.frbsf.org/economic-research/indicators-data/total-factor-productivity-tfp/

[34] I am grateful to the author for sharing his data with me.

through 2019.[35] Expected S&P500 stock returns are calculated from the same source. To both series the annual CPI inflation is then subtracted. The equity premium is defined as the average difference between expected stock returns and risk-free rate. As explained by the author, expected returns are computed by a Dividend Discount Model (DDM) which accounts for the free cash flow to equity (FCFE) as a measure of potential dividends. Specifically, this measure considers potential dividends instead of actually paid dividends to calculate the cash flow to equity left after taxes, reinvestment and debt repayments. This method therefore addresses the fact that in the last decade firms have paid out only about half of their FCFE as dividends. In other words, this measure adds stock buybacks to the dividends actually paid in order to gauge a more accurate estimate for the total cash flow. Finally, the annual S&P500 real dividends and price to dividend ratio are taken from Lansing (2015)[36], who constructs both series from Robert Shiller's database for the period 1871-2012.

The direct and indirect stock-ownership rates are taken from the 2016 Survey of Consumer Finances.[37] Figure A.1 reports several measures of this rate. In the main text, and for the calibration of the model, I consider the raw participation rate (in blue). Indeed, the raw participation rate appears to be the least arbitrary measure of participation. According to this measure, a household is defined as stockowner if holding any positive amount of stocks. However, such measure could hide substantial heterogenity in stock-holdings, since many of the new entrants in the stock market might hold very low amounts of stocks. In this sense, the raw participation rate could in principle increase even if the large majority of new stockowners hold negligible amounts of the assets.

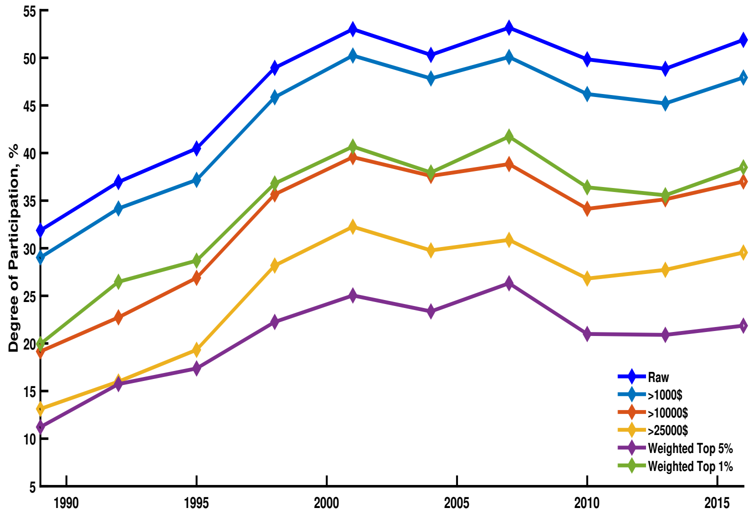
To address this concern, in Figure A.1 I also report additional measures of stock market participation. For example, one can see that even when raising the dollar-amount threshold to 1000, 10000 or 25000$ (light blue, orange and yellow lines, respectively) the participation rate experienced a quite strong upward trend, although obviously the estimated rate is lower than the baseline. Similar trends can be observed when measuring wealth-weighted rates. Wealth-weighted participation rates take into account the (well-documented) possibility that a large part of total stock-holdings are concentrated at the top of the stock-wealth distribution. For example, following Lettau et al. (2018), the top 5% wealth-weighted participation rate (purple line) is computed as:

$$5\% \times w^{5\%} + (rpr - 5\%) \times (1 - w^{5\%})$$

where $w^{5\%}$ is the share of total stocks held by the household falling in the top 5% of

[35] Available at: http://pages.stern.nyu.edu/~adamodar/
[36] Available at: http://dx.doi.org/10.1257/mac.20110130
[37] Available at: https://www.federalreserve.gov/econres/scfindex.htm

Figure A.1: Direct and indirect stock ownership: additional measures



Notes: The figure documents the upward trend in several measures of stock market participation, based on different dollar-amount thresholds or different weighting schemes.

the stock-wealth distribution and rpr is the raw participation rate. The top 1% wealth-weighted participation rate (green line) is computed similarly. Even in these cases, however, it is clear that the participation rate trended upward over the sample period.

B Comparison to earlier models

In this Appendix I show that in existing models with concentrated ownership of capital an increase in financial participation lowers the equity premium and stock market volatility while raising the standard deviation of aggregate consumption growth and the mean level of the risk-free rate. Indeed, the level of participation only affects the quantity of risk borne by investors. As a result, these models fail to provide a plausible explanation for the behavior of U.S. macroeconomic and financial variables over the last few decades. Specifically, I consider the case where capitalists exhibit external habit preferences as in Lansing (2015) or a different specification for distribution shocks, as proposed by Danthine and Donaldson (2002) and De Graeve et al. (2010).

External habit utility as in Lansing (2015)

I start with the case where both agents supply labor inelastically but capitalists have external habit utility preferences specified in a similar fashion as Lansing (2015). In particular, the utility function of capitalists now depends on the past value of the representative capitalist's consumption, not on the past aggregate per capita consumption:

$$U^c\left(C_t^c\right) = \frac{\left(C_t^c - \chi_c H_t\right)^{1-\sigma}}{1-\sigma}$$

where χ_c denotes the weight of the external habit in the capitalists' utility function and the habit stock H_t is

$$H_t = mH_{t-1} + (1-m)\, C_{t-1}^c$$

For comparability with the baseline results, this specification of the habit stock includes the parameter m, controlling the degree of persistence in the habit stock and hence allowing for a slow-moving process. The marginal utility of consumption is therefore given by $\Lambda_t = \left(C_t^c - \chi_c H_t\right)^{-\sigma}$ and the consequent stochastic discount factor reads $M_{t,t+1} = \beta \frac{E_t[\Lambda_{t+1}]}{\Lambda_t}$.

Notice that in the steady state the level of relative risk-aversion is equal to

$$RRA \equiv -\frac{C^c U^c\left(C_t^c\right)''}{U^c\left(C_t^c\right)'} = \frac{\sigma}{1-\chi_c}$$

which depends only on the deep preference parameters and is therefore independent of the gap between capitalists' and aggregate perc-capita consumption. In other words, in this version of the model the *risk-aversion channel* is shut down.

The calibration adopted (reported in Table B.1) is kept as close as possible to the baseline model, in order to generate comparable macro-financial moments for the pre-Great Moderation period, as can be seen in the first panel of Table B.2. Nevertheless, the implications of higher participation are now opposite to the baseline case. In particular, an increase in participation entails higher (lower) consumption (investment) growth volatility and an increase in the average risk-free rate. Conversely, the standard deviation of asset returns declines as well as the average equity premium, which drops by more than 1 percentage point. This result stems from the fact that the capitalist's steady state average risk-aversion now does not depend on the distribution of consumption, being determined only by exogenous parameters. Hence, the degree of participation does not affect the average risk-aversion but only the quantity of risk borne by investors, i.e. the model features only the standard *risk-sharing channel.*

Operating leverage

Danthine and Donaldson (2002) and De Graeve et al. (2010) focus on the asset pricing implications of the "operating leverage", i.e. the riskiness of dividends deriving from the priority status of wage claims. If the wage share is not constant over the cycle, then

Table B.1: External habit utility as in Lansing (2015) - calibration

Description	Parameter	Value
Discount Rate	β	0.957
Capital Share of Income	α	0.37
Depreciation Rate	δ	0.115
Technology Persistence	ρ_a	0.97
Distribution Shock Persistence	ρ_ν	0.99
IST Persistence	ρ_{p^I}	0.92
IST Shock Volatility	σ_{p^I}	0.251
Distribution Shock Volatility	σ_ν	0.04
Technology Shock Volatility	σ_a	0.017
Local Utility Curvature	σ	3
Habit Stock Persistence	m	0.8
Capital Adjustment Cost	χ_k	0.45
Leverage Factor	χ	0.186
Weight of Habit in Utility	χ_c	0.66

Notes: Calibration of the model with external habit utility as in Lansing (2015).

wages represent an insurance device between workers and firms (hence, shareholders). A countercyclical wage share exacerbates the procyclicality of dividends by simultaneously smoothing the income of workers. The mechanism analyzed by these authors is clearly similar in spirit to the distribution shock featured in the baseline model.

I study the implications of increasing participation in the two-agent version of the model proposed by De Graeve et al. (2010) featuring workers and shareholders.[38] In this model, both agents exhibit GHH preferences and supply labor elastically, but workers have lower EIS than shareholders. As a consequence, workers have a stronger consumption smoothing motive than capitalists. This stronger consumption smoothing motive is satisfied through a long-term labor contract that "*guarantees an optimal risk-sharing between workers and shareholders on a period-by-period basis for a given realization of the exogenous bargaining weight*" (De Graeve et al., 2010, p. 1683). The contract therefore makes workers' consumption smooth at the expenses of higher volatility in capitalists' consumption.

The second panel of Table B.2 shows that higher participation produces the same effects counterfactual results highlighted earlier. Note that, for comparability of the results, the calibration is kept at the frequency[39] and the values employed by the authors.[40]

[38]This is the "T1-T3 with correlated shocks" specification in De Graeve et al. (2010), see Tables 3 and 4 in the paper. I am grateful to Ferre De Graeve for sharing the replication codes with me.

[39]In particular, the model is calibrated at the quarterly frequency. Macroeconomic variables are detrended with the Hodrick-Prescott filter, while asset pricing moments are reported in annualized terms.

[40]The only difference lies in the volatility of the redistribution shocks, which is lowered from 3.1675%

Table B.2: Comparison to earlier models

$(1-\gamma)$	$\sigma(\Delta y)$	$\sigma(\Delta c)$	$\sigma(\Delta i)$	$\sigma(\Delta d)$	$E(R^b)$	$E(R^{e,s}-R^b)$	$\sigma(R^s)$	$\sigma(R^b)$
External Habit Utility (Lansing, 2015)								
20%	2.68	1.72	6.77	6.15	1.46	4.42	18.45	2.88
50%	2.67	1.75	6.01	8.52	2.60	3.08	16.37	2.30
Operating Leverage (De Graeve et al., 2010)*								
20%	1.72	1.36	3.19	23.86	1.69	4.58	21.28	1.87
50%	1.72	1.46	2.73	20.77	2.85	2.56	16.70	1.57

Notes: This table reports the results of two variants of the baseline model, for two different degrees of participation (1-γ). *This case is based on the workers-shareholders version of the model proposed by De Graeve et al. (2010). The calibration is kept at the frequency (quarterly) and values employed by the authors. Moments are reported in percent. Macroeconomic variables are detrended with the HP filter. The moments are therefore referred to the detrended variables (not the growth rates). Asset pricing moments are reported in annualized terms.

In this case, the standard deviation of dividends drops as the fraction of capitalists rises. This is due to the fact that higher participation implies a less broad application of the wage contract, since shareholders supply labor at the spot (perfectly competitive) wage. Wages become less rigid which in turn stabilizes dividends. This mechanism further enhances the *risk-sharing channel* through which higher participation reduces the equity premium while raising the risk-free rate.

Effect of the Great Moderation on the participation rate

The analysis in this paper focuses on exogenous variations in the participation rate and their effects on both the macroeconomic and the financial market environments. These exogenous variations are based on the underlying assumption, shared with most of the literature on the topic (see Bilbiie and Straub, 2013; Favilukis, 2013, for a discussion), that the upward trend in the participation rate was due to regulatory reforms that facilitated the access to financial instruments for most households, thereby reducing the implicit participation costs.

A concern, however, regards the possible existence of a two-way relationship between the degree of participation and the macro-financial environment. For example, it is possible to suppose that a lower macroeconomic volatility, as observed in the U.S. since the mid-1980s, could have in turn caused the trend in participation registered over the same years, by reducing the background risk faced by most households. To account for this reasonable concern, in this section I present a simple back-of-the-envelope exercise

to 2.2%. This generates essentially the same macroeconomic moments (compare with Table 4 in the article) but asset pricing results that are more directly comparable to my baseline model. The effect of higher participation does not depend on the size of the standard deviation of the shocks.

that shows how a decrease in macro volatility, which by itself would shrink the equity premium as clearly shown in Section 4.5, actually reduces the incentive to enter the stock market. Therefore, the Great Moderation would not be sufficient to explain the upward participation trend.

Following the working paper version of Bilbiie (2008), suppose that each time a household goes to the stock market a household-specific proportional transaction cost has to be paid. For simplicity, assume an extreme, binomial distribution of costs such that only workers have to pay the transaction cost F, while capitalists do not. In this case, the capitalists' (who face a zero cost) Euler equation for stock shares is $1 = E_t[M_{t,t+1}R^s_{t+1}]$. On the other hand, workers choose not to hold stocks if

$$1 + F > E_t[M^w_{t,t+1}R^s_{t+1}] > (1 + F)^{-1}$$

where $M^w_{t,t+1}$ is the worker's SDF and the second inequality holds for the case in which the household shorts the asset.

For the case of log-utility (with no habit formation) and taking second-order approximations under the assumption of joint conditional lognormality and homoskedasticity, the author shows that a lower bound for the transaction cost in the stock market is

$$F \approx \left| E_t\Delta c^c_{t+1} - E_t\Delta c^w_{t+1} + 0.5(\sigma^2_w - \sigma^2_c) + \sigma_{r^s_{t+1},c} - \sigma_{r^s_{t+1},w} \right|$$

where $\sigma^2_j \equiv var(c^j_{t+1} - E_t c^j_{t+1})$ and $\sigma_{r^s_{t+1},j} \equiv cov(r^s_{t+1} - E_t r^s_{t+1}, c^j_{t+1} - E_t c^j_{t+1})$ for $j = w, c$. Therefore, F can be calculated by computing the moments involved in the equation in the general equilibrium model for given values of the volatility of the exogenous shocks and the participation rate.

Table B.3 reports the simulated implicit participation costs for the baseline model with log-utility and no habit formation. The table reports F in absolute value and in percentage of the average wage in the economy for the two volatility regimes (pre-Great Moderation and Great Moderation) defined in the baseline model. In both cases, the participation rate is set to the baseline value of 20%. First, it can be noticed that even a very low participation cost, smaller than 0.5% of the aggregate per capita wage in both volatility regimes, is (implicitly) able to prevent 80% of households from accessing financial markets (notice that the presence of a positive equity premium entails that, if households prefer to stay away from the stock market, they also implicitly avoid the bond market). Second, the participation cost decreases both in absolute and relative (to the wage) terms when shifting from a high-volatility (pre-GM) to a low-volatility (GM) regime. Therefore, in the low-volatility regime the incentive to enter the stock

Table B.3: Implicit participation costs

Regime	F	$\frac{F}{E(W_t)}$
	Participation=20%	
Pre-GM	5.275×10^{-3}	0.47%
GM	4.866×10^{-3}	0.4372%

market is lower than the high-volatility regime, since a lower participation cost is needed to rationalize why the large majority of households do not invest in stocks. Indeed, as discussed in Section 4.5 lower aggregate risk by itself determines a reduction in the equity premium, which can be interpreted as the foregone gain from investing in stocks. Overall, this simple exercise shows that the Great Moderation is unlikely to have caused the rise in asset market participation rates.

C Inspecting the mechanism: additional results

The model produces a high relative volatility between capitalists' and workers' consumption growth, in line with the empirical evidence reported in Guvenen (2009). How does the model achieve this result? I demonstrate that the volatility of dividend growth and the covariance between dividend and wage growth are crucial in order to make capitalists' consumption growth sufficiently volatile. The distribution shock helps make dividends strongly procyclical even in presence of habit preferences, which induce a strong consumption smoothing motive for capitalists. This in turn makes capitalists' consumption growth volatile even if the covariance between dividend and wage growth is slightly negative.

The role of the distribution shock

Recall that in equilibrium, with fixed labour supply, workers simply consume their wage, as in equation (1), while the individual capitalist's consumption is given by equation (26). We can therefore rewrite

$$C_t^c = C_t^w + \frac{D_t}{1-\gamma}$$

Now, consider a log-linearized version of these budget constraints. Denoting the logs with lower case letters, workers' consumption growth is

$$\Delta c_t^w = \Delta w_t$$

For capitalists, it can be shown[41] that a log-linear approximation of their budget constraints leads to

$$\Delta c_t^c \approx \lambda_1 \Delta c_t^w + \lambda_2 \Delta d_t \tag{C.1}$$

where λ_1 and λ_2 are a convolution of the deep parameters of the model.[42] In particular, λ_1 and λ_2 are a decreasing and increasing function of γ, respectively, and are both positive but smaller than 1, for any value $\gamma \in [0, 1)$. I can now obtain the variance of capitalists' consumption growth

$$Var(\Delta c_t^c) \approx \lambda_1^2 Var(\Delta c_t^w) + \lambda_2^2 Var(\Delta d_t) + 2\lambda_1\lambda_2 Cov(\Delta c_t^w, \Delta d_t) \tag{C.2}$$

implying

$$\frac{Var(\Delta c_t^c)}{Var(\Delta c_t^w)} \approx \lambda_1^2 + \lambda_2^2 \frac{Var(\Delta d_t)}{Var(\Delta c_t^w)} + 2\lambda_1\lambda_2 \frac{Cov(\Delta c_t^w, \Delta d_t)}{Var(\Delta c_t^w)} \tag{C.3}$$

Equation (C.3) reveals the requirements for $\frac{Var(\Delta c_t^c)}{Var(\Delta c_t^w)} > 1$, i.e. to have capitalists' consumption growth to be more volatile than workers'. Specifically, capitalists' consumption growth will be more volatile than workers' if dividend growth is sufficiently volatile relative to wage growth and/or comoves sufficiently strongly, and positively, with wage growth. As I show in the next section, the model delivers a negative correlation (implying $\frac{Cov(\Delta c_t^w, \Delta d_t)}{Var(\Delta c_t^w)} < 0$) as in the data. Hence, the distribution shock plays a key role in ensuring that capitalists' consumption growth is more volatile than workers'. As displayed in Figures 2 and 5, conditional to neutral or investment-specific technology shocks, dividend growth volatility would be insufficient to guarantee a volatile capitalists' consumption. However, the distribution shock makes dividends sufficiently volatile and procyclical, while affecting only mildly workers' consumption. This in turn delivers a high relative volatility between capitalists' and workers' consumption growth.

Steady state level of capital

In the steady state, the parameters a_1 and a_2 ensure that $\Phi\left(\frac{I}{K}\right) = \delta$, since $(I/K) = \delta$. This in turn implies that $\Phi' \frac{I}{K} = 1$. Moreover, $P^I = 1$. Therefore, considering equation (21), in steady state we have

$$1 = \beta[\alpha K^{\alpha-1} + (1 - \delta + \delta - \delta)]$$

$$\frac{1}{\beta} - (1 - \delta) = \alpha K^{\alpha-1}$$

[41] All the proofs are provided in this Appendix.

[42] I define $\lambda_1 \equiv \frac{(1-\gamma)(1-\alpha)(r+\delta)}{[1-\gamma(1-\alpha)]r+(1-\gamma)(1-a)\delta}$ and $\lambda_2 \equiv \frac{\alpha r}{[(1-\gamma(1-\alpha)]r+(1-\gamma)(1-a)\delta}$, with $r \equiv -\log(\beta)$.

thus

$$K = \left[\frac{\alpha}{1/\beta - (1-\delta)}\right]^{1/1-\alpha}$$

From the same FOC, we know that in s.s. $R = (1/\beta)$. Therefore, following Campbell (1994), by setting $R \approx 1 + r$, where $r = \log(R)$, we finally get

$$K = \left[\frac{\alpha}{r+\delta}\right]^{1/1-\alpha} \tag{C.4}$$

and all the other variables' s.s. values follow from K.

Proof for equation (C.1)

Consider capitalists' budget constraint in equilibrium

$$C_t^c = W_t + \frac{D_t}{1-\gamma} = C_t^w + \frac{D_t}{1-\gamma}$$

taking logs on both sides (lower case letters denote the logarithm of upper case letters)

$$c_t^c = \log[\exp(c_t^w) + \frac{1}{1-\gamma}\exp(d_t)]$$

taking a first-order Taylor expansion of the RHS around the non-stochastic steady state I can write (the absence of time subscripts denote steady state values)

$$c_t^c \approx \frac{C^w}{C^w + \frac{D}{1-\gamma}} c_t^w + \frac{\frac{D}{1-\gamma}}{C^w + \frac{D}{1-\gamma}} d_t$$

i.e.

$$c_t^c \approx \frac{(1-\gamma)C^w}{(1-\gamma)C^w + D} c_t^w + \frac{D}{(1-\gamma)C^w + D} d_t \tag{C.5}$$

Note that the steady state value of workers' consumption is

$$C^w = W = (1-\alpha)Y = (1-\alpha)K^\alpha \tag{C.6}$$

while dividends

$$D = Y - W - I = \alpha Y - \delta K = K^\alpha[\alpha - \delta K^{1-\alpha}]$$

Substituting these steady state values into equation (C.5)

$$c_t^c \approx \frac{(1-\gamma)(1-\alpha)K^\alpha}{(1-\gamma)(1-\alpha)K^\alpha + K^\alpha[\alpha - \delta K^{1-\alpha}]} c_t^w + \frac{K^\alpha[\alpha - \delta K^{1-\alpha}]}{(1-\gamma)(1-\alpha)K^\alpha + K^\alpha[\alpha - \delta K^{1-\alpha}]} d_t$$

$$c_t^c \approx \frac{(1-\gamma)(1-\alpha)K^\alpha}{K^\alpha\left\{(1-\gamma)(1-\alpha)+\alpha-\delta K^{1-\alpha}\right\}}c_t^w + \frac{K^\alpha\left[\alpha-\delta K^{1-\alpha}\right]}{K^\alpha\left\{(1-\gamma)(1-\alpha)+\alpha-\delta K^{1-\alpha}\right\}}d_t$$

$$c_t^c \approx \frac{(1-\gamma)(1-\alpha)}{[1-\gamma(1-\alpha)]-\delta K^{1-\alpha}}c_t^w + \frac{\alpha-\delta K^{1-\alpha}}{[1-\gamma(1-\alpha)]-\delta K^{1-\alpha}}d_t$$

and plugging the expression for the steady state level of capital, given by equation (C.4), after some simple algebra one obtains

$$c_t^c \approx \frac{(1-\gamma)(1-\alpha)(r+\delta)}{[1-\gamma(1-\alpha)](r+\delta)-\delta a}c_t^w + \frac{\alpha r}{[1-\gamma(1-\alpha)](r+\delta)-\delta a}d_t$$

Therefore, noticing that $[1-\gamma(1-\alpha)](r+\delta)-\delta a = [1-\gamma(1-\alpha)r]+(1-\gamma)(1-a)\delta$

$$c_t^c \approx \frac{(1-\gamma)(1-\alpha)(r+\delta)}{[1-\gamma(1-\alpha)]r+(1-\gamma)(1-a)\delta}c_t^w + \frac{\alpha r}{[1-\gamma(1-\alpha)]r+(1-\gamma)(1-a)\delta}d_t$$

implying that

$$\Delta c_t^c \approx \frac{(1-\gamma)(1-\alpha)(r+\delta)}{[1-\gamma(1-\alpha)]r+(1-\gamma)(1-a)\delta}\Delta c_t^w + \frac{\alpha r}{[1-\gamma(1-\alpha)]r+(1-\gamma)(1-a)\delta}\Delta d_t$$

as in equation (C.1).

Log-Linear CRRA Euler equation

Consider the pricing equation

$$1 = E_t\left\{M_{t,t+1}R_{t+1}^s\right\}$$

that in case of simple CRRA preferences becomes

$$1 = E_t\left\{\beta\left(\frac{C_{t+1}^c}{C_t^c}\right)^{-RRA}R_{t+1}^s\right\}$$

which note depends on capitalists', not aggregate, consumption growth. Using joint log-normality and homoskedasticity, with the latter implying constant variance and co-variance, I can rewrite

$$0 = \log\left[E_t\{\cdot\}\right] = E_t\left\{\log\beta - RRA\times\left(\Delta c_{t+1}^c\right)+r_{t+1}^s\right\}$$

yielding (ignoring constants)

$$RRA\times E_t\left(\Delta c_{t+1}^c\right) = E_t\left(r_{t+1}^s\right)$$

D Construction of consumption series from CEX

In this appendix I describe the dataset and preliminaries used to construct annual time series of real aggregate per capita consumption (representative agent) and real average consumption for households who own stocks (representative stockholder) and who do not own stocks (representative non-stockholder) over the period 1984-2017 from the U.S. Consumer Expenditure Survey.

Description of the dataset

The CEX is a national survey collecting household-level data on detailed consumption expenditures together with income, financial and demographic information on a sample that is designed to represent the non-institutionalized civilian population of the U.S. The survey is divided in two parts: Interview Survey and Diary Survey. The analysis developed here focuses on the Interview Survey. Data from CEX are available from the start of 1980 to the end of 2017. The survey is a rotating panel containing interviews of about 4,500 households per quarter before 1999, increasing to about 7,500 thereafter. About 20% of the sample is replaced each quarter. In each interview, households report detailed expenditures made in the previous three months. Households are interviewed every three months for a maximum of 5 interviews. The first interview is for practice and is not publicly available, while financial information is collected only in the last interview.

Sample choice, consumption definition and stockholder status

Consumption definition

The analysis employs the data available for the whole sample 1980Q1-2018Q1. The consumption measure employed consists of nondurable goods and some services aggregated from the disaggregated expenditure categories reported in the monthly expenditure files (MTAB and MTBI files) of the CEX. Following Malloy et al. (2009), I exclude the services categories that have substantial durable components, such as housing expenses (except for household operations and utilities), medical care costs and education costs. More specifically, the categories included are food, alcoholic beverages, household operations, utilities, apparel and services, gasoline and motor oil, public transportation, fees and admissions, reading, tobacco and personal care products.

Stock-holding status from the CEX

Regarding the distinction between stockholders and non-stockholders, similarly to Malloy et al. (2009) I define the stock-holding status based on holdings at the beginning of period t, since the standard Euler equation links the consumption growth rate between t and $t+1$ with stock returns at time $t+1$. The FMLY/FMLI files report household-level financial information on holdings of "stocks, bonds, mutual funds and other such securities". For the period 1980-2012, I use the same variables as the authors to define the stock-holding status. Recall that financial information is collected only in the fifth (last) interview. The first variable, SECESTX, reports whether the household holds (at the last day of the month preceding the interview) positive amounts of the aforementioned asset categories; the second variable, COMPSEC, asks whether the household holds the same amount, more or less of those assets compared to the same day of the previous year; the third variable, COMPSECX, quantifies, in dollar values, the change reported in the variable COMPSEC. Therefore, a household is defined as stockholder at the beginning of period t if: 1) holds a positive amount of the assets at the time of the interview and reports having the same amount as last year; 2) reports having lower holdings compared to last year; 3) reports an increase compared to last year, but by a dollar amount lower than the current holdings.[43]

From 2013 the variables SECESTX, COMPSEC and COMPSECX have been removed from the survey. However, at the same time two new variables, STOCKYRX and STOCKYRB, were added. The latter variable reports the "range which best reflects the total value of all directly-held stocks, bonds, and mutual funds one year ago today", while the former indicates the "median value of bracket range for STOCKYRB". Therefore, these two variables can be directly used to determine stock-holding status at the beginning of period t. In particular, for the period 2013 through 2017 I define as stockholders those households who report: 1) a positive value for STOCKYRX; 2) a positive range for STOCKYRB when the response for STOCKYRX is flagged as nonvalid (type "B" or "C" responses).

Exclusions and replication of Malloy et al. (2009)

In a first step, I replicate the quarterly stockholders' and non-stockholders' consumption growth series constructed by Malloy et al. (2009) for the sample March 1982 to November 2004 and available on Tobias Moskowitz's personal webpage.[44] This ensures

[43]Similarly to the authors, I also define as stockholders those households who report an increase in their asset holdings but do not specify either the current amount or the dollar difference from last year. Indeed, these few households are likely to have held these assets the previous year.

[44]Available at: https://faculty.som.yale.edu/tobymoskowitz/research/data/

that the consumption and stock-holding status definitions, together with the exclusions applied, are in line with previous literature.

I calculate average quarterly consumption growth rates for both groups of households, namely stockholders and non-stockholders, as

$$\frac{1}{H_t^g}\sum_{h=1}^{H_t^g}(c_{t+1}^{h,g} - c_t^{h,g})$$

where $c_t^{h,g}$ is quarterly log-real consumption of household h in group g (stockholders or non-stockholders) for quarter t and H_t^g is the number of households in group g at quarter t. Notice that this quantity is conceptually different from the growth rate in the average consumption of a certain group, which would be more in line with the concept of a representative agent. As discussed by the authors, the representative agent specification, which is the one employed in the analysis in the main text, assumes perfect risk-sharing within each group of households and thus ignores uninsurable idiosyncratic consumption shocks. To the contrary, the above definition sums the household-specific growth rates cross-sectionally, thus being able to capture comovements of asset returns with the within-group cross-household inequality. Nevertheless, the authors show that such comovements play a minor role in explaining risk-premia.

Quarterly consumption is constructed as the sum of real monthly expenditures reported in each of the four interviews, with nominal values being deflated by the monthly BLS Consumer Price Index for All Urban Consumers: Nondurables, Index 1982=100, Not Seasonally Adjusted (CUUR0000SAN from FRED). Hence, for each household at most four (three) quarterly consumption (consumption growth) observations are available. However, while the same household is interviewed every three months, interviews across households are made every month. Hence, household-level quarterly growth rates can be constructed at the monthly frequency. Changes in log-consumption are regressed over changes in log-family size and 12 monthly seasonal dummies at the household level and separately for each group of households. The residuals from this regression constitute the consumption growth measure.

I apply the same exclusions as the authors. To construct household-specific consumption growth rates it is necessary to match households across quarters. Only households who completed the survey, i.e. for which four interviews are available in the FMLY/FMLI files, are kept in the sample. Indeed, financial information is collected only in the fifth (i.e. the last) interview. Matching households across quarters is not possible around changes in sample design, which happened at the beginning of 1986, 1996, 2005 and

2015.[45] Such changes imply new household identification numbers. Therefore, all the households who did not finish their interviews before the ID changes are dropped. This boils down to treating the full sample 1980-2017 as five independent samples 1980-1986Q1, 1986-1996Q1, 1996-2005Q1, 2005-2015Q1 and 2015-2018Q1.[46]

Observations for which the consumption growth ratio C^h_{t+1}/C^h_t is less than 0.2 or more than 5 are dropped, as these could reflect reporting or coding errors. Negative or missing consumption observations are also dropped. Non-urban households, households residing in student housing and households with incomplete income responses are excluded form the sample. Regarding the latter exclusion, for the period 1980-2013 the variable RESPSTAT is used, which indicates whether the household is a complete or incomplete income reporter. From 2014 such variable is no longer available. Hence the variable ERANKH, which measures the weighted cumulative percent expenditure outlay ranking of the household to total population and is left blank for incomplete income reporters, is used. Moreover, all consumption observations for households interviewed in the years 1980 and 1981 are dropped as the food question was changed in 1982 leading to a drop in reported food expenditures.[47] Finally, all households who report a change in the household head's age different from 0 or 1 between any two interviews are excluded.

The final sample for the period March 1982 (the month for which the first quarterly consumption growth observation is available) to November 2004 consists of $196,813$ quarterly consumption growth observations across $75,346$ households, 21.91% of which are classified as stockholders (implying 78.09% as non-stockholders). These numbers are very close, respectively, to $206,067$, $76,568$ and 22.7 as reported by Malloy et al. (2009), who consider the same sample period. As shown in Figure D.1, the consumption growth rate series obtained for stockholders and non-stockholders track quite closely the authors', with correlation coefficients of 81% and 87%, respectively, over the whole sample.

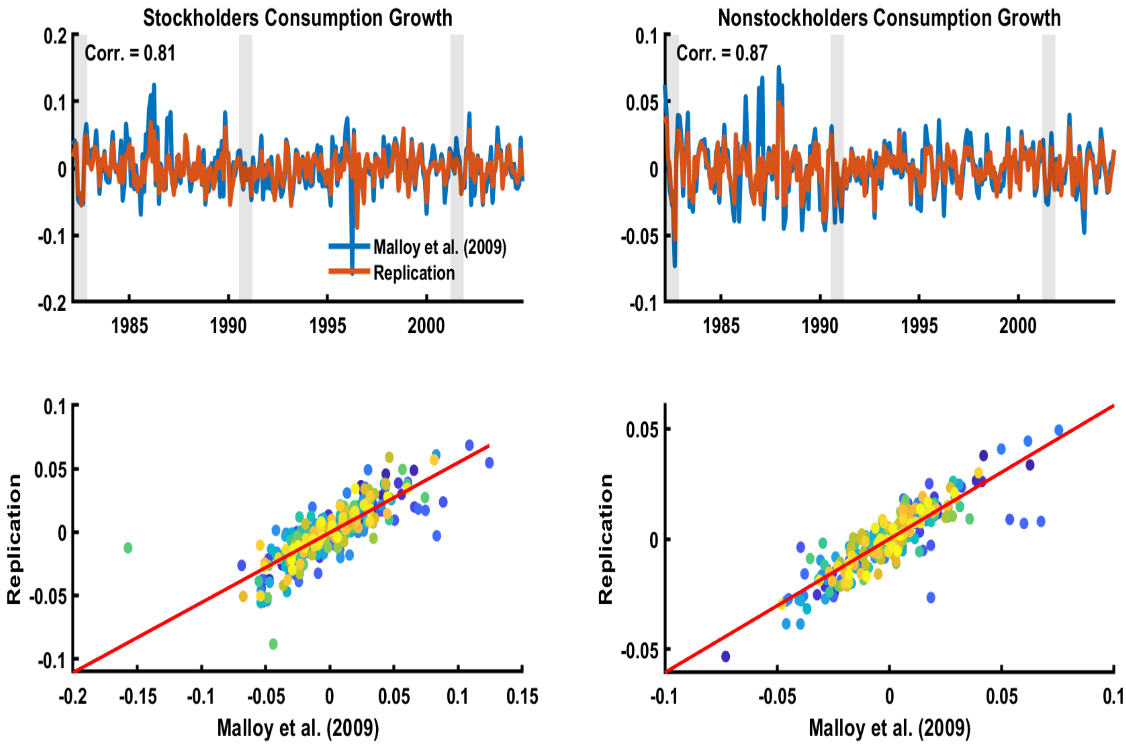
The left panel of Figure D.2 shows the population-weighted stock market participation rate obtained from the CEX (in red) in comparison to the one obtained from the SCF (in blue). Clearly, the two measures of participation substantially differ. Indeed, while the SCF includes both direct and indirect stock-ownership, the latter cannot be retrieved from

[45]The year-specific documentation files report this type of information. These files can be found at: http://www.nber.org/ces

[46]It is important to note that each year of the survey includes five quarters, as the first quarter of the following year is necessary to calculate average expenditures for the year of interest. Regarding the sample design changes, and taking 1986 as an example, the data for 1986Q1 reported in the 1985 survey will be different from the 1986Q1 data for the 1986 survey, as the two surveys will employ different sample designs. Therefore, in my analysis the sample 1980-1986Q1 includes 1986Q1 as reported in the 1985 survey, while the sample 1986-1996Q1 will include the 1986Q1 as reported in the 1986 survey. Same reasoning for all the other breaks in the sample design.

[47]As noted by the authors, the food question was changed back to the initial one in 1988 but there is no sensible way to solve this issue without losing a substantial number of observations.

Figure D.1: Replication of Malloy et al. (2009)



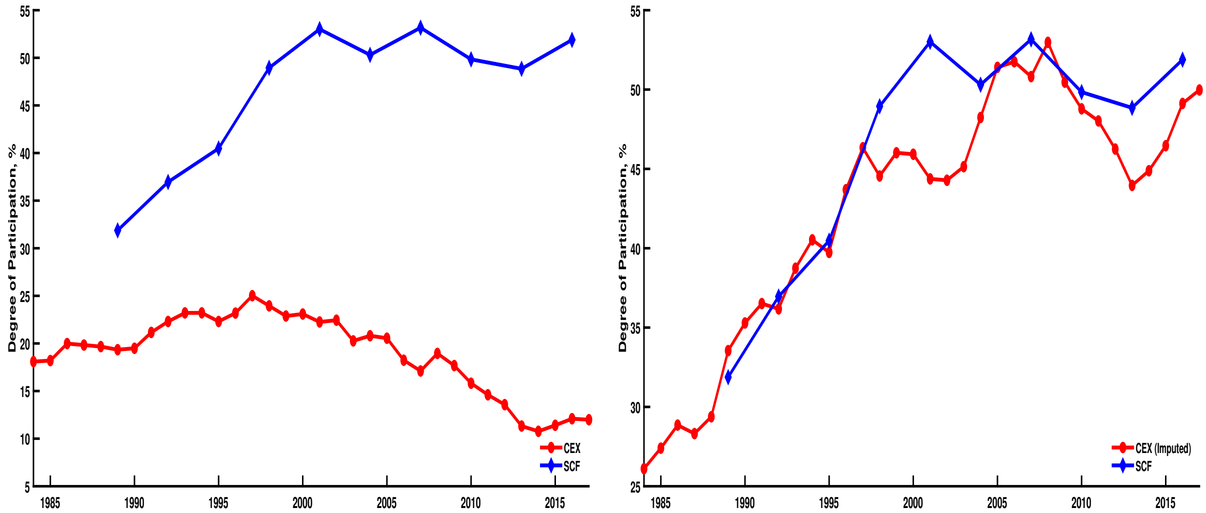
Notes: The figure reports the average quarterly growth rates for stockholders (top-left panel) and non-stockholders (top-right panel) at the monthly frequency for the period March 1982 to November 2004. The series obtained here (in orange) closely tracks the authors' (in blue), with correlation coefficients of 81 and 87 percent for stockholders and non-stockholders, respectively. The bottom panels depict the corresponding scatterplots.

the CEX, as also noted by Malloy et al. (2009). Consistent with the authors, I find that the participation rate in the CEX is somewhat upward trending until the early 2000s, from around 18% to 25%. Nevertheless, the same rate substantially drops from those years until 2017, when only about 11% of the sample is classified as stockholders. This result could be consistent with the evidence from the SCF if interpreted as a decrease in the rate of direct stock-ownership. Indeed, in 2013 the financial assets question about was changed to consider only direct holdings. Also, Lettau et al. (2018) argue that the CEX provides inferior measures for financial holdings compared to other surveys, including the SCF.

Imputation procedure from the Survey of Consumer Finances

The replication of the consumption growth series constructed by Malloy et al. (2009) ensures that the exclusions applied on the dataset and the consumption and stockholder status definitions are in line with previous literature. However, the participation rate

Figure D.2: Stock-ownership rate: CEX vs SCF



Notes: Comparison between the stock market participation rate from the CEX (red lines) and from the SCF (blue lines). The left panel reports the raw participation rate from the CEX. The right panel reports the one estimated from the CEX by imputation.

estimated from the CEX variables does not include the indirect stock-ownership, which likely determined a good part of the upward trend observed in the SCF dataset. To refine the stock-holding status definition, I follow the imputation procedure proposed by Attanasio et al. (2002) and Malloy et al. (2009).

Specifically, I employ a probit analysis from the SCF, since this dataset contains wealth information on both direct and indirect stock-holdings (the variable "equity"), to predict the probability that a household holds stocks directly or indirectly in the CEX. I use the SCF 1989 through 2016 (the last year available). I generate a dummy variable equal to 1 if such holdings are positive. Following Malloy et al. (2009), I then estimate a probit model where the dependent variable is the stock-holding dummy and the regressors are observable characteristics that are available also in the CEX: age, age squared, an indicator for the head of household having education of > 12 but < 16 years ("highschool"), and one for > 16 years ("college"), an indicator for race not being white/Caucasion, year dummies, log real total household income before taxes, log of real dollar amount in checking and saving accounts (put to 0 if the sum of checking and savings equals zero), an indicator for checking+savings accounts equal to zero, an indicator for positive interest+dividend income, and a constant. SCF weights are employed in the probit model to have population estimates. The estimated coefficients (with t-statistics in parentheses) from the probit regression are:

$$
\begin{aligned}
x'_{SCF}b = & \underset{(-126.74)}{-8.68} + \underset{(20.24)}{0.0269}age + \underset{(-22.88)}{-0.0003}age^2 + \underset{(-24.35)}{0.319}\,highschool + \underset{(-45.28)}{0.583}\,college \\
& + \underset{(-37.53)}{-0.323}nonwhite + \underset{(10.78)}{0.225}Y_{1992} + \underset{(19.25)}{0.393}Y_{1995} + \underset{(30.5)}{0.619}Y_{1998} + \underset{(35.43)}{0.713}Y_{2001} + \underset{(32.38)}{0.649}Y_{2004} \\
& + \underset{(37.35)}{0.758}Y_{2007} + \underset{(38.11)}{0.717}Y_{2010} + \underset{(36.64)}{0.698}Y_{2013} + \underset{(40.35)}{0.767}Y_{2016} + \underset{(90.06)}{0.587}log(income) \\
& + \underset{(34.83)}{0.082}log(chk + sav) + \underset{(12.17)}{0.264}(chk + sav = 0) + \underset{(63.80)}{0.599}log(int + div > 0)
\end{aligned}
$$

The estimated coefficients are very similar to the ones estimated by Malloy et al. (2009). I then use these coefficients to predict the probability that a household in the CEX holds stocks as $\Phi(x'_{CEX}b)$, where Φ is the CDF of the standard normal distribution and x_{CEX} is the vector of the same regressors as in the SCF. When predicting the stock-holding probability for a household in the CEX I use the dummy 1992 coefficient for the years 1990-1993, the dummy coefficient 1995 for the years 1994-1996, the dummy 1998 coefficient for the years 1997-1999 and so on. Similar to the SCF, dollar amounts for the variables in the regression in year t are multiplied by the absolute variation between year $t-1$ and year t in the (yearly average of the monthly) current-methods version of the CPI for all urban consumers (CPI-U-RS).[48]

In the baseline measure used in the main text, a household in the CEX is classified as indirect or direct stockowner if the predicted probability is greater that 41%, which represents a mid-value in the trend observed in the SCF. Specifically, I use the probability predicted for the last month of observation for the households, since financial information is reported only in the last interview. Notice that this imputation procedure is applied only to those households who 1) are classified as non-stockholders according to the baseline CEX definition and 2) have non-missing responses to the checking and savings account questions. Therefore, households who are classified as stockholders based on the CEX definition remain classified as such with probability 1; and households who are non-stockholders in the CEX but have no valid responses to the checking and savings accounts receive probability 0 of being stockholders.

The result of the procedure is depicted in the right panel of Figure D.2, which compares the rate of direct and indirect stock-ownership from the SCF (in blue) and the one imputed in the CEX (in red) for the sample 1984-2017. The imputed series closely tracks the SCF one especially in the first part of the sample, where the rates are essentially identical. However, since the end of the 1990s the two series slightly diverge, although from the late 2000s the two series follow very similar dynamics. The difference in the

[48]Available at: https://www.bls.gov/cpi/research-series/home.htm

levels, rather than dynamics, of the two participation rates could be due to differences in the design of the two surveys. As discussed in Lettau et al. (2018), the SCF is designed to measure the wealthiest households and has high quality financial information. On the other hand, the CEX has notorious limitations when measuring the top-end of the wealth distribution due to under-reporting, with very wealthy households being more likely to hold stocks. Moreover, Bee et al. (2012) document that such under-reporting increased since the 2000s, suggesting that the imputation based on income and financial observables can be expected to underestimate the true participation rate.

Nevertheless, the result of the imputation is quite satisfactory. For example, it is worth noting that the participation rate estimated for 1984 is about 25%, which justifies the 20% adopted in the calibration of the model for the pre-Great Moderation period.[49] Also, the maximum participation rate estimated in the CEX is around 53% as in the SCF, with both values occurring right before the financial crisis (2007 in the SCF, 2008 in the CEX). Moreover, both series capture a U-shaped pattern in the stock-ownership rate following the crisis and display a strong upward trend until the early 2000s, when they reach a new plateau. Overall, the imputed series captures the key properties of the stock-ownership rate in the U.S.

Annual consumption estimates

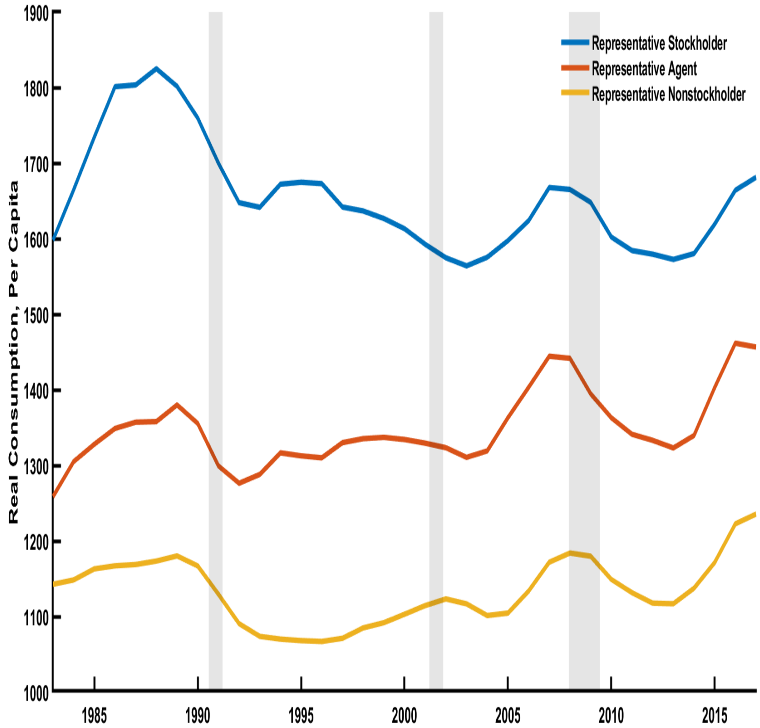
The ultimate aim of my analysis is to obtain a time series of real consumption for a representative stockholder and a representative non-stockholder over the sample 1984-2017, by employing the stock-holding status definition obtained from the imputation procedure described above. To do so, I compute population (weighted) annual mean expenditure estimates aggregated from monthly expenditures[50] across 120,934 households, following the formulae provided in the CEX documentation.[51] Nominal expenditure values are deflated by the CPI for nondurables, and divided by family size in order to obtain per capita expenditures. Mean estimates are calculated for a representative agent, i.e. over the whole sample of households; for a representative stockholder, i.e. when considering only the group of households who own stocks; and for a representative non-stockholder, i.e. when considering only the group of households who do not own stocks, according to the imputed participation rate. Similarly to Cloyne et al. (2019), to eliminate some of

[49] Recall that Poterba et al. (1995) estimate a participation rate of 20% for 1962. Hence, this rate was clearly quite stable until the mid-1980s.

[50] In particular, I focus on calendar months. Calendar periods are the periods (months, quarters or years) when expenditures were actually made, while collection periods correspond to the periods when expenditures where reported in the interview. See the CEX documentation for a detailed discussion.

[51] In particular, I employ the example codes provided at the link: https://www.bls.gov/cex/pumd-getting-started-guide.htm#section5

Figure D.3: Annual consumption series



Notes: This figure plots the real consumption for representative agent (orange), representative stockholder (blue) and representative non-stockholder (yellow) as estimated from the CEX with stock-holding status imputed from the SCF.

the noise inherent to survey data, the consumption series are smoothed with a backward looking (current and previous year) moving average.

Figure D.3 shows the results. The annual consumption series for representative agent, representative stockholder and representative non-stockholder are compared. As we can see, the representative stockholder (non-stockholder) consumes more (less) than the average. This is consistent with the evidence that only richer households tend to invest in the stock market. Moreover, stockholders' consumption process appears to be less smooth than non-stockholders'.

Cross-sectional evidence at the state-level

The CEX reports the variable "STATE", which identifies the state in which a particular household resides at the time of the interview. This variable therefore allows me to conduct the empirical validation also in the cross-section of the different U.S. states. As explained in the survey manual, the state identifier is however not reported for all states because of topcoding reasons. Moreover, the population weights in the survey are designed to be representative for the entire U.S. population and not for the state-level population. Therefore, in the cross-sectional evidence section all estimates are unweighted. Finally, the STATE variable is available only since 1993.

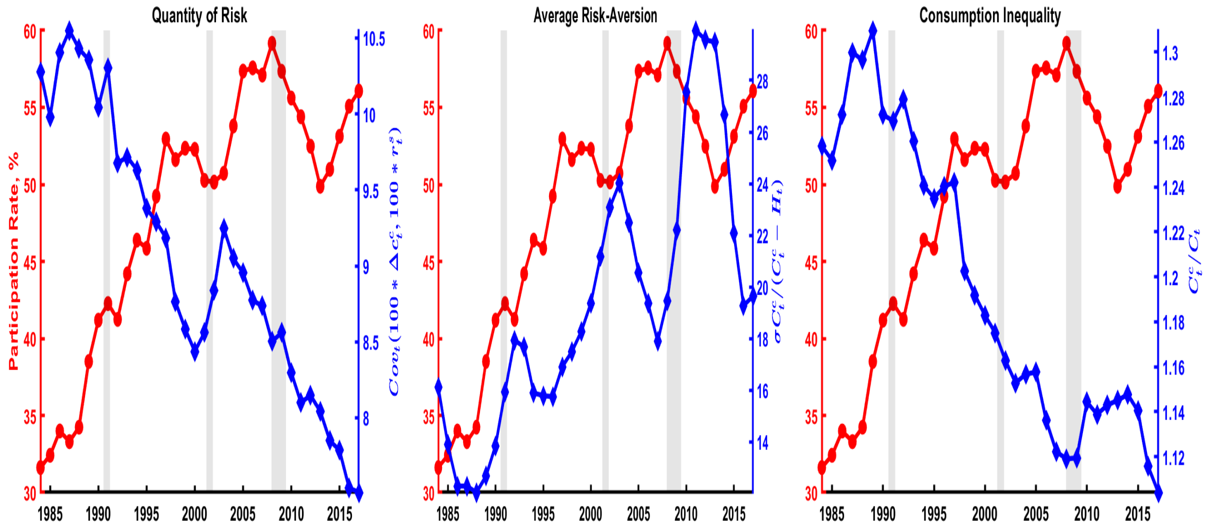
The state-level sample is restricted to states for which the data is available over the period 1993-2017 and in which at least 100 (household-level) observations are available on average over the sample. These restrictions are imposed in order to ensure that the state-level sample size is not too small. Taking into account that not all states are available in the survey, the final sample comprises 27 states. The consumption series and stockholding status dummy are constructed exactly as in the time-series analysis.

E Evidence from the CEX: robustness

Participation probability threshold = 0.32

In this case, the household is classified as stockholder if the predicted stock-holding probability is greater than 32%, i.e. the SCF participation rate in 1989.

Figure E.1: Validating the mechanism: robustness (I)



Notes: Evidence from the CEX for a participation probability threshold equal to 0.32 instead of 0.41.

Table E.1: Long-run covariation with participation rate: robustness (I)

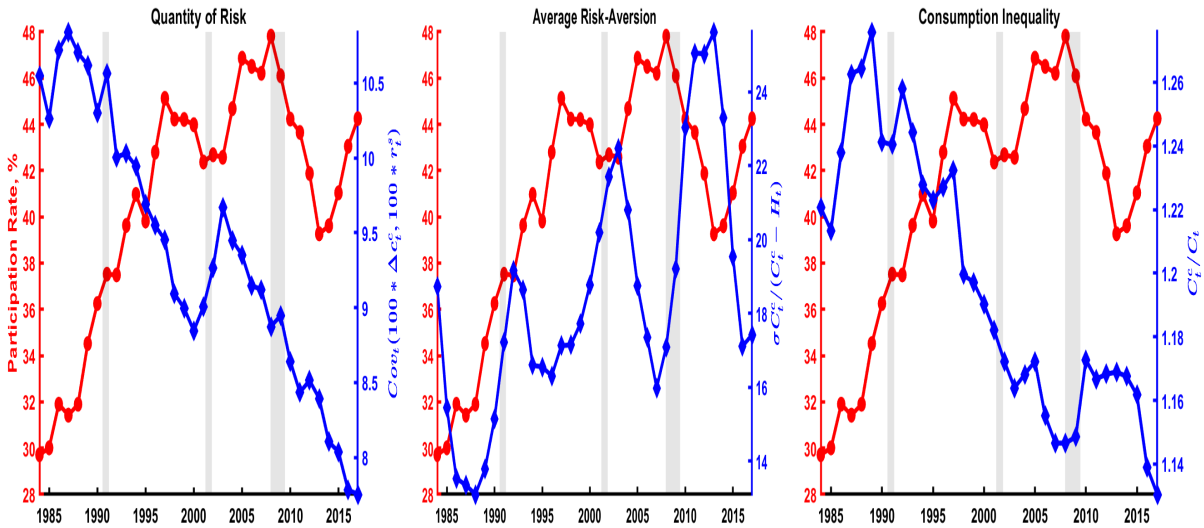
Variable	ρ_{LR}		β_{LR}	
	$\hat{\rho_{LR}}$	67% CI	$\hat{\beta_{LR}}$	67% CI
$Cov_t(\Delta c_t^c, r_t^s)$	-0.68	$[-0.92, -0.198]$	-0.094	$[-0.143, -0.021]$
$\sigma C_t^c/(C_t^c - H_t)$	0.682	$[0.15, 0.916]$	0.421	$[0.064, 0.643]$
C_t^c/C_t	-0.704	$[-0.92, -0.209]$	-0.007	$[-0.01, -0.002]$

Notes: Long-run covariation measures for a participation probability threshold equal to 0.32 instead of 0.41.

Probability-weighted participation and consumption

Here, I employ a "continuous" measure of participation. To compute the participation rate and the representative stockholder's consumption, each household's population weight and consumption expenditure is multiplied by the imputed probability of being stockholder.

Figure E.2: Validating the mechanism: robustness (II)



Notes: Evidence from the CEX for the probability-weighted participation and consumption case.

Table E.2: Long-run covariation with participation rate: robustness (II)

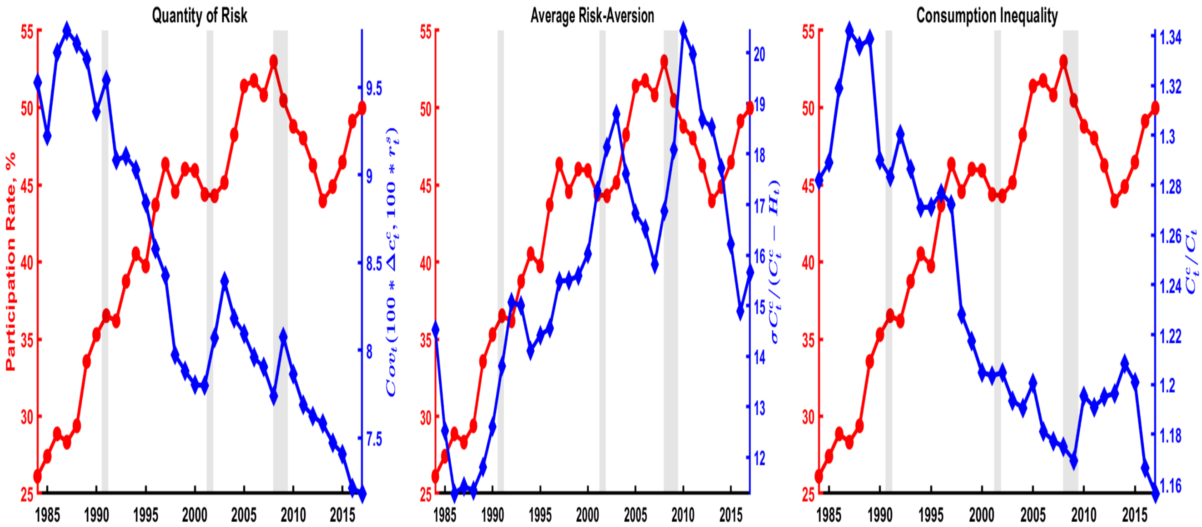
Variable	ρ_{LR}		β_{LR}	
	$\hat{\rho_{LR}}$	67% CI	$\hat{\beta_{LR}}$	67% CI
$Cov_t(\Delta c_t^c, r_t^s)$	-0.46	$[-0.85, 0.013]$	-0.118	$[-0.228, 0.025]$
$\sigma C_t^c/(C_t^c - H_t)$	0.569	$[0.063, 0.874]$	0.323	$[0.007, 0.546]$
C_t^c/C_t	-0.456	$[-0.85, 0.013]$	-0.005	$[-0.01, 0.001]$

Notes: Long-run covariation measures for the probability-weighted participation and consumption case.

Different consumption definition

Nondurables and services are constructed as in Wong (2019). The categories included are the same as in the baseline consumption definition plus health and education expenses. Moreover, each category is now deflated by the respective BLS CPI index, following the classification in Krueger and Perri (2006).

Figure E.3: Validating the mechanism: robustness (III)



Notes: Evidence from the CEX with different consumption definition.

Table E.3: Long-run covariation with participation rate: robustness (III)

Variable	ρ_{LR}		β_{LR}	
	$\hat{\rho_{LR}}$	67% CI	$\hat{\beta_{LR}}$	67% CI
$Cov_t\left(\Delta c_t^c, r_t^s\right)$	-0.841	$[-0.97, -0.35]$	-0.095	$[-0.127, -0.045]$
$\sigma C_t^c/(C_t^c - H_t)$	0.892	$[0.5, 0.98]$	0.266	$[0.195, 0.331]$
C_t^c/C_t	-0.829	$[-0.97, -0.35]$	-0.006	$[-0.009, -0.003]$

Notes: Long-run covariation measures with different consumption definition.