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ASSET MARKET PARTICIPATION, REDISTRIBUTION, AND ASSET PRICING

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JEL Classification: D31, E13, E21, E25, E32, E44, G12, G51

Keywords: Consumption, Income, Heterogeneity, Limited participation, Asset pricing

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Asset Market Participation, Redistribution, and Asset Pricing*

Francesco Saverio Gaudio† Ivan Petrella‡ Emiliano Santoro§

March 7, 2023

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1 Introduction

A long-standing tradition in macro-finance has sought to provide a unified explanation of macroeconomic and financial fluctuations (see, e.g., Cochrane, 2017, and references therein). While this literature has originally taken a representative agent perspective, which presumes aggregate (average) consumption growth to be an appropriate measure of systematic risk, recent advances have emphasized how neglecting household heterogeneity may severely limit our ability to understand the connection between fluctuations in asset prices and the macroeconomy. Indeed, the representative agent assumption stands in contradiction with the most basic observation about asset ownership—about a third of U.S. households do not own any form of liquid assets, on average—along with implying a poor performance in explaining stylized asset-pricing facts (Brunnermeier et al., 2021). In response to such a discrepancy, various contributions have stressed the need to (re)consider limited asset market participation as a crucial dimension of consumer heterogeneity (see, e.g., Mankiw and Zeldes, 1991; Attanasio et al., 2002; Guvenen, 2009; Malloy et al., 2009).

An inherent property of economies where a substantial share of the population has no access to financial investment is the emergence of a wedge between the consumption of the average market participant and aggregate consumption, which is in turn related to a metric of consumption inequality: the consumption of assetholders *relative* to that of non-assetholders.[1] In this paper, we argue that studying fluctuations in relative consumption is essential to understand how allowing for heterogeneity in the access to financial markets may improve upon the representative agent benchmark, thus reconciling salient macroeconomic and asset-pricing facts. To this end, we document how different aggregate shocks induce varying degrees of cyclicality in relative consumption, and explain how this novel evidence is strategic to validating production-based asset-pricing frameworks.

We highlight important differences in the response of relative consumption (and income) to technology shocks—both neutral and investment-specific—as well as to factor-share shocks—i.e., shocks that reallocate the rewards of production towards capital and away from labor income. All of these have been considered in the macro-finance literature as common drivers of macroeconomic and financial fluctuations (see, e.g., Jermann, 1998; Guvenen, 2009; Papanikolaou, 2011; Lansing, 2015; Greenwald et al., 2019). Unlike technology shocks, factor-share shocks are able to generate strongly procyclical relative consumption, thus contributing to the emergence of a siz-

[1] In the remainder, when referring to this measure, we will interchangeably use the expressions relative consumption and consumption inequality.

able wedge between the volatility of assetholders' consumption growth and that of aggregate consumption growth. As a result, fluctuations in consumption inequality triggered by shifts in the income shares of factors of production induce both qualitatively and quantitatively meaningful time-variation in future expected excess returns. This evidence can be exploited to discriminate among theories that are seemingly consistent at the *aggregate* level, but that may bear very different implications at the *household* level (i.e., in terms of consumption and income inequality). A model with concentrated capital ownership that replicates the conditional properties of relative consumption identifies factor-share shocks as key drivers of consumption inequality. Being redistributive in nature, though, these shocks play a limited role in explaining macroeconomic fluctuations. Yet, they represent the main source of risk priced in financial markets.

Using the U.S. Consumption Expenditure Survey (CEX) and the Survey of Consumer Finances (SCF), we construct the consumption and income series pertaining to two distinct groups of households, based on their holdings of financial assets (including stocks), in line with the two-agent structure embodied by models featuring limited asset market participation (e.g., Mankiw and Zeldes, 1991; Mankiw, 2000). Thus, we retrieve the dynamic responses of both aggregate and household-level variables to the shocks of interest. While all these shocks induce similar and significant expansionary effects on real GDP, investment, and aggregate consumption, they bear different income and consumption redistribution properties across households sorted with respect to their assetholdings. Neutral technology shocks attenuate households' consumption and income inequality—again, as captured by the gap between average assetholders' and non-assetholders' consumption and income, respectively—while factor-share shocks amplify inequality along these dimensions. By contrast, investment-specific shocks display limited capacity to stimulate household inequality. Moreover, the behavior of both assetholders' and non-assetholders' consumption tends to be predominantly shaped by the conditional behavior of their incomes—rather than by heterogeneity in their propensities to consume—especially in response to technology shocks, for these tend to induce permanent-income effects.

Differences in the conditional cyclicality of relative consumption are key to understanding fluctuations in asset prices. Together with standard measures of aggregate risk—such as changes in aggregate consumption—cyclical movements in consumption inequality capture substantial time-variation in expected excess stock returns. In fact, factor-share shocks are ultimately responsible for the predictive capacity of changes in relative consumption. Altogether, these facts set some clear targets for designing and validating structural models that can account for cyclical inequality in household consumption and income, as well as for core macroeconomic and asset-pricing facts.

A production-based asset-pricing model with limited asset market participation replicates the behavior of the consumption of assetholders relative to that of non-assetholders, conditional on each of the three shocks identified in the empirical analysis. Moreover, it highlights how different sensitivities of wages and dividends to a given shock rationalize the conditional behavior of consumption inequality, and how this reflects into the pricing of risky assets. The model highlights a deep disconnect between asset-pricing and macroeconomic fundamentals. On one hand, technology shocks—both neutral and investment-specific—are mainly responsible for driving macroeconomic fluctuations, exerting a relatively limited impact on relative consumption and, in light of this, displaying little grip on the level and volatility of asset returns. On the other hand, fluctuations in relative consumption—particularly those fuelled by shifts in the factor shares, which redistribute resources from labor to capital income in good times—emerge as a significant predictor of expected excess stock returns. Relatedly, we show how varying degrees of participation in financial markets have small influence on the volatility of macroeconomic aggregates, while exerting substantial impact on asset-pricing moments.

Related literature This paper relates to several strands of the literature. First and foremost, we speak to the large body of studies exploring the potential of limited asset market participation to tackle a number of financial puzzles within general-equilibrium frameworks (Danthine and Donaldson, 2002; Guvenen, 2009; De Graeve et al., 2010; Lansing, 2015). We contribute to this broad line of inquiry by showing how replicating the conditional dynamics of relative consumption represents an essential input for the design of production-based asset-pricing models. In particular, our work suggests that previous contributions employing technology-neutral shocks as a primary source of economic fluctuations could match asset-pricing moments only by implying counterfactual dynamics at the household level. In these settings, sizeable equity premia are generated by embedding specific mechanisms, such as operating leverage (Danthine and Donaldson, 2002) or preference heterogeneity (Guvenen, 2006, 2009), that entail stronger sensitivity of assetholders' consumption to aggregate fluctuations, relative to that of non-assetholders and, as a byproduct, procyclical household inequality. However, a main takeaway from our empirical analysis is that relative consumption is markedly countercyclical, conditional on technology-neutral shocks, while it is procyclical in response to factor-share shocks.[2]

Our work complements recent contributions highlighting redistribution between factors of production as a source of risk being priced in the stock market (Lansing, 2015; Lettau et al., 2019), and as a driver of fluctuations in stock valuations (Greenwald et al., 2019). By disproportionately affecting financial income relative to labor income, factor-share shocks imply a marked redistribution of resources between assetholders and non-assetholders. In fact, empirically relevant risk premia and volatility are shown to primarily emerge from fluctuations in the factor shares that are orthogonal to technology shocks. In doing so, we stress the importance of accounting for dynamic interaction between factor shares and TFP, in line with Ríos-Rull and Santaeulalia-Llopis (2010) and Santaeulalia-Llopis (2011). Moreover, we report a macro-finance disconnect, in that shocks affecting business-cycle moments are not equally important drivers of asset-pricing moments (and *vice versa*), within a standard real business cycle framework with concentrated capital ownership.[3] Relatedly, our analysis adds to recent literature seeking to establish the role of aggregate-risk measures (e.g., Atanasov et al., 2020), or of wealth inequality (Toda and Walsh, 2019) in terms of stock-return predictability. In this respect, we show how consumption (and income) inequality, whose short-run fluctuations mostly reflect the impact of factor-share shocks, do bear sizeable predictive power on expected excess returns.

Finally, we speak to recent developments in the macroeconomic literature that examine the role of household heterogeneity for the transmission of monetary and fiscal policy (see Mankiw, 2000; Galí et al., 2007; Bilbiie, 2008; Debortoli and Galí, 2017; Broer et al., 2019; Bilbiie, 2020; Cantore and Freund, 2021; Bilbiie et al., 2022, among others). Unlike these, we focus on technology and factor-share shocks, which have been considered as key drivers in production-based asset-pricing models. We highlight that, while not playing a major role in amplifying business fluctuations, agent heterogeneity and the conditional dynamics of consumption inequality are key to reproducing volatile stock returns, along with a sizable equity premium and a sufficiently low risk-free rate. While much is known about the empirical transmission of the shocks we consider on macroeconomic aggregates, less is understood about their impact on different groups of households. In this respect, we complement the work of Cloyne and Surico (2017) and Cloyne et al. (2019), who highlight that the transmission of mon-

[2]Otherwise, investment-specific shocks—which play a key role in Justiniano and Primiceri (2008), Papanikolaou (2011), Kogan and Papanikolaou (2013), Garlappi and Song (2017), and Kogan et al. (2020)—do not induce sizeable cyclicality in relative consumption.

[3]A related work, in this sense, is Bianchi et al. (2018): while retaining a representative-agent perspective, they stress that shocks driving the business cycle are unlikely to account for the volatility of stock prices.

etary and fiscal shocks mainly hinges on their impact on the disposable income of consumers who are financially/liquidity constrained. By contrast, we stress that heterogeneous consumption and income responses to the shocks we consider map into the asymmetric reaction of labor and financial income.

Structure The rest of the paper is organized as follows. In Section 2 we present the aggregate and the household survey data employed in the analysis. Section 3 presents the identification of the aggregate shocks of interest, the responses of macroeconomic and household-level variables, and connects this evidence to time-variation in asset prices. Section 4 examines the role of different shocks and that of household heterogeneity within a quantitative setting with concentrated capital ownership. Section 5 concludes.

2 Data

In this section, we describe the aggregate macroeconomic and financial data, as well as the survey data and the procedure to aggregate them into representative household variables.

2.1 Aggregate macro and financial data

The identification of the structural (technology and factor-share) shocks employs data on TFP and the relative price of investment from Fernald (2014), together with the labor share of income series from the Bureau of Labor Statistics (BLS). The impulse-response analysis with macroeconomic data employs NIPA quarterly aggregate series on Consumption (non-durable goods and services, as well as durables), Gross Domestic Product (GDP) and Total Investment, in addition to the Consumer Price Index (CPI) for all items, from the BLS. Per-capita real measures are obtained by dividing their aggregate counterparts by the U.S. total population (NIPA) and by the CPI. We also investigate the responses of labor and dividend income, both being collected by the Bureau of Economic Analysis (BEA). As for financial data, both the stock returns and risk-free rate quarterly series are retrieved from Amit Goyal's webpage (see Welch and Goyal, 2008). In all cases, the sample spans over the 1982Q4-2017Q4 period, in line with the availability of household-level data. Further details on these sources are reported in Appendix A.

2.2 Household survey data

To estimate consumption expenditure and income at the household level, we rely on the U.S. CEX over the sample 1980-2017. Produced by the BLS, the CEX is a national survey featuring household-level data on consumption expenditure—along with income and other financial and demographic information—on a sample that is designed to represent the non-institutionalized civilian population. This section summarizes the main steps to obtain the consumption and income series for the two representative household groups of interest, namely assetholders and non-assetholders.[4]

2.2.1 Assetholding status definition and imputation

In the baseline analysis, we focus on a key dimension of household heterogeneity, defined by housholds' holdings of financial assets (including stocks). In line with a wide set of macroeconomic two-agent models (Bilbiie, 2008; Lansing, 2015; Debortoli and Galí, 2017, among others), we distinguish between assetholders and non-assetholders. Unlike assetholders, non-assetholders typically hold very little liquid financial assets. To accommodate this sorting criterion, we rely on both the CEX and the SCF.[5]

As in Mankiw and Zeldes (1991), we define a household to be an assetholder if the dollar value of held assets plus liquid accounts exceeds 1000$. The CEX collects information on whether a household holds "stocks, bonds, mutual funds and other such securities", along with checking and savings accounts. However, the CEX does not encompass indirect assetholdings, with the likely implication of underestimating households' participation in financial markets. To refine the assetholding status definition, we follow an imputation procedure similar to the one employed by Attanasio et al. (2002) and Malloy et al. (2009).[6]

Using SCF data over the 1989-2016 sample, we estimate a probit model for the probability of a household holding assets, directly or indirectly, based on a set of observables that are also available through the CEX. We include age and education (as well as the interaction term between the two), race (white or non-white), year dummies, (log) income, and a dummy variable capturing whether the household earns any financial income (defined as dividend plus interest income). The assetholding status

[4] Appendix B.1 and Appendix B.2 provide a detailed description of the dataset and the restrictions applied to the sample.

[5] The SCF is an independent triennial survey run by the Federal Reserve that collects detailed information on income and wealth holdings of U.S. households, while not including consumption expenditure.

[6] Further details on the definition of the assetholding status are available in Appendix B.3.

is captured by a dummy taking value 1 if (direct or indirect) holdings of stocks, bonds, and liquid accounts exceed the threshold of 1000$. The estimated coefficients are then used to predict the probability that a household in the CEX holds assets. In the baseline analysis, we construct a 'continuous' measure of asset-market participation. To obtain such a measure for the representative assetholder, each household's population weight is multiplied by the imputed probability of holding assets in amounts that exceed the threshold, and then divided by the total population. As for the representative non-assetholder, we employ the complement to one of such probability. The imputation only applies to households who have valid responses to questions connected with all variables used in the regression with CEX data. If this prerequisite is not met, the household is imputed an assetholding probability equal to zero.[7]

Following the outlined procedure, we obtain a series of the participation rate that closely tracks the one based on the SCF, especially in the last part of the sample, where the two rates are essentially identical.[8] Even in the first half of the sample, where the imputed rate is relatively lower, the difference amounts to few percentage points, and the imputation captures the upward trend observed in SCF data. The level discrepancy between the two participation rates mainly reflects differences in survey design. As stressed by Lettau et al. (2019), the SCF contemplates relatively wealthy households. On the other hand, the CEX has some well-known limitations, when trying to measure the top-end of the wealth distribution, mostly due to under-reporting. From a quantitative perspective, our procedure classifies between 25% and 40% (35%) of the households as non-assetholders in the CEX (SCF). These values are very close to the range considered in the existing literature. For example, Kaplan et al. (2018) estimate that around a third of the U.S. population consists of hand-to-mouth households, while Aguiar et al. (2020) estimate such percentage to be around 40%.

Some more comments are in order. First, it is important to stress that our focus is on households' liquid financial assets, rather than on their total net wealth, as in Kaplan et al. (2018), Aguiar et al. (2020), and Kehoe et al. (2020). Crucially, our approach allows us to speak to both the macroeconomic and the asset-pricing literature. In the latter, the emphasis is typically on the dichotomy between stockholders and non-stockholders, given the focus on the determination of risky asset prices (Mankiw and Zeldes, 1991). In this sense, our classification can easily accommodate such distinction, and throughout the analysis we verify that all our results also hold in this case. Second, it is well known that, at present, no comprehensive data on consumption, income, and wealth

[7] A comprehensive description of the imputation protocol can be found in Appendix B.4.

[8] See Figure B.1 in Appendix B.4, which compares the two rates of asset-ownership over the 1982Q4-2017Q4 sample.

at the household level are available for the US. Our imputation procedure allows us to combine wealth information from the SCF with consumption and income data from the CEX. Nevertheless, an implicit assumption is that households with the same demographic and income characteristics in the two datasets are seen as equally likely to have sufficient liquid wealth. In connection with this, we see our continuous measure of participation—which weighs household-level variables by the imputed probability of being an assetholder, rather than defining the assetholding status univocally by assigning a zero or a one—as a reliable option to deal with the potential imprecision of the classification. In any case, it is possible to adapt our sorting strategy to rely on the financial information available from the CEX, and apply the imputation procedure only residually (i.e., on households whose financial information is not available). Finally, it will be important to test the sensitivity of our results to other relevant aspects of household portfolios—most prominently housing—so as to account for the wealthy vs. poor hand-to-mouth distinction, as in Cloyne et al. (2019). All of these checks are performed at the end of Section 3, and lead to no qualitatively different results.

2.2.2 Household-level consumption and income

We focus on household expenditure on both non-durable goods and services, as well as on durable goods, together with after-tax income. To this end, we compute quarterly consumption expenditure (and income) based on calendar periods for the representative agent of each category (i.e., assetholders and non-assetholders) as the population-weighted expenditure (and income) within the group.[9] Spending and income variables are expressed in per-capita real terms by dividing nominal dollar amounts by family size and the CPI. More in detail, we construct a raw measure of per-capita assetholders' consumption (and income) by multiplying households' population-weighted consumption (and income) by the imputed probability of holding assets in amounts that exceed the (1000$) threshold, and divide this by the total population of assetholders. Per-capita non-assetholders' variables are constructed in a symmetric way, using the complement to one of the imputed assetholding probability. Thus, in every quarter the group-level series of expenditure and income are adjusted by the ratio of the corresponding NIPA national account aggregate to the corresponding aggregate from the CEX. The adjusted series are then smoothed through a backward-looking moving average,[10] so as to deal with seasonal adjustment and the

[9]Calendar periods are intended as quarters in which spending actually takes place, while collection periods refer to the quarters in which spending is reported. See the CEX documentation at `https://www.bls.gov/cex/`, for a detailed discussion.

[10]This includes both the current and the previous three quarters.

noise that typically characterizes survey data.[11]

3 Identification and transmission of aggregate shocks

In this section, we first discuss the strategy for identifying the shocks of interest. Thus, we examine the conditional dynamics of different macroeconomic variables, as well as that of assetholders' and non-assetholders' consumption and income. Finally, we explore the extent to which the (unconditional and conditional) dynamics of household consumption inequality help predict U.S. excess stock returns. For all these exercises, several robustness checks are performed at the end of the section.

3.1 Identification

We consider three shocks that have been widely regarded as key drivers of both macroeconomic and asset-pricing variables, namely neutral and investment-specific technology shocks, as well as redistribution shocks in the form of shifters to the factor shares of income. A long-standing literature (see Gali, 1999; Fisher, 2006, among others) has studied the transmission of technology shocks to the macroeconomy. However, these contributions typically assume that factor shares are constant over time. Recently, several studies (Ríos-Rull and Santaeulalia-Llopis, 2010; Santaeulalia-Llopis, 2011; Choi and Ríos-Rull, 2020) provide evidence that accounting for the observed time-variation in the factor shares profoundly modifies the propagation of technology shocks to aggregate variables. Moreover, Lettau et al. (2019) demonstrate that fluctuations in the factor shares also have the potential to explain the observed risk premia in the stock market.

Our identification strategy follows the procedure outlined by Santaeulalia-Llopis (2011). We specify a trivariate Vector Autoregression (VAR) model with four lags, where the growth rate of the (inverse) relative price of investment to that of consumption goods ($\Delta \log(\mu_t)$), the growth rate of total factor productivity ($\Delta \log(z_t)$) and the linearly detrended (log) labor share of income ($\log(ls_t)$) are the endogenous variables. The choice of detrending the labor share follows Choi and Ríos-Rull (2020), and is intended to deal with the secular decline observed over the last few decades. Specifically, we define the system

$$\boldsymbol{y}_t = \boldsymbol{\alpha} + \sum_{j=1}^{4} \boldsymbol{\Gamma}_j \boldsymbol{y}_{t-j} + \boldsymbol{\epsilon}_t, \tag{1}$$

[11] Further details on the construction of household-level consumption and income series can be found in Appendix B.5. The series are depicted in Figure B.2.

where $\boldsymbol{y}_t = [\Delta \log(\mu_t), \Delta \log(z_t), \log(ls_t)]'$, $\boldsymbol{\alpha}$ is a vector of constant terms, $\boldsymbol{\Gamma}_j$ (with $j = 1, ..., 4$) are the matrices of dynamic coefficients and $\boldsymbol{\epsilon}_t \sim N(0, \Sigma)$ is a vector of normally-distributed innovations with mean zero and variance-covariance matrix Σ.

We estimate the reduced-form system (1) over the 1981Q4-2017Q4 sample.[12] Since the innovations $\boldsymbol{\epsilon}_t$ are contemporaneously correlated, to obtain the orthogonal shocks $\boldsymbol{u}_t$ we exploit the relationship $\boldsymbol{\epsilon}_t = \boldsymbol{H}\boldsymbol{u}_t$, where $\boldsymbol{H}$ is 3×3 matrix that we identify through standard long-run restrictions. The identification strategy we adopt imposes that shocks to the factor share of income do not affect the long-run levels of TFP and the relative price of investment, and are therefore purely redistributive. As for the remaining shocks, we follow Fisher (2006) in assuming that neutral technology shocks do not affect the relative price of investment in the long run.[13] Thus, investment-specific technology shocks are the only ones capable to permanently affect the relative price of investment. The identified factor-share (FS) shocks capture innovations to the factor shares of income that are orthogonal to (neutral and investment-specific) technology shocks. Bergholt et al. (2022) discuss in detail potential explanations for exogenous movements in the factor shares. We interpret a FS shock as an unexpected shift of resources between the factors of production; this is labeled as "distribution shock" in Lansing (2015), or as "factor-share shock" in Ríos-Rull and Santaeulalia-Llopis (2010) and Greenwald et al. (2019).[14] In addition, by isolating unexpected variation in the labor share, our shock implicitly bundles together shifts in both the capital and the profit shares of income (Barkai, 2020).

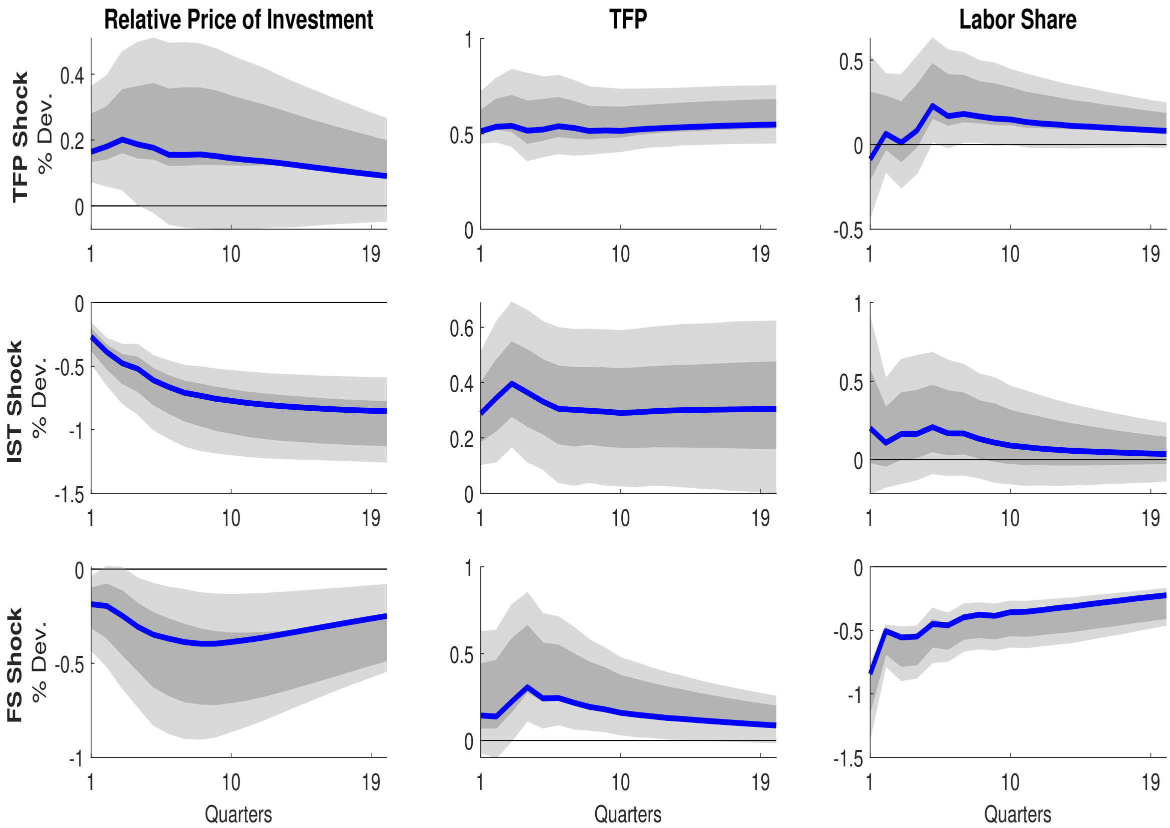
The impulse-response functions are displayed in Figure 1. The identified shocks induce some marked interaction among the three variables we consider. A neutral technology (TFP) shock persistently increases the relative price of investment, while the labor share falls on impact, to then display a temporary increase above the trend (see Ríos-Rull and Santaeulalia-Llopis, 2010; Choi and Ríos-Rull, 2020). An investment-specific (IST) shock is associated with a permanent fall in the relative price of investment and a permanent increase in TFP, while the labor share displays a mild and short-lived expansion. On the other hand, a FS shock is associated with a prolonged decline in the labor share, while contracting the relative price of investment and expanding

[12] This sample is chosen for three main reasons. First, given that the household-level data are available over the 1982Q4-2017Q4 time-window, and we use a VAR(4) model, we need to consider that the first 4 time-series observations will be discarded. Second, Fisher (2006) documents the presence of a structural break in the trend of the relative price of investment in 1982. Finally, the sample is consistent with a large literature focusing on the Great Moderation period (e.g., Stock and Watson, 2002).

[13] The resulting series for the three structural shocks are displayed by Figure C.1 in Appendix C.

[14] We have implicitly assumed that all the three variables in the system are fully explained by the identified shocks. In Section 3.5 we extend the VAR system to capture additional shocks, and highlight how our results remain robust to this variation.

Figure 1: Responses of the relative price of investment, TFP, and income share of labor



Notes: The figure displays the impulse-response functions, estimated from the VAR in equation (1), to the identified neutral technology (TFP, top panel), investment-specific technology (IST, middle panel), and factor-share (FS, bottom panel) shocks over the sample 1982Q4-2017Q4. Light-grey (dark-grey) shaded areas represent the 90% (68%) confidence intervals. The latter are computed using the moving block bootstrap (Bruggemann et al., 2016), with small-sample bias correction (Kilian, 1998).

TFP.

The quantitative relevance of dynamic interdependence among the endogenous variables under scrutiny also emerges from the forecast error variance decompositions (FEVD) reported in Table C.1 in Appendix C. For instance, FS shocks explain a non-negligible fraction of the fluctuations in the relative price of investment—both over short and medium horizons—whereas IST shocks account for almost a quarter of the variance of TFP, even in the long run (i.e., for $h = \infty$). Finally, technology shocks—both neutral and investment-specific—jointly account for a non-negligible share of the FEV of the labor share, at both business-cycle and low frequencies.

3.2 Responses of main macroeconomic aggregates

We estimate the following autoregressive distributed-lag model, so as to retrieve the impact of the identified shocks on selected macroeconomic variables:

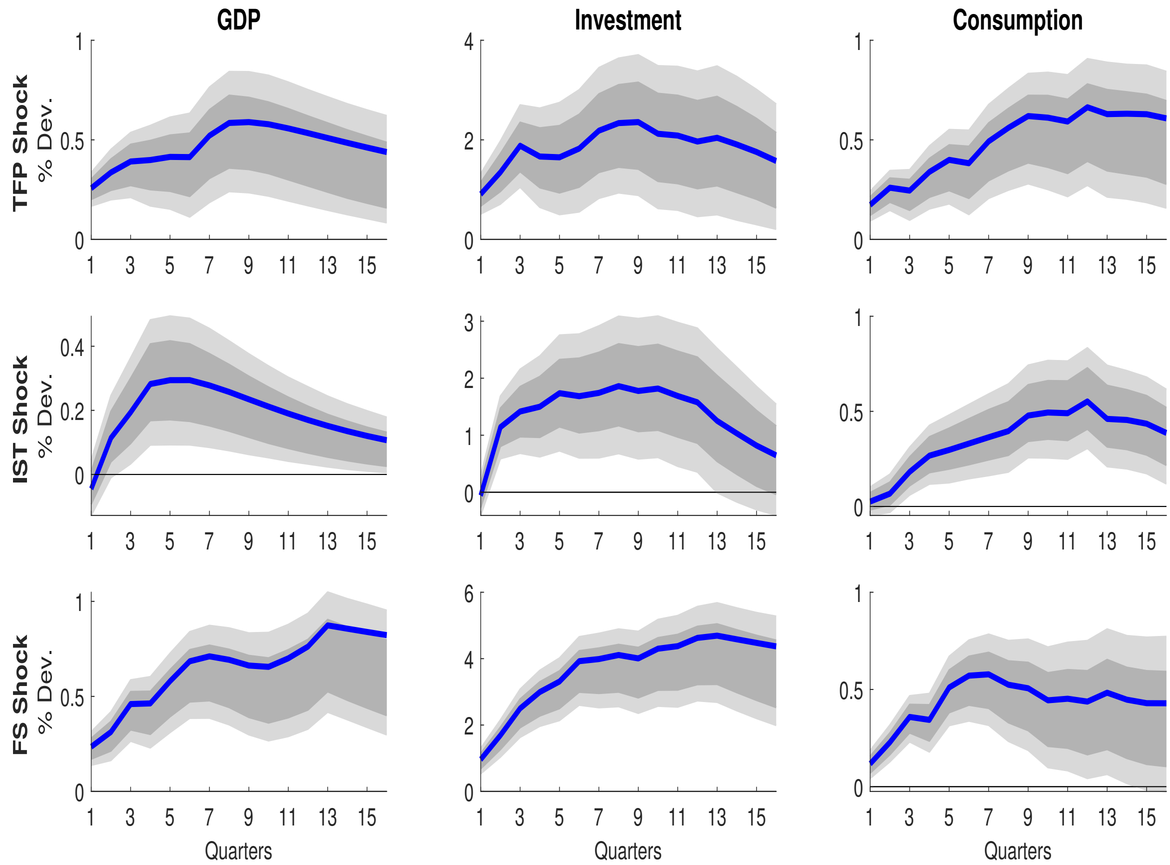
$$x_t = \alpha_0 + \alpha_1 t + \sum_{r=0}^{R} \beta_r u_{j,t-r} + \sum_{p=1}^{P} \delta_p x_{t-p} + e_t, \tag{2}$$

where t denotes the time trend, while x_t denotes the (log) aggregate variable for which we compute the impulse-response function to either of the three shocks, as captured by $u_{j,t}$ where $j \in \{TFP, IST, FS\}$. We control for R lags of the shock and P lags of the endogenous variable, with both R and P being optimally determined by a corrected-Akaike information criterion, for each regression separately. Finally, heteroskedasticity-consistent standard errors are computed using the wild bootstrap methodology of Goncalves and Kilian (2004).

Figure 2 reports the responses of output, investment and consumption. Shocks are set so that a positive TFP shock increases TFP, whereas positive IST and FS shocks decrease the relative price of investment and the labor share of income, respectively (in line with Figure 1). All shocks are associated with strong positive comovement among the three macroeconomic aggregates. A TFP shock generates a simultaneous increase in GDP, consumption, and investment, with the full impact of the shock taking roughly two years to be fully reflected into a persistent increase, in all variables. All of these display a more hump-shaped response following an IST shock, with the impact on output and investment being somewhat transitory. This is consistent with the view that the expansionary effects of an improvement in investment-specific technology unfold through the formation of new capital (in line with Greenwood et al., 1988).

While the business-cycle implications of IST and TFP shocks have been widely studied by both the theoretical and the empirical literature, less is known about the macroeconomic consequences of exogenous shifts in the factor shares of income. The third row of Figure 2 shows that FS shocks are expansionary, being characterized by particularly delayed and protracted responses. This type of shock is also associated with a very large reaction of investment, the peak response being almost twice as large as that induced by a TFP shock. Indeed, a FS shock renders physical capital more productive, thus exerting a sustained expansionary force on output. On the other hand, the response of consumption is more muted, reaching its peak after about 6 quarters, to then steadily decline back to the trend level.

Figure 2: Macroeconomic aggregates

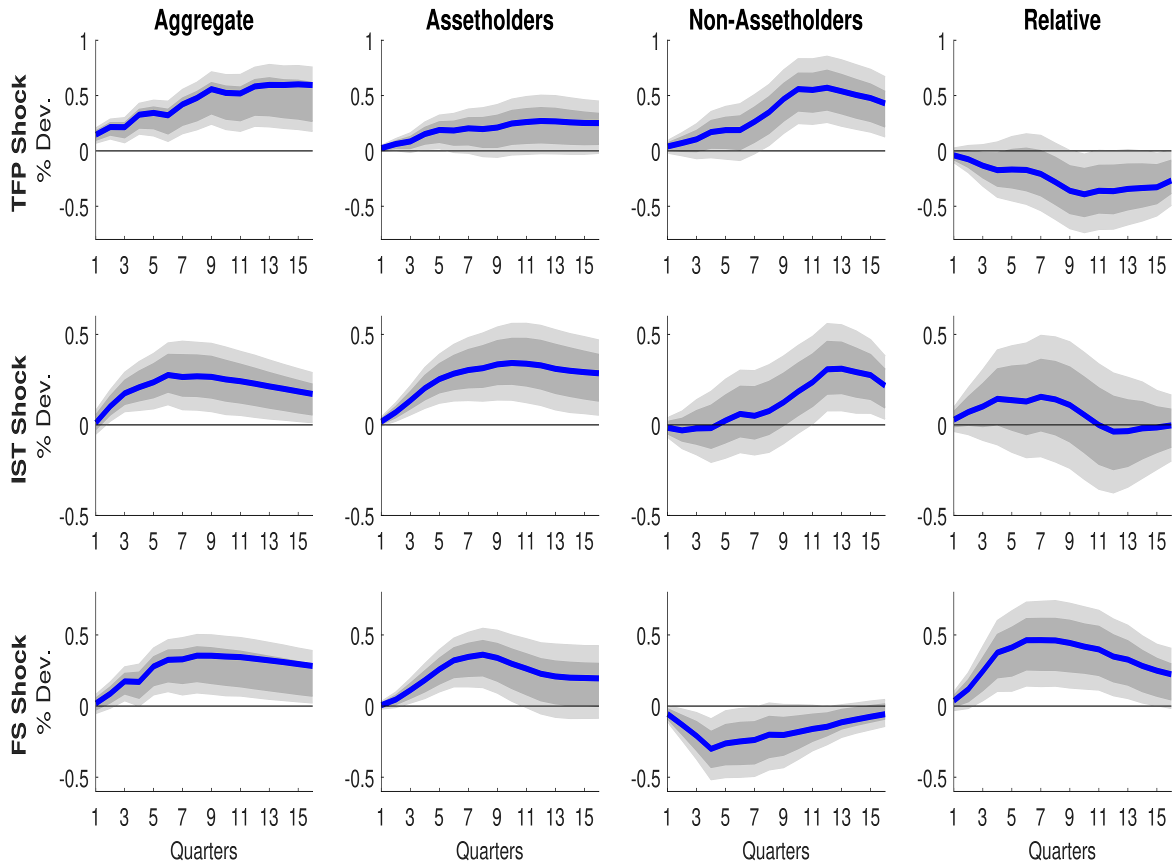


Notes: The figure displays the IRFs of GDP, investment and consumption to the identified neutral technology (TFP, top row), investment-specific technology (IST, middle row), and factor-share (FS, bottom row) shocks, estimated over the sample 1982Q4-2017Q4. Dark and light-grey shaded areas represent the 68% and 90% confidence intervals, respectively.

3.3 Consumption and income responses at the household level

While the responses of main macroeconomic aggregates display strong positive comovement—conditional on each of the three aggregate shocks—we document major differences in the responses of consumption and income of households sorted based on their asset ownership. Figures 3 and 4 report the IRFs of non-durable and services expenditure, as well as of net income, respectively, both of which are obtained by estimating (2) with household-level variables. We focus on the response of: *i*) the economy-wide representative household (first column), *ii*) the representative assetholder (second column), *iii*) the representative non-assetholder (third column), and *iv*) the ratio between the consumption (or income) of the representative assetholder and that of the non-assetholder (fourth column), which is taken as a metric to account for inequality between the consumption (or income) responses of the two representative households.

Figure 3: Non-durables and services expenditure



Notes: The figure displays the IRFs of non-durables and services expenditures for the representative agent (first column), the representative assetholder (second column), the representative non-assetholder (third column), and the ratio between assetholder's and non-assetholder's (fourth column) to the identified neutral technology (TFP, top row), investment-specific technology (IST, middle row), and factor-share (FS, bottom row) shocks, estimated over the sample 1982Q4-2017Q4. Dark and light-grey shaded areas represent the 68% and 90% confidence intervals, respectively.

Looking at Figure 3, we infer that the responses of different macroeconomic aggregates were hiding substantial heterogeneity. Both types of technology shocks induce positive comovement between the consumption of the two representative households. However, facing a TFP shock, non-assetholders' consumption rises relatively more than that of the assetholders, thus implying a contraction in relative consumption. The latter tends to expand, instead, following an expansionary IST shock, although the overall response is not statistically significant: on impact, and for the first few quarters, non-assetholders' consumption response is flat and insignificantly different from zero, whereas assetholders' IRF is strongly significant and positive, displaying a hump-shaped pattern. As for the FS shock, we document a strong contraction in non-assetholders' expenditure, as opposed to the surge displayed by assetholders. Thus, a positive FS shock inevitably widens the gap between the two agents' consumption.[15]

An important consideration should be made, at this stage. Unconditionally, we measure a correlation of 0.15 between the (quarter-on-quarter) growth rate of relative consumption and that of GDP. [16] In this respect, FS shocks have the potential of inducing procyclicality in consumption inequality, while TFP shocks run counter this empirical fact. Therefore, our evidence does not support models where TFP shocks induce procyclical variation in consumption inequality, through a variety of mechanisms (e.g., Danthine and Donaldson, 2002; Guvenen, 2009). This aspect will be at the center stage of our strategy, when disciplining the structural framework in Section 4.

As for household (after-tax) income, the IRF associated with non-assetholders peaks at almost 1%, in response to a positive TFP shock, as compared with the 0.4% estimated for assetholders, which implies a contraction in relative income. Conversely, IST shocks induce a significant and positive response in assetholders' income, while leaving that of non-assetholders almost unaffected, resulting in a relative income response that is insignificantly different from zero. Finally, we document negative comovement between households' respective income, thus implying an expansion in relative income, following a FS shock. In principle, heterogeneity in consumption responses may reflect different propensities to consume out of disposable income—for given and comparable income responses—or heterogeneous responses of income itself. Looking at Figure 4, we see that conditional income dynamics, both at the aggregate and at the household level, are broadly in line with the behavior of the corresponding measures of household expenditure. This is particularly true in connection with technology shocks—i.e., shocks that tend induce permanent-income effects—thus suggesting that, at least in this case, asymmetries in consumption responses are not to be ascribed to heterogeneity in households' marginal propensities to consume.

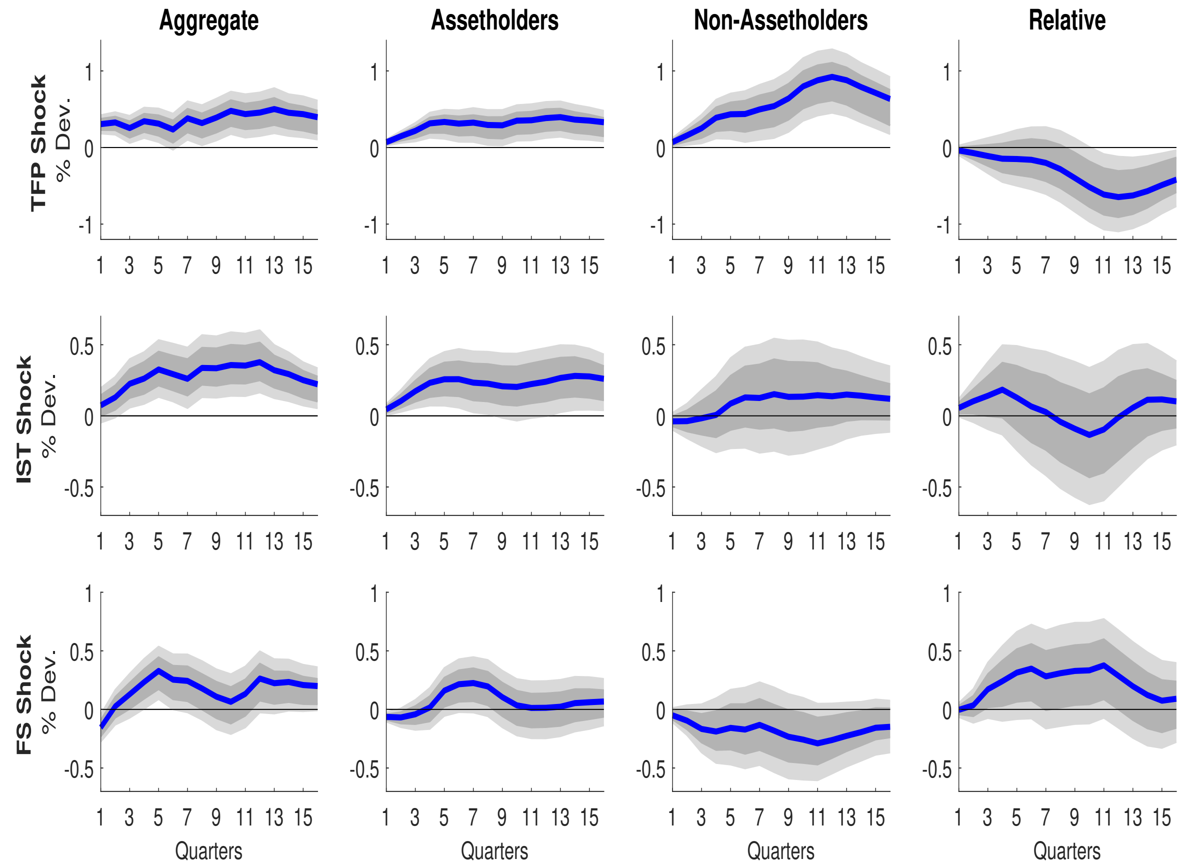
To summarize our findings so far, Table 1 reports the cumulative response of different measures of household-level consumption and income over 16 quarters, following the shock of interest.[17] According to Panel A, following a positive TFP shock non-assetholders increase their spending on non-durables and services, as well as total

[15] While classifying households into assetholders and non-assetholders, we implicitly assume that the transition between groups is reason for no particular concern and that the shocks do not trigger significant endogenous changes in assetholding status. This condition is required to interpret the consumption responses as actual changes in expenditures, rather than as mere compositional effects. Figure D.1 in Appendix D supports this view. Despite the conditional behavior of the share of assetholders being in line with that of relative consumption—as expected on theoretical grounds—little variation emerges, regardless of the specific shock we consider.

[16] This finding echoes the evidence about the cyclical behavior of consumption and income inequality in Lansing (2015).

[17] Figure D.2 shows that the IRFs of total expenditure are close in line with those of non-durables and services.

Figure 4: Net income



Notes: The figure displays the IRFs of net income for the representative agent (first column), the representative assetholder (second column), the representative non-assetholder (third column), and the ratio between assetholder's and non-assetholder's (fourth column) to the identified neutral technology (TFP, top row), investment-specific technology (IST, middle row), and factor-share (FS, bottom row) shocks, estimated over the sample 1982Q4-2017Q4. Dark and light-grey shaded areas represent the 68% and 90% confidence intervals, respectively.

consumption expenditure, by a statistically significant 5.35% and 6.59%, respectively, as compared to the 3.05% and 3.36% increase in assetholders' spending. Thus, consistent with the IRF analysis on relative consumption, TFP shocks exert long-lasting and large effects that favor, in relative terms, non-assetholders. Non-assetholders also denote a more marked rise in net income (8.82%, compared to 4.73% for assetholders).[18] As for the IST shock (Panel B), this triggers a rise in assetholders' total consumption (4.34%), which exceeds, albeit marginally, the overall upward adjustment in their net income (3.42%). At the same time, the cumulative response of non-assetholders' income is statistically indistinguishable from zero, while their consumption rises by 3.84%. While this should imply an expansion of both consumption and inequality,

[18] However, when looking at total consumption for each representative agent, this adjusts relatively less, as compared with their net income, suggesting that part of the increase in the disposable income is actually saved.

Table 1: Cumulative responses

	Non-Durables and Services	Total Consumption	Net Income
	Panel A: TFP Shock		
Assetholders	3.05	3.36	4.73
	[0.99,4.25]	[1.19,4.69]	[2.73,5.72]
Non-Assetholders	5.35	6.59	8.82
	[3.07,6.82]	[3.95,8.42]	[5.28,10.69]
	Panel B: IST Shock		
Assetholders	4.02	4.34	3.42
	[2.55,5.32]	[2.46,5.87]	[1.9,4.75]
Non-Assetholders	2.01	3.84	1.46
	[0.67,3.52]	[2.28,5.76]	[-0.96,3.81]
	Panel C: FS Shock		
Assetholders	3.49	6.2	0.99
	[1.5,4.6]	[3.64,7.75]	[-0.75,2.23]
Non-Assetholders	-2.64	-2.87	-2.85
	[-3.71,-1.29]	[-4.33,-1.23]	[-4.68,-0.3]

Notes: Cumulative responses over 16 quarters to the identified neutral technology (TFP, Panel A), investment-specific technology (IST, Panel B), and factor-share (FS, Panel C) shocks, estimated over the sample 1982Q4-2017Q4. Bootstrapped 68% confidence intervals reported in brackets. The cumulative responses are computed as the present discounted value (given an average annual real interest rate equal to 1%) of the relative change in expenditure or income over the 16 quarters following the shock.

there is substantial overlapping between the household-specific confidence bands associated with different variables, in line with Figures 3 and 4. Finally, following a FS shock (Panel C), the cumulative response of non-assetholders' consumption and income is negative and statistically meaningful.

A word of caution is in order, at this stage. As discussed earlier, a decline (increase) in relative consumption and income indicates a relatively stronger conditional response for non-assetholders (assetholders). Nevertheless, one should consider that these household types feature different average consumption and income levels. In particular, assetholders are richer and consume more than non-assetholders, on average. Therefore, an increase in the relative measure mechanically translates into a stronger adjustment, also in *monetary* terms, for assetholders. However, the same is not necessarily true for non-assetholders, when relative consumption and income decline. Our estimates stress the potential emergence of such a discrepancy in the case of TFP shocks. To check whether this is actually the case, Table D.1 in Appendix D reports the cumulative responses expressed in dollar values (adjusted for the group-specific means). According to this, also the responses of assetholders' net income and consumption to a positive TFP shock are lower than their non-assetholders' counterparts.

3.4 Relative consumption and stock-return predictability

The impulse-response analysis highlights markedly different responses of relative consumption (and income) to different aggregate shocks. This evidence is crucial to understanding how and why household heterogeneity plays a central role in explaining the connection between asset prices and the macroeconomy. Indeed, Guvenen (2009) highlights how reconciling an empirically plausible equity premium with a smooth aggregate consumption process requires mechanisms that induce the volatility of assetholders' consumption be greater than that of aggregate consumption. This aspect is strictly related to the dynamic properties of relative consumption, which we show to vary markedly in response to different shocks. To see this, it is convenient to recall that the definition of aggregate consumption growth implies

$$Var(g_{c^a,t}) = Var(g_{c,t}) + \kappa^2 Var(g_{rc,t}) + 2\kappa Cov(g_{c,t}, g_{rc,t})$$

where g stands for the growth rate of the index variable, while c, c^a, c^{na}, and $rc \equiv c^a/c^{na}$ denote aggregate, assetholders', non-assetholders', and relative consumption, respectively (with κ indicating the share of non-assetholders' consumption in the economy).[19] To obtain a sizable equity premium, one needs $Var(g_{rc,t}) \gg -2Cov(g_{c,t}, g_{rc,t})/\kappa$. As non-assetholders have a smaller share in aggregate consumption (i.e., κ is relatively low), procyclicality of consumption inequality is essential to induce a substantially higher volatility of the growth rate of assetholders' consumption. Quite crucially, we document this to be the case only conditional to FS shocks.

While this argument pertains to the origin of the average equity premium, a different but related question is whether cyclical fluctuations in relative consumption do capture *time-variation* in expected stock returns. The first step in addressing this point consists of considering that assetholders' consumption, once again, can be expressed as a function of aggregate and relative consumption. Thus, we can test whether changes in consumption inequality act as systematic drivers of expected excess stock returns, while controlling for aggregate consumption growth. Specifically, we run the following predictive regressions:

$$r^{ex}_{t,t+h} = \alpha + \boldsymbol{\beta}'\boldsymbol{x}_t + \epsilon_{t+h}, \tag{3}$$

where h denotes the time horizon in quarters, $r^{ex}_{t,t+h}$ denotes annualized excess returns

[19] As anticipated in the introduction, this property simply derives from the definition of aggregate consumption as a weighted average of assetholders' and non-assetholders' consumption, which implies (up to a first-order approximation) $g_{c^a,t} = g_{c,t} + \kappa(g_{c^a,t} - g_{c^{na},t})$.

between time t and $t + h$, and $\boldsymbol{x_t}$ contains time t growth in aggregate consumption and alternative measures of relative consumption. Table 2 reports the estimated coefficients, along with their t-statistics (in parentheses) and p-values (in square brackets), for different forecast horizons over the sample 1982Q4-2017Q4.

Panel A reports the model specification including the growth rates of aggregate consumption and of unconditional relative consumption (i.e, $g_{rc,t}$).[20] Aggregate consumption growth is found to bear predictive power only from 8 quarters onwards, being somewhat complementary to the growth rate of the unconditional relative consumption, which displays predictive power up to 12 quarters. Interestingly, we find that the two slope coefficients have opposite signs. While positive aggregate consumption growth predicts lower future expected returns, as in Atanasov et al. (2020), higher-than-average consumption inequality predicts higher expected excess returns. On one hand, the coefficient on aggregate consumption growth tends to pick up the well-understood countercyclicality of the equity premium (Fama and French, 1989). On the other hand, we need to consider that relative consumption is (unconditionally) procyclical. Therefore, the positive coefficient reflects the transitory nature of the rise in assetholders' relative to non-assetholders' consumption during expansions, thus predicting future periods of lower-than-average growth in relative consumption (due to mean-reversion), in tandem with lower expected excess returns.

The transmission of aggregate shocks discussed in the previous section suggests that procyclical fluctuations in relative consumption should mostly be driven by FS shocks (recall Figure 3). In light of this, we re-estimate the predictive regression by replacing relative consumption growth with its *conditional* counterparts.[21] In line with our conjecture, Panel B shows how only variation in relative consumption associated with FS shocks (i.e., $g^{FS}_{rc,t}$) proves significant, from $h = 4$ to $h = 20$. Moreover, the coefficients attached to $g^{FS}_{rc,t}$ are by far the largest, consistent with the idea that this shock drives the bulk of the procyclicality in relative consumption. Most of the predictive power of relative consumption stems from fluctuations induced by factor-share shocks, rather than by technology shocks (either neutral or investment-specific). Hence, the positive relationship between relative consumption fluctuations and future expected returns is driven by shifts in factors shares that, in times of low expected returns (i.e., during expansions), *temporarily* move resources towards assetholders but are expected to dissipate over time, thus implying future expected negative growth

[20] Following Hamilton (2018), in the baseline analysis we capture business-cycle fluctuations by computing growth rates as 8-quarters log-differences

[21] This is constructed as (the growth rate of) the part of relative consumption that is explained by each shock in isolation. Specifically, for $j \in \{TFP, IST, FS\}$ we reconstruct relative consumption as $\log(rc_t)^j \equiv \sum_{r=0}^{R} \hat{\beta}_r u_{j,t-r} + \sum_{p=1}^{P} \hat{\delta}_p \log(rc_{t-p})$, where $\hat{\beta}_r$ and $\hat{\delta}_r$ are the estimated coefficients from (2).

Table 2: Predictive regressions

	Panel A		Panel B			
	$r^{ex}_{t,t+h} = \alpha + \beta_1 g_{c,t} + \beta_2 g_{rc,t}$		$r^{ex}_{t,t+h} = \alpha + \beta_1 g_{c,t} + \beta_2 g^{TFP}_{rc,t} + \beta_3 g^{IST}_{rc,t} + \beta_4 g^{FS}_{rc,t}$			
h	β_1	β_2	β_1	β_2	β_3	β_4
1	−1.38	**2.07**	−1.52	−0.40	0.45	3.65
	(1.43)	(1.00)	(1.39)	(1.92)	(3.27)	(2.30)
	[0.34]	[0.04]	[0.27]	[0.83]	[0.89]	[0.11]
4	−1.39	**1.45**	−1.72	0.16	0.71	**3.21**
	(1.11)	(0.75)	(1.03)	(1.36)	(2.35)	(1.64)
	[0.22]	[0.06]	[0.10]	[0.91]	[0.76]	[0.05]
8	**-2.21**	**1.21**	**-2.70**	−0.22	1.02	**3.03**
	(0.79)	(0.60)	(0.70)	(0.84)	(1.35)	(0.96)
	[0.01]	[0.05]	[0.00]	[0.79]	[0.45]	[0.00]
12	**-2.45**	**0.96**	**-2.95**	−0.97	0.82	**2.65**
	(0.67)	(0.41)	(0.62)	(0.76)	(1.19)	(0.77)
	[0.00]	[0.02]	[0.00]	[0.20]	[0.49]	[0.00]
16	**-2.25**	0.42	**-2.80**	−0.93	0.56	**2.02**
	(0.51)	(0.42)	(0.52)	(0.69)	(1.07)	(0.73)
	[0.00]	[0.32]	[0.00]	[0.18]	[0.60]	[0.01]
20	**-1.97**	0.09	**-2.44**	−0.64	0.49	**1.22**
	(0.41)	(0.41)	(0.43)	(0.79)	(1.19)	(0.60)
	[0.00]	[0.83]	[0.00]	[0.42]	[0.68]	[0.04]

Notes: The table presents results of predictive regressions of the form $r^{ex}_{t,t+h} = \alpha + \boldsymbol{\beta x_t} + \epsilon_{t+h}$, where h denotes the horizon in quarters and $r^{ex}_{t,t+h}$ denotes annualized excess returns between period t and $t+h$. $\boldsymbol{x_t}$ represents the matrix of (demeaned) predictors, which includes: in Panel A, aggregate and relative consumption growth; in Panel B, aggregate consumption growth and relative consumption growth conditioned on each shock at a time. Growth rates are computed as 8-quarters log-differences. For each regression, Newey-West corrected standard errors (4 lags) appear in parentheses below the coefficient estimate, while p-values are reported in square brackets. Significant coefficients at the ten percent level are highlighted in bold. The sample covers the period 1982Q4-2017Q4.

rates in consumption inequality.

To summarize, this body of evidence supports the view that asset prices are driven not only by fluctuations in aggregate variables, but also by changes in the relative consumption between assetholders and non-assetholders. Furthermore, FS shocks—which generate markedly procyclical consumption inequality—subsume the bulk of the predictive power of relative consumption on future expected excess returns.

3.5 Robustness

This section briefly summarizes some robustness exercises on the empirical evidence reported so far. The motivation and further details on these can be found in Appendices D.4 (regarding the conditional behavior of household consumption and income) and D.5 (regarding the predictive regressions).

We first conduct a number of exercises with the aim of ensuring that what highlighted so far about the behavior of relative consumption and income is robust to different household sorting criteria, features of the raw data being employed, and shock identification. Specifically, we opt for three alternative household-sorting strategies: first, we distinguish between assetholders and non-assetholders that are otherwise identical along different socio-economic characteristics (such as age, gender, education, and housing-tenure status), so as to enhance comparability; thus, we sort households based on stockholdings, rather than on their holdings of liquid assets; finally, we devise a sorting according to which a household is classified as an assetholder either if it fulfills this requirement in the CEX data, or if its SCF-based probability to be an assetholder exceeds a given threshold. All these exercises lead to results that are very similar to the evidence reported in this section. Analogously, the evidence is virtually unchanged when identifying the shocks of interest in VAR settings where we adopt a utilization-adjusted measure of TFP, or where we add log per-capita hours as a fourth variable (so as to account for the potential role of additional shocks).

We also perform further tests on the robustness of relative consumption as a significant predictor of future expected excess returns. Specifically, we verify this to be the case when relying on a sorting based on stockholdings, when controlling for a well-established stock-return predictor—the aggregate consumption-wealth ratio (cay) proposed by Lettau and Ludvigson (2001)—and when computing the growth rates of the tested predictors on a quarter-to-quarter basis.

4 Framing the empirical analysis

To rationalize the main insights from the empirical analysis we consider a production-based asset-pricing framework with limited asset market participation. The model allows us to shed light on the implications of household heterogeneity for the transmission of aggregate shocks and the relationship between asset prices and the macroeconomy.

4.1 Setup

The model features concentrated capital ownership. Non-assetholders, who constitute a fraction γ of the unit-mass population, are assumed to be excluded from the bond and the stock markets, thus behaving in a hand-to-mouth fashion, and consuming labor income in every period. Assetholders, who represent the complementary fraction $1 - \gamma$ of the population, own firms through equity shares, and smooth consumption intertemporally by trading one-period bonds. Both agents are assumed to inelastically supply their entire time-endowment to the firms. The economy's supply side is standard, with firms producing according to a Cobb-Douglas technology and facing capital adjustment costs. Importantly, assetholders feature external habit preferences, which generate countercyclicality in risk-aversion and excess stock returns (Campbell and Cochrane, 1999). Combined with capital adjustment costs, these preferences sensibly improve the asset-pricing performance of models with endogenous production (Jermann, 1998; Chen, 2017). Appendix E contains all the analytical details about the model economy.

The model features the same aggregate shocks considered in the empirical analysis. The dynamics of the three exogenous state variables are governed by the trivariate VAR outlined in Section 3.1.[22] Inspired by Ríos-Rull and Santaeulalia-Llopis (2010), we see the resulting VAR specification as a flexible tool to capture (potentially) endogenous dynamic interactions between TFP, IST, and the labor share.[23] This modeling strategy contrasts with the traditional approach of assuming independent autoregressive processes for the shocks, thus calibrating them to match the dynamics of macroeconomic data. In fact, our approach is in line with Chari et al. (2007) in that TFP, the relative price of investment and the labor share are conceived as "wedges".

[22]Given the permanent nature of IST and TFP shocks, the model exhibits non-stationary dynamics. See Appendix E.2 for additional details.

[23]For instance, Choi and Ríos-Rull (2020) show that a combination of putty-clay technology, time-bias—whereby shocks may affect newer firms in a stronger way than older firms—and competitive wage setting, can rationalize the overshooting property of the labor share, following a TFP shock.

4.2 A key prediction of the model

Before calibrating the model, we find it important to test a key prediction it conveys about the behavior of consumption inequality, providing us with a basis of interpretation of macroeconomic and (especially) asset-pricing outcomes. Recall that, being excluded from financial markets, non-assetholders consume their wage every period, $c_t^{na} = w_t$. On the other hand, assetholders have access to both the bond and the stock market. It can be shown that, in equilibrium, the consumption of the representative assetholder reads as (see Appendix E)

$$c_t^a = w_t + \frac{d_t}{1-\gamma}, \tag{4}$$

meaning that the representative asset-market participant consumes the wage plus the dividends accruing from firm ownership. Notice that dividends are multiplied by the number of stocks held by the assetholder, $q_t^s = \frac{1}{1-\gamma}$, which derives from the stock market equilibrium condition, $(1-\gamma)q_t^s = 1$, where the supply of stocks is normalized to 1. On the other hand, as bonds are in zero net supply and perfect risk-sharing within groups applies, bond-holdings are always equal to 0, in equilibrium.

Given the equilibrium consumption levels for both agents, it is easy to see that relative consumption can be expressed as:

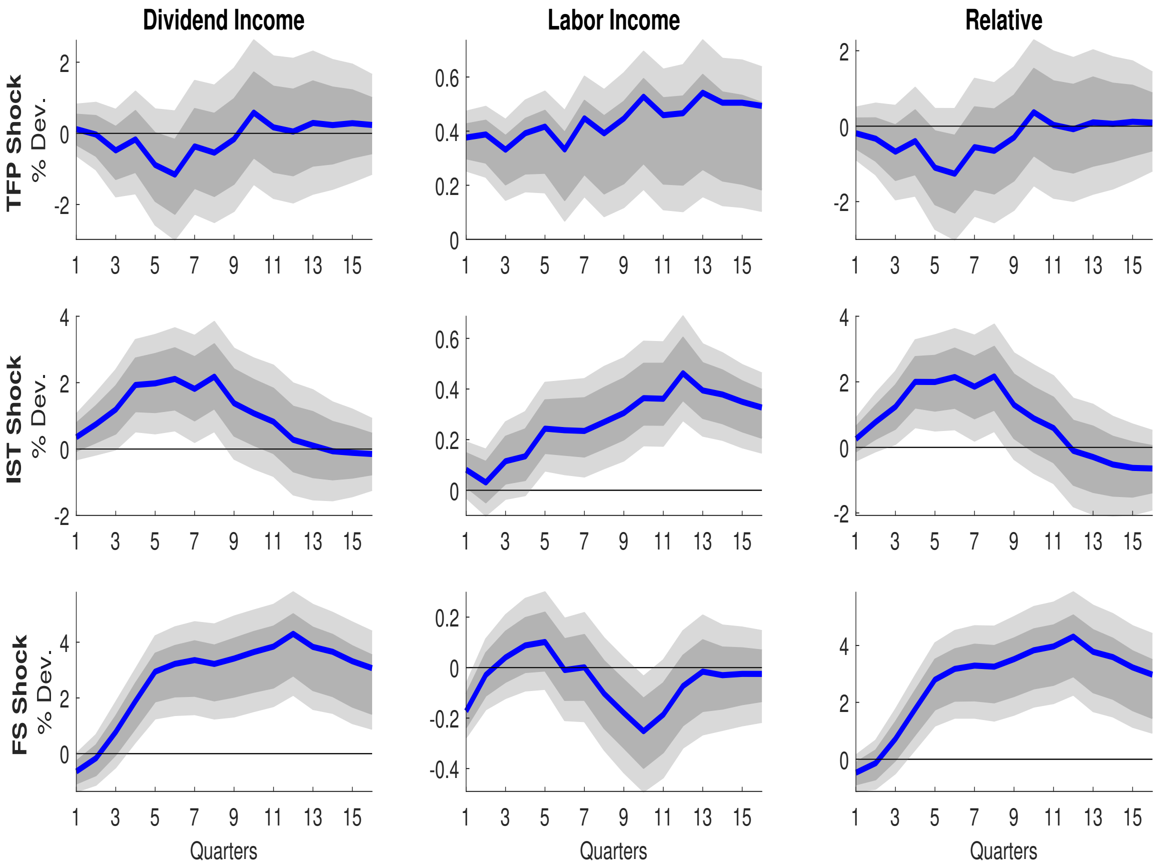
$$rc_t = 1 + \frac{1}{1-\gamma}\frac{d_t}{w_t}, \tag{5}$$

implying that asymmetries in household-specific per-capita consumption depend on the income distribution between labor and capital. In other words, the dynamics of relative (per-capita) consumption in response to different shocks reflect the dynamics of dividend relative to wage income.[24]

We test this prediction in the data. Figure 5 graphs the empirical response of after-tax dividend and labor income, as well as the response of (the log of) the ratio between the two. TFP shocks disproportionately affect labor income, with a peak response of about 0.5% after 13 quarters, while the IRF of dividends is not significant at any horizon. This, in turn, reflects into their ratio declining below trend over the first 9 quarters. By contrast, both IST and FS shocks tend to favor dividend income more than labor income, implying a significant expansion in their ratio. In particular, FS

[24] In this respect, Appendix F provides some analytical insights from a simplified, two-period version of the model, showing how each supply shock affects the sources of income and, thus, relative consumption, in line with the empirics. In particular, while expansionary IST and FS shocks disproportionately benefit the productivity of capital investment and, thus, dividend income, a TFP shock produces a more balanced impact on labor and capital income, with the former displaying higher reactiveness.

Figure 5: Aggregate dividend and labor income



Notes: The figure displays the IRFs of after-tax dividend income, labor income and the ratio between the two to the identified neutral technology (TFP, top row), investment-specific technology (IST, middle row), and factor-share (FS, bottom row) shocks, estimated over the sample 1982Q4-2017Q4. Dark and light-grey shaded areas represent the 68% and 90% confidence intervals, respectively.

shocks produce very sizeable and lingering effects on dividends, which rise by more than 4% after 12 quarters. Conversely, except for a significant drop at impact, the response of labor income is almost muted for the first two years after the shock, to then decline by about 0.2% after 10 quarters.[25]

As shown in Figure 3, conditional movements in the dividend-to-labor income ratio are in line with those of relative consumption. These results also echo the evidence reported in Table 1, where heterogeneity in the response of households' consumption mostly reflects that of their income responses (especially in relation to technology shocks). Assetholders' net income—which comprises a relatively large fraction of financial income—expands less (more), as compared with that of non-assetholders—which predominantly reflects the behavior of labor income—in the face of an expan-

[25]Notably, the responses of the ratio essentially inherit the shape and magnitude of dividend responses, as dividends are much more volatile than labor income, conditional on the shocks under scrutiny.

sionary TFP (IST or FS) shock.

4.3 Setting the model to work

Appendix E details our calibration strategy. The model is solved using second-order perturbation methods. Its parameters are split in two groups. The first group is calibrated to match targeted long-run relationships, while the second group is estimated both via impulse-response matching, as well as by matching a subset of selected unconditional macroeconomic moments. Specifically, the estimated coefficients include the capital adjustment cost parameter, the consumption utility curvature parameter, the parameter capturing the persistence of the habit stock, as well as the parameters of the VAR governing the dynamics of the exogenous process for TFP, the relative price of investment, and the labor share. The estimates are obtained so as to match the responses of TFP, the relative price of investment, and the labor share to the TFP, IST, and FS shocks, as we report in Figure 1. We also target the volatility of (the growth rate of) output, consumption, investment, and dividends, as well as the correlation between the growth rates of dividends and output.

Unconditional moments The framework does a fairly good job at replicating both the targeted unconditional moments, as well as some non-targeted moments, such as the unconditional volatility of relative consumption (0.45 in the model vs. 0.68 in the data), and key asset-pricing moments.[26] Focusing on the latter, we need to recall that restricting access to financial investment to a limited number of households raises the equity premium they demand, through the connection between their consumption growth and financial income, which is intrinsically more volatile. In fact, we can reproduce plausible excess stock returns, both in level and volatility. The average equity premium is 4.59 (vs. 4.39 in the data), while its volatility is 19.94 (vs. 15.67 in the data). The empirical equity premium is estimated following Fama and French (2002), rather than by using average historical excess returns. As argued by the authors, using the latter would largely overestimate the true equity premium, especially when considering the post-WWII period.[27] We are also successful at reproducing a plausible risk-free rate (1.17, vs. 1.07 in the data), though this denotes a certain excess volatility, as compared with the data. As in Jermann (1998) and Lansing (2015), consumption habits and capital adjustment costs, while necessary to generate sufficiently volatile

[26]See Appendix E.1 for further details on the calibration exercise, the resulting estimates, and matched moments. It is worth recalling that, given that the shock structure is imported from our VAR estimates, the calibration entirely relies on the habit parameter and the capital-adjustment cost, to target 5 unconditional macroeconomic moments.

[27]In our sample, the average historical excess return amounts to 8.47%.

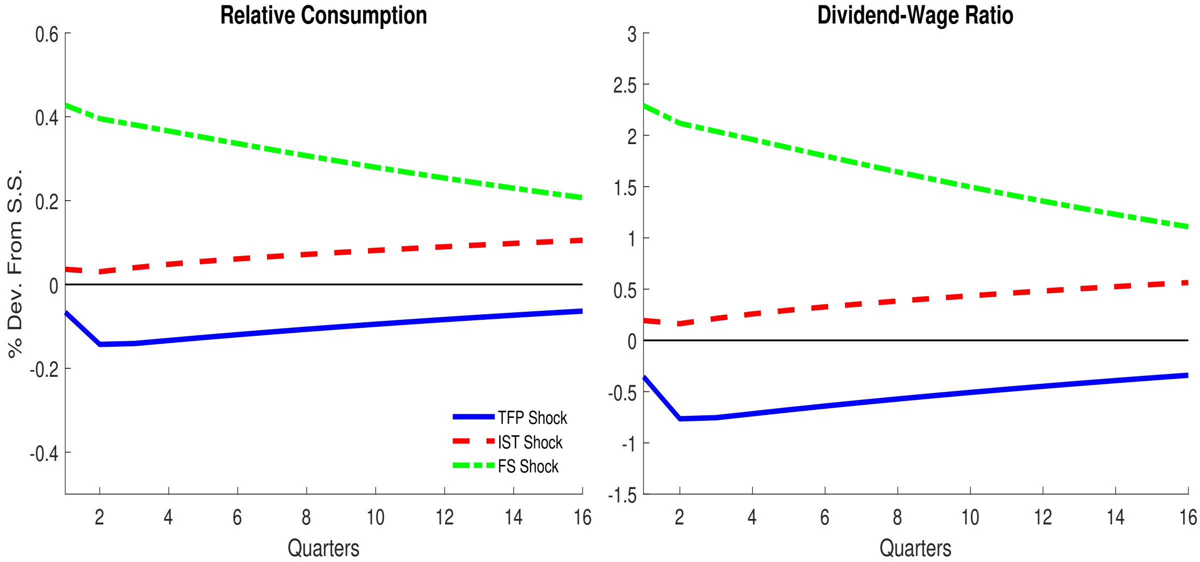
stock returns, concurrently induce strong fluctuations in investors' marginal utility, which inevitably reflects into a rather volatile risk-free rate.

Conditional dynamics All shocks are associated with broad comovement of output, aggregate consumption and investment, in line with the evidence we provide in Figure 2 (compare with Figure E.2 in Appendix E). Through Figure 6, we evaluate the capacity of the model to reproduce the cyclical properties of consumption and income redistribution between the two representative households, conditional on each shock. Expansionary FS shocks are associated with a positive response of relative consumption, as well as with a stronger response of dividends with respect to labor income. This is also the case for IST shocks, albeit to a more limited extent. Conversely, expansionary TFP shocks induce a countercyclical change in relative consumption, which reflects higher sensitivity of labor income with respect to dividend income. Notably, the model-implied IRFs—for both relative consumption and different income sources—are quantitatively consistent with their empirical counterparts, if one abstracts from the absence of a gradual buildup of the responses. For instance, the peak response (which is reached on impact, in the model) of relative consumption to the FS (TFP) shock is about 0.42% (-0.15%), which is comparable with the IRFs displayed in Figure 3. In addition, the responses of both relative consumption and the dividend-to-labor ratio to the IST shock are relatively more muted, in line with the empirical evidence of Section 3. From a quantitative viewpoint, allowing for dynamic interaction among TFP, the relative price of investment, and the labor share turns out to be important to reproduce results in line with the empirical findings. Without such interaction, dividends would otherwise increase persistently after an increase in TFP. By contrast, in Section 3.1 we have documented that an exogenous TFP increase is associated with a rise in the relative price of investment and the labor share of income (after the first period, in this second case), with both effects exerting a negative force on dividends, in line with Figure 5. Coherently, the model produces a relatively muted response of dividends to a TFP shock.

Stock-return dynamics Figure 7 depicts the IRFs of the stochastic discount factor (SDF)—which is a function of assetholders' consumption—divided growth, as well as realized and expected excess stock returns.[28] All shocks are expansionary and raise assetholders' consumption above the habit level, as dividend income increases (although

[28]As it is well-known, a higher-order approximation is required to produce time variation in expected excess returns (see, e.g., Jaccard, 2014). Therefore, the results for Figure 7 and the predictive regressions with simulated data that we report below are obtained through a third-order approximation of the model solution.

Figure 6: Relative consumption and dividend-wage ratio - IRFs

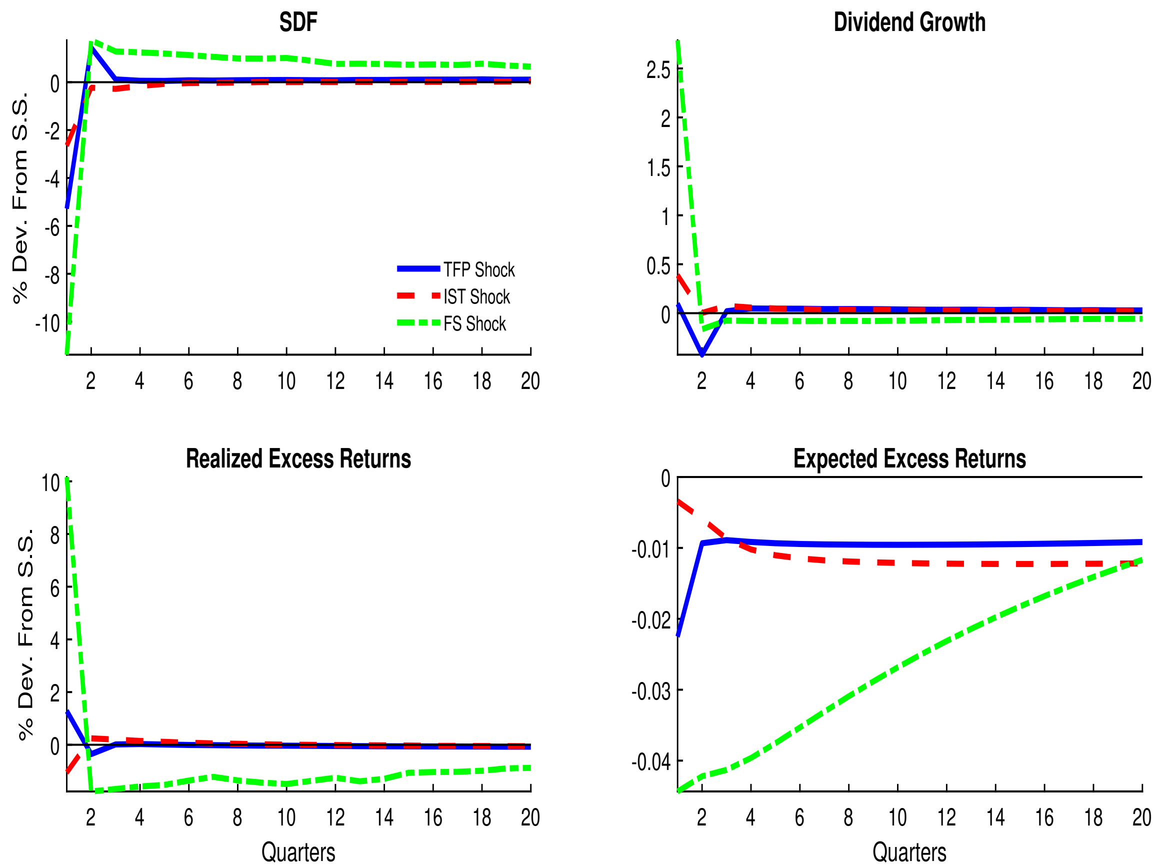


Notes: Responses of relative consumption and dividend-to-wage ratio to TFP, IST, and FS shocks.

very mildly so, in face of TFP shocks). Assetholders' marginal utility of consumption declines, as indicated by the on-impact drop in the SDF. However, especially in face of FS shocks, those gains are expected to revert, as the SDF overshoots from the second period onward. Indeed, FS shocks trigger a redistribution of resources in favor of assetholders that is only temporary in nature (as we show in Figure 6). On the other hand, technology shocks of either type are associated with much smaller fluctuations in investors' marginal utility and dividend growth.

The strong impact of FS shocks on assetholders' consumption and dividends is reflected in the sizeable positive response of realized excess stock returns. In comparison with technology shocks, FS shocks make stocks particularly risky, as they command high payoffs in good times, i.e. when the marginal utility of consumption is relatively low. Clearly, the excess return response largely exceeds that of dividend growth. Through the lens of the Campbell-Shiller decomposition, this property suggests that also stock valuations substantially rise, which is consistent with the countercyclicality of risk-aversion implied by habit preferences. Moreover, FS shocks are associated with falling expected excess returns, thus reproducing the classical "mean reversion"—i.e., realized and expected stock returns are negatively correlated (see, e.g. Campbell and Shiller, 1988; Campbell, 1991). This result also holds, although to a much smaller extent, in the face of TFP shocks. By contrast, IST shocks lead to a modest increase in dividends without exerting a large impact on prices, therefore being associated with (mildly) negative returns, a fall in the price-to-dividend ratio, and protractedly negative expected returns.

Figure 7: Excess stock returns - IRFs



Notes: Stochastic discount factor, dividend growth and realized and expected excess stock returns responses to TFP, IST, and FS shocks.

4.4 Macroeconomic and asset-pricing drivers

Our quantitative setting allows us to ask which shock acts as the main driver of macroeconomic and asset-pricing variables. We approach this question in two ways. First, we quantify the contribution of each shock to generating the moments of interest. Second, we repeat the predictive regression analysis conducted in Section 3.4, using simulated data from the model.

Shock contribution Table 3 reports the relative contribution of each shock to the macroeconomic (top panel) and asset-pricing (bottom panel) moments of interest. As for the volatility of macroeconomic variables, the shock decomposition is performed both over the short run (up to 16 quarters) and over the long run. As for the asset-pricing variables, instead, we decompose only their long-run moments. The short-run variance decomposition is performed as in den Haan (2000), while the shock contribu-

Table 3: Shock contribution

Moment		TFP	IST	FS
Macro aggregates				
$\sigma^2_{\log(\tilde{y})}$	SR	16.5	56.8	26.6
	LR	12.9	57.4	29.7
$\sigma^2_{\log(\tilde{c})}$	SR	25.9	59.1	14.9
	LR	17.7	55.5	26.8
$\sigma^2_{\log(\tilde{inv})}$	SR	6.8	49.5	43.6
	LR	8.5	58.5	33
$\sigma^2_{\log(\tilde{rc})}$	SR	9.4	4.9	85.6
	LR	7.6	53.1	39.3
Financial moments				
$E(r^b)$		16.8	4.1	79.1
$E(r^s - r^b)$		5.6	-2.2	96.6
$\sigma^2_{r^b}$		28.1	5	66.9
$\sigma^2_{(r^s-r^b)}$		1.5	1	97.5

Notes: Each entry indicates the (percentage) contribution of the corresponding shock to a specific macroeconomic or asset-pricing moment. Along each row, the sum of the three shock contributions amounts to 100. For the macroeconomic variables, the decomposition is presented over both the short run (SR) and the long run (LR). For the asset-pricing variables, the decomposition is only presented in terms of long-run moments.

tion to long-run moments is obtained by following Jensen et al. (2018).[29]

Table 3 highlights a clear disconnect between asset-pricing and macroeconomic drivers. Technology shocks (both investment-specific and neutral) are responsible for a large part of business fluctuations, jointly accounting for roughly 70% of the (unconditional) volatility of output, investment and consumption. In fact, both the short-run and the long-run decompositions consistently identify IST shocks as the main drivers of macroeconomic fluctuations, in line with Justiniano and Primiceri (2008). Turning our focus on the equity premium, IST shocks account for a negligible fraction of its volatility, consistent with the fact that such shocks emerge as being rather neutral, in terms of consumption redistribution between the two classes of households, as indicated by the empirical analysis of Section 3. Though to a lesser extent, this is also the case for TFP shocks, which have traditionally been considered a key source of risk in production-based asset-pricing models. Therefore, TFP and IST shocks play a very marginal role when it comes to reproducing fluctuations in asset prices, whereas the

[29]Specifically, for the generic stationary variable x and the corresponding moment $\mathcal{M}(x)$, the relative contribution of shock ξ to the moment of interest is defined as $\mathcal{M}(x)_\xi = \frac{\mathcal{M}(x)-\mathcal{M}(x)_{-\xi}}{\sum_\xi[\mathcal{M}(x)-\mathcal{M}(x)_{-\xi}]}$ for $\xi = u^\mu, u^z, u^\alpha$, where $\mathcal{M}(x)_{-\xi}$ is the unconditional moment of x when shock ξ is turned off.

dynamics of (the first and second moment of) the equity premium and, to a lesser extent, the risk-free rate, can almost entirely be attributed to FS shocks.

Focusing on relative consumption is key for understanding the emergence of a disconnect between the drivers of financial volatility and macroeconomic fundamentals. The long-run volatility of relative consumption is mostly accounted for by FS and IST shocks, with the former explaining almost the entire the short-run volatility. Once again, this sheds light on the close connection between the consumption of assetholders relative to that of non-assetholders, and how assets are priced in this economy. As shown in Figure 7, the average equity premium mainly reflects risk stemming from FS shocks, which redistribute resources in favor of assetholders in a markedly procyclical manner. In light of the mapping from the dividend-to-wage ratio to relative consumption, this fact can be rationalized through the lens of stronger sensitivity of dividends, as compared to labor income, to FS shocks (recall Figures 3 and 6 for the data and the model, respectively). Otherwise, TFP shocks mainly affect labor income, as compared with dividends, thus implying a relatively muted response of assetholders' consumption, and a more compressed equity premium.

Predictive regressions with simulated data Table 3 reports the decomposition of unconditional (long-run) asset-pricing moments. To study the model-implied drivers of short-run fluctuations in asset prices—with a special focus on excess stock returns—we replicate the estimation of regression (3) with simulated data. Panel A of Table 4 reports the estimated coefficients when featuring as regressors the growth in aggregate and relative consumption. Interestingly, the slope coefficients are quite consistent—both in terms of sign and size—with those reported in Table 2, suggesting that the dynamics of relative consumption are significantly connected with fluctuations in future expected returns, even after controlling for aggregate consumption. Again, a natural question is to what extent this result is driven by technology rather than by redistribution shocks. In this respect, Panel B confirms that the model attributes to fluctuations induced by FS shocks the entirety of relative consumption's predictive power on future expected excess returns.

In line with the empirical estimates reported in Table 2, also the simulated predictive regressions indicate that higher aggregate (relative) consumption growth predicts lower (higher) expected returns in the future. The theoretical setup allows us to provide a transparent interpretation of the sign of the coefficients. As commented for Figure 7, the countercyclicality of expected excess returns—as implied by the negative coefficient on aggregate consumption growth—stems from external habit preferences: booms are times of relatively low risk-aversion, as the consumption of in-

Table 4: Predictive regressions - Simulated data

	Panel A		Panel B			
	$r^{ex}_{t,t+h} = \alpha + \beta_1 g_{c,t} + \beta_2 g_{rc,t}$		$r^{ex}_{t,t+h} = \alpha + \beta_1 g_{c,t} + \beta_2 g^{TFP}_{rc,t} + \beta_3 g^{IST}_{rc,t} + \beta_4 g^{FS}_{rc,t}$			
h	β_1	β_2	β_1	β_2	β_3	β_4
1	**-1.12**	**2.91**	**-1.92**	-1.22	6.46	**4.42**
	(0.54)	(1.04)	(0.78)	(3.07)	(4.69)	(1.54)
	[0.04]	[0.01]	[0.01]	[0.69]	[0.17]	[0.00]
4	-0.39	**1.68**	**-1.11**	-2.15	3.54	**3.13**
	(0.42)	(0.82)	(0.60)	(2.44)	(3.87)	(1.18)
	[0.35]	[0.04]	[0.06]	[0.38]	[0.36]	[0.01]
8	-0.49	**1.37**	**-1.17**	-2.33	2.53	**2.79**
	(0.31)	(0.67)	(0.47)	(1.80)	(3.20)	(0.97)
	[0.11]	[0.04]	[0.01]	[0.20]	[0.43]	[0.00]
12	**-0.60**	**1.01**	**-1.24**	-2.42	2.20	**2.34**
	(0.26)	(0.56)	(0.41)	(1.55)	(2.81)	(0.85)
	[0.02]	[0.07]	[0.00]	[0.12]	[0.43]	[0.01]
16	**-0.59**	**0.88**	**-1.10**	-1.99	1.54	**1.95**
	(0.23)	(0.48)	(0.35)	(1.38)	(2.44)	(0.74)
	[0.01]	[0.07]	[0.00]	[0.15]	[0.53]	[0.01]
20	**-0.58**	**1.03**	**-0.93**	-1.12	0.85	**1.80**
	(0.21)	(0.44)	(0.32)	(1.21)	(2.13)	(0.67)
	[0.01]	[0.02]	[0.00]	[0.36]	[0.69]	[0.01]

Notes: The table presents results of predictive regressions, estimated on simulated data, of the form $r^{ex}_{t,t+h} = \alpha + \boldsymbol{\beta x_t} + \epsilon_{t+h}$, where h denotes the horizon in quarters and $r^{ex}_{t,t+h}$ denotes annualized excess returns between period t and $t+h$. $\boldsymbol{x_t}$ represents the matrix of (demeaned) predictors, which includes: in Panel A, aggregate and relative consumption growth; in Panel B, aggregate consumption growth and relative consumption growth conditioned on each shock at a time. Growth rates are computed as 8-quarters log-differences. For each regression, Newey-West corrected standard errors (4 lags) appear in parentheses below the coefficient estimate, while p-values are reported in square brackets. Significant coefficients at the ten percent level are highlighted in bold. The regressions are estimated over the last 2000 observations of a simulated sample of $500,000$ periods. Simulated time-series data are obtained by solving the model by third-order perturbation methods.

vestors rises relative to habit, which in turn commands lower required stock returns in the following periods (Campbell and Cochrane, 1999). On the other hand, booms are also periods that temporarily benefit financial relative to labor income, implying higher consumption inequality. Indeed, relative consumption growth is procyclical also in the model economy (see Table E.2). As documented by Figure 6 and Table 3, such procyclicality mostly comes from the temporary redistributive consequences of FS shocks. Therefore, the positive coefficient on changes in relative consumption is driven by shifts in the rewards of production (towards capital owners) that are however expected to mean-revert, implying that lower expected returns are associated with negative relative consumption growth in the periods following the initial expansion.[30]

4.5 On the role of household heterogeneity

Armed with these insights on the determinants of macroeconomic and asset-pricing moments, we finally examine the role of concentrated capital ownership in driving the aggregate results. To this end, we perform a simple comparative-statics exercise, whereby we compare a RA benchmark to two alternative TA economies that only differ in the degree of access to the asset market participation.

In the existing literature, the calibration of the share of hand-to-mouth households, γ, primarily depends on how this is interpreted. On one hand, most contributions in the macroeconomic literature see γ as capturing the share of the population with limited or no access to financial markets *lato sensu* and, as such, they set it within the $[0.2 - 0.4]$ interval (e.g., Bilbiie, 2008; Debortoli and Galí, 2017; Aguiar et al., 2020). This view is compatible with the share of U.S. households holding very few liquid assets. On the other hand, some look at $1 - \gamma$ as the share of the population with direct exposure to the stock market, hence holding the ultimate ownership of the productive assets in the economy (e.g., Lansing, 2015; Lansing and Markiewicz, 2017). In this case, γ is calibrated to higher values—typically between 0.75 and 0.9—a choice that allows matching the striking heterogeneity between wealthy households and the rest of the population.

In light of this, Table 5 contrasts a set of macroeconomic and asset-pricing moments for two different levels of asset ownership: our baseline value ($\gamma = 0.33$), and a relatively high value of 0.8. In both cases, we report the percentage deviation from the corresponding moment in the RA economy (i.e., at $\gamma = 0$). Moreover, we report both unconditional statistics and their counterparts conditional on each shock at a time. It

[30]This mechanism can alternatively be gauged from Figure 7: following a FS shock, the SDF overshoots after the initial drop, while expected excess returns decline.

Table 5: Effects of household heterogeneity

	Macro aggregates	Baseline $\gamma = 0.33$	High $\gamma = 0.8$		Asset prices	Baseline $\gamma = 0.33$	High $\gamma = 0.8$
$\sigma_{\log(\tilde{y})}$	unc.	0.5	0.4	$E(r^b)$	unc.	-23.5	-85.1
	TFP	1.75	4.72		TFP	0.96	2.8
	IST	-1.45	-6.6		IST	-0.41	-2.4
	FS	4.08	12.4		FS	-17.1	-59.1
$\sigma_{\log(\tilde{c})}$	unc.	0.15	0.96	$E(r^s - r^b)$	unc.	15.9	57.1
	TFP	-0.32	-0.76		TFP	-11.3	-36.6
	IST	-0.15	-0.52		IST	12.1	48.3
	FS	1.1	5.1		FS	17.9	64.1
$\sigma_{\log(\tilde{inv})}$	unc.	0.87	0.95	σ_{r^b}	unc.	10.9	37
	TFP	3.9	10.2		TFP	14.1	46.4
	IST	-3	-13.7		IST	2.9	12
	FS	8.1	25.3		FS	10.3	35.2
				$\sigma_{(r^s - r^b)}$	unc.	8.1	27.6
					TFP	-9	-31.8
					IST	4.6	8.1
					FS	8.5	28.8

Notes: Each entry indicates the percent variation in the macroeconomic or asset-pricing moment obtained in the TA economy relative to the RA economy. Results are shown for both the baseline value of the fraction of non-assetholders ($\gamma = 0.33$) and for $\gamma = 0.80$. Both unconditional (unc.) and conditional percentage variations are reported.

is striking how the degree of asset market participation has an extremely muted impact on the volatility of the three macroeconomic aggregates, regardless of considering conditional vs. unconditional moments. In fact, moving from the RA benchmark to $\gamma = 0.33$ and $\gamma = 0.8$ only implies a sizable increase in the volatility of investment and—to a lesser extent—output, conditional on FS shocks.

On the asset-pricing front, instead, restricting access to financial markets further allows the model to jointly reduce the average risk-free rate and increase the average equity premium, while increasing the volatility of both moments. Quantitatively, the impact of γ becomes pronounced only at relatively high values: at $\gamma = 0.8$, the average equity premium (the risk-free rate) is higher (lower) by 57.1% (85.1%), relative to the RA case, with the conditional analysis showing that such tendency is chiefly driven by IST and FS shocks.

The marked curvature of the dilution factor, $1/(1-\gamma)$, is key to explaining why the equity premium only expands at relatively low levels of participation in asset markets.[31] With this in mind, there are two elements we need to account for when ex-

[31] Relatedly, it is worth stressing that the dilution factor dominates the impact of γ on total dividends,

amining the impact of γ on the conditional dynamics of relative consumption and, thus, the equity premium. First, increasing γ necessarily leads to a more concentrated distribution of stocks: thus, the smaller the share of assetholders, the larger the contribution of the dividend component (relative to the wage component) to their total income. Second, as we suggest in Section 4.4, the equity premium mainly reflects assetholders' exposure to risk associated with changes in consumption and income that are induced by FS shocks. This type of disturbances is particularly effective at generating large procyclical swings in relative consumption and, therefore, the ratio between dividend and labor income. Combining these two facts implies that, as asset ownership progressively becomes less diluted through an increase in γ—so that dividends account for a larger share of assetholders' income—the equity premium factors in higher risk emanating from FS shocks. Conditional on TFP shocks, which typically affect dividends less than labor income, instead, the equity premium drops as the fraction of non-assetholders increases. Once again, the analysis points at FS shocks as being chiefly important to obtaining plausible cyclicality in household inequality and, as a byproduct, sizable equity premia. To this end, household heterogeneity—as captured by concentrated capital ownership—appears as an essential ingredient, while (much) less so in order to capture stylized business-cycle facts.

5 Conclusion

Technology and redistributive shocks induce markedly different responses of the consumption and income of households sorted according to their assetholdings. Only shocks to the income share of production factors generate sizable procyclicality in relative consumption and income, which is closely associated with time variation in expected excess returns. The stylized facts we highlight are useful to distinguish among different theories that are seemingly consistent at the aggregate level, while implying very different properties at the household level. A model with concentrated capital ownership is able to account for our empirical evidence, with the propagation of each type of shock resting on its capacity to stimulate dividend *vis-à-vis* labor income, a prediction that is robustly confirmed by the data. In this setting, household inequality is quantitatively irrelevant to macroeconomic volatility, while being central to the understanding of asset prices.

d_t, thus shaping the unconditional volatility of per-capita dividends, $(1/(1-\gamma))d_t$, which expand as γ increases.

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Appendices

A Data sources

Below is reported the list of sources for the macroeconomic data employed in the empirical analysis. Unless otherwise noted, the data are provided by the Bureau of Economic Analysis (NIPA) and retrieved from the FRED website. Real per-capita measures are obtained by dividing nominal values by the U.S. population and the end-of-the-quarter monthly Consumer Price Index for all items computed by the Bureau of Labor Statistics.

- GDP: Gross Domestic Product, Billions of Dollars, Quarterly, Seasonally Adjusted at Annual Rate. Code: GDP.
- Investment: Gross Private Domestic Investment, Billions of Dollars, Quarterly, Seasonally Adjusted at Annual Rate. Code: GPDI.
- Non-durables: Personal Consumption Expenditures: Nondurable Goods, Billions of Dollars, Quarterly, Seasonally Adjusted at Annual Rate. Code: PCEND.
- Services: Personal Consumption Expenditures: Services, Billions of Dollars, Quarterly, Seasonally Adjusted at Annual Rate. Code: PCES.
- Durables: Personal Consumption Expenditures: Durable Goods, Billions of Dollars, Quarterly, Seasonally Adjusted at Annual Rate. Code: PCEDG.
- Total Consumption: Non-durables + Services + Durables.
- CPI: Consumer Price Index for All Urban Consumers: All Items in U.S. City Average, Index 1982-1984=100, Monthly, Seasonally Adjusted. Code: CPIAUCSL. Aggregated to quarterly frequency by taking the end-of-quarter value. CPI Inflation is computed as the first log-difference in the quarterly series.
- Gross Income: Personal Income, Billions of Dollars, Quarterly, Seasonally Adjusted at Annual Rate. NIPA, Table 2.1, Line 1.
- Net Income: Disposable Personal Income, Billions of Dollars, Quarterly, Seasonally Adjusted at Annual Rate. NIPA, Table 2.1, Line 27.
- Wages: Compensation of Employees, Billions of Dollars, Quarterly, Seasonally Adjusted at Annual Rate. NIPA, Table 2.1, Line 2.

- Financial Income: Personal Income Receipts on Assets, Billions of Dollars, Quarterly, Seasonally Adjusted at Annual Rate. NIPA, Table 2.1, Line 13.
- Interest Income: Personal Interest Income, Billions of Dollars, Quarterly, Seasonally Adjusted at Annual Rate. NIPA, Table 2.1, Line 14.
- Dividend Income: Personal Dividend Income, Billions of Dollars, Quarterly, Seasonally Adjusted at Annual Rate. NIPA, Table 2.1, Line 15.
- Population: Population, Thousands, Quarterly, Not Seasonally Adjusted. Code: B230RC0Q173SBEA.
- Relative Price of Investment: Price of "equipment" relative to price of "consumption", Quarterly, Annualized Growth Rates (400$\times$log-difference), from Fernald (2014).
- TFP: Business Sector (both not utilization-adjusted and utilization-adjusted) Total Factor Productivity, Quarterly, Annualized Growth Rates (400$\times$log-difference), from Fernald (2014).
- Labor Share of Income: Nonfarm Business Sector: Labor Share, Index 2012=100, Quarterly, Seasonally Adjusted. Code: PRS85006173.
- Aggregate Hours: Index/Level and Office of Productivity And Technology and Work Hours: Hours Worked, Nonfarm Business. BLS. Code: PRS85006033. The per-capita measure is obtained by dividing over Population 16+.
- Population 16+: Civilian noninstitutional population, Level (in thousands), 16 years and over. BLS. Code: LNU00000000.
- Quarterly financial data are sourced from Amit Goyal's website (as discussed in Welch and Goyal, 2008), and the equity premium is computed from the average dividend yield and dividend growth following Fama and French (2002).

After-tax dividend and labor income. The definition of after-tax aggregate dividend and labor income series employed for the IRFs in Figure 5 in the main text follows Lettau and Ludvigson (2013) and relies on data from the NIPA, Table 2.1. Specifically, after-tax labor income is defined as compensation of employees (Line 2) + transfer payments (Line 16) − employee contributions for social insurance (Line 25) − taxes. Taxes are defined as [wages and salaries (Line 3) / (wages and salaries + proprietors' income with IVA and Ccadj (Line 9) + rental income (Line 12) + personal income receipt on assets (Line 13))] times personal current taxes (Line 26). After-tax dividend income is defined similarly as personal dividend income (Line 15) – taxes, where the latter are defined as above, but replacing dividend income at the numerator.

B Construction of household-level series from the CEX

In this appendix, we describe the dataset and preliminaries used to construct quarterly time series of consumption and income at the household level over the period 1982-2017 from the U.S. CEX.

B.1 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 US. The survey is divided into two parts: the Interview Survey and the Diary Survey. The analysis developed in this paper focuses on the first one. Data from the 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 3 months, for a maximum of 5 interviews. The first interview is just for practice, and as such is not made publicly available, while financial information is collected only in the last interview.

B.2 Sample choice

Our analysis employs data available for the whole sample (1980Q1-2018Q1). Standard restrictions are applied to the sample. Only households who completed the survey, i.e. for which five interviews are available in the FMLY/FMLI files, are included. Matching households across quarters is not possible around changes in sample design, which happened at the beginning of 1986, 1996, 2005 and 2015.[32] Such changes imply new household ID numbers. Therefore, all the households who did not finish their interviews before their ID changed are dropped.

Households with negative net income or incomplete income responses are excluded from the sample. Regarding the latter restriction, for the period 1980-2013 the variable

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

RESPSTAT is used, which indicates whether the household is a complete or an incomplete income reporter. Since 2014 this variable is no longer available. Hence, we use the variable ERANKH, which measures the weighted cumulative percent expenditure outlay ranking of the household to total population is left blank for incomplete income reporters. 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.[33] Finally, we exclude all households who denote a change in the household head's age between any two consecutive interviews that is different from either 0 or 1.

B.3 Assetholding status definition

The FMLY/FMLI files report household-level financial information on holdings of "stocks, bonds, mutual funds and other such securities" and of liquid accounts such as savings and checking accounts.

For the period 1980-2012, we use the following variables: SECESTX, which reports the amount of the household holdings in the aforementioned asset categories (on the last day of the month preceding the interview); CKBKACTX, which reports the amounts (at the last day of the month preceding the interview) "in checking accounts, brokerage accounts, and other similar accounts"; SAVACCTX, which asks "On the last day of (last month), what was the total amount your CU had in savings accounts in banks, savings and loans, credit unions, and similar accounts?". Since 2013, these three variables were removed from the survey. However, at the same time a new variable STOCKX was added, which asks "As of today, what is the total value of all directly-held stocks, bonds, and mutual funds?". Similarly, the new variable LIQUIDX was introduced, which measures the amounts invested in "checking, savings, money market accounts, and certificates of deposit or CDs".

Given these variables, we define a household as an assetholder if the sum of SECESTX, CKBKACTX and SAVACCTX or STOCKX and LIQUIDX exceeds the threshold of 1000$. To keep comparability with the SCF variables, dollar amounts 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).[34]

Crucially, indirect holdings cannot be retrieved from the CEX, as also noted by Malloy et al. (2009). In fact, the stock-market participation rate that we retrieve from

[33] As noted by Malloy et al. (2009), 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.

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

this survey trends up until the early 2000s, to then stabilize around 10%, which is way below the actual share of US households that are typically classified as stockholders. Moreover, in 2013 the 'financial assets' question was changed to consider only direct holdings. In fact, Lettau et al. (2019) argue that the CEX provides inferior measures for financial holdings, as compared with other surveys such as the SCF, which can potentially explain the lower estimated rates.

B.4 Imputation procedure

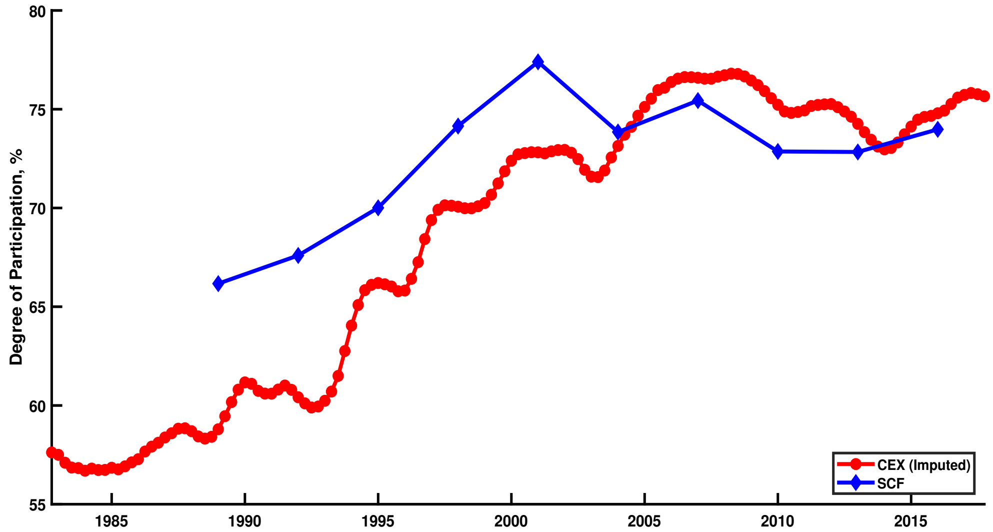
To refine the assetholding status definition to account for indirect holdings, we follow the imputation procedure proposed by Attanasio et al. (2002) and Malloy et al. (2009). Specifically, we perform a probit analysis based on the SCF. This dataset contains wealth information on both direct and indirect stock or assetholdings that can be used to predict the probability that a household holds assets, directly or indirectly, in the CEX. We use the SCF, from 1989 through 2016 (i.e., the last available year). For the asset definition we generate a dummy variable equal to 1 if the sum of (direct and indirect) holdings in equity, bonds, savings accounts, and checking accounts exceeds the threshold of 1000$.

Following Malloy et al. (2009), we then estimate a probit model where the dependent variable is the assetholding dummy and the regressors are the observable characteristics that are also available in the CEX: age, age squared, an indicator for the household head with education of > 12 but < 16 years (highschool), one for education > 16 years (college), an indicator for race not being white/caucasian, year dummies, (log) real total household income before taxes, an indicator for positive interest$+$dividend income, and a constant. We also include interaction terms between age and highschool (agehs) and between age and college (ageco).[35] Here are the estimated coefficients (with t-statistics in parentheses) from the probit regression for assetholdings:

$$\begin{aligned} x'_{SCF}b_{asst} = & \underset{(-56.72)}{-5.07} + \underset{(13.72)}{0.022}age - \underset{(-5.96)}{0.00008}age^2 + \underset{(14.75)}{0.51}\,highschool + \underset{(35.86)}{1.22}\,college \\ & \underset{(-2.92)}{-0.002}agehs \underset{(-13.07)}{-0.008}ageco \underset{(-45.76)}{-0.38}nonwhite + \underset{(1.57)}{0.03}Y_{1992} + \underset{(9.27)}{0.20}Y_{1995} \\ & + \underset{(15.93)}{.35}\,Y_{1998} + \underset{(20.19)}{0.43}\,Y_{2001} + \underset{(14.65)}{0.31}\,Y_{2004} + \underset{(17.50)}{0.37}\,Y_{2007} + \underset{(16.67)}{0.33}\,Y_{2010} + \underset{(16.30)}{0.32}\,Y_{2013} \\ & + \underset{(18.42)}{0.37}\,Y_{2016} + \underset{(44.36)}{0.37}\,log(income) + \underset{(73.13)}{0.95}\,(int + div > 0). \end{aligned}$$

[35]Importantly, SCF weights are employed to map household-level estimates into population estimates.

Figure B.1: Direct and indirect asset-ownership rates

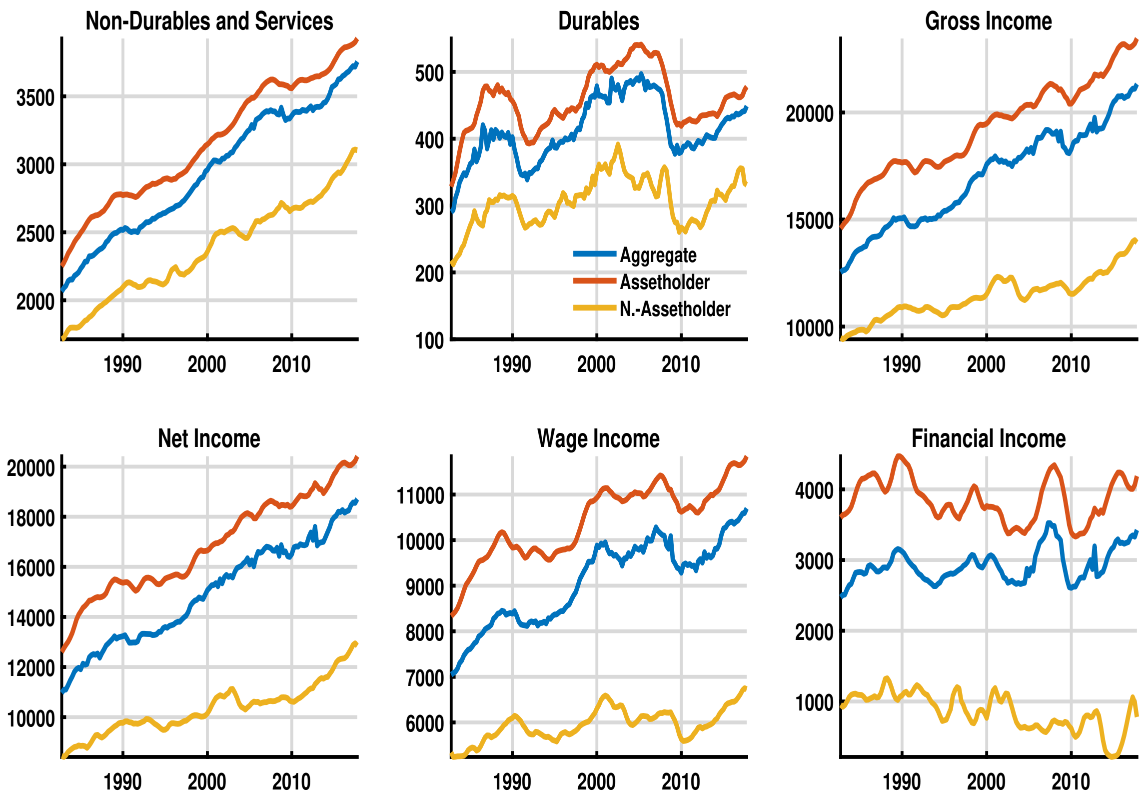


Notes: The figure compares the rates of direct and indirect asset-ownership, as measured from the SCF (blue line) and the CEX (red line).

We then use these coefficients to predict the probability that a household in the CEX holds assets as $\Phi(x'_{CEX} b_{asst})$, 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 assetholding probability for a household in the CEX, we use the dummy *1992* coefficient for the years 1990-1993, the dummy *1995* coefficient for the years 1994-1996, the dummy *1998* coefficient for the years 1997-1999, and so on.

We employ a 'continuous' measure of participation, whereby every household contributes to the population weight, consumption and income of the representative assetholder, according to the predicted probability. Specifically, we use the probability predicted for the last month each household is observed, since financial information is reported only in the last interview. Notice that this imputation procedure is applied only to those households who have non-missing responses to all the questions involved in the imputation procedure. Otherwise, the household receives a probability 0 of being an assetholder. Figure B.1 compares the resulting participation rate compared to the one from the SCF. As for the resulting consumption series, the participation rates in the CEX are smoothed through a backward-looking 4-quarters moving average filter.

Figure B.2: Household-level consumption and income



Notes: Selected consumption and income variables for the representative household (blue line) from the NIPA, together with the representative assetholder (orange line) and the representative non-assetholder (yellow line), as estimated from the CEX, based on the probability-weighted assetholding status imputed from the SCF.

B.5 Household-level consumption and income series

We compute consumption of non-durable goods and services and durable goods aggregated from the disaggregated expenditure categories reported in the monthly expenditure files (MTAB and MTBI files) of the CEX. Non-durables and services consist of food, alcoholic beverages, apparel and services, gasoline and motor oil, household operations, utilities, tobacco, public transportation, fees and admissions, personal care products, reading, other vehicle expenses, and other entertainment supplies, equipment, and services. Durable goods include purchases of vehicles, house furnishings, and tv and audio equipment. Finally, gross and net income are defined as before and after-tax income, respectively, while financial income is computed as the sum of dividend and interest income. Wage income is given by the sum of wages and salaries.

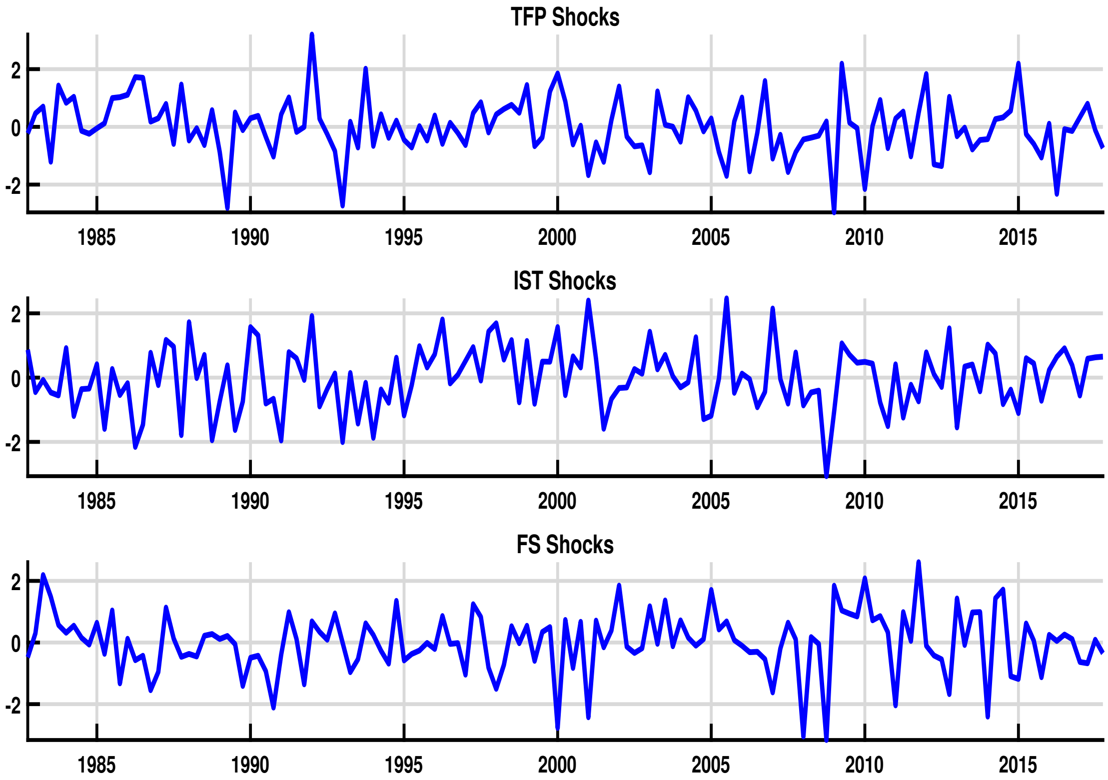
The ultimate aim of the analysis is to obtain time series of both consumption and income—for a representative assetholder and a non-assetholder—by employing the assetholding status definition obtained from the imputation procedure described above. To do so, we compute population-weighted quarterly mean expenditure and income by aggregating from monthly data, and following the formulae provided in the CEX documentation.[36]. Nominal expenditure values are deflated by the end-of-the-quarter CPI for all items, and divided by family size in order to obtain per-capita expenditures.

In line with Cloyne et al. (2019), the group-specific consumption expenditure and income series are adjusted every quarter by the ratio between the corresponding aggregate NIPA series and the estimated CEX aggregate. Finally, to limit some of the noise inherent to survey data and to seasonally adjust the consumption and income series, these are smoothed through a backward-looking moving average encompassing the current and the previous three quarters. Figure B.2 displays the results based on the chosen sorting criterion. Mean estimates are also calculated for the representative household, i.e. over the whole sample and for all households, so as to obtain an aggregate consumption estimate from the CEX. The final quarterly consumption and income series cover the sample 1982Q4-2017Q4.

[36]In particular, we employ the example codes provided at the link: `https://www.bls.gov/cex/pumd-getting-started-guide.htm#section5`. These codes allow one to compute calendar period estimates.

C Identified shocks

Figure C.1: Structurally-identified shocks



Notes: The figure displays the time series of the identified neutral technology (TFP, top panel), investment-specific technology (IST, middle panel), and factor-share (FS, bottom panel) shocks over the sample 1982Q4-2017Q4.

Table C.1: Forecast error variance decomposition

	Relative Price of Investment	TFP	Labor Share
		$h = 4$	
TFP Shock	0.12	0.63	0.01
	[0.00,0.32]	[0.28,0.85]	[0.01,0.43]
IST Shock	0.67	0.28	0.06
	[0.16,0.95]	[0.07,0.56]	[0.00,0.50]
FS Shock	0.21	0.09	0.93
	[0.00,0.71]	[0.00,0.38]	[0.33,0.95]
		$h = 8$	
TFP Shock	0.06	0.65	0.06
	[0.00,0.21]	[0.32,0.88]	[0.03,0.37]
IST Shock	0.73	0.25	0.07
	[0.24,0.98]	[0.06,0.53]	[0.01,0.53
FS Shock	0.21	0.10	0.87
	[0.00,0.65]	[0.00,0.34]	[0.35,0.92]
		$h = 16$	
TFP Shock	0.03	0.68	0.07
	[0.00,0.14]	[0.36,0.91]	[0.03,0.37]
IST Shock	0.79	0.24	0.06
	[0.40,0.99]	[0.05,0.53]	[0.02,0.53]
FS Shock	0.17	0.07	0.87
	[0.00,0.50]	[0.00,0.24]	[0.35,0.91]
		$h = 20$	
TFP Shock	0.03	0.69	0.08
	[0.00,0.12]	[0.37,0.92]	[0.03,0.37]
IST Shock	0.81	0.24	0.06
	[0.49,0.99]	[0.05,0.54]	[0.02,0.53]
FS Shock	0.16	0.07	0.86
	[0.00,0.43]	[0.00,0.20]	[0.35,0.91]
		$h = \infty$	
TFP Shock	0.00	0.77	0.08
	[0.00,0.00]	[0.39,0.98]	[0.02,0.37]
IST Shock	1.00	0.23	0.06
	[0.97,1.00]	[0.01,0.61]	[0.02,0.53]
FS Shock	0.00	0.00	0.86
	[0.00,0.02]	[0.00,0.01]	[0.35,0.91]

Notes: Forecast error variance decomposition at different quarterly horizons (h) estimated from the trivariate VAR in equation (1) over the sample 1982Q4-2017Q4. Bootstrapped 90% confidence intervals reported in brackets.

D Additional results and robustness

In this appendix, we report additional evidence on the response of consumption and income inequality in Section 3.3, as well as on the predictive regressions in Section 3.4, together with all the details—including figures and tables—about the robustness exercises discussed in Section 3.5.

D.1 Compositional change

As discussed in the main text, the interpretation of changes in consumption and income by assetholders and non-assetholders as a causal effect of exogenous shocks requires that the same shocks do not cause a sizeable transition of households from one group to the other. To address this point, Figure D.1 reports the responses of the assetholders' population share to TFP, IST and FS shocks. All the three shocks generate statistically significant responses, with a peak response to TFP (IST and FS) shocks of about -0.4% (0.5%). Nevertheless, we argue that their economic significance is negligible. To see this, recall that assetholders constitute, on average, 67% of the population. Therefore, the IRF to a TFP (IST and FS) shock implies that the assetholding rate decreases (increases) from 67% to about 66.7% (67.3%) at the peak. Clearly, these fluctuations are extremely small, thus allowing us to interpret our estimated household-level consumption and income responses as the causal effect of exogenous shocks.

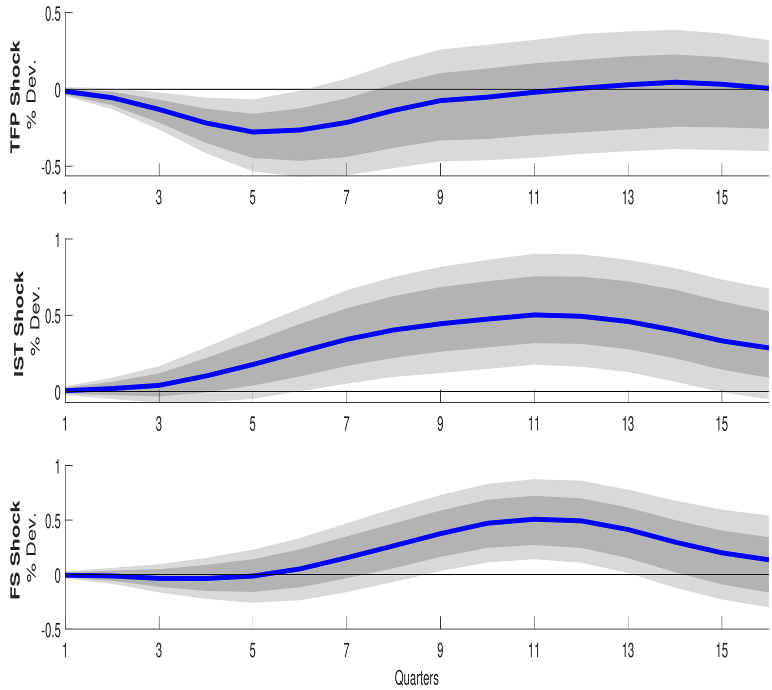
D.2 Total consumption responses

Figure D.2 reports the aggregate and household-level IRFs of total consumption, defined as the sum of non-durables and services and durable goods, for the baseline analysis. The inclusion of durables does not affect the conditional dynamics of relative consumption, which declines (rises) following TFP (FS) shocks and is not significantly affected by IST shocks.

D.3 Cumulative responses: Dollar values

To provide an idea of the magnitudes entailed by the cumulative responses reported in Table 1, we report the corresponding dollar-value responses in Table D.1. According to Panel A, following a positive neutral technology shock non-assetholders increase their spending on non-durables and services, as well as total consumption expenditure, by a statistically significant dollar amount of $787\$$ and $1085\$$, respectively, as compared to the $598\$$ and $755\$$ expenditure increase by the assetholders. Consis-

Figure D.1: Assetholders' population share



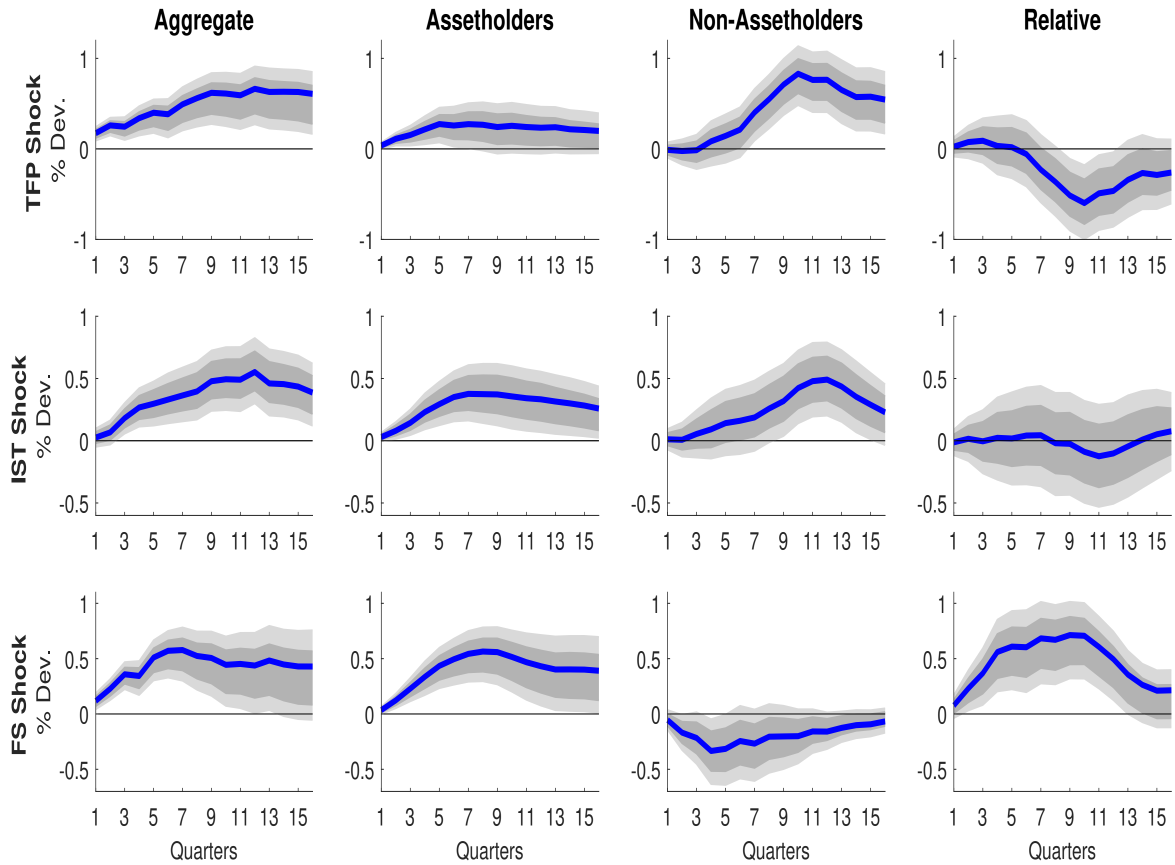
Notes: The figure displays the IRF of the assetholders' population share.

tent with the IRF analysis, the larger consumption adjustment by non-assetholders reflects a more marked rise in net income (1411\$, compared to 1236\$). By contrast, an investment-specific technology shock (Panel B) triggers a remarkable cumulative rise in assetholders' total consumption (977\$), which is in the ballpark of the dollar-amount upward adjustment in net income (892\$). At the same time, the cumulative response of non-assetholders' consumption and income are relatively smaller. Finally, similar conclusions apply for the factor-share shock (Panel C), although the cumulative responses of the hand-to-mouth consumers' consumption and income are now significantly negative.

D.4 Robustness: consumption and income inequality

Controlling for observable heterogeneity Our first robustness exercise aims at controlling for households' observable heterogeneity. Most heterogeneous agent models assume that households are ex-ante identical, and therefore do not differ by dimensions other than their income history, or the ability to access financial markets. Nevertheless, it is well known that the composition of households' portfolios is strongly correlated with demographic characteristics such as age, education, and gender (Guiso

Figure D.2: Total consumption



Notes: The figure displays the IRFs of total expenditure.

and Sodini, 2013). Moreover, recent work has shown that housing tenure is a key determinant of the responsiveness of households' consumption and income to demand shocks (see Cloyne and Surico, 2017; Cloyne et al., 2019, among the others). To control for such potentially relevant dimensions of heterogeneity, we follow Kehoe et al. (2020). Based on CEX data, we partition the population into twenty-four groups for all possible combinations of the following classifications: gender (male and female), age (young-up to 40 years, and old-above 40 years), education (college and no college) and housing tenure status (renter, mortgagor and outright owner). We then compute the average consumption and income series for assetholders and non-assetholders (based on the baseline sorting criterion) within each group. We then reweigh each group by the respective population share, and compute the consumption and income series for the representative assetholder or non-assetholder. As a consequence, after the reweighting the two groups are equally balanced in terms of age, gender, education or housing tenure status. More specifically, for the variable x (e.g., consumption) we

Table D.1: Cumulative responses - Dollar values

	Non-Durables and Services	Total Consumption	Net Income
	Panel A: TFP Shock		
Assetholders	597.85	755.4	1236.39
	[195.38,815.98]	[268.17,1075.58]	[724.76,1493.01]
Non-Assetholders	787.17	1085.9	1410.72
	[448.73,1005.65]	[657.95,1400.41]	[842.27,1719.76]
	Panel B: IST Shock		
Assetholders	787.31	977.64	892.22
	[518.8,1054.1]	[561.48,1308.57]	[495.17,1251.88]
Non-Assetholders	296.49	632.46	233.68
	[101.61,526.37]	[359.87,932.27]	[-170.24,625.87]
	Panel C: FS Shock		
Assetholders	683.32	1394.85	258.4
	[303.32,918.93]	[808.03,1726.97]	[-188.78,591.49]
Non-Assetholders	-388.18	-472.89	-456.09
	[-560.03,-206.21]	[-722.14,-204.71]	[-779.82,-68.57]

Notes: Cumulative responses over 16 quarters expressed in dollar values. To obtain a total expenditure/income effect at the household level in 2017 dollars, the magnitude is multiplied by the group-specific average expenditure over the sample, an average household size of 2.5, and by a price-adjustment factor equal to 2.48 (recall that the CPI for all items is expressed in 1982-1984 basis.)

compute:

$$\bar{x}_t^a = \sum_k x_{k,t}^a \times \omega_{k,t} \quad \text{and} \quad \bar{x}_t^{na} = \sum_k x_{k,t}^{na} \times \omega_{k,t},$$

for the representative assetholder and non-assetholder, respectively, where k indicates the group (for example, male-no college-renter or female-college-outright owner), $x_{k,t}^{a,na}$ denotes the within-group k average assetholder or non-assetholder variable, and $\omega_{k,t}$ represents the population share of group k at time t.

Figures D.3 and D.4 show that the relative responses of household-level consumption and income are essentially invariant, with respect to the original specification. Table D.2 also reassures us of the size and significance of the cumulative responses remaining essentially unchanged.

Sorting based on stockholdings Our analysis has focused on a assetholders vs. non-assetholders dichotomy. However, not only the distinction between stockholders and non-stockholders has traditionally received wide consideration in the asset-pricing literature (Malloy et al., 2009). Thus, it seems appropriate to verify that the conditional cyclical properties of relative consumption and income also apply to this type of household groups. The sorting procedure is exactly symmetric to the baseline presented in the main text. The only difference lies in the types of assets we consider. In this case, we sort households only based on their (direct or indirect) holdings of stocks. Specifically, we re-estimate a probit regression where the dependent variable is a dummy taking value one if the variable EQUITY in the SCF is positive. The variable equity summarizes the value of stocks held directly, in mutual funds or pension schemes, by the household. Therefore, this sorting criterion is much more in line with most of the asset-pricing literature. Consistently, we estimate that only about 20% of the households participated in the stock market at the beginning of the sample. At the end of the sample, instead, the participation rate is estimated around 50%.

As displayed by Figures D.5 and D.6, the responses remain in line with the baseline analysis. In particular, over a 16-quarters horizon the cumulative response of both agents' non-durables and services consumption is very similar, in the face of both TFP and IST shocks (see Table D.3). Furthermore, non-stockholders' cumulative consumption response to a positive FS shock is still negative, yet not statistically indistinguishable from zero.

Different sorting method The representative household-specific series are constructed using a 'continuous' measure of participation to the asset market. While we deem this method appropriate to deal with the uncertainty entailed by the imputation procedure, it involves two unappealing features. First, it ignores the information on assetholdings provided in the CEX (as the probability of being an assetholder is computed based on SCF data). Second, it implies that the same household's consumption (income) simultaneously contributes to the representative assetholder's and non-assetholder's consumption (income), according to the imputed probability. Therefore, as a robustness check we employ a method whereby: *i*) the imputation based on the SCF is applied only to those households who cannot be defined as assetholders, according to the financial information in the CEX; *ii*) a household is univocally classified as an assetholder or a non-assetholder. According to this, we predict the probability of a household being an assetholder only for those households who are not defined as such based on the CEX variables, using the same probit coefficients as for the baseline analysis. Next, to uniquely partition households between the two groups, we apply a threshold method. In particular, households are classified as assetholders for sure (hence, with probability 1) if the predicted probability exceeds 70%. By contrast, households are defined as non-assetholders for sure (thus receiving a probability 0 of being assetholders) if the predicted probability is below 70%. In other words, according to this method a household is defined as an assetholder either if it fulfills the requirement in the CEX data, or if the imputed probability exceeds 70%. The fraction of hand-to-mouth households estimated according to this sorting criterion is essentially unchanged, compared to the baseline case.

Figures D.7 and D.8 and Table D.4 show that, based on this sorting procedure, the results are even more clearcut, compared to the baseline. For example, the IRFs of relative consumption and income to the IST shock are now statistically significant, and the negative comovement between the two agents' consumption responses is further exacerbated, conditional on a FS shock.

Utilization-adjusted TFP We also check the robustness of our results to changing the series for total factor productivity. Specifically, we employ a measure of utilization-adjusted TFP (Fernald, 2014) in the VAR system (1), rather than a non-utilization-adjusted measure. Figures D.9 and D.10, and Table D.5, show that this departure from the baseline analysis is essentially inconsequential for the household-level responses we report.

Extended VAR Finally, we repeat the empirical analysis by extending the VAR system in equation (1) to include (log) per-capita hours as a fourth variable. This allows us to control for the potential impact of additional shocks on TFP, the relative price of investment, and the labor share. The identification assumptions on the purely redistributive effects of FS shocks remain intact also in this quadrivariate version of the VAR. We then use the structurally identified IST, TFP and FS shocks to compute household-level consumption and income responses. Figures D.11 and D.12 show that the responses of household-level consumption and income—as well as those of their relative measures—maintain the same dynamic properties as in the baseline analysis. Also, Table D.6 reports cumulative responses that are very close to the baseline estimates.

Table D.2: Cumulative responses - Observable heterogeneity

	Non-Durables and Services	Total Consumption	Net Income
	Panel A: TFP Shock		
Assetholders	2.86	3.6	3.75
	[1.01,3.97]	[1.3,5.32]	[2.02,4.8]
Non-Assetholders	5.7	6.89	5.86
	[3.17,6.83]	[3.77,8.21]	[2.89,7.61]
	Panel B: IST Shock		
Assetholders	3.88	4.43	3.39
	[2.64,5.11]	[2.87,6.1]	[1.79,4.74]
Non-Assetholders	2.91	3.79	2.64
	[1.58,4.39]	[2.34,5.47]	[0.71,4.31]
	Panel C: FS Shock		
Assetholders	4.66	6.73	1.7
	[2.83,5.65]	[4.29,8.05]	[-0.05,2.87]
Non-Assetholders	-0.26	1.88	-2.16
	[-2.19,1.25]	[-0.68,3.81]	[-4.15,-0.23]

Notes: Cumulative responses over 16 quarters, controlling for observable heterogeneity.

Table D.3: Cumulative responses - Sorting based on stockholdings

	Non-Durables and Services	Total Consumption	Net Income
	Panel A: TFP Shock		
Stockholders	5.99	4.15	5.63
	[3.03,6.71]	[1.77,5.86]	[3.75,6.99]]
Non-Stockholders	5.88	7.04	8.77
	[4.14,6.83]	[5.45,8.16]	[6.24,10.02]
	Panel B: IST Shock		
Stockholders	2.48	4.14	2.03
	[1,3.7]	[2.17,5.89]	[0.18,3.75]
Non-Stockholders	1.26	2.64	1.27
	[0.21,2.53]	[1.52,3.97]	[-1.01,3.29]
	Panel C: FS Shock		
Stockholders	2.52	4.76	-0.18
	[0.65,3.67]	[1.99,6.15]	[-2.04,1.57]
Non-Stockholders	-0.99	-0.46	-2.18
	[-2.29,0.26]	[-2.42,1.12]	[-3.75,-0.16]

Notes: Cumulative responses over 16 quarters for households sorted based on stockholdings.

Table D.4: Cumulative responses - Different sorting method

	Non-Durables and Services	Total Consumption	Net Income
	Panel A: TFP Shock		
Assetholders	2.83	4.47	3
	[0.69,3.72]	[1.82,5.81]	[1.33,3.83]
Non-Assetholders	4.53	4.62	9.27
	[2.02,6.54]	[1.22,7.05]	[4.15,11.94]
	Panel B: IST Shock		
Assetholders	3.9	3.64	2.37
	[2.54,5]	[2.04,5.11]	[1.41,3.41]
Non-Assetholders	-0.57	3.6	-0.82
	[-2.27,1.45]	[1.13,6.48]	[-3.76,2.48]
	Panel C: FS Shock		
Assetholders	2.62	4.55	-0.14
	[0.76,3.67]	[1.95,5.28]	[-1.38,0.95]
Non-Assetholders	-4.36	-6.94	-3.66
	[-6.09,-2.88]	[-9.03,-4.54]	[-6.78,-0.47]

Notes: Cumulative responses over 16 quarters for households sorted according to the probability-threshold method.

Table D.5: Cumulative responses - Utilization-adjusted TFP

	Non-Durables and Services	Total Consumption	Net Income
	Panel A: TFP Shock		
Assetholders	0.55	0.54	3.63
	[-1.19,2.2]	[-1.21,2.47]	[1.85,4.94]
Non-Assetholders	4.57	5.45	8.96
	[2.67,6.09]	[3.13,7.61]	[5.95,10.96]
	Panel B: IST Shock		
Assetholders	3.54	3.7	2.85
	[1.85,5.16]	[1.61,5.19]	[1.01,4.56]
Non-Assetholders	1.28	3.38	-0.56
	[-0.43,3.08]	[1.27,5.71]	[-3.24,2.16]
	Panel C: FS Shock		
Assetholders	1.94	3.5	1.32
	[0.09,3.2]	[1.12,4.66]	[-0.4,2.56]
Non-Assetholders	-2.65	-3.81	-1.65
	[-4.24,-1.14]	[-5.91,-1.45]	[-3.64,0.72]

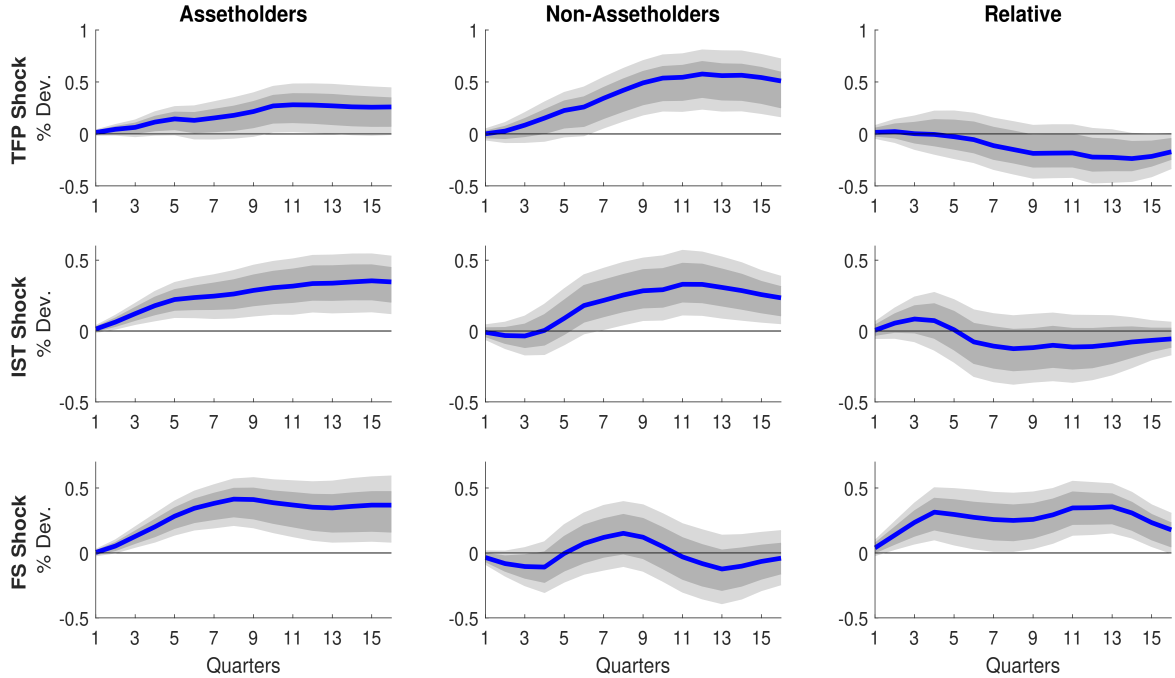
Notes: Cumulative responses over 16 quarters to the shocks identified in the VAR with utilization-adjusted TFP.

Table D.6: Cumulative responses - Extended VAR

	Non-Durables and Services	Total Consumption	Net Income
	Panel A: TFP Shock		
Assetholders	3.16	3.27	4.32
	[1.24,4.25]	[1.26,4.67]	[2.5,5.28]
Non-Assetholders	4.66	5.81	8.02
	[2.48,6.2]	[3.26,7.73]	[4.45,9.96]
	Panel B: IST Shock		
Assetholders	4.37	4.88	2.95
	[2.8,5.59]	[2.93,6.27]	[1.24,4.24]
Non-Assetholders	2.19	2.38	1.31
	[0.69,3.59]	[0.28,3.86]	[-1.08,3.43]
	Panel C: FS Shock		
Assetholders	2.31	3.65	2.19
	[0.11,3.74]	[1.15,4.93]	[0.11,3.63]
Non-Assetholders	-1.81	-2.86	-1.96
	[-3.67,-0.21]	[-5.4,-0.34]	[-4.36,0.49]

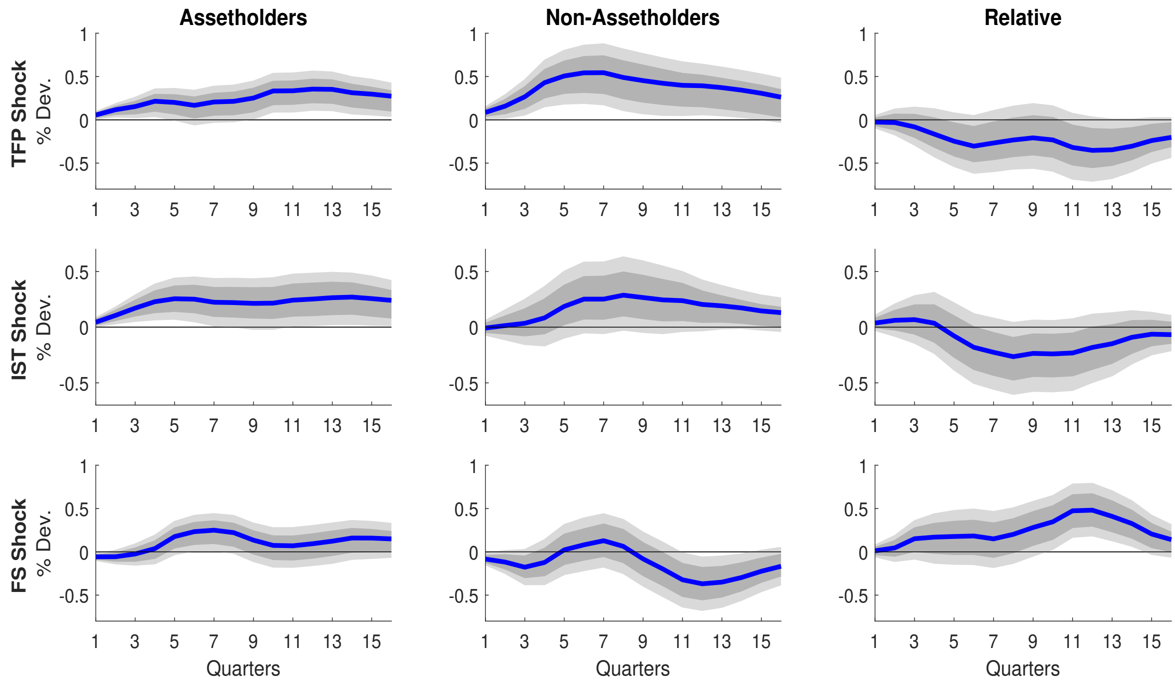
Notes: Cumulative responses over 16 quarters to the shocks identified in the extended VAR.

Figure D.3: Non-durables and services expenditure - Observable heterogeneity



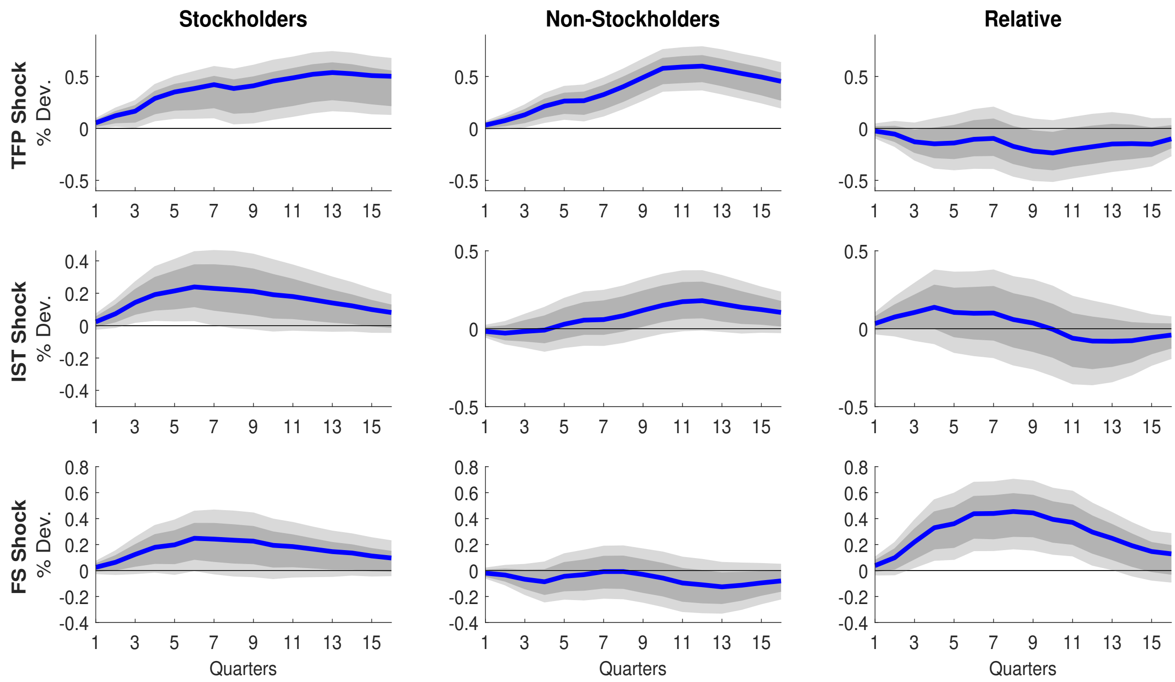
Notes: The figure displays the IRFs of non-durables and services expenditures, controlling for observable heterogeneity.

Figure D.4: Net Income - Observable heterogeneity



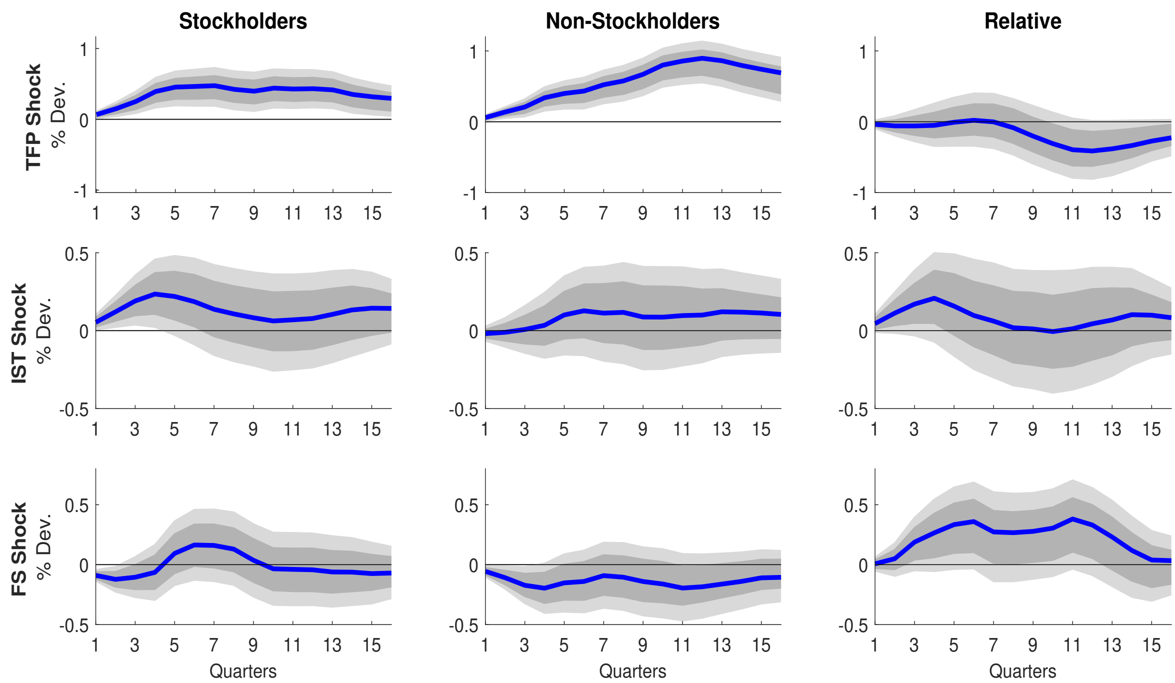
Notes: The figure displays the IRFs of net income, controlling for observable heterogeneity.

Figure D.5: Non-durables and services expenditure - Sorting based on stockholdings



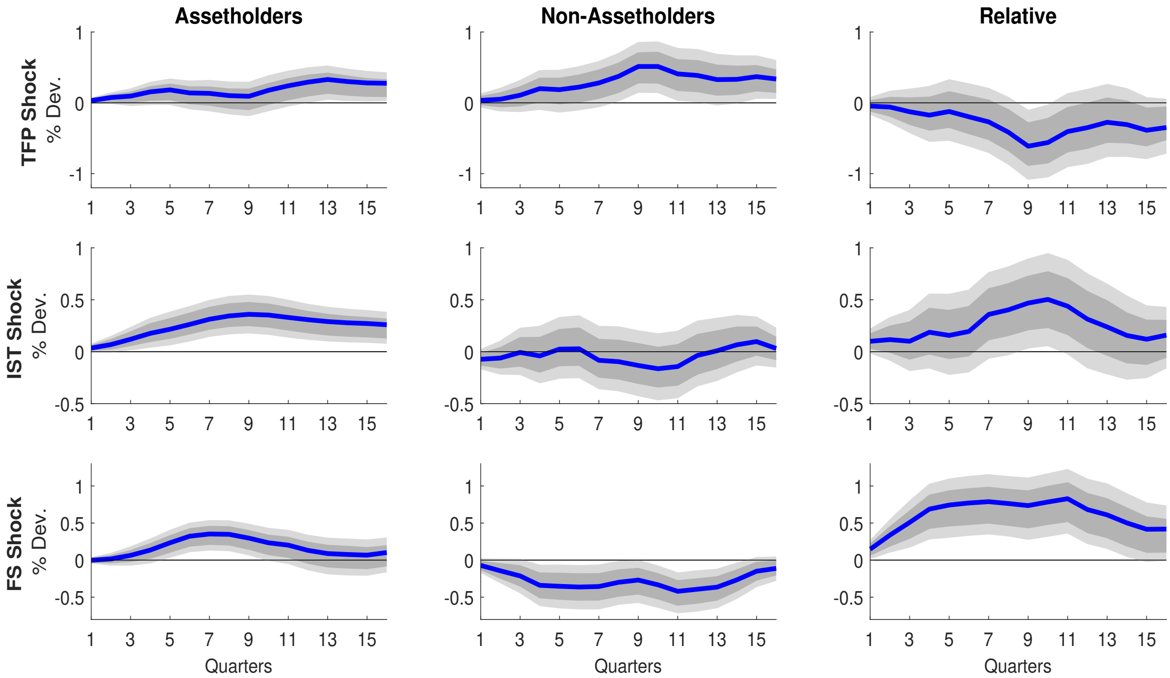
Notes: The figure displays the IRFs of non-durables and services expenditures for households sorted based on stockholdings.

Figure D.6: Net Income - Sorting based on stockholdings



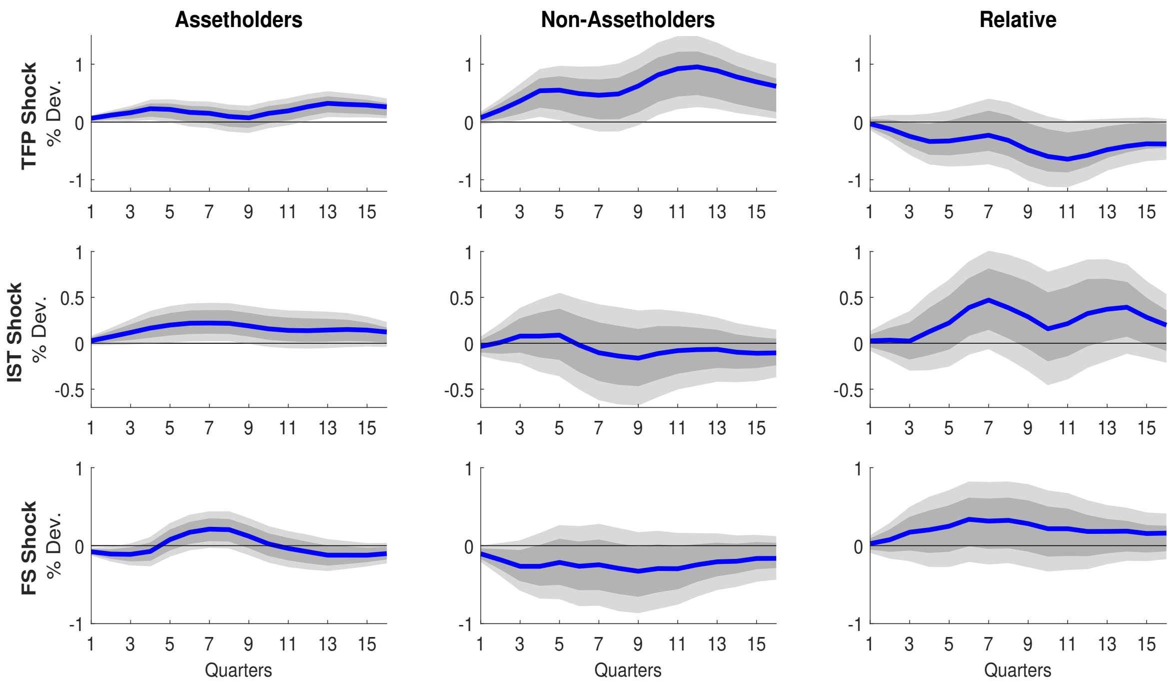
Notes: The figure displays the IRFs of net income for households sorted based on stockholdings.

Figure D.7: Non-durables and services expenditure - Different sorting method



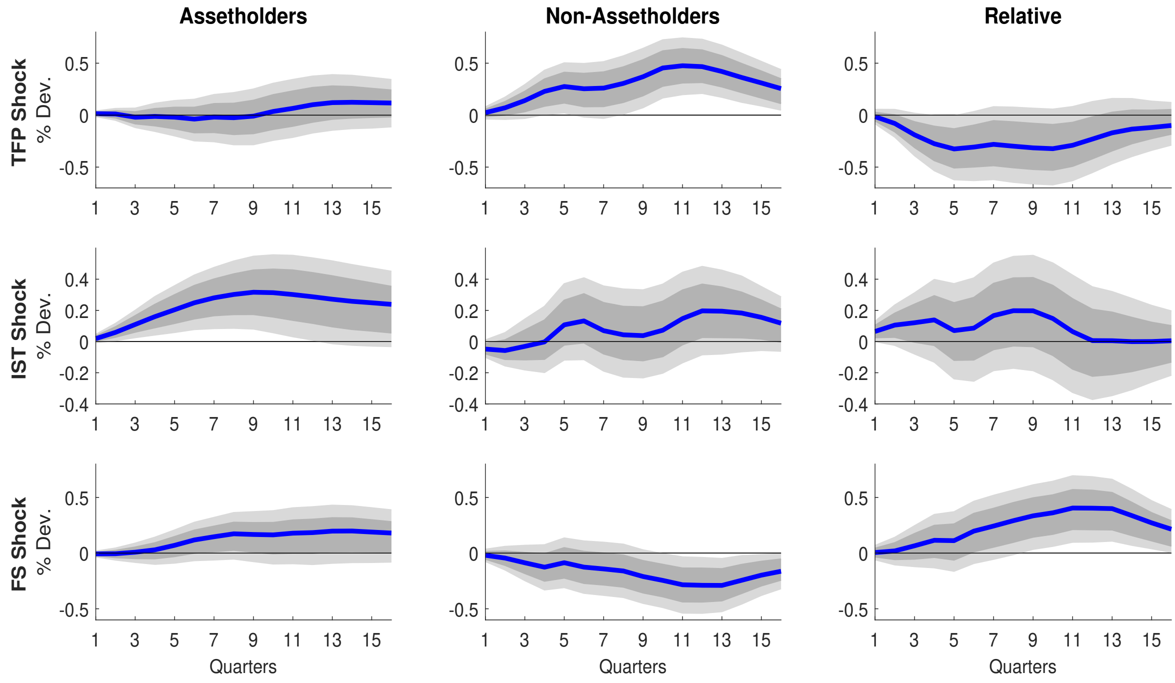
Notes: The figure displays the IRFs of non-durables and services expenditures for households sorted according to the probability-threshold method.

Figure D.8: Net Income - Different sorting method



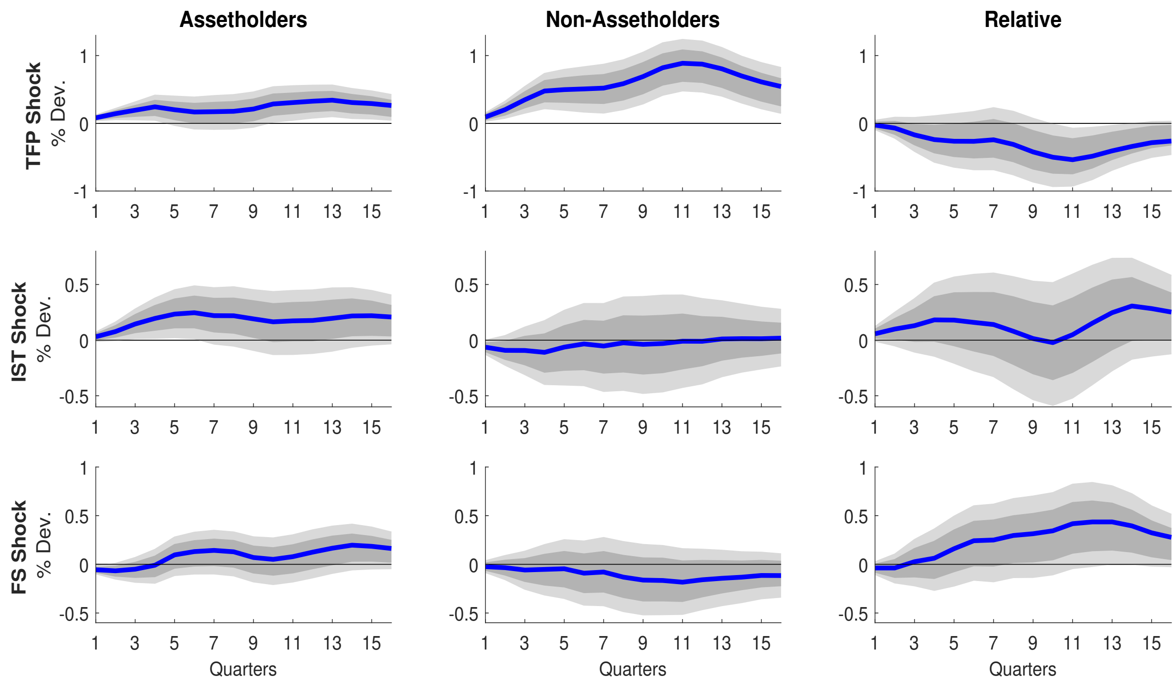
Notes: The figure displays the IRFs of net income for households sorted according to the probability-threshold method.

Figure D.9: Non-durables and services expenditure - Utilization-adjusted TFP



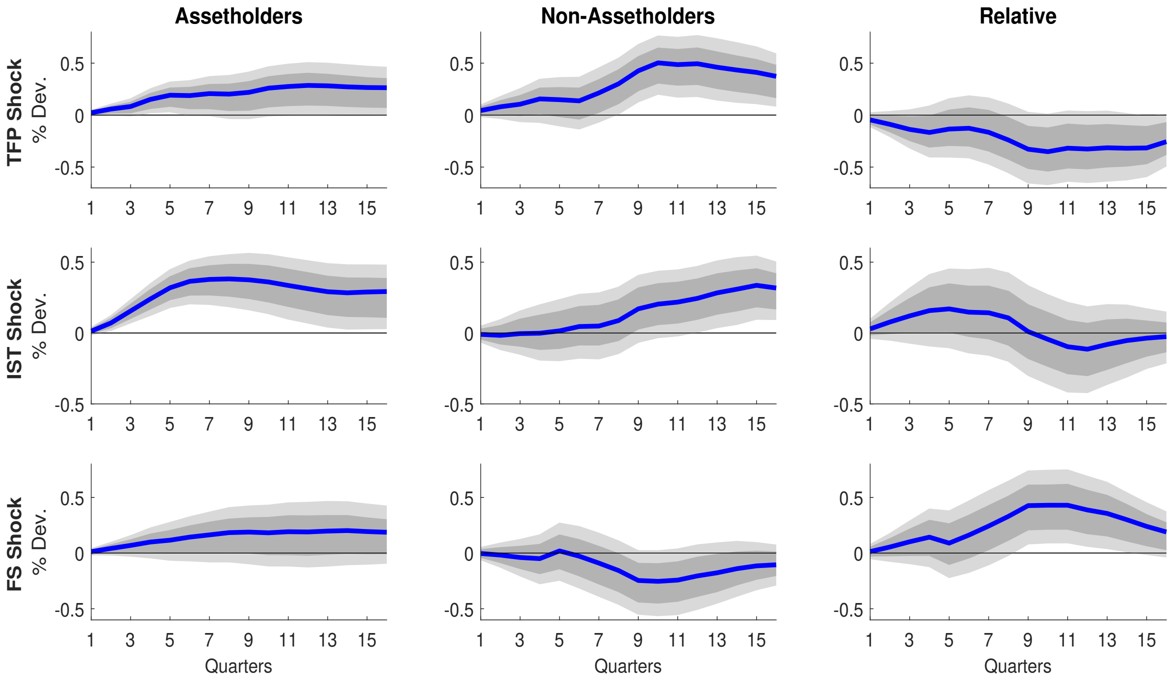
Notes: The figure displays the IRFs of non-durables and services expenditures to the shocks identified in the VAR with utilization-adjusted TFP.

Figure D.10: Net Income - Utilization-adjusted TFP



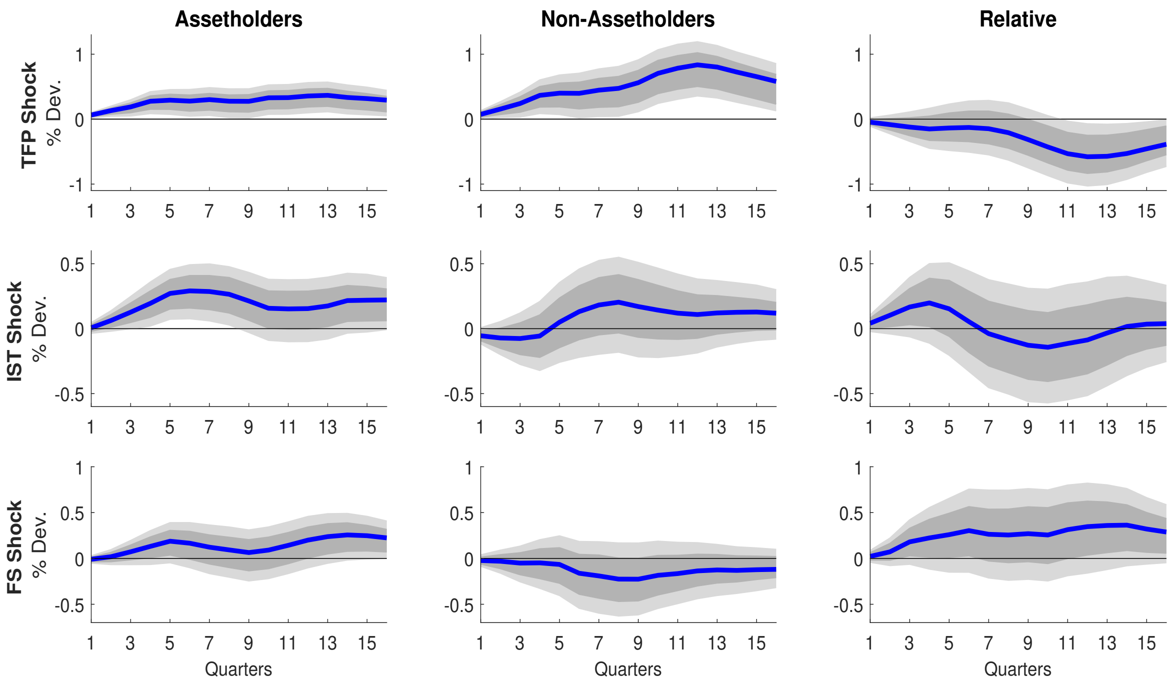
Notes: The figure displays the IRFs of net income to the shocks identified in the VAR with utilization-adjusted TFP.

Figure D.11: Non-durables and services expenditure - Extended VAR



Notes: The figure displays the IRFs of non-durables and services expenditures to the shocks identified in the extended VAR.

Figure D.12: Net Income - Extended VAR



Notes: The figure displays the IRFs of net income to the shocks identified in the extended VAR.

D.5 Robustness: predictive regressions

We now provide further tests on the robustness of relative consumption as a predictor of future expected excess returns.

Relative consumption based on stockholdings First, we estimate regression (3) by using the consumption of stockholders relative to that of non-stockholders (rather than assetholders vs. non-assetholders). As reported in Table D.7, the results are essentially identical when using this alternative measure of relative consumption—in fact, the unconditional relative measure is now found to significantly predict future returns even at the 16-quarters horizon.

Controlling for cay Second, we make sure that relative consumption remains significant even when controlling for an additional, well-known stock return predictor—namely, the aggregate consumption-wealth ratio (cay) proposed by Lettau and Ludvigson (2001). This variable represents the residual from a cointegrating relationship between consumption, asset wealth, and labor income.[37] Such residual captures fluctuations in aggregate consumption relative to aggregate (human and non-human) wealth, and can therefore be seen as an additional measure of aggregate risk. Table D.8 shows that the coefficient on this variable is positive and significant at several horizons, in line with Lettau and Ludvigson (2001). Even in this case, the unconditional relative consumption (Panel A), as well as relative consumption conditional on FS shocks (Panel B) remain strongly significant.

QoQ growth filter Finally, Table D.9 displays the results from considering quarter-on-quarter (QoQ) growth rates for aggregate and relative consumption, rather than 8-quarters log-differences as in Hamilton (2018). Since there is no *a priori* theoretical guidance on what procedure is best suited to isolate cyclical fluctuations, it is instructive to compare the ability of changes in relative consumption to predict future excess returns using a different filter. Clearly, the results remain unaltered in terms of statistical significance, although (obviously) the size of the coefficients varies (given the larger noise inherent to the first-difference filter). Overall, we conclude that the emergence of changes in consumption inequality as a stock return predictor is a reliable feature of the data.

[37]The series is available at the quarterly frequency from the authors' webpage.

Table D.7: Predictive regressions - Stockholders

	Panel A		Panel B			
	$r^{ex}_{t,t+h} = \alpha + \beta_1 g_{c,t} + \beta_2 g_{rc,t}$		$r^{ex}_{t,t+h} = \alpha + \beta_1 g_{c,t} + \beta_2 g^{TFP}_{rc,t} + \beta_3 g^{IST}_{rc,t} + \beta_4 g^{FS}_{rc,t}$			
h	β_1	β_2	β_1	β_2	β_3	β_4
1	−1.47 (1.46) [0.31]	**2.20** (1.06) [0.04]	−1.41 (1.37) [0.30]	0.29 (3.19) [0.93]	−0.19 (3.89) [0.96]	3.29 (2.43) [0.18]
4	−1.46 (1.18) [0.22]	**1.56** (0.85) [0.07]	−1.65 (1.02) [0.11]	1.07 (2.19) [0.62]	0.07 (3.16) [0.98]	2.91 (1.74) [0.10]
8	**-2.37** (0.79) [0.00]	**1.51** (0.66) [0.02]	**-2.66** (0.71) [0.00]	0.64 (1.40) [0.65]	0.54 (1.83) [0.77]	**2.87** (1.01) [0.01]
12	**-2.66** (0.65) [0.00]	**1.35** (0.45) [0.00]	**-3.02** (0.62) [0.00]	−1.26 (1.05) [0.23]	1.60 (1.49) [0.28]	**2.52** (0.80) [0.00]
16	**-2.42** (0.48) [0.00]	**0.73** (0.40) [0.07]	**-2.86** (0.50) [0.00]	−1.19 (0.91) [0.19]	1.32 (1.23) [0.28]	**1.91** (0.78) [0.02]
20	**-2.06** (0.40) [0.00]	0.26 (0.43) [0.55]	**-2.47** (0.41) [0.00]	−0.79 (0.97) [0.42]	0.82 (1.28) [0.52]	**1.25** (0.70) [0.08]

Notes: The table presents results of predictive regressions of the form $r^{ex}_{t,t+h} = \alpha + \boldsymbol{\beta x_t} + \epsilon_{t+h}$, where h denotes the horizon in quarters and $r^{ex}_{t,t+h}$ denotes annualized excess returns between period t and $t+h$. $\boldsymbol{x_t}$ represents the matrix of (demeaned) predictors, which includes: in Panel A, aggregate and relative consumption growth; in Panel B, aggregate consumption growth and relative consumption growth conditioned on each shock at a time. Growth rates are computed as 8-quarters log-differences. In this robustness check, relative consumption is computed using the consumption series of stockholders and non-stockholders, rather than assetholders and non-assetholders. For each regression, Newey-West corrected standard errors (4 lags) appear in parentheses below the coefficient estimate, while p-values are reported in square brackets. Significant coefficients at the ten percent level are highlighted in bold. The sample covers the period 1982Q4-2017Q4.

Table D.8: Predictive regressions - CAY

	Panel A			Panel B				
	$r^{ex}_{t,t+h} = \alpha + \beta_1 g_{c,t} + \beta_2 cay_t + \beta_3 g_{rc,t}$			$r^{ex}_{t,t+h} = \alpha + \beta_1 g_{c,t} + \beta_2 cay_t + \beta_3 g^{TFP}_{rc,t} + \beta_4 g^{IST}_{rc,t} + \beta_5 g^{FS}_{rc,t}$				
h	β_1	β_2	β_3	β_1	β_2	β_3	β_4	β_5
1	−0.95 (1.48) [0.52]	−1.86 (1.39) [0.18]	**2.08** (1.03) [0.05]	−1.06 (1.45) [0.47]	**-2.45** (1.44) [0.09]	−0.57 (2.09) [0.78]	0.36 (3.34) [0.91]	**4.04** (2.39) [0.09]
4	−1.57 (1.14) [0.17]	0.78 (1.16) [0.50]	**1.47** (0.74) [0.05]	−1.77 (1.06) [0.10]	0.27 (1.11) [0.81]	0.17 (1.33) [0.90]	0.71 (2.34) [0.76]	**3.18** (1.65) [0.06]
8	**-2.78** (0.97) [0.00]	2.06 (1.40) [0.14]	**1.30** (0.58) [0.03]	**-3.13** (0.81) [0.00]	1.71 (1.23) [0.17]	−0.23 (0.74) [0.76]	1.05 (1.19) [0.38]	**2.98** (0.91) [0.00]
12	**-3.63** (0.83) [0.00]	**3.35** (1.14) [0.00]	**1.27** (0.38) [0.00]	**-3.87** (0.67) [0.00]	**2.93** (0.89) [0.00]	−0.96 (0.72) [0.18]	0.79 (1.10) [0.48]	**2.75** (0.64) [0.00]
16	**-3.48** (0.54) [0.00]	**3.40** (0.74) [0.00]	**0.85** (0.31) [0.01]	**-3.71** (0.47) [0.00]	**3.05** (0.57) [0.00]	−0.68 (0.63) [0.29]	0.05 (0.93) [0.95]	**2.28** (0.55) [0.00]
20	**-2.84** (0.46) [0.00]	**2.52** (0.78) [0.00]	0.41 (0.35) [0.24]	**-3.07** (0.44) [0.00]	**2.35** (0.64) [0.00]	−0.33 (0.65) [0.61]	0.01 (1.05) [1.00]	**1.38** (0.46) [0.00]

Notes: The table presents results of predictive regressions of the form $r^{ex}_{t,t+h} = \alpha + \boldsymbol{\beta x_t} + \epsilon_{t+h}$, where h denotes the horizon in quarters and $r^{ex}_{t,t+h}$ denotes annualized excess returns between period t and $t+h$. $\boldsymbol{x_t}$ represents the matrix of (demeaned) predictors, which, in this robustness check, includes: in Panel A, aggregate consumption growth, the variable CAY, and relative consumption growth; in Panel B, aggregate consumption growth, the variable CAY, and relative consumption growth conditioned on each shock at a time. Growth rates are computed as 8-quarters log-differences. For each regression, Newey-West corrected standard errors (4 lags) appear in parentheses below the coefficient estimate, while p-values are reported in square brackets. Significant coefficients at the ten percent level are highlighted in bold. The sample covers the period 1982Q4-2017Q4.

Table D.9: Predictive regressions - First differences

	Panel A		Panel B			
	$r^{ex}_{t,t+h} = \alpha + \beta_1 g_{c,t} + \beta_2 g_{rc,t}$		$r^{ex}_{t,t+h} = \alpha + \beta_1 g_{c,t} + \beta_2 g^{TFP}_{rc,t} + \beta_3 g^{IST}_{rc,t} + \beta_4 g^{FS}_{rc,t}$			
h	β_1	β_2	β_1	β_2	β_3	β_4
1	1.11	**8.10**	−0.34	−4.10	−3.19	13.21
	(4.18)	(4.51)	(4.91)	(11.92)	(14.55)	(13.38)
	[0.79]	[0.07]	[0.95]	[0.73]	[0.83]	[0.33]
4	−0.26	**6.65**	−1.67	−4.45	2.23	11.76
	(3.45)	(3.18)	(3.40)	(8.13)	(10.00)	(7.26)
	[0.94]	[0.04]	[0.62]	[0.58]	[0.82]	[0.11]
8	−1.09	**4.24**	−2.75	−1.11	−2.22	**15.68**
	(2.19)	(2.41)	(1.96)	(5.38)	(7.63)	(6.16)
	[0.62]	[0.08]	[0.16]	[0.84]	[0.77]	[0.01]
12	−2.42	2.83	**-3.62**	1.28	−0.87	**11.85**
	(1.50)	(1.97)	(1.43)	(4.39)	(5.50)	(3.59)
	[0.11]	[0.15]	[0.01]	[0.77]	[0.87]	[0.00]
16	**-2.47**	2.31	**-3.57**	−1.84	−0.85	**9.51**
	(1.28)	(1.68)	(1.34)	(3.61)	(5.28)	(3.60)
	[0.06]	[0.17]	[0.01]	[0.61]	[0.87]	[0.01]
20	**-3.06**	0.68	**-3.77**	−0.20	−3.12	**7.33**
	(1.17)	(1.43)	(1.21)	(3.06)	(3.64)	(3.28)
	[0.01]	[0.64]	[0.00]	[0.95]	[0.39]	[0.03]

Notes: The table presents results of predictive regressions of the form $r^{ex}_{t,t+h} = \alpha + \boldsymbol{\beta x_t} + \epsilon_{t+h}$, where h denotes the horizon in quarters and $r^{ex}_{t,t+h}$ denotes annualized excess returns between period t and $t+h$. $\boldsymbol{x_t}$ represents the matrix of (demeaned) predictors, which includes: in Panel A, aggregate and relative consumption growth; in Panel B, aggregate consumption growth and relative consumption growth conditioned on each shock at a time. In this robustness check, growth rates are computed as 1-quarter log-differences, instead of 8-quarters log-differences. For each regression, Newey-West corrected standard errors (4 lags) appear in parentheses below the coefficient estimate, while p-values are reported in square brackets. Significant coefficients at the ten percent level are highlighted in bold. The sample covers the period 1982Q4-2017Q4.

E A model with concentrated capital ownership

This appendix details the model employed in Section 4, as well as its calibration and ability to match macroeconomic and asset-pricing moments.

Households Assetholders own firms through equity shares, and smooth consumption intertemporally by trading one-period bonds. Non-assetholders are assumed to be excluded from the bond and the stock markets. Both agents are assumed to inelastically supply their entire time-endowment to the firms. Households are equally productive and, therefore, all earn the same wage, regardless of their type. The fraction of assetholders in the total population of consumers equals $1-\gamma$.

The utility of the representative assetholder reads as

$$E_0 \sum_{t=0}^{\infty} \beta^t \frac{(c_t^a - \chi_c h_t)^{1-\sigma} - 1}{1-\sigma}, \tag{E.1}$$

where we assume assetholders to exhibit external habits in utility, with the habit stock, h_t, weighing on per-period utility by the parameter χ_c, and evolving according to the following law of motion (Jaccard, 2014):

$$h_t = m h_{t-1} + (1-m) c_{t-1}^a, \tag{E.2}$$

where c_{t-1}^a denotes assetholders' per-capita consumption at time $t-1$. The parameter m allows us to introduce a slow-moving component in habit formation. Similar to Campbell and Cochrane (1999), $1-m$ captures how sensitive the reference level is to changes in assetholders' per-capita consumption.

Consumption and saving decisions are limited by the following budget constraint

$$c_t^a + p_t^s q_{t+1}^s + p_t^b q_{t+1}^b = (p_t^s + d_t) q_t^s + q_t^b + w_t n_t^a. \tag{E.3}$$

which states that consumption and the purchase of equity shares (in quantity q_{t+1}^s at the price p_t^s) as well as of one-period bonds (in quantity q_{t+1}^b at the price p_t^b) must be financed by labor income, $w_t n_t^a$ (where $n_t^a = 1$), and the returns on the financial investments. Shares purchased in the previous period yield a dividend d_t, while one-period bonds yield a single consumption unit per-bond in the following period.

The two agents differ only for their ability to access financial markets.[38] Being

[38]Since non-assetholders do not price securities, they can in principle have exactly the same preferences as assetholders, without affecting the equilibrium conditions.

unable to smooth consumption intertemporally, non-assetholders consume their labor income hand-to-mouth, so that

$$c_t^{na} = w_t n_t^{na}, \tag{E.4}$$

where w_t is the wage and $n_t^{na} = 1$.

Asset prices The first-order conditions for assetholders' optimization problem with respect to c_t^a, q_{t+1}^s, and q_{t+1}^b are:

$$\lambda_t = (c_t^a - \chi_c h_t)^{-\sigma}, \tag{E.5}$$

$$p_t^s = E_t m_{t,t+1}(p_{t+1}^s + d_{t+1}), \tag{E.6}$$

$$p_t^b = E_t m_{t,t+1}, \tag{E.7}$$

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 assetholder's stochastic discount factor. The first-order conditions (E.6) and (E.7) govern asset-pricing dynamics. In particular, the risk-free rate is given by $r_{t+1}^b = 1/p_t^b = 1/E_t m_{t,t+1}$, while the stock return is $r_{t+1}^s = \frac{p_{t+1}^s + d_{t+1}^s}{p_t^s}$. Asset prices depend on the preferences of the marginal investor: the assetholder, in our case.

Firms Firms operate under perfect competition and produce according to a standard Cobb-Douglas technology:

$$y_t = A z_t n_t^{1-\alpha_t} k_t^{\alpha_t}, \ \alpha_t \in (0,1), \tag{E.8}$$

where n_t is aggregate employment, k_t is aggregate capital, z_t is total factor productivity and A is a scaling factor (to be discussed in Section E.1). The labor share of income, $ls_t \equiv 1 - \alpha_t$, is allowed to fluctuate over time.

Following Jermann (1998), capital accumulation follows a law of motion featuring capital adjustment costs:

$$k_{t+1} = (1-\delta)k_t + \phi\left(\frac{i_t}{k_t}\right)k_t, \tag{E.9}$$

where δ is the depreciation rate and

$$\phi\left(\frac{i_t}{k_t}\right) = \left[\frac{a_1}{1-1/\chi_k}\left(\frac{i_t}{k_t}\right)^{1-1/\chi_k} + a_2\right] \tag{E.10}$$

is a concave adjustment-cost function. In particular, $\chi_k \to 0\ (\infty)$ implies higher (lower) adjustment costs.

The firm's problem consists of choosing labor, capital, and investment to maximize

$$\max_{i_t, n_t, k_{t+1}} E_0 \sum_{t=0}^{\infty} m_{t,t+1} \left\{ d_t - q_t[k_{t+1} - (1-\delta)k_t - \phi(i_t/k_t)k_t], \right\} \tag{E.11}$$

subject to the constraints (E.8), (E.9), and (E.10), where q_t is the shadow price of capital.

Dividends are defined as

$$d_t = y_t - w_t n_t - \frac{i_t}{\mu_t}, \tag{E.12}$$

where, following Greenwood et al. (1988) and Liu et al. (2013), μ_t accounts for investment-specific technological change. Profit maximization leads to:

$$w_t = (1 - \alpha_t) y_t / n_t, \tag{E.13}$$

implying that dividends can be rewritten as

$$d_t = \alpha_t y_t - \frac{i_t}{\mu_t}, \tag{E.14}$$

whereas the first-order condition with respect to capital investment is

$$\phi^{'} \left(\frac{i_t}{k_t} \right) = \frac{1}{\mu_t q_t}, \tag{E.15}$$

with

$$\phi^{'} \left(\frac{i_t}{k_t} \right) = a_1 \left(\frac{i_t}{k_t} \right)^{-1/\chi_k}. \tag{E.16}$$

Finally, the firm's optimal decision regarding capital yields

$$q_t = E_t \left\{ m_{t,t+1} \left[\alpha_{t+1} \frac{y_{t+1}}{k_{t+1}} + q_{t+1} \left((1-\delta) + \phi \left(\frac{i_{t+1}}{k_{t+1}} \right) - \phi^{'} \left(\frac{i_{t+1}}{k_{t+1}} \right) \frac{i_{t+1}}{k_{t+1}} \right) \right] \right\}. \tag{E.17}$$

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 (E.4), (E.5), (E.6), (E.7), (E.13), (E.15) and (E.17) hold, all constraints are satisfied, and all markets clear. More specifically, labor-market clearing requires that

$$n_t = \gamma n_t^{na} + (1 - \gamma) n_t^a = 1, \tag{E.18}$$

while equilibrium in the good market implies

$$y_t = c_t + i_t, \tag{E.19}$$

where

$$c_t = \gamma c_t^{na} + (1-\gamma) c_t^a \tag{E.20}$$

defines 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

$$(1-\gamma) q_t^s = 1, \tag{E.21}$$

where the left side of the equality represents the aggregate demand of stocks, since only a fraction $(1-\gamma)$ of the population participates in the stock market. Therefore, in equilibrium the budget constraint (E.3) for the representative assetholder reads as

$$c_t^a = w_t n_t^a + \frac{d_t}{1-\gamma}. \tag{E.22}$$

Finally, plugging (E.4) and (E.22) into equation (E.20) yields

$$c_t = \gamma w_t n_t^{na} + (1-\gamma)\left(w_t n_t^a + \frac{d_t}{1-\gamma}\right), \tag{E.23}$$

which, given the assumption that both non-assetholders and assetholders supply all their time-endowment to firms ($n_t^{na} = n_t^a = 1$), becomes $c_t = w_t + d_t$; that is, aggregate consumption consists of labor income plus dividends.

Exogenous state variables The dynamics of the three exogenous state variables in the model, namely investment-specific technology μ_t, total factor productivity z_t and the labor share ls_t, are governed by the trivariate VAR estimated as in equation (1). Given the permanent nature of IST and TFP shocks, the model exhibits non-stationary dynamics. Thus, in Appendix F.2 we rewrite it in stationary form. In the remainder, '$\sim$' will be used to denote variables in log-deviation from their trend.

Unlike most of the extant literature (Justiniano and Primiceri, 2008; Papanikolaou, 2011; Lansing, 2015, among the others), we do not assume that exogenous processes are independent. Relatedly, Ríos-Rull and Santaeulalia-Llopis (2010), Santaeulalia-Llopis (2011) and Choi and Ríos-Rull (2020) emphasize the dynamic effects of technology shocks on the labor share, and how this bears important implications for the propagation of the shocks to aggregate variables. To take this into account we assume that TFP, the relative price of investment, and the labor share follow a VAR process whose parameters are estimated, as we describe in the next section.

Table E.1: Baseline parameter values

Description	Parameter	Value
Calibrated		
Fraction of non-assetholders	γ	0.3300
Depreciation rate	δ	0.0271
Capital share of income	α	0.3500
Discount rate	β	0.9893
Local utility curvature	σ	4.0000
Estimated		
Capital adjustment cost	χ_k	0.2089
Habit weight	χ_c	0.6936
Habit stock persistence	m	0.9498

Notes: The model is simulated at a quarterly frequency.

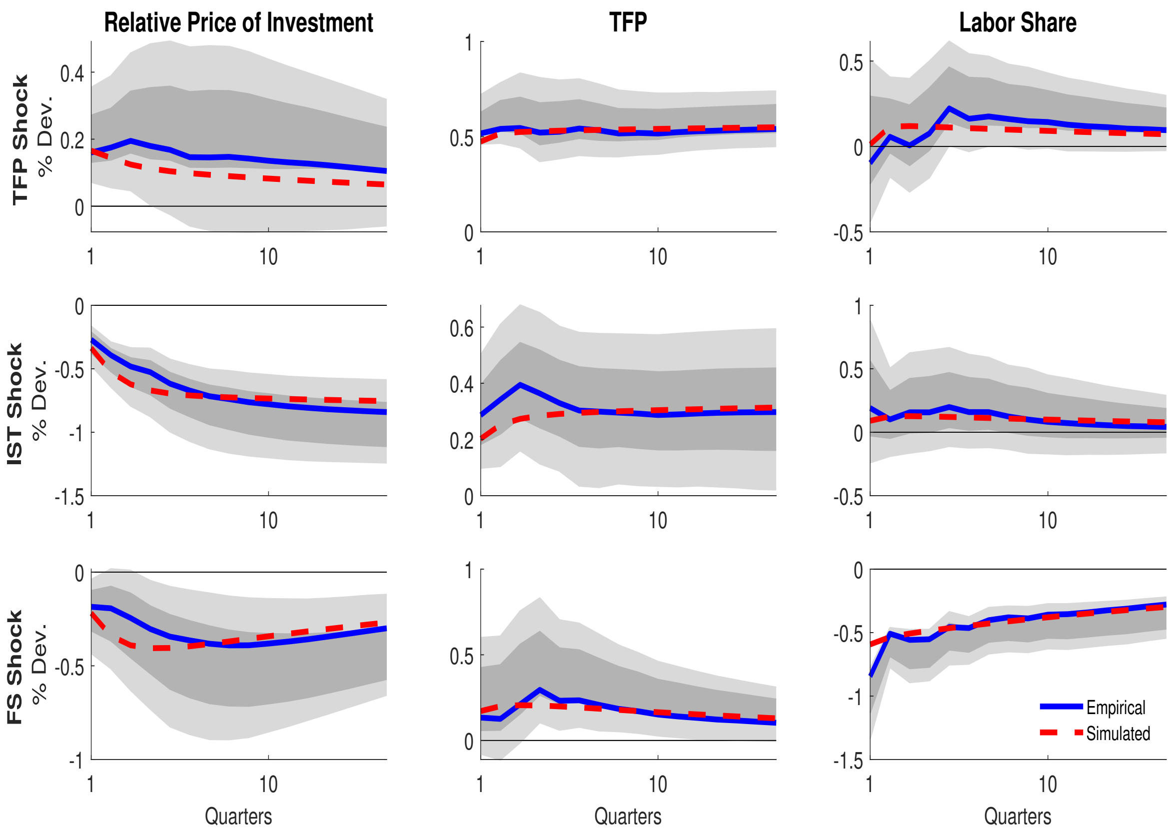
E.1 Calibration

The model is solved by second-order perturbation methods. A time period in the model is taken to be one quarter. We split the parameters into two groups. The first group of parameters is calibrated to match targeted long-run relationships, while the second group is estimated both via impulse-response matching, as well as by matching a subset of selected unconditional macroeconomic moments. The baseline parameter values are summarized in Table E.1.

E.1.1 Calibrated parameters

The fraction of non-assetholder, γ, is set to 0.33, which represents a mid-value over the sample 1982Q4-2017Q4. The calibration strategy for the depreciation rate (δ), the discount rate (β), and the unit parameter in the production function (A) follows Ríos-Rull and Santaeulalia-Llopis (2010). We target the capital-to-output ratio in yearly terms $k/y = 2.31$, and the investment-output ratio $i/y = 0.25$. Given these targets, from the relationship $i/y = \delta k/y$, we retrieve $\delta = 0.0271$. After evaluating equation (E.17) at the steady state and setting the capital share $\alpha = 0.35$—as in Choi and Ríos-Rull (2020)—we obtain $1 = \beta(1 - \delta + \alpha y/k)$, which yields $\beta = 0.9893$. Without loss of generality, we normalize steady-state output to one, thus solving equation (E.8) for $A = 1/n(k/n)^{-\alpha}$. Finally, the local utility curvature parameter, σ, is set to 4, which is in line with standard calibrations in the production-based asset-pricing literature, lying within the range of values adopted in Lansing (2015) (3.3) and Jermann (1998) (5).

Figure E.1: IRFs Matching



Notes: The figure displays the structural impulse-response functions estimated from the VAR in equation (1) (blue solid lines) together with the 90% and 68% confidence intervals (light-grey and dark-grey areas, respectively); and the corresponding IRFs generated by the estimated model (red dashed lines).

E.1.2 Estimated parameters

The remaining coefficients include the capital adjustment cost parameter, χ_k,[39] the consumption utility curvature parameter, χ_c, the parameter capturing the persistence of the habit stock, m, as well as the parameters of the VAR governing the dynamics of the exogenous process for TFP, the relative price of investment, and the labor share.[40] These are estimated by matching both some empirical impulse-responses (e.g., Christiano et al., 2005; Iacoviello, 2005, among others), as well as a selected number of unconditional macroeconomic moments.

Specifically, we match the responses of TFP, the relative price of investment, and the labor share to the TFP, IST, and FS shocks. Figure E.1 reports the estimated impulse-

[39] Both a_1 and a_2 in equation (E.10) are constructed so that capital adjustment costs do not affect the steady state of the economy. Thus, we set $a_1 = \delta^{1/\chi_k}$ and $a_2 = \delta - \frac{\delta}{1-1/\chi_k}$, which implies that $\phi\left(\frac{i}{k}\right) = \delta$, $\frac{i}{k} = \delta$ and $\phi'\left(\frac{i}{k}\right) = 1$ in the steady state.

[40] In line with the VAR estimated in Section 3.1, we select a VAR(4). The results are robust to choosing a VAR(1).

response functions from the model, alongside their empirical counterparts from the VAR model. We also target the unconditional volatility of the growth rates of output, consumption, investment, and dividends, as well as the unconditional correlation between the growth rate of output and that of dividends.[41] The matched moments are reported, alongside their empirical counterparts, in the upper panel of Table E.2.

We estimate $\chi_k = 0.21$, which is in line with Jermann (1998), Guvenen (2009) and Chen (2017). As for χ_c, this is estimated at 0.69, which can be considered in line with standard calibrations in the production-based asset-pricing literature, lying within the range of values adopted in Lansing (2015) (0.2) and Jermann (1998) (0.82). Finally, $m = 0.95$, in line with Cochrane (2017), and close to the persistence of the surplus-consumption ratio considered by Campbell and Cochrane (1999).

Moment matching The theoretical business-cycle statistics, together with their data counterparts, are reported in Table E.2. The framework does a fairly good job at replicating the unconditional targeted moments, returning output and consumption growth volatilities above their data counterparts, while the opposite holds true for investment and dividend growth.[42] The limited participation economy is also able to replicate a number of non-targeted moments, such as the unconditional volatility of relative consumption, whose dynamics are central to our narrative. Moreover, the model shares a typical feature of RBC frameworks, namely a rather high correlation of all macroeconomic aggregates with output. On the other hand, the output correlations of the exogenous drivers (TFP, IST, and the labor share) compare fairly well with the point estimates.

As shown in Table E.3, the two-agent economy is also able to account for plausible stock excess returns, both in terms of mean and standard deviation. The close mapping between relative consumption and the dividend-to-labor income ratio is of key importance, in this respect. Restricting access to financial investment to a limited number of assetholders raises the equity premium they demand, through the connection between their consumption growth and financial income, which is intrinsically more volatile. Along with these properties, the model returns a risk-free rate that is close in line with the data, though it appears rather volatile. As in Jermann (1998) and Lansing (2015), the combination of habit utility and (high) capital adjustment costs that generates sufficiently volatile stock returns induces, at the same time, strong fluctuations in investors' marginal utility, which reflects into the volatility of the risk-free rate.

[41] The latter is particularly informative to pin down the capital adjustment cost parameter.

[42] Guvenen (2009) and Chen (2017) have extensively discussed how selecting the parameters characterizing household utility and the capital adjustment costs typically entails some distinctive trade-offs when trying to match the volatility of investment, dividends and consumption.

Table E.2: Macroeconomic moments

Variable	Empirical	Simulated
	Targeted	
σ_{g_y}	0.71 [0.58,0.80]	1.14
σ_{g_c}	0.52 [0.42,0.60]	0.88
σ_{g_i}	3.16 [2.46,3.81]	2.18
σ_{g_d}	4.98 [3.13,7]	2.82
$corr_{g_d,g_y}$	0.25 [0.1,0.44]	0.92
	Implied	
$\sigma_{g_{rc}}$	0.68 [0.56,0.79]	0.45
$corr_{g_c,g_y}$	0.74 [0.64,0.81]	0.98
$corr_{g_i,g_y}$	0.69 [0.6,0.75]	0.97
$corr_{g_{rc},g_y}$	0.15 [$-$0.03,0.26]	0.84
$corr_{g_z,g_y}$	0.49 [0.33,0.6]	0.65
$corr_{g_\mu,g_y}$	-0.06 [$-0-16$,0.1]	0.30
$corr_{\log(ls),g_y}$	-0.08 [$-$0.27,0.07]	-0.16

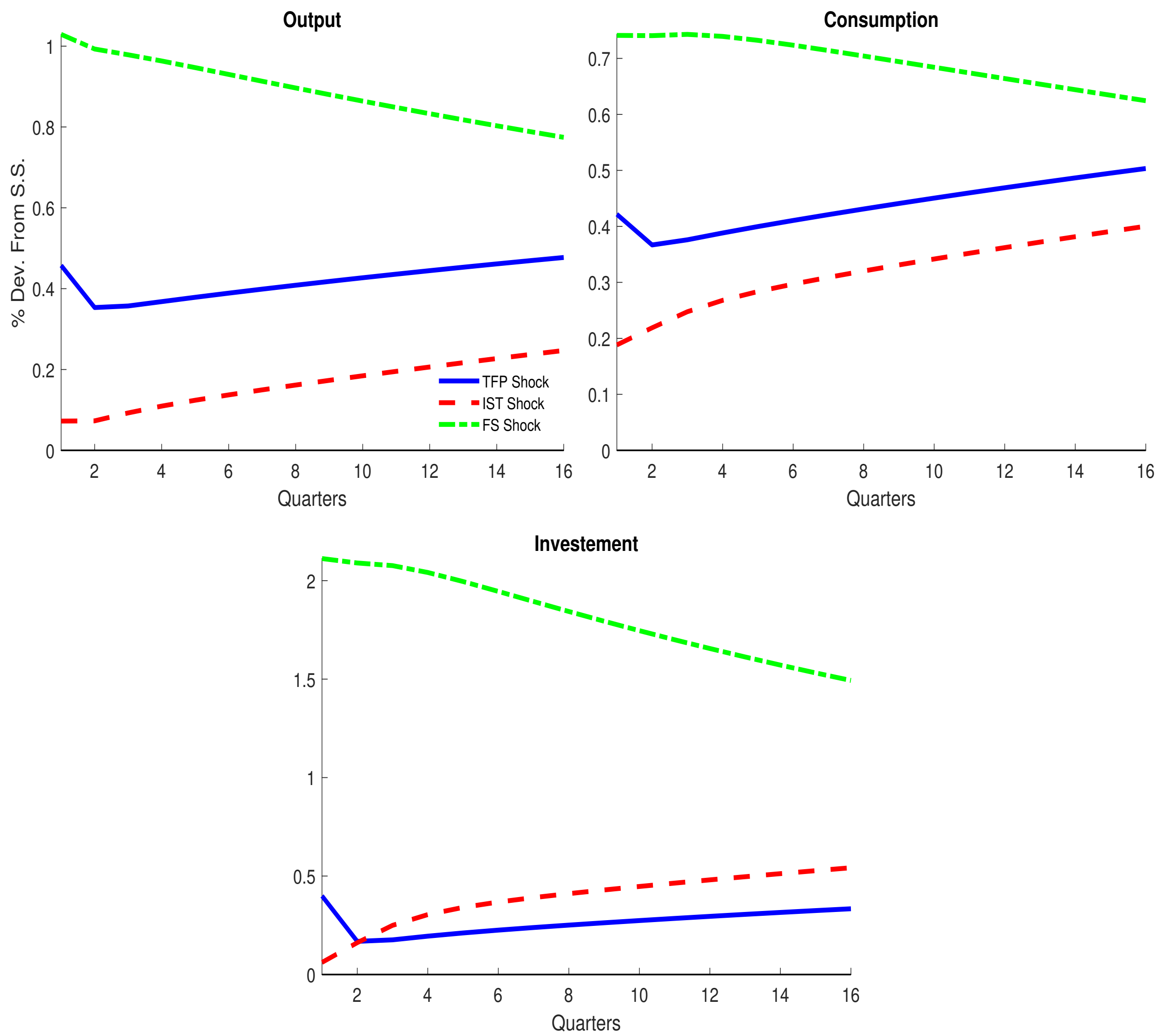
Notes: Bootstrapped 90% confidence intervals in brackets. All moments refer to quarterly variables. g_x denotes the first-differenced logarithm of a generic variable x.

Table E.3: Asset-pricing moments

Variable	Empirical	Simulated
$E(r^b)$	1.07 [0.19,1.84]	1.17
$E(r^s - r^b)$	4.39 [2.56,6.83]	4.59
σ_{r^b}	1.50 [1.00,1.78]	4.54
$\sigma_{r^s - r^b}$	15.67 [14.43,17.89]	19.94

Notes: Bootstrapped 90% confidence intervals in brackets. All moments refer to annualized variables.

Figure E.2: Macroeconomic aggregates - IRFs



Notes: Output, consumption and investment responses to TFP, IST, and FS shocks.

E.2 Stationary representation of the model

Given the permanent nature of TFP and IST shocks, the model exhibits non-stationary dynamics. As such, it needs to be rewritten in stationary form by appropriately transforming the growing variables. Define $\Gamma_t \equiv (z_t \mu_t^{\alpha})^{\frac{1}{1-\alpha}}$, and the associated growth rate $g_{\Gamma,t} \equiv \Delta \log(\Gamma_t) = \frac{1}{1-\alpha}\left[g_{z,t} + \alpha g_{\mu,t}\right]$, where $g_{z,t} \equiv \Delta \log(z_t)$ and $g_{\mu,t} \equiv \Delta \log(\mu_t)$ denote the growth rates of TFP and IST, respectively. We apply the following transformations:

$$\tilde{y}_t \equiv \frac{y_t}{\Gamma_t}, \quad \tilde{k}_t \equiv \frac{k_t}{\Gamma_{t-1}\mu_{t-1}}, \quad \tilde{i}_t \equiv \frac{i_t}{\Gamma_t \mu_t}, \quad \tilde{q}_t \equiv q_t \mu_t, \quad \tilde{d}_t \equiv \frac{d_t}{\Gamma_t}, \quad \tilde{w}_t \equiv \frac{w_t}{\Gamma_t}, \quad \tilde{c}_t \equiv \frac{c_t}{\Gamma_t},$$
$$\tilde{c}_t^{na} \equiv \frac{c_t^{na}}{\Gamma_t}, \quad \tilde{c}_t^{a} \equiv \frac{c_t^{a}}{\Gamma_t}, \quad \tilde{h}_t \equiv \frac{h_t}{\Gamma_t}, \quad \tilde{\lambda}_t \equiv \lambda_t \Gamma_t^{\sigma}.$$

Then, the stationary equilibrium is the solution to the following system of equations:

$$\tilde{c}_t^{na} = \tilde{w}_t, \tag{E.24}$$

$$\tilde{c}_t^{a} = \tilde{w}_t + \frac{\tilde{d}_t}{1-\gamma}, \tag{E.25}$$

$$\tilde{c}_t = \tilde{w}_t + \tilde{d}_t, \tag{E.26}$$

$$\tilde{h}_t = \exp\left(-g_{\Gamma,t}\right)[m\tilde{h}_{t-1} + (1-m)\tilde{c}_{t-1}^{a}], \tag{E.27}$$

$$\tilde{\lambda}_t = (\tilde{c}_t^{a} - \chi_c \tilde{h}_t)^{-\sigma}, \tag{E.28}$$

$$m_{t,t+1} = \beta E_t \left(\frac{\tilde{\lambda}_{t+1}}{\tilde{\lambda}_t}\right) \exp\left(-\sigma g_{\Gamma,t+1}\right), \tag{E.29}$$

$$p_t^s = E_t m_{t,t+1}(p_{t+1}^s + \tilde{d}_{t+1}), \tag{E.30}$$

$$p_t^b = E_t m_{t,t+1}, \tag{E.31}$$

$$\tilde{y}_t = \exp\left[-\frac{\alpha}{1-\alpha}(g_{z,t} + g_{\mu,t})\right] A n^{1-\alpha_t} \tilde{k}_t^{\alpha_t}, \tag{E.32}$$

$$\tilde{k}_{t+1} = \exp\left(-g_{\Gamma,t} - g_{\mu,t}\right)\left[(1-\delta)\tilde{k}_t + \phi\left(\frac{\tilde{i}_t}{\tilde{k}_t}\right)\tilde{k}_t\right], \tag{E.33}$$

$$\phi\left(\frac{\tilde{i}_t}{\tilde{k}_t}\right) = \frac{a_1}{1-1/\chi_k}\left[\frac{\tilde{i}_t}{\tilde{k}_t}\exp\left(g_{\Gamma,t} + g_{\mu,t}\right)\right]^{1-1/\chi_k} + a_2, \tag{E.34}$$

$$\tilde{d}_t = \tilde{y}_t - \tilde{w}_t n_t - \tilde{i}_t, \tag{E.35}$$

$$\tilde{w}_t = (1-\alpha_t)\frac{\tilde{y}_t}{n_t}, \tag{E.36}$$

$$\phi'\left(\frac{\tilde{i}_t}{\tilde{k}_t}\right) = a_1\left[\frac{\tilde{i}_t}{\tilde{k}_t}\exp\left(g_{\Gamma,t} + g_{\mu,t}\right)\right]^{-1/\chi_k}, \tag{E.37}$$

$$\phi'\left(\frac{\tilde{i}_t}{\tilde{k}_t}\right) = \frac{1}{\tilde{q}_t}, \tag{E.38}$$

$$\tilde{q}_t = E_t m_{t,t+1}\Bigg\{\alpha_{t+1}\frac{\tilde{y}_{t+1}}{\tilde{k}_{t+1}}\exp\left(g_{\Gamma,t+1}\right) + \tilde{q}_{t+1}$$
$$\left[(1-\delta)\exp\left(-g_{\mu,t+1}\right) + \phi\left(\frac{\tilde{i}_{t+1}}{\tilde{k}_{t+1}}\right)\exp\left(-g_{\mu,t+1}\right) - \phi'\left(\frac{\tilde{i}_{t+1}}{\tilde{k}_{t+1}}\right)\frac{\tilde{i}_{t+1}}{\tilde{k}_{t+1}}\exp\left(g_{\Gamma,t+1}\right)\right]\Bigg\}. \tag{E.39}$$

F A two-period model

For illustrative purposes, we consider a two-period economy with a representative firm and a continuum of households of unit size. We restrict the analysis to a perfect-foresight scenario, without loss of generality.

Households are either non-assetholders (indexed by *na*), whose share in the total population of households equals γ, or assetholders (indexed by *a*), whose population share amounts to $1-\gamma$. The two types of households have the same utility function, $U^i = \log(c_1^i) + \beta \log(c_2^i)$, where $\beta \in (0,1)$ is a common discount factor and c_t^i denotes household-specific consumption of a generic perishable good, for $i = a, na$ and $t = 1, 2$.

Non-assetholders cannot access financial markets. Being unable to smooth consumption intertemporally, the representative non-assetholder consumes her labor income hand-to-mouth, $c_t^{na} = w_t$, where w_t is the wage rate and where we have implicitly assumed that non-assetholders supply their entire time-endowment, which is normalized to one, in both periods of life. Assetholders, instead, purchase stocks of the representative firm in period one, and inelastically supply their labor to the firm in both periods. The resulting period budget constraint is $c_t^a = w_t + d_t s_1$, where d_t denotes firm profits and s_1 is the number of shares purchased in period one, to be held over the entire lifetime. We assume stocks to be in unit net supply, so that $(1-\gamma)s_1 = 1$.

Production of the non-durable consumption good is carried out by the representative firm through the constant return-to-scale technology $y_t = z_t k_{t-1}^{\alpha_t} n_t^{1-\alpha_t}$, (with $\alpha_t \in (0,1)$), where k_{t-1} denotes the existing capital stock, n_t is the total labor input, z_t is TFP, and α_t is the income share of capital. Firm profits in periods 1 and 2 read as $d_1 = y_1 - w_1 - i_1/\mu_1$ and $d_2 = y_2 - w_2$, respectively. Following Greenwood et al. (1997), we interpret μ_1 as capturing investment-specific technological change. Finally, to retain analytical tractability, we assume full capital depreciation, so that the effective investment taking place in period one equals the capital stock.

F.1 Solution

The firm maximizes the discounted flow of profits choosing labor input and the capital input:

$$\max_{k_1, n_1, n_2} \left[d_1 + \beta \frac{\lambda_2}{\lambda_1} d_2 \right], \tag{F.1}$$

where λ_t denotes assetholders' period marginal utility, for $t = 1, 2$.

The first-order conditions from firm optimization read as

$$r_2^k = \frac{1}{\beta} \frac{\lambda_1}{\mu_1 \lambda_2}, \tag{F.2}$$

$$w_1 = (1-\alpha_1) z k_0^{\alpha_1}, \tag{F.3}$$

$$w_2 = (1-\alpha_2) z k_1^{\alpha_2}, \tag{F.4}$$

where $r_2^k \equiv \frac{dy_2}{dk_1}$.

By plugging the assetholder's period budget constraints into (F.2), we obtain (F.5); thus, we plug the latter into the expressions for the wage rate in each of the two periods, (F.3) and (F.4), to obtain (F.6), (F.8), (F.7), and (F.9):

$$k_1 = \frac{\beta\alpha_2[1-\gamma(1-\alpha_1)]}{1-\gamma(1-\alpha_2)+\beta\alpha_2}\mu_1 z k_0^{\alpha_1}, \tag{F.5}$$

$$c_1^{na} = (1-\alpha_1) z k_0^{\alpha_1}, \tag{F.6}$$

$$c_2^{na} = (1-\alpha_2) z k_1^{\alpha_2}, \tag{F.7}$$

$$c_1^{a} = \frac{1-\gamma(1-\alpha_2)}{1-\gamma}\frac{1-\gamma(1-\alpha_1)}{1-\gamma(1-\alpha_2)+\beta\alpha_2} z k_0^{\alpha_1}, \tag{F.8}$$

$$c_2^{a} = \frac{1-\gamma(1-\alpha_2)}{1-\gamma} z k_1^{\alpha_2}. \tag{F.9}$$

Thus, after aggregating over the two household types:

$$c_1 + i_1 = y_1, \tag{F.10}$$

$$c_2 = y_2. \tag{F.11}$$

F.2 Heterogeneity in the transmission of aggregate shocks

We are now ready to examine the response of relative consumption, conditional on each of the three shocks. Unlike the quantitative model, the two-period framework mainly serves as a device to frame the propagation of the different shocks in a setting featuring household heterogeneity in the access to a saving technology. Nevertheless, to maintain adherence to our identification strategy, we consider TFP as featuring a permanent shift over the two periods of life, while FS shocks will be temporary, and take place in the first period only. As for investment-specific technological change, this can only occur in period one, by construction.

Relative consumption assumes a central role in examining the responsiveness of different household types' consumption in the face of the shocks we study in Section 3.3. We consider rc as the ratio between the total discounted consumption of the representative assetholder relative to that of the representative non-assetholder (i.e., $c^i = c_1^i + \beta c_2^i$, for $i = a, na$). After substituting each agent's consumption by the period budget constraints, it can be shown that the relative consumption reads:

$$rc = 1 + \frac{1}{1-\gamma}\frac{d_1+\beta d_2}{w_1+\beta w_2}, \tag{F.12}$$

implying that asymmetries in household-specific per-capita consumption depend on the income distribution between labor and capital.

The next step consists of computing the derivative of rc with respect to each shock.[43] We start with the IST shock, reporting the following proposition:

Proposition 1. *An expansionary IST shock determines an expansion in relative consumption, i.e.*

$$\frac{drc}{d\mu_1} > 0. \tag{F.13}$$

Proof. *See Appendix F.3.*

To see why this is the case, we can express equation (F.12) in terms of primitives:

$$rc = \frac{1}{(1-\gamma)(1-\alpha)} \frac{z\left(k_0^\alpha + \beta k_1^\alpha\right) - i_1/\mu_1}{z\left(k_0^\alpha + \beta k_1^\alpha\right)} - \frac{\gamma}{1-\gamma}. \tag{F.14}$$

Ceteris paribus, ameliorating the efficiency at which the final good can be transformed into physical capital limits the negative effect induced by capital investment on assetholders' income—and, thus, consumption—while expanding period-1 equilibrium capital stock. Thus, relative consumption ultimately expands.

This line of reasoning helps us understand the behavior of relative consumption with respect to TFP shocks, as detailed by the next proposition:

Proposition 2. *An expansionary TFP shock induces a contraction in relative consumption:*

$$\frac{drc}{dz} < 0, \tag{F.15}$$

where $z_t = z$, for $t = 1, 2$. ***Proof.*** *See Appendix F.3.*

As shown in Figures 3 and 4, an expansionary TFP shock redistributes resources from assetholders to non-assetholders, thus contracting relative consumption. This relationship indicates that capital investment acts as a drag on the increase in assetholders' income, restricting it below that of non-assetholders. In turn, such tendency maps into agents' consumption choices.

Finally, the next proposition delves into the relative consumption effects of a FS shock that redistributes resources from labor to the capital, thus favoring dividend income over labor income, and ultimately expanding relative consumption.

Proposition 3. *An expansionary FS shock determines an expansion in relative consumption, i.e.*

$$\frac{drc}{d\alpha_1} > 0 \tag{F.16}$$

[43] In each exercise of comparative statics, we set the shocks that are not being investigated to their steady-state values.

as long as the following sufficient condition is met: $1 \leq k_0/y_1 < \exp(1/\beta\alpha)$. ***Proof.*** *See See Appendix F.3.*

Notably, at standard calibrations of the capital-to-output ratio, the lower-bound inequality is always satisfied. To provide some intuition on the economic principle behind the upper bound to the capital-to-output ratio, it is useful to recall that the latter is inversely related to the marginal product of capital (MPk). When the capital-to-output ratio is high enough and exceeds the upper bound in the sufficient condition, a positive FS shock has limited capacity to stimulate MPk—at standard calibrations of the capital share—while expanding the marginal product of labor ($MPn = (1-\alpha_1)k_0^{\alpha_1}$), with the ultimate effect of favoring labor income over capital income. Therefore, the FS shock generates a decline in relative consumption. For standard calibrations of the capital-to-output ratio that are typically within the range $[1, \exp(1/\beta\alpha))$, instead, the expansionary effect on wages is too weak to overcome that on dividends, so that relative consumption expands.

F.3 Proofs

Proof of Proposition 1. We take (F.12) and set $\alpha_t = \alpha$ and $z_t = 1$, for $t = 1, 2$. Thus, we compute:

$$\frac{drc}{d\mu_1} = \beta \frac{1-\alpha}{1-\gamma} r_2^k \frac{dk_1}{d\mu_1} \frac{k_1}{\mu_1},$$

and check when this is positive, which is always the case.

Proof of Proposition 2. We take (F.12) and set $\mu_1 = 1$ and $\alpha_t = \alpha$, for $t = 1, 2$. Taking $\frac{drc}{dz}$ and imposing it to be negative amounts to prove the following inequality:

$$rc < \frac{1-\gamma(1-\alpha)}{(1-\gamma)(1-\alpha)},$$

which can be reduced to $\beta\alpha > 0$, the latter always holding true, under the restrictions imposed to β and α.

Proof of Proposition 3. We take (F.12) and set $\mu_1 = 1$, as well as $\alpha_2 = \alpha$ and $z_t = 1$, for $t = 1, 2$. Thus, we compute:

$$\begin{aligned} \left.\frac{drc}{d\alpha_1}\right|_{\alpha_1=\alpha} = & \left[y_1 \log(k_0) - \frac{dk_1}{d\alpha_1} + \beta r_2^k \frac{dk_1}{d\alpha_1}\right] [(1-\alpha)y_1 + \beta(1-\alpha)y_2] \\ & - \left[(1-\alpha)y_1 \log(k_0) - y_1 + \beta(1-\alpha) r_2^k \frac{dk_1}{d\alpha_1}\right] [y_1 - k_1 + \beta y_2], \end{aligned}$$

which can be simplified into

$$\frac{drc}{d\alpha_1}|_{\alpha_1=\alpha} = \Psi y_1 + \beta y_2 + (1-\alpha)\log(k_0)$$

where

$$\Psi \equiv \frac{1-\gamma(1-\alpha)}{[1-\gamma(1-\alpha)+\beta\alpha]} - (1-\alpha)\log(k_0)\frac{[1-\gamma(1-\alpha)]\,\beta\alpha}{[1-\gamma(1-\alpha)+\beta\alpha]}.$$

We note that y_1 and y_2 are always positive, and that $log(k_0)$ is non-negative as long $k_0 \geq 1$. Therefore, $\left.\frac{drc}{d\alpha_1}\right|_{\alpha_1=\alpha} > 0$ as long as $\Psi > 0$. This is verified for

$$k_0 < \exp\left(\frac{1}{1-\alpha}\frac{1}{\beta\alpha}\right),$$

Therefore, we conclude that $\left.\frac{drc}{d\alpha_1}\right|_{\alpha_1=\alpha} > 0$ if the following sufficient condition holds:

$$1 \leq k_0 \leq \exp\left(\frac{1}{1-\alpha}\frac{1}{\beta\alpha}\right).$$

Equivalently, for $\kappa_0 \equiv \frac{k_0}{y_1} = k_0^{1-\alpha}$:

$$1 \leq \kappa_0 \leq \exp\left(\frac{1}{\beta\alpha}\right)$$

needs to be met.

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