Human Capital Risk, Stockholder Consumption, and Asset Returns

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ABSTRACT

The success of long-run consumption risk in pricing assets is tied to labor income risk. Labor income shocks are disproportionately borne by stockholders and stockholder ultimate consumption risk captures the cross-sectional variation in asset returns, including the size and value premia. Under Epstein-Zin preferences, risk aversion as low as 6.5 is sufficient to match both the cross-sectional price of risk and the equity premium for the wealthiest stockholders. In addition, the stockholder share of aggregate consumption captures time-variation in stock and bond market returns and mirrors the dynamics of the aggregate consumption-to-wealth ratio. The results highlight the importance of stockholder consumption risk and its ties to human capital risk for asset pricing.

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The central insight of asset pricing theory is that differences in expected returns across assets and time can be explained by differences in risk. A cornerstone of asset pricing theory focuses on consumption risk (i.e., the consumption CAPM) and the role of human capital in making consumption and portfolio decisions. Recent evidence suggests that long-run or ultimate aggregate consumption risk, the covariance of returns with the present value of long-run changes in aggregate consumption growth, captures cross-sectional and aggregate stock returns (Parker and Julliard (2004), Bansal, Dittmar, and Lundblad (2004), Parker (2001), Bansal and Yaron (2004), Hansen, Heaton, and Li (2005)).

The empirical success of long-run consumption risk raises several questions. Why does consumption take so long to adjust to news in stock returns? What are the underlying shocks driving stock returns and consumption? Do households that actually own and trade stocks care about these risks? If so, does stockholder consumption risk price assets even better?

In this paper, we tie ultimate consumption risk to labor income risk and show that the consumption risk of households who hold financial assets is particularly relevant for asset pricing. We find that ultimate aggregate labor income risk, or the covariance of returns with the present value of long-run changes in labor income growth, is equally able to explain cross-sectional variation in average stock returns as ultimate aggregate consumption risk. This finding suggests that the reason aggregate consumption growth responds slowly to news in asset returns is that aggregate labor income (and total income) growth responds slowly. We then argue that analyzing who bears aggregate labor income risk is important for understanding why it is priced. Using micro-level household data from the Consumer Expenditure Survey (CEX) for the period 1982 to 2001, we show that long-run consumption growth rates of the set of agents who bear stock market risk is highly sensitive to the long-run growth rates in aggregate labor income, and exhibits significant non-linearities. As a result, the ultimate consumption risk of stockholders is even better at capturing the cross-sectional variation of average stock returns. There are statistically and economically significant differences across assets in their covariation with long-run stockholder consumption growth. In particular, small and value stocks earn low returns when the future consumption growth of stockholders is low, requiring a premium for stockholders to bear this risk.

Following Bansal and Yaron (2004) and Hansen, Heaton, and Li (2005) we interpret our findings within a framework that adopts Epstein-Zin (1989) and Weil (1989) preferences, allowing for the separation of the elasticity of intertemporal substitution (EIS) from risk aversion. In this setting, households care about covariances of asset returns with the present value of current and *future* consumption growth.[1] We show that the value of risk aversion implied by the cross-sectional reward for ultimate consumption risk of stockholders is around 10 to 15, and as low as 6.5 for the wealthiest third of stockholders with the largest holdings of equity. These implied risk aversion estimates are significantly smaller than those obtained from aggregate or non-stockholder consumption growth.

Since CEX data is limited to the period 1982 to 2001, we project consumption growth on a set of asset returns and use the coefficients from this regression to construct a factor-mimicking portfolio for ultimate stockholder consumption growth dating back to 1926 and extending through 2004. This factor portfolio can explain 80 to 85 percent of the variation in average returns across the 25 Fama and French (1996) size and book-to-market equity portfolios, performing about as well as the three factor model of Fama and French (1993), even over the out of sample period prior to CEX data availability from which the factor portfolio was constructed.

To further explore the scope of our findings, we examine the entire cross-section of individual stocks. Stockholder consumption risk, whether measured by actual stockholder consumption from the CEX (1982 to 2001) or by the consumption growth factor portfolio (1926 to 2004), captures cross-sectional variation in the average returns of individual stocks, even after controlling for market, size, book-to-market equity, and momentum premia.

We then consider the implications of our framework for the dynamics of aggregate stock and bond market returns. Theoretical work by Basak and Cuoco (1998) and Guvenen (2004) shows that the fraction of total consumption consumed by stockholders is a state variable in general equilibrium models of asset pricing with limited stock market participation. The equity and term premium should be low when the stockholder consumption share is high.[2] Consistent with this prediction, we find that the ratio of stockholder consumption to total consumption is strongly negatively correlated with both stock and bond market yields and has predictive power for the excess return on stocks and long-term government bonds. We also find that stockholder consumption shares vary over time in a manner that mirrors the dynamics of the consumption-to-wealth ratio. The correlation between the consumption share of stockholders and Lettau and Ludvigson (2001a,b)'s aggregate consumption-to-wealth ratio cay is -0.5. This finding is consistent with micro-level evidence that

[1]Under constant relative risk aversion (CRRA) preferences, investors care only about the covariance with contemporaneous consumption growth. Hence, Parker and Julliard (2004) offer explanations for the importance of long-run consumption growth that range from measurement error in consumption to costs or constraints of adjusting consumption.

[2]This prediction holds even if stockholders and non-stockholders have identical preferences, though the effects are naturally stronger if stockholders are less risk averse than non-stockholders.

savings rates increase with wealth (Dynan, Skinner, and Zeldes (2004), Carroll (2000)) and offers a new interpretation for the empirical success of *cay*.

Finally, we analyze whether human capital risk and stock market participation play a role in understanding the equity premium. The value of risk aversion implied by the equity premium for ultimate stockholder consumption risk matches the magnitudes we find in the cross-section: around 10 for all stockholders and 7 for the wealthiest third of stockholders.

Our results highlight the importance of consumption risk, its relation to human capital risk, and who bears this risk for asset pricing. Ultimate consumption risk is tied to human capital risk and is disproportionately borne by stockholders. The fact that it is the consumption and labor sensitivity of those households that actually own financial assets and participate in capital markets that drives these relations is comforting and implies fairly reasonable risk aversion estimates. Stock market participation provides a mechanism amplifying risk sharing in the economy that appears important for asset prices.

The rest of the paper is organized as follows. Section I cites some of the vast literature on consumption and human capital risk in asset pricing. Section II outlines a theoretical framework under Epstein-Zin preferences linking asset prices to ultimate labor income and consumption risk. Section III summarizes the data sources and construction of our variables. Section IV examines the relation between ultimate labor income risk, stockholder consumption, and the cross-section of returns. Section V examines aggregate stock and bond market return predictability and the equity premium, linking the stockholder consumption share to the consumption-to-wealth ratio. Section VI concludes.

I. Related Literature

Our work relates to the vast literature on consumption growth and asset pricing which finds a modest relation between returns and contemporaneous consumption growth (Mankiw and Shapiro (1996), Breeden, Gibbons, and Litzenberger (1989), Cochrane (1996), and Lettau and Ludvigson (2001b)) but a stronger relation for long-run consumption growth (Bansal, Dittmar, and Lundblad (2004), Parker and Julliard (2004), and Hansen, Heaton, and Li (2005) consider the cross-section of stocks, while Parker (2001) and Bansal and Yaron (2004) consider the equity premium). Other studies find some success using conditioning variables in the CCAPM such as the consumption-to-wealth ratio *cay* of Lettau and Ludvigson (2001a, 2001b) and *cay* combined with labor income growth lr of Julliard (2005). However, there is debate over whether these conditioning variables can give rise to adequate dispersion in risks across assets (Lewellen and Nagel (2004)), resulting in large risk aversion parameters needed to match observed return premia. Recent work also examines measures of consumption risk using different components of consumption, such as durables (Yogo (2005) and Pakos (2005)), housing (Piazzesi, Tuzel, and Schneider (2005), Lustig and Van Nieuwerburgh (2004)), and luxury goods (Ait-Sahalia, Parker, and Yogo (2004)) or computes covariances with aggregate consumption differently as in Jagannathan and Wang (2005), who employ fourth-quarter-to-fourth-quarter contemporaneous consumption growth. These perturbations yield slightly stronger relations with returns but also typically require high risk aversion. Although beyond the scope of this paper, it would be interesting to assess how much overlap exists between these measures and long-run, low frequency consumption risk measures.

Our findings relate to the literature on human capital and asset returns. Jagannathan and Wang (1996) and Jagannathan, Kubota, and Takehara (1996) find some ability of covariances of returns with contemporaneous labor income growth to explain the size effect. Campbell (1996) uses a vector autoregressive model to construct a measure of the innovation in the present value of labor income, but covariances with this measure are not significant in cross-sectional regressions. Santos and Veronesi (2004) show that the slow moving component of income measured by the ratio of wages to consumption has predictive power for the expected returns and conditional betas of the 25 Fama-French portfolios. Although, risk aversion is required to be large in their model. However, we find that the stockholder consumption share is *negatively* correlated with their wages-to-consumption ratio, implying that the stockholder consumption share is low when wages-to-consumption is high. Hence, it is likely these two variables are capturing different effects.[3] It might be interesting to incorporate limited participation into their general equilibrium framework and analyze the relation between wages-to-consumption, stockholder consumption shares, and asset prices.

Our work also complements the existing literature on limited stock market participation, beginning with Mankiw and Zeldes (1991) who document higher covariance between consumption growth and excess stock returns for stockholders than non-stockholders. Parker (2001), Vissing-Jørgensen (2002), Brav, Constantinides, and Geczy (2002), and Attanasio, Banks, and Tanner (2002) also em-

[3]The reason stockholder consumption shares are negatively correlated with wages-to-consumption is that even though stockholder consumption is more sensitive to labor income than non-stockholder consumption, stockholder consumption is even more sensitive to capital income relative to non-stockholder consumption, and this effect tends to dominate. Since we find that the capital component of income does not explain asset returns, we do not focus on this relation.

phasize, within a constant relative risk aversion setting, that stockholder consumption growth leads to more reasonable risk aversion estimates. Vissing-Jørgensen (2002) and Attanasio and Vissing-Jørgensen (2003) find that elasticity of intertemporal substitution estimates are much higher for stockholders than non-stockholders and Attanasio and Vissing-Jørgensen (2003) show that fairly low risk aversion coefficients are sufficient to reconcile the equity premium with stockholder consumption growth and Epstein-Zin preferences. We supplement these findings using long-run stockholder consumption growth and our factor-mimicking portfolio for consumption growth over the much longer period 1926 to 2004. Brav, Constantinides, and Geczy (2002) find that a stochastic discount factor (SDF) based on contemporaneous stockholder consumption is able to reconcile the value premium for low values of risk aversion, if they either do not log-linearize the SDF or use a third-order Taylor expansion of the SDF. They therefore emphasize the role of skewness in consumption growth rates for explaining the value premium. We focus on long-run growth rates and log-linearize the SDF. We also examine the size premium and the entire cross-section of individual stock returns and link long-run consumption growth back to labor income growth.

Finally, we test the implications of general equilibrium models of limited stock market participation (e.g., Basak and Cuoco (1998), Guvenen (2004)) that suggest the stockholder consumption share has forecasting power for aggregate returns. To our knowledge these tests are novel.

II. Asset Pricing with Epstein-Zin Preferences

Our theoretical setup follows Hansen, Heaton, and Li (2005). The innovation of our study is in the empirical implementation of this setup.

Each household has Epstein-Zin preferences of the form

$$V_t = \left[(1-\beta) C_t^{1-\frac{1}{\sigma}} + \beta \left[E_t \left(V_{t+1}^{1-\gamma} \right) \right]^{\frac{1-\frac{1}{\sigma}}{1-\gamma}} \right]^{\frac{1}{1-\frac{1}{\sigma}}} \tag{1}$$

where C_t is consumption, σ is the elasticity of intertemporal substitution, γ is relative risk aversion, and β is the discount factor. We focus on the special case where σ equals one,[4] where the recursion

[4]Attanasio and Vissing-Jørgensen (2003) estimate conditional Euler equations for stockholders in the Consumer Expenditure Survey (CEX) using after-tax returns and find a value for the elasticity of intertemporal substitution around 1.4 when using the after-tax T-bill return as the asset return, and around 0.4 when using the after-tax stock return. These are the same values documented in Vissing-Jørgensen (2002) with the exception that they adjust for the effect of taxes. Vissing-Jørgensen (2002, p. 851) notes that adjusting for taxes increases the elasticity of intertemporal substitution estimates. The corresponding estimates for the top third of stockholders are 2.3 and 0.7, respectively. This evidence suggests that an elasticity of intertemporal substitution of one is not unreasonable.

becomes,

$$V_t = C_t^{1-\beta}\left[E_t\left(V_{t+1}^{1-\gamma}\right)\right]^{\frac{\beta}{1-\gamma}}.$$

Following Hansen, Heaton, and Li (2005), log consumption growth follows a moving-average process[5]

$$\begin{aligned} c_t - c_{t-1} &= \mu_c + \alpha(L)\, w_t \\ &= \mu_c + \left(\Sigma_{s=0}^{\infty}\alpha_s L^s\right) w_t = \mu_c + \Sigma_{s=0}^{\infty}\alpha_s w_{t-s} \end{aligned}$$

where $\{w_t\}$ is an *iid* standard normal process. The log of the household's stochastic discount factor is then given by

$$\begin{aligned} s_{t+1,t} &= \ln\beta - \left[\mu_c + \alpha(L)\, w_{t+1}\right] + (1-\gamma)\,\alpha(\beta)\, w_{t+1} - \frac{1}{2}(1-\gamma)^2\,\alpha(\beta)^2 \\ &= \ln\beta - \left[c_{t+1} - c_t\right] + (1-\gamma)\left(\Sigma_{s=0}^{\infty}\alpha_s\beta^s\right) w_{t+1} - \frac{1}{2}(1-\gamma)^2\left(\Sigma_{s=0}^{\infty}\alpha_s\beta^s\right)^2. \end{aligned}$$

(See Hansen, Heaton, and Li (2005) for derivations and proofs.) Written in terms of consumption growth rates, the term $\left(\Sigma_{s=0}^{\infty}\alpha_s\beta^s\right) w_{t+1}$ equals $(E_{t+1} - E_t)\,\Sigma_{s=0}^{\infty}\beta^s\left(c_{t+1+s} - c_{t+s}\right)$, representing the innovation in the present value of consumption growth rates. The term $\left(\Sigma_{s=0}^{\infty}\alpha_s\beta^s\right)^2$ is the variance of this innovation. The term $(E_{t+1} - E_t)\,\Sigma_{s=0}^{\infty}\beta^s\left(c_{t+1+s} - c_{t+s}\right)$ can be rewritten as $(1-\beta)(E_{t+1} - E_t)\left[\Sigma_{s=0}^{\infty}\beta^s c_{t+s}\right]$, or $(1-\beta)$ times the change in the expected present value of future consumption.

Consider the pricing of financial asset i held (and traded) by this household, whose return follows a moving-average representation of the form,

$$\begin{aligned} r_t^i &= \mu_r^i + k^i(L)\, w_t \\ &= \mu_r^i + \Sigma_{s=0}^{\infty} k_s^i w_{t-s} \end{aligned}$$

where $\{w_t\}$ is the same process as above and r_t^i is the log return of asset i. Using the stochastic discount factor $s_{t+1,t}$, Hansen, Heaton, and Li (2005) show that the expected excess log return on asset i, over and above the log return on the (conditionally) riskless asset, is given by

$$E_t\left[r_{t+1}^i\right] - r_{t+1}^f = -\frac{1}{2}\left(k_0^i\right)^2 + \left[\alpha_0 + (\gamma-1)\left(\Sigma_{s=0}^{\infty}\alpha_s\beta^s\right)\right] k_0^i.$$

The last term $\left(\Sigma_{s=0}^{\infty}\alpha_s\beta^s\right) k_0^i$ is the conditional covariance of the log return and the present value of consumption growth rates, $cov_t\left(r_{t+1}^i, (E_{t+1} - E_t)\,\Sigma_{s=0}^{\infty}\beta^s\left(c_{t+1+s} - c_{t+s}\right)\right)$. For large risk aversion,

[5] Hansen, Heaton, and Li (2005) emphasize the difficulty in measuring long-run relations between consumption growth and asset returns. They suggest one reasonable approach is to use a highly structured but interpretable model of long-run growth variation. Hence, we adopt their moving-average process for consumption growth.

this term will be much larger than the term $\alpha_0 k_0^i$, the conditional covariance of the log asset return and the log consumption growth rate, $cov_t\left(c_{t+1}-c_t, r_{t+1}^i\right)$. We therefore follow Hansen, Heaton, and Li (2005) and ignore the term $\alpha_0 k_0^i$. The conditional variance of the log asset return, $V_t\left[r_{t+1}^i\right]$, is $\left(k_0^i\right)^2$. The main asset pricing relation used in our empirical analysis is thus,

$$E_t\left[r_{t+1}^i\right]-r_{t+1}^f+\frac{1}{2}V_t\left[r_{t+1}^i\right] \simeq (\gamma-1)\, cov_t\left(r_{t+1}^i,\left(E_{t+1}-E_t\right)\Sigma_{s=0}^{\infty}\beta^s\left(c_{t+1+s}-c_{t+s}\right)\right)$$
$$= (\gamma-1)\, cov_t\left(r_{t+1}^i,\Sigma_{s=0}^{\infty}\beta^s\left(c_{t+1+s}-c_{t+s}\right)\right). \qquad (2)$$

If $\beta=1$ (no discounting), and if we end the sum at $s=11$ (in quarterly data) rather than ∞, then $cov_t\left(r_{t+1}^i,\left(E_{t+1}-E_t\right)\Sigma_{s=0}^{\infty}\beta^s\left(c_{t+1+s}-c_{t+s}\right)\right)$ reduces to $cov_t\left(r_{t+1}^i,\Delta c_{t+1}+...+\Delta c_{t+12}\right)$, the measure of risk considered by Parker and Julliard (2004). Hence, Epstein-Zin utility provides one possible explanation for the success of the risk measure used by Parker and Julliard (2004) in pricing the cross-section of expected returns. We follow Hansen, Heaton, and Li (2005) and assume a discount factor of 5% per annum, implying $\beta=0.95^{1/4}$ in quarterly data.

Restoy and Weil (1998) and Bansal and Yaron (2004) employ models with Epstein-Zin preferences that allow the elasticity of intertemporal substitution to be greater than one. In this setting, consumption is not the only variable that enters the Euler equation. Allowing both the elasticity of intertemporal substitution and risk aversion to be greater than one, Bansal and Yaron (2004) justify simultaneously the high equity premium, low risk-free rate, and volatilities of the market and real riskless rate. We focus on the special case in which the elasticity of intertemporal substitution is one and emphasize the importance of using the consumption of households who actually hold the asset in analyzing asset pricing relations. Empirically, we find more reasonable risk aversion estimates can match the data when employing the long-run consumption growth of assetholders, without appealing to an elasticity of intertemporal substitution greater than one. Allowing the elasticity of intertemporal substitution to be greater than one and employing the consumption growth of stockholders may provide even lower risk aversion estimates. In addition to aggregate market returns, which is the focus of Bansal and Yaron (2004), we also examine the cross-section of expected returns, and trace differences in consumption patterns of assetholders (and non-assetholders) back to their sensitivity to aggregate labor income shocks.

III. Data

We briefly describe the data sources and variables used in the analysis.

A. Labor income and aggregate consumption

Quarterly data from 1947Q1 to 2004Q4 on labor income, capital income, and total national income are obtained from Table 1.12 of the National Income and Product Accounts (NIPA). Labor income is the sum of compensation of employees [line 2] and proprietors' income [line 9]. Capital income is the sum of corporate profits [line 13], rental income [line 12], and net interest [line 18]. National income is the sum of all income [line 1]. All values are seasonally adjusted by NIPA. We calculate real per capita income growth rates by subtracting the inflation rate calculated from the consumer price index for all urban consumers (from the Bureau of Labor Statistics), and subtracting the population growth rate using population data from NIPA Table 2.1 [line 38].

As part of our analysis we also use seasonally-adjusted monthly labor income data from NIPA Table 2.6 [line 2 plus line 7] and aggregate consumption of non-durables from NIPA Table 2.8.3 [line 3]. Both are available beginning January, 1959. Real per capita growth rates are calculated by subtracting the CPI inflation rate and population growth rates, using monthly population from NIPA Table 2.6 [line 29].

B. Household level consumption: Stockholders and non-stockholders

We calculate consumption growth rates for stockholders and non-stockholders separately using data from the Consumer Expenditure Survey (CEX), which covers the period 1980Q1 to 2002Q1. Before 1999 about 4,500 households are interviewed per quarter. The sample size increases to about 7,500 households per quarter after 1999. Each household is interviewed five times. The first time is practice and the results are not in the data files. The interviews are three months apart and households are asked to report consumption for the previous three months. Financial information is gathered in the fifth quarter only. Aside from attrition, with only about 60 percent of households making it through all five quarters, the sample is representative of the U.S. population.

The definition of stockholder status, the consumption measure, and the sample selection follow Vissing-Jørgensen (2002), with the exception that 12 quarter consumption growth rates, as opposed to semiannual growth rates in Vissing-Jørgensen (2002), are used and data through 2002 are employed. We also compute consumption separately for the wealthiest third of stockholders based on their beginning quarter dollar amount of holdings. We consider both a narrow definition of stockholders based on positive holdings of "stocks, bonds, mutual funds and other such securities" and a broader definition which also includes households with positive holdings of savings bonds, which we refer to as assetholders. The motivation for this broader category is that stockholding is known to be strongly increasing in financial wealth (see Vissing-Jørgensen (2003)). Therefore, large holdings of savings bonds may be a good indicator of holding stocks indirectly through pension plans which may not be included in responses. Appendix A details the construction of these series, the number of households in the survey classified as stockholders and non-stockholders, as well as compares the level and trend of stock market participation rates calculated from the CEX to those from other well-known sources such as the Survey of Consumer Finances. There are 145,815 quarterly consumption growth observations across 48,605 households. Stockholders (assetholders) comprise 22.8 (31.4) percent of these households. The average number of consumption observations per month is 141 (195) for stockholders (assetholders) and 479 (426) for non-stockholders (non-assetholders). The top third of stockholders (assetholders) hold a minimum of \$25,974 (\$13,614) in stocks (stocks and savings bonds) in 1982 dollars.[6]

B.1. Average group consumption growth rates

The panel dimension for each household in the CEX allows us to calculate consumption growth rates at the household level. To construct a time series of average consumption growth for a particular group of households (e.g., for stockholders), we average the (log) consumption growth rates for households in a particular group as follows:

$$\frac{1}{H_t^g}\sum_{h=1}^{H_t^g}\left(c_{t+3}^{h,g}-c_t^{h,g}\right)$$

where $c_t^{h,g}$ is the quarterly log consumption of household h in group g for quarter t, $c_{t+3}^{h,g}$ is the log consumption of the same household for the following quarter, and H_t^g is the number of households in group g in quarter t. This cross-sectional summation *does not* assume a representative agent within each group, an assumption that would be violated in an incomplete market setting with uninsurable idiosyncratic labor income shocks. Rather, the Epstein-Zin asset pricing relation for holders of asset i is,

$$E_t\left[r_{t+1}^i\right]-r_{t+1}^f+\frac{1}{2}V_t\left[r_{t+1}^i\right]\simeq(\gamma-1)\,cov_t\left(r_{t+1}^i,\Sigma_{s=0}^{\infty}\beta^s\left[\frac{1}{H_t^g}\sum_{h=1}^{H_t^g}\left(c_{t+1+s}^{h,g}-c_{t+s}^{h,g}\right)\right]\right). \quad (3)$$

This equation is simply the Epstein-Zin asset pricing relation based on the consumption of a particular household, summed cross-sectionally across households in the group. The log linearization

[6]The CEX tends to underweight the super wealthy (see Bosworth, Burtless, and Sabelhaus (1991)). Thus, our results likely understate the importance of stockholders and the top third of stockholders.

allows us to sum the individual Euler equations across households.

C. Consumption-to-wealth ratio

We use Lettau and Ludvigson's (2001a,b) consumption-to-wealth ratio cay obtained from Sydney Ludvigson's webpage. This variable is at the quarterly frequency from 1951Q4 to 2003Q2. We also augment the consumption-to-wealth ratio following Julliard (2005) by including expected future labor income growth rates, lr, estimated from an ARIMA(0,1,2) model for quarterly log per capita labor income.[7]

D. Asset returns

The full cross-section of all NYSE, AMEX, and NASDAQ stock returns with beginning of month share prices above $5 is obtained from the Center for Research in Security Prices (CRSP) from July, 1926 to December, 2004. We also obtain the market capitalization and book-to-market equity ratio of every stock over this same time period from CRSP, Compustat, and Kenneth French's website, which contains book values of NYSE firms prior to June, 1962 from Moody's. We obtain the monthly returns on the 25 size and book-to-market equity sorted portfolios of Fama and French (1996) as well as the Fama and French (1993) factors $RMRF$ (excess return on the CRSP value-weighted index), SMB (small minus big portfolio), and HML (high minus low book-to-market equity portfolio) from Kenneth French's web page from July, 1926 to December, 2004.

Monthly returns on 10-year U.S. Government bonds from January, 1947 to December, 2004 and on 30-day T-bills from July, 1926 to December, 2004 are from CRSP. The 30-day T-bill rate is employed as the riskless rate of interest. Monthly yields on 10-year U.S. government bonds are from the National Bureau of Economic Research from January, 1951 to March, 1953 and from Global Insight from April, 1954 to December, 2003. The earnings-to-price ratio on the S&P500 is calculated based on data from Robert Shiller's web page.

IV. Labor Income Risk, Stockholder Consumption, and the Cross-Section of Expected Returns

We begin by examining growth rates in labor income over various horizons and find that long-run growth rates have substantial predictive power for the cross-section of returns.

[7] Julliard (2005) experiments with various specifications for the labor income process in the ARIMA class and performs the standard set of Box-Jenkins selection procedures.

A. Ultimate labor income risk and the cross-section

Table I reports cross-sectional regression results of the average log excess returns of the 25 Fama-French portfolios against the covariance of returns with future labor income growth from quarter t to $t+S$. Specifically, we estimate the following cross-sectional regression,

$$E[r_{i,t+1}] - r_f + \sigma_i^2/2 = c + \lambda Cov(r_{i,t+1}, \sum_{s=0}^{S-1} \beta^s (y_{t+s+1} - y_{t+s})) + e_i \qquad (4)$$

where $E[r_{i,t}] - r_f$ is the sample average of the log excess return on the 25 Fama-French portfolios $(i = 1, ..., 25)$, σ_i is the standard deviation of the log excess returns to portfolio i, y_t is the log of real per capita labor income, and $\beta = 0.95^{1/4}$. Although the length of the sample period depends on S, since mean returns are notoriously difficult to estimate we employ the entire time-series of returns from July, 1926 to December, 2004 to estimate the mean and standard deviation of returns on the 25 Fama-French portfolios, irrespective of S.

As Panel A of Table I shows, discounted labor income growth rates over 1, 2, 4 or 8 quarters have little ability to explain the cross-sectional variation in average returns across the 25 Fama-French portfolios. However, 12, 16, 20 and 24 quarter growth rates in labor income have significant power to explain cross-sectional average returns. The cross-sectional R^2 from long-run labor income growth is a little more than 40 percent and the t-statistics on labor growth covariance range from around 2.7 to 4.3. Accounting for cross-correlation in the residuals and for first-stage estimation error (equivalent to Fama and MacBeth (1973) standard errors with the Shanken (1982) correction), the t-statistics drop to between 1.99 and 3.20, but are still significant at the 5% level. For reference, Figure 1 plots the results for the 12 and 24 quarter ahead growth rates, with the average returns on the 25 Fama-French portfolios on the vertical axis and the covariance of returns with future discounted labor income growth on the horizontal axis. As the figure shows, the average returns of the portfolios line up nicely with their labor income growth covariances with the exception of the most extreme small growth portfolio. The cross-sectional fit improves slightly when we force the model to also price the riskless asset, by adding the risk-free asset as a test asset. Since the magnitude of the coefficients does not change much, we do not report these results for brevity (available upon request).

Panels B and C show results based on national income and capital income, respectively. The national income results are fairly similar to those based on labor income, but results based on capital income are generally much weaker. These results are consistent with the poor performance of the unconditional CAPM. This evidence implies the importance of human capital and labor income relative to financial capital and financial asset income. Size and value seem to be correlated with average returns because they forecast long-run labor income risk. Low returns on small and value stocks tend to be associated with negative shocks to long-run labor income growth rates.

The results in Panels A, B, and C are the second-stage regression estimates of average returns on covariances with income growth, which themselves are estimated in a first stage. A typical concern is that the first-stage estimates may not differ substantially across the test assets and may be imprecise. Although we adjust for first-stage estimation error in computing second-stage test statistics, it is useful to examine the dispersion in covariances across the test assets and their reliability. Panel E reports the first-stage covariance estimates and t-statistics for each of the 25 Fama-French portfolios with ultimate labor income growth from quarter t to $t+12$. Standard errors employ a Newey-West adjustment of 12 quarter lags. As the panel indicates, many of the covariance estimates are at least two standard errors from zero. Moreover, there appears to be wide dispersion in covariances across the portfolios. An F-test on the equality of the first-stage covariances across the 25 portfolios is clearly rejected at less than the 1% significance level. The same significant spread in covariances across the portfolios holds for $S = 16$, 20, and 24. However, for shorter horizons, the spread in covariances is much smaller and often insignificant, consistent with the lack of significance in the second stage at shorter horizons.

Since the choice of horizon is arbitrary and does not seem to make much difference after 12 quarters, we subsequently define long-run growth rates at $S = 12$ to capture long-run risk yet provide enough time-series to estimate reasonably precise covariances.

Our results on long-run labor income risk are similar to those for aggregate consumption growth reported by Parker and Julliard (2004). Panel D reports results for aggregate nondurable consumption for reference. At short-horizons, consumption growth has little explanatory power for cross-sectional average returns, but approximately three year consumption growth captures a large fraction of the returns associated with size and value. Our evidence on long-run labor income growth complements these findings, suggesting there is something important about long-horizon risks. The fact that long-run labor income growth also captures the same effects suggests that the reason aggregate consumption growth responds slowly to news in asset returns is that labor income (and total income) growth responds slowly.[8] Moreover, the poor explanatory power of the capital

[8]Although movements in investments, government spending, or net exports could cause aggregate consumption and income to diverge, our findings suggest they are substantially linked.

component of income further highlights the importance of the labor component.

B. Stockholder sensitivity to labor income

In considering the importance of labor income for asset prices, we ask whether wealthy individuals, who own and trade most of the equity, actually care about labor risk. Do households that participate in capital markets care about aggregate labor income risk, or is this risk borne primarily by poorer households who typically do not own financial assets? Even if stockholders are sensitive to their own labor income shocks, they may be immune to aggregate shocks. We first show that stockholder consumption is more sensitive to aggregate labor income shocks than non-stockholder consumption and then show that labor income is a substantial portion of total income, even for the wealthiest stockholders.

B.1. Sensitivity of stockholder consumption to aggregate labor income

Table II shows the sensitivity of stockholder consumption to aggregate labor income growth. We focus on the results for assetholders, but results for stockholders are similar. We regress the discounted long-run consumption growth of financial assetholders from quarter t to $t + 12$ on long-run discounted aggregate labor income growth over the same period. The results are reported in the first column of Panel A of Table II. Newey-West standard errors are calculated allowing for 12 quarter lags with t-statistics in parentheses. The second column of Panel A reports results for non-assetholder consumption growth on labor income growth. As the table shows, assetholders are much more sensitive to labor income shocks than non-assetholders, by more than a factor of two. Column 3 reports results for the top third of assetholders who are more than 3 times as sensitive as non-assetholders. The differences in sensitivities across the three groups are statistically significant as indicated in the last two columns of Panel A. The top right graph in Figure 2 plots the predicted values from these regressions, indicating that the consumption of assetholders, and particularly the wealthiest assetholders, bears a larger amount of aggregate labor income risk.[9]

The relation between assetholder consumption growth and labor income growth is also nonlinear. Panel A reports results from a piecewise linear regression that allows the coefficient on labor income to vary across three even categories. An F-test rejects the null of linearity at less than the 1% significance level. The bottom right graph in Figure 2 plots the predicted value of consumption

[9]Not surprisingly, assetholder consumption is also more sensitive to capital income growth in unreported results. However, since capital income growth fails to explain much cross-sectional variation in returns (Table I), we do not pursue this line of inquiry further, and focus on the importance of labor income.

growth from the nonlinear specification for assetholders and non-assetholders against labor income growth. As the figure shows, the strongest difference in sensitivities is in the middle third of the labor income growth range.

Panel B and the left half of Figure 2 repeat the regressions using aggregate consumption shocks instead of labor income shocks. The results are similar. Assetholders and the top third of assetholders are much more sensitive to aggregate consumption shocks, and in a nonlinear fashion, than non-assetholders.

Guvenen (2004) provides a general equilibrium model with limited participation, preference heterogeneity, and precautionary savings in response to aggregate labor income shocks. In the model, limited participation amplifies the effect of aggregate risk by concentrating it among those with capital income. More importantly, limited participation also creates different consumption smoothing opportunities for non-stockholders, since they can only use the bond market to respond to aggregate shocks. In the model risk sharing is further exacerbated by preference differences in the elasticity of intertemporal substitution between stock and non-stockholders, and by the fact that the bond market is not as effective in smoothing aggregate shocks. Because consumption risk is driven by aggregate (labor) income and not idiosyncratic shocks, non-stockholder consumption smoothing via the bond market comes at the expense of higher stockholder consumption volatility which is also procyclical. In addition to bond market trading transferring risk from non-stockholders to stockholders, other possible explanations for why stockholder consumption is more sensitive to aggregate labor income are that stockholders have more risky human capital or perhaps the tax system transfers some labor income risk from non-stockholders to stockholders.[10]

B.2. Fraction of stockholder wealth tied to their own labor income

Table III shows the fraction of stockholder wealth tied to their own labor income using data from the 1989, 1992, 1995, 1998, and 2001 *Survey of Consumer Finances*. Panel A reports the ratio of labor income to total income across all stockholders, the top third of stockholders, and the top one percent of stockholders based on dollar stockholdings. Stockholdings include both direct and indirect (through pension and mutual funds) holdings. Labor income is defined as income from wages and salaries plus income from unemployment benefits and worker's compensation. Panel B reports the ratio of labor and business income to total income. Business income refers to income

[10]We are in the process of figuring out the importance of each of these mechanisms in explaining why stockholder consumption is more sensitive to aggregate shocks and how important each of these mechanisms are for asset pricing.

or losses from a professional practice, business or farm.[11] All income measures are pre-tax. All numbers in Table III are calculated using SCF weights and based on data that are averaged across SCF imputations.

As Panels A and B show, the fraction of total income accounted for by labor or labor and business income is substantial even among stockholders. The average (median) stockholder has between 79 to 82 (96 to 98) percent of his total income tied to labor income, across the five survey years. Including business income (Panel B) this fraction rises to between 84 to 92 percent for the average stockholder and 97 to 100 percent for the median stockholder. Clearly, labor income is a huge component of total income among those who own stocks.

Among the top third wealthiest stockholders (based on dollar value of stockholdings), labor income comprises between 64 to 72 percent for the mean household and between 75 and 91 percent for the median household across the survey years. This fraction increases to roughly 80 (92) percent when including business income for the mean (median) household. For the top 1 percent of stockholders, the share of labor income declines to about 40 percent for the mean household and about half of that for the median household. Including business income raises the labor share substantially to 58 (38) percent for the average (median) household in the top 1 percent stockholder category.

Panel C shows the fraction of the total stock market owned for the top third and top 1 percent of stockholders. The top 1 percent own about a third of all stock owned by SCF households, while another 60 percent of stocks are owned by the next wealthiest 32 percent of stockholders for whom labor income comprises more than half of their income. For reference, Panel D shows that stockholdings comprise a large fraction of total financial wealth for stockholders, the top third of stockholders, and the top 1 percent of stockholders.

Finally, Panel E of Table III reports a measure of risk aversion for stock and non-stockholders. The SCF poses the following question to each respondent: "Which of the statements on this page comes closest to the amount of financial risk that you are willing to take when you save or make investments?"

(1) Take substantial financial risks expecting to earn substantial returns.

(2) Take above average financial risks expecting to earn above average returns.

[11] The other income categories in total income are income from non-taxable investments (e.g., municipal bonds), other interest income, dividend income, income or losses from net gains or losses from the sale of stocks, bonds, or real estate, income from rent, trusts, or royalties, income from child support or alimony, welfare payments, income from social security and other pensions, annuities or disability or retirement programs, and other income (including settlements, gambling winnings, education scholarships or grants, inheritances and gifts).

(3) Take average financial risks expecting to earn average returns.
(4) Not willing to take any financial risks.

Panel E summarizes the responses to this question across stockholders, the top third and top 1 percent of stockholders, and non-stockholders separately. Non-stockholders appear more risk averse than stockholders, who in turn are more risk averse than the top third of stockholders, who are more risk averse than the top 1 percent of stockholders. The differences in responses are also statistically significant. An F-test for equal risk aversion of stockholders and non-stockholders is clearly rejected at less than the 1% significance level. The difference in risk aversion across stock and non-stockholders is important for general equilibrium models of asset pricing with limited participation (e.g., Basak and Cuoco (1998) and Guvenen (2004)) since returns are driven both by the distribution of wealth between the two groups and differences in preferences.

Our findings that stockholders (and the wealthiest stockholders) are more sensitive to aggregate income shocks, have higher risk tolerance than non-stockholders, and have more consumption volatility (as documented by Mankiw and Zeldes (1991), Attanasio, Banks, and Tanner (2002), and Vissing-Jørgensen (2002)), suggests that using the consumption risk of stockholders may provide a better fit for the cross-section of average returns than either labor income or aggregate consumption risk and lower risk aversion estimates.

C. Ultimate stockholder consumption risk and the cross-section of returns

Table IV examines the relation between ultimate asset and stockholder consumption risk and the cross-section of expected stock returns.

C.1. The 25 Fama and French portfolios

Following equation (3), we run cross-sectional regressions of the average log excess returns on the 25 Fama-French portfolios (measured quarterly from July, 1926 to December, 2004) against the covariance of returns with long-run discounted consumption growth of financial assetholders, the top third of assetholders, non-assetholders, stockholders, the top third of stockholders, and non-stockholders. The regressions are run separately for each of the household groups. We define long-run consumption growth from quarter t to $t+12$ with a discount factor $\beta = 0.95^{1/4}$. Specifically, we run the following cross-sectional regression,

$$E[r_{i,t+1}] - r_f + \frac{\sigma_i^2}{2} \quad = \quad c + (\gamma - 1)cov\left(r_{i,t+1}, \sum_{s=0}^{11} \beta^s \left[\frac{1}{H_t^g} \sum_{h=1}^{H_t^g} \left(c_{t+1+s}^{h,g} - c_{t+s}^{h,g}\right)\right]\right) + e_i \qquad (5)$$

where γ is the implied risk aversion coefficient from the model and $r_{i,t+1}$ is the log return on asset i between time t and $t+1$.[12]

Since consumption is measured with error, the coefficient estimate in equation (5) may be biased toward zero if measurement error in consumption is uncorrelated with the error in expected returns. To address attenuation bias, we also run the regression in reverse,

$$cov\left(r_{i,t+1}, \sum_{s=0}^{11} \beta^s \left[\frac{1}{H_t^g} \sum_{h=1}^{H_t^g} \left(c_{t+1+s}^{h,g} - c_{t+s}^{h,g}\right)\right]\right) = c + \frac{1}{\gamma - 1}\left(E[r_{i,t+1}] - r_f + \frac{\sigma_i^2}{2}\right) + u_i. \quad (6)$$

Table IV reports the results from regressions (5) and (6) for each consumption series under the heading CEX. For comparability we convert the coefficients from the regressions into implied risk aversion estimates (γ) and report those along with t-statistics, adjusted for cross-correlated residuals and first-stage estimation error, computed using the delta method. Assetholder and stockholder ultimate consumption risk captures the variation in average returns across the 25 Fama-French portfolios quite well. The implied risk aversion from regression model (5) is about 20 for both assetholder consumption and stockholder consumption with R^2 values of 0.40 and 0.59, respectively. Estimating equation (6), which places consumption risk on the left-hand side of the regression, there is still a significant relation between consumption risk of asset and stockholders and the cross-section of expected returns, but implied risk aversion estimates jump to 47.8 and 33.6, respectively. Using the consumption risk of the top third of asset and stockholders, the cross-sectional fit remains about the same, but implied risk aversion estimates fall to 13.8 (11.0) under regression (5) and 23.6 (19.7) under the reverse regression (6) for asset (stock) holders.

Also reported for each set of regressions is an F-test on the equality of the first-stage covariances with ultimate consumption growth across the 25 portfolios. The F-tests are rejected at the 1% significance level for asset and stockholders and at the 10% level for the top third of asset and stockholders, indicating that there is substantial variation in the estimated first-stage covariances with consumption growth that is statistically different across the tests assets.

Table IV also reports results using non-assetholder and non-stockholder consumption risk. Although the Euler equation may not hold for these households, it is still interesting to examine the estimates obtained for these households for comparison, even if we do not think they are valid measures of their risk aversion. There appears to be no significant relation between long-run non-

[12]Since households in the CEX are not tracked for 12 quarters, these growth rates are not for the same household, but rather for the cohort of households who own stocks. Given the linear relation between returns and consumption growth and assuming that households are representative of their cohort, we interpret equation (5) as holding for each individual household.

assetholder and non-stockholder consumption risk and expected returns. The cross-sectional R^2's are only about 5 percent and the coefficients from the regressions are unreliably different from zero. Hence, the implied risk aversion coefficients from the model are very noisy. For instance, under regression equation (5) we estimate γ to be 20.3 for non-assetholders and 23.1 for non-stockholders with standard errors so large that we cannot reject they are different from zero. When we reverse the regression (equation (6)), we estimate γ to be 479 for non-assetholders and 413 for non-stockholders, which again are not reliably different from zero! These results indicate stronger covariation between returns and consumption of assetholders relative to non-assetholders. Consistent with there being a negligible relation between non-stockholder consumption and returns, the first-stage covariance estimates are also not different from each other (and often not different from zero) across the test assets, as we fail to reject the F-test.

By computing implied risk aversion estimates directly from the model, these regressions provide a direct economic measure of the plausibility of the model. For instance, if the covariance between consumption growth and returns is too small to capture cross-sectional return premia (e.g., Hansen and Singleton (1982), Hansen and Jagannathan (1991), and Lewellen and Nagel (2004)), the regression will produce an implausibly large risk aversion estimate. Another benefit from this structural approach is the use of covariances rather than betas of consumption growth, which reduces the impact of measurement error in consumption since covariances are not scaled by the variance of consumption growth.

Figure 3 shows the positive relation between consumption risk and average returns for asset and stockholders and virtually no relation for non-asset and non-stockholders. It is also evident from the figure that the top third of asset and stockholders generate an even greater spread in consumption risk across the Fama-French portfolios.

C.2. Measurement error

We make the standard assumption in the literature that measurement error is unrelated to the error in expected returns. However, if measurement error in consumption is correlated with error in expected returns, which seems plausible since survey response rates and responses are likely related to the performance of the market, then the coefficient estimates from these regressions will be inconsistent. One solution would be to find instruments for expected returns and consumption growth covariances, which seems difficult. Alternatively, to check whether bias induced by measurement error is a serious issue, we run the regressions using estimates of average returns from a period that does not overlap with the consumption data used to compute covariances. Specifically, we compute average returns on the portfolios from July, 1926 to December, 1981 and consumption growth covariances from 1982 to 1999. Measurement error in expected returns from one period should be uncorrelated with measurement error in consumption from another period. The coefficient estimates from regressions (5) and (6) are remarkably similar to those using the full sample of returns (which partially overlaps with the CEX consumption data) and equally similar to coefficient estimates obtained when estimating expected returns and consumption covariances over the same sample period. These results are not reported for brevity (available upon request) and indicate that potential bias due to correlated measurement error is not a serious concern.

C.3. Consumption growth factor-mimicking portfolios

The CEX covers only the period 1982 to 2002, which Parker (2001) notes, and we will show in Section V, is a period when the covariance of aggregate consumption growth with equity returns is weak. Thus, if the CEX happens to exist over an unfortunate period when the covariance between consumption and returns is low, our estimates of risk aversion may be inflated.

To construct a longer time-series for stockholder consumption growth, we project the present value of consumption growth on a constant, and the excess returns of a set of tradeable assets. Since stock return data is available for much longer periods than the CEX consumption data, we can use the factor portfolio weights estimated in sample to project a time-series of returns for the factor portfolio out of sample, obtaining a factor portfolio designed to mimic long-run assetholder and stockholder consumption growth over a much longer period. The longer time-series may help improve the accuracy of our findings and may cover a period when consumption growth is more strongly related to returns. In addition, if measurement error in consumption is uncorrelated with the asset returns used to construct the factor-mimicking portfolio, then the factor portfolio will contain less measurement error than the CEX consumption growth measures themselves.

Following Breeden, Gibbons, and Litzenberger (1989) and Lamont (2001), we create the factor-mimicking consumption growth portfolio, CGF, by estimating the following regression for asset and stockholder consumption and the top third of asset and stockholders,

$$\sum_{s=0}^{11} \beta^s (c_{t+s+1} - c_{t+s}) = a + b' R_t + \eta_t \tag{7}$$

where c_t is the log of per capita consumption, R_t are the excess returns on the base assets, and $\beta = 0.95^{1/4}$. We use returns, as opposed to log returns, in this regression so that the coefficients

b will be easily interpretable as the weights in a zero-cost portfolio. The return on the portfolio CGF is then,

$$CGF_t = b'R_t \tag{8}$$

which mimics innovations in long-run consumption growth. The resulting factor portfolio is the minimum variance combination of assets that is maximally correlated with ultimate stockholder consumption growth in sample. Equation (8) is not limited to the CEX sample period if b is relatively stable over time.[13] Hence, we create the CGF over the entire period for which reliable returns data is available on the test assets. The test assets we employ are the value-weighted small growth (intersection of the smallest 40% size, lowest 40% BE/ME stocks, based on NYSE breakpoints), large growth (intersection of the largest 40% size, lowest 40% BE/ME stocks), small value (intersection of the smallest 40% size, highest 40% BE/ME stocks), and large value (intersection of the largest 40% size, highest 40% BE/ME stocks) portfolios, whose returns are available from July, 1926 to December, 2004.

Regression (7) is estimated from June, 1982 to May, 1999 for which 12-quarter discounted CEX asset and stockholder consumption growth rates are available. The results from this first-stage regression (with Newey-West adjusted t-statistics employing 36 month lags) are reported in Panel B of Table IV. Consistent with our earlier findings, ultimate consumption growth for asset and stockholders is strongly negatively related to small growth and strongly positively related to small value. These relations are even stronger for the top third of asset and stockholders and are negligible for non-asset and non-stockholder consumption.

Repeating regressions (5) and (6) by replacing actual CEX consumption growth with the factor-mimicking portfolio CGF over July, 1926 to December, 2004, Table IV shows a marked improvement of the cross-sectional fit and significance of consumption risk as well as considerably lower implied risk aversion estimates. For assetholders, the cross-sectional R^2 jumps from 40 percent to 80 percent when using the longer CGF data and implied risk aversion coefficients drop to between 12 and 15. For the top third of assetholders the cross-sectional R^2 increases to 85 percent and risk aversion estimates range from 7.4 to 8.6. The same improvement is evident for stockholder consumption, exhibiting an 80 percent cross-sectional R^2 and even lower risk aversion estimates of 10.4 and 12.7.

[13] Assuming, of course, the relation between consumption growth and asset returns is relatively constant over time. Increases over time in the fraction of the population who hold stocks could have lowered covariances of stock returns with stockholder consumption growth via increased risk sharing (we return to this issue in Section V on aggregate stock returns). If so, results from our factor-mimicking portfolios may be weaker than if these portfolios could be constructed based on a longer time series of stockholder consumption.

For the top third of stockholders the cross-sectional fit improves further and implied risk aversion coefficients are as low as 6.4 (7.5 when running the reverse regression equation (6)). F-tests on the joint equality of the first-stage covariance estimates are rejected at the 5% significance level.

The improvement in cross-sectional fit may be driven by both stronger covariation between asset returns and consumption growth in the period prior to CEX data and by less measurement error in the CGF which captures only the part of consumption growth that is correlated with returns. To examine the relative importance of these effects, we reran the regressions on two subsamples: before CEX data availability (July, 1926 to December, 1981) and during the period for which CEX data exists. Although we omit the results for brevity, we find that risk aversion estimates prior to the CEX period are even lower than those shown in Table IV. For example, the top third of assetholders (stockholders) only require risk aversion between 5.9 and 7.4 (5.2 and 6.4) to explain the cross-sectional price of risk, suggesting the covariance between returns and consumption growth is stronger in the early sample period, consistent with Parker (2001).

Panel C repeats the analyses in Panels A and B using aggregate nondurable consumption from NIPA over the CEX sample period 1982 to 1999, the entire NIPA sample period from 1947 to 2004, and the 1926 to 2004 period using a CGF portfolio constructed for aggregate consumption growth. As Panel C shows, risk aversion estimates become smaller as the sample period extends back in time, indicating stronger covariation between consumption and returns in earlier periods. During the same period for which CEX data exists, we obtain slightly worse cross-sectional fits and slightly larger risk aversion estimates using the CGF instead of actual consumption, suggesting little improvement from measurement error. Hence, the improvement in results from using the CGF over the longer period, is likely due to consumption covarying more with asset returns in the pre-CEX period.

We also report results for the CGF of non-asset and non-stockholder consumption growth, obtaining a poorer fit with cross-sectional average returns and higher risk aversion estimates. Any significance between the CGF of non-asset and non-stockholders and cross-sectional returns should be viewed with caution because the first-stage F-tests on the joint equality of covariances across the 25 Fama-French portfolios are not even close to being rejected. Moreover, the first-stage regression used to construct the CGF for non-assetholders exhibits negligible coefficients on the base assets that are not reliably different from zero. Hence, the results for non-asset and non-stockholder CGF's are likely just noise.

C.4. The entire cross-section of individual stocks

Thus far, our testing ground has been the 25 Fama-French portfolios that capture the size and value spread in average returns. Use of these portfolios is motivated by a large empirical literature claiming size and value premia capture a large fraction of the cross-sectional variation in returns (see Fama and French (1993, 1996)) and for comparability with a vast literature that employs these portfolios exclusively in asset pricing tests. However, the theory applies to all assets in the economy. Hence, in Table V we investigate whether ultimate asset and stockholder consumption risk helps capture the entire cross-section of all individual stock returns. This analysis also helps alleviate the concern that we may be "overfitting" the 25 Fama-French portfolios through specification searches in the literature (see Daniel and Titman (2005)).

Each month, we estimate the covariance of returns with consumption growth for each individual stock traded on the NYSE, AMEX, and Nasdaq with beginning of month share prices above \$5. Since individual stock returns are highly noisy, we follow the procedure of Fama and French (1992) to reduce estimation error. The basic idea is to look for an instrument for the true covariance that we can combine with information from the pre-ranking covariance estimate to rank stocks and form portfolios to compute post-ranking measures. A natural instrument is a characteristic of the stock that we believe is correlated with the true covariance. If this characteristic is also related to average returns, then this procedure will increase variation in the spread in average returns across our portfolios and will enable more powerful asset pricing tests. Both the size of the firm and its BE/ME ratio seem to fit these two criteria based on our previous results. Fama and French (1992) follow a similar procedure by using size and pre-ranking beta sorted portfolios to form post-ranking betas in testing the CAPM following this rationale. Specifically, for each individual stock, we estimate the covariance of its returns with consumption growth using the past 120 months of quarterly overlapping log excess returns before July of year t. Stocks are then sorted in June into 3 size categories (based on the bottom 30, middle 40, and top 30 percentile NYSE breakpoints), 3 book-to-market categories (based on the bottom 30, middle 40, and top 30 percentile NYSE breakpoints), and 10 pre-ranking covariance deciles (based on universal breakpoints). Based on these rankings, each stock is assigned to one of these 90 groups. We then compute the equal-weighted quarterly log excess returns on these 90 portfolios over the next 12 months, from July to June. This procedure is repeated every year, forming a time-series of returns on these 90 portfolios. We then reestimate covariances for the portfolios formed from the size, BE/ME, and pre-ranking sorts using the full sample of returns to obtain *post-ranking* covariances. The post-ranking covariance estimate for a given group is then assigned to each stock in the group, with group assignments updated each June. Even though the post-ranking covariances themselves do not change over time, as an individual stock moves into and out of one of the 90 portfolios due to either its size, BE/ME, or pre-ranking covariance changing, that stock will receive a different post-ranking covariance. This procedure reduces estimation error by effectively shrinking the individual covariance estimates to a portfolio average and also employing the full sample of data.

Using the post-ranking covariance estimates, we run Fama and MacBeth (1973) month-by-month cross-sectional regressions from June, 1982 to May, 1999 of the entire cross-section of log excess stock returns on their covariance with long-run consumption growth, COV_{cg}. To assess the marginal impact of COV_{cg} controlling for other known determinants of returns, we also include COV_{RMRF}, the covariance of each stock's return with the excess return on the CRSP value-weighted index, the log of market capitalization, the log of BE/ME, and the past year return of the stock, $ret_{-12:-2}$, skipping the most recent month as regressors. The time-series average of the monthly coefficient estimates and their t-statistics are reported in the style of Fama and MacBeth (1973).

Table V Panel A indicates that covariance with long-run consumption growth of asset and stockholders captures significant cross-sectional variation in average returns, even when controlling for the market, size, BE/ME, and past year returns. For the top third of assetholders, results are even stronger and for non-asset and non-stockholder consumption, there appears to be no significant relation to the cross-section of returns. These results are consistent with those from the 25 Fama-French portfolios and provide additional evidence supporting ultimate asset and stockholder consumption risk as a significant variable explaining the entire cross-section of returns.

We omit implied risk aversion estimates from Table V because covariance estimates for individual stocks are likely measured with much more noise than those for the 25 Fama-French portfolios, leading to unreliable risk aversion estimates. For example, we obtain very small risk aversion estimates when we run regression equation (5) and very large ones when we run equation (6). Thus, Table V focuses on statistical significance and whether ultimate consumption risk has explanatory power in the full cross-section after known determinants of returns are controlled for.

Panel B of Table V employs CGF in place of actual CEX consumption over the longer sample period July, 1926 to December, 2004. Covariance with the CGF of asset and stockholders carries a significant positive risk premium in the entire cross-section of stocks even in the presence of size, BE/ME, and momentum premia. Covariance with the non-asset and non-stockholder CGF exhibits little relation to the cross-section.

D. Asset Pricing Tests

Another useful feature of the CGF is that we can use it in asset pricing tests to compare its performance against other factor-mimicking portfolios such as those of Fama and French (1993). Equations (7) and (8) used to form the CGF can be thought of as a way to form a factor related to size and value that employs consumption growth as an economic guide to determine the weights that should be placed in these asset classes. For instance, rather than simply going long a dollar in value and short a dollar in growth, as in HML, the Fama and French (1993) value factor, we let the data tell us the relative weights that should be placed in each asset category, motivated by the underlying theory of Section II. If the theory is correct, then the consumption growth factor should perform as well as, or even better than, the Fama and French (1993) factors SMB and HML. While this is likely to be the case in sample, we are particularly interested in the performance of the CGF out of sample.

Before turning to the asset pricing tests, Panel A of Table VI reports the mean returns (and t-statistics) of the CGF and the Fama and French (1993) factors $RMRF$, SMB, and HML, as well as the momentum factor-mimicking portfolio UMD (taken from Ken French's website, and only available from January, 1931 to December, 2004). The premium on the asset and stockholder CGF is 46 and 47 basis points per month, respectively, and is highly statistically different from zero. For the top third of asset and stockholders, the premium jumps to 77 and 87 basis points per month, respectively. The Sharpe ratios of the CGF's (which can be judged from their t-statistics) are higher than those of the market, SMB, HML, or UMD. Importantly, despite its high Sharpe ratio, the CGF does not appear to be "too good a deal," since roughly 38 percent of the time it delivers negative returns, which is about the same frequency exhibited for the market. The correlations between the CGF and the Fama-French factors are as expected given the results from the first-stage factor portfolio weights from Table IV Panel B. That is, the CGF is moderately positively correlated with size (SMB) and highly correlated with value (HML). The CGF is also highly positively correlated with the market, which we will explore in Section V.

Panel B of Table VI reports regression results of the CGF on the Fama and French (1996) 4-factor model, which consists of $RMRF$, SMB, HML, and UMD. As Panel B indicates, the intercepts or α's from these time-series regressions are significant and range from 24 to 44 basis points per month across the CGF's. Hence, over the entire July, 1926 to December, 2004 time period, the 4-factor model fails to account for the CGF premium. We also conduct the same time-series regression over the period prior to the existence of CEX consumption data, July, 1926 to December, 1981, as an out of sample test. The α's over this out of sample period are remarkably similar and range from 22 to 41 basis points per month, which are also highly statistically significant.

Since the CGF is not fully captured by the 4-factor model, we now investigate whether the CGF can capture the Fama-French factors and how it performs relative to the Fama-French factors in asset pricing tests. Panel C of Table VI examines the ability of various factor models to capture the returns on the 25 Fama-French portfolios. We regress the excess returns on the 25 Fama-French portfolios on $RMRF$, $RMRF$ in combination with SMB and HML (both separately and together), and a two factor model consisting of $RMRF$ and the CGF, as well as a single factor model consisting of the CGF alone. Formally, for each of the 25 Fama-French portfolios we run the following regression,

$$r_{i,t} - r_{f,t} = \alpha_i + \beta'_{i,F} F_t + \epsilon_{i,t} \tag{9}$$

where F_t is a set of factor portfolios. We assess the joint significance of the α's using the Gibbons, Ross, and Shanken (1989) F-statistic (GRS) and report the average absolute α and R^2 across the 25 time-series regressions. The first four columns report results over the entire sample period from July, 1926 to December, 2004. The first row reports results for the unconditional CAPM (i.e., $RMRF$ only). As is well-known, the market portfolio fails to capture the variation in returns across the 25 size and BE/ME portfolios, generating a 3.18 F-statistic that rejects the null that the intercepts are jointly zero (p-value < 0.0001) and leaves an absolute α of 21 basis points per month on average across the portfolios. Adding SMB and HML as factors improves the fit, with the best fit coming from including both factors, which is the Fama and French 3-factor model. However, the GRS F-test is still rejected (p-value $= 0.001$), although the absolute α is smaller at 13 basis points per month.[14]

The remaining rows report results for the CGF's. As the table shows, the GRS F-test fails to reject when using the CGF in place of SMB and HML (p-values range from 0.07 to 0.11). In part, rejection of the GRS test is driven by the CGF having slightly lower R^2's (about 0.80) than the Fama-French 3-factor model (0.90), but it is also driven by the smaller economic magnitudes of

[14] As Fama and French (1996) point out, part of the reason their model rejects the F-test is that it captures a huge fraction of the variation in these portfolios (the average R^2 is 0.90), leaving a small residual covariance matrix that when inverted produces a large F-statistic.

the intercepts. The average absolute α is only 9 basis points when using the assetholder CGF and 11 basis points when using the top assetholder CGF. Even when we remove $RMRF$ and examine the CGF by itself, the average absolute α's remain small (13 to 17 basis points).

For robustness, we perform the same time-series tests over the subperiod July, 1926 to December, 1981 that preceded the availability of CEX consumption data to provide an out of sample test of the consumption growth factors, whose weights were constructed from June, 1982 to May, 1999. As the last four columns of Panel C indicate, the success of CGF is even better in this out of sample period. In general, all of the models are better able to price the 25 size and BE/ME portfolios in this period. However, the CGF still produces lower F-statistics (with p-values typically > 0.30), and the economic magnitudes of the pricing errors are slightly smaller in this out of sample period. The average absolute α under the Fama-French model in this period is 13 basis points, but the CGF's generate average absolute α's that range from 10 to 12 basis points per month. Even out of sample, the CGF performs no worse than the Fama and French (1993) three factor model in explaining the returns on the 25 Fama-French portfolios.

In addition to examining whether the CGF prices the cross-section of the 25 size and BE/ME-sorted portfolios, we also examine whether the CGF can capture the Fama-French factors themselves. The bottom half of Panel C reports results from time-series tests of SMB and HML directly on the CGF. The lack of significance of the intercepts (α's) from these regressions indicates that the CGF effectively captures the premia associated with SMB and HML. Both economically and statistically, the α's for SMB and HML are negligible, even over the out of sample period prior to CEX data availability from which the factors were constructed. The CGF, which itself is not fully captured by SMB and HML (Panel B), appears to capture the size and value premia associated with these two factors.

V. Aggregate Returns

In this section we examine the relation between stockholder consumption and aggregate returns on stocks and bonds.

A. Stock and bond market predictability

A central theme in models of limited stock market participation (Basak and Cuoco (1998), Guvenen (2004)) is that the consumption (or wealth) share of stockholders is a state variable that should predict excess returns on stocks. This prediction holds even if stockholders and non-stockholders have identical preferences, though the effects are stronger if stockholders are less risk averse than non-stockholders. As suggested in Table III from responses to the SCF, stockholders seem to have more risk tolerance than non-stockholders. Of course, for this mechanism to be powerful there needs to be substantial variation over time in stockholder consumption shares, possibly induced endogenously by preference differences and resulting differences in sensitivity to labor and financial income shocks (Table II and Figure 2). The question as to whether these movements are large enough to generate significant return predictability is an empirical one.

The issue of whether stockholder consumption shares predict excess stock and bond returns is also closely related to two recent papers by Lettau and Ludvigson (2001a) and Julliard (2005) who show that measures of the consumption-to-wealth ratio have strong predictive power for future excess stock returns. Both papers include financial wealth and a measure of human capital in their wealth calculations. The findings are interpreted based on a representative agent's intertemporal budget set. High consumption-to-wealth this period reflects either high expected returns or low expected consumption growth. The latter channel is shown to be less relevant empirically.

Economically, the predictive power of the consumption-to-wealth ratio is less well understood. Lettau and Ludvigson (2001a) suggest that their findings are consistent with the habit-formation framework of Campbell and Cochrane (1999). Julliard (2005) suggests that a representative agent framework with Epstein-Zin utility is consistent with the predictive power of the consumption-to-wealth ratio if there is a predictable component in consumption growth (as in the model of Bansal and Yaron (2002)). We offer evidence in favor of a third explanation based on limited stock market participation and time-variation in the stockholder consumption share. A link between stockholder wealth shares and the consumption-to-wealth ratio can also be directly motivated by micro-level evidence on savings rates that wealthy households (i.e., stockholders) have higher savings rates (Dynan, Skinner, and Zeldes (2004), Carroll(2000)). Bosworth, Burtless, and Sabelhaus (1991) show that savings rates for stock and bondholders in the CEX are much higher than for non-assetholders. Thus, when stockholder wealth increases relative to non-stockholders (or participation rises), we expect to see a lower aggregate consumption-to-wealth ratio and vice versa.

Table VII reports results from predictability regressions of future stock and bond market excess returns and yields on Lettau and Ludvigson's (2001a) cay, Julliard's (2005) expected future labor income growth rate lr combined with cay, and the ratio of quarterly consumption of assetholders, the top third of assetholders, stockholders, and the top third of stockholders to total aggregate quarterly consumption in the CEX, calculated using the CEX survey weights. Results are similar if we employ the ratio of stockholder to non-stockholder consumption instead (these results are available upon request). Regressions are estimated from June, 1982 to February, 2002 for which we have CEX consumption data. We use quarterly CEX ratios, available at the monthly frequency. To address the effect of measurement error, we run regressions in both directions with consumption ratios appearing on the left hand side of the regression as well.

Panel A of Table VII reports results for predicting 12 quarter ahead stock market returns in excess of the Treasury bill rate. Newey-West standard errors are computed with 12-quarter lags. Confirming the results in Lettau and Ludvigson (2001a) and Julliard (2005), *cay* and *caylr* forecast future stock market returns. Since savings rates are higher among the wealthy, we expect stockholder consumption (a proxy for wealth) shares to be negatively related to *cay* and forecast market returns with opposite sign. Repeating these regressions by replacing *cay* (or *caylr*) with asset and stockholder consumption shares, we find a negative but insignificant relation between stockholder consumption shares and future market returns. However, when we reverse the regression to mitigate measurement error in consumption, we find a significant negative relation (except for all assetholders) that is stronger for the wealthiest asset and stockholders. The last column of Panel A employs annual charitable donations as a fraction of personal disposable income from Giving USA (2003) from 1962 to 2002 (results are similar dividing by aggregate consumption) as another proxy for the wealth share of assetholders. This ratio also exhibits a significant negative relation with future stock returns.

Since returns are noisy, we also employ the earnings-to-price ratio on the aggregate stock market in excess of the yield on 30-day T-bills as a measure of future value in Panel B. Yields are forward looking and may have less noise. As Panel B shows, *cay* correlates positively with stock market yields, though *caylr* is negatively related to stock market yields. The consumption shares of asset and stockholders are strongly negatively related to stock market yields. In addition, charitable donations are also negatively related to the stock market yield.

Panel C examines whether the various ratios can forecast aggregate bond excess returns and yields (over the T-bill rate) as well. Since theory pertains to all tradable assets, it is interesting to examine other asset classes besides equities. In addition, this examination will help alleviate any concerns of sample specific issues such as data mining. As Panel C indicates, *cay* and *caylr* are weakly related to bond market excess returns, defined as the term spread or difference in returns between long-term (10-year) government bonds and the T-bill rate. Asset and stockholder consumption shares, including donations, are strongly negatively related to aggregate bond excess returns. Panel D shows similar results for bond market yields rather than returns.

Finally, Panel E reports regression results of the log of each of the consumption shares and donations on *cay* and *caylr*, as well as the correlations between them. The relation between our consumption share measures and the consumption-to-wealth ratio is strongly negative. Figure 5 plots the consumption-to-wealth ratio against stockholder consumption shares (and donations), highlighting the strong negative relation between the two.[15] This evidence suggests that the consumption-to-wealth ratio may be linked to stock market participation, possibly providing an economic story for *cay*'s empirical success in pricing assets.

B. The equity premium

Finally, we address to what extent a model emphasizing limited stock market participation and ultimate consumption growth can help resolve the equity premium puzzle. The main added value of this section over previous work by Parker (2001), Attanasio, Banks, and Tanner (2002), Brav, Constantinides, and Geczy (2002) and Vissing-Jørgensen (2002) is to use our factor-mimicking portfolios to construct longer (estimated) time series of stock and assetholder consumption.

Table VIII reports implied measures of risk aversion for the equity premium across various measures of ultimate consumption growth using aggregate consumption (of nondurables) from NIPA and CEX consumption for assetholders, the top third of assetholders, and non-assetholders. The ultimate consumption growth measures are, as before, 12-quarter discounted growth rates available at a monthly frequency. Implied risk aversion coefficients are computed from equation (2), with the (quarterly) excess log return on the CRSP value-weighted index used as the market proxy. For the longest time-period considered, we also report a 95% confidence interval around each risk aversion estimate using a block bootstrap method that resamples 10,000 times, 12 quarters at a time and uses the percentile method.

[15] Santos and Veronesi (2004) provide a general equilibrium model emphasizing the role of wages-to-consumption as a state variable that is also negatively correlated with the consumption-to-wealth ratio. They find empirical success for this ratio in predicting the cross-section of returns on the Fama-French portfolios and the time-series of aggregate market returns, though require a much larger risk aversion coefficient than our setting. However, we find that the stockholder consumption share is *negatively* correlated with their wages-to-consumption ratio, implying that the stockholder consumption share is low when wages-to-consumption is high. Hence, it is likely that these two variables are capturing different effects.

We report results over three sample periods: the period over which CEX data are available June, 1982 to May, 1999; the sample for which monthly NIPA aggregate consumption data are available January, 1959 to December, 2004; and the period over which the factor-mimicking portfolio CGF is available July, 1926 to December, 2004. For reference, we also show the results based on the factor-mimicking portfolios for the two shorter time periods. To minimize the effect of estimation uncertainty in the equity premium, implied risk aversion coefficients are calculated using quarterly excess stock return data for the full July, 1926 to December, 2004 period in all cases. Differences in risk aversion estimates across periods are therefore driven by differences in the covariance of the excess return on stocks with the ultimate consumption growth measures. These covariances are also reported in the table.

For comparison, we first compute the implied risk aversion for the equity premium using aggregate consumption from NIPA. As Table VIII shows, implied risk aversion is around 78 over both the CEX sample period and the longer period of NIPA data availability and whether we use the ultimate consumption growth measure or its factor-mimicking portfolio. However, prior to 1959, the covariance between the aggregate consumption growth CGF and market returns is higher, implying an estimated risk aversion coefficient of about 22.7 over the full sample from 1926 to 2004 with a confidence interval of 8.2 to 59.2.

Turning to the CEX-based ultimate assetholder consumption growth, however, we obtain much smaller risk aversion estimates. Over the CEX sample period, we obtain a point estimate of 24.4 based on ultimate assetholder consumption growth and 37.6 based on ultimate non-assetholder consumption growth (results for stockholders versus non-stockholders are similar). For the top assetholders the risk aversion estimate is 17.8. Parker (2001) also obtains lower risk aversion estimates based on CEX data than those based on NIPA data, and suggests that this could be partly due to the fact that the Euler equations can be aggregated correctly across households in the CEX (taking the cross-sectional average of the change in log consumption growth, rather than the log of per capita consumption).

Since we do not have CEX data prior to 1982, we employ the factor-mimicking consumption growth factor, CGF, dating back to 1926. As in the aggregate data, using the full 1926 to 2004 period results in higher covariances and thus much lower risk aversion estimates. For the top third of assetholders, the risk aversion estimate is as low as 6.6 with a 95% confidence interval of 1.5 to 15.3. These risk aversion estimates are similar to those found previously in the cross-section of returns. These results suggest that fairly reasonable risk aversion estimates can be obtained using the full 1926 to 2004 period and focusing on ultimate assetholder consumption risk.

VI. Conclusion

We find empirical support for consumption-based asset pricing and the role of human capital. Ultimate labor income and stockholder consumption risk helps capture the return premia associated with size and value as well as time-varying risk premia in aggregate stock and bond markets. This evidence supplements recent findings of long-run aggregate consumption risk having explanatory power for asset prices. We highlight the importance of distinguishing assetholder from non-assetholder consumption and link the role of ultimate consumption risk back to ultimate labor income risk.

The fact that the consumption and labor sensitivity of households that actually own financial assets drives these relations is comforting and suggests that recent evidence on the success of long-run consumption growth is unlikely to be due to chance. Future research may well investigate further why labor income (and stockholder consumption) growth responds slowly to the news contained in asset returns and uncover what is driving these long-run relations.

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Appendix A: CEX Consumption Measures

We detail below our classification of households in the CEX as well as our consumption measure and sample selection criteria in this data set.

Stockholder status

The CEX contains information about four categories of financial assets. Households are asked for their holdings of "stocks, bonds, mutual funds and other such securities", "U.S. savings bonds", "savings accounts", and "checking accounts, brokerage accounts and other similar accounts."

We refer to households with positive responses to the category "stock, bonds, mutual funds and other such securities" as stockholders and those with zero holdings as non-stockholders. It is known, for instance from the Survey of Consumer Finances, that many households hold stocks or bonds only in their pension plan. Unfortunately, it is not possible to determine whether households with defined contribution plans report their stockholdings and bondholdings in these plans when answering the CEX questions. The percent of stockholders (documented below) in the CEX is smaller than in other sources. However, consistent with other data sources, the proportion of stockholders is upward trending during the sample, with the exception of the last couple of years. The lower participation rates in the CEX may indicate that some households with stockholdings in pension plans do not report these, leading them to be miscategorized as non-stockholders. We therefore also consider a broader definition of who likely holds stocks by including households who own U.S. savings bonds and refer to this broader category as assetholders. Households with zero holdings of both of these categories are classified as non-assetholders. The motivation for this broader category is that stockholding is known to be strongly increasing in financial wealth (see Vissing-Jørgensen (2003)) and therefore large holdings of savings bonds may be a good indicator of holding stocks indirectly through pension plans. We do not include households with positive holdings in only savings or checking accounts in the assetholder category. Results are similar, however, if we also include these households as assetholders.

The Euler equation involving consumption in period t and $t+1$ should hold for those who hold the asset as of date t. Therefore, holding status must be defined based on holdings at the beginning of period t (when considering the consumption growth between period t and $t+1$). Two additional CEX variables are used for this purpose. The first variable reports whether the household holds the same amount, more, or less of the asset category compared to a year ago. The second variable reports the dollar difference in the estimated market value of the asset category held by the household last month compared to the value of the asset category held a year prior to last month. We define a household as holding an asset category at the beginning of period t if it 1) reports holding the same amount of the asset as a year ago and holds a positive amount at the time of the interview (the fifth interview) or 2) reports having lower holdings of the asset than a year ago, or 3) reports having had an increase in its holdings of the asset but by a dollar amount less than the reported holdings at the time of the question.[16] Based on these criteria we classify

[16] Around 1,121 households in our final sample of 48,605 households report an increase in their holdings of stocks, bonds and mutual funds but do not report their current holdings. Most of these households are likely to have held these assets a year ago and are therefore placed into the stockholder category. Similarly, around 543 households report an increase in their holdings of U.S. savings bonds but do not report their current holdings. We classify these as assetholders.

A small number of households report an increase in their holdings of stocks, bonds or mutual funds or of U.S. savings bonds larger than the value of the reported end of period holdings. We classify these as non-assetholders.

22.78 percent of households in our sample as stockholders and 31.38 percent as assetholders.

In addition, the set of stockholders and assetholders are each split into three layers of approximately equal size based on dollar amounts of initial holdings, calculated as current holdings minus the change in holdings during the current period. We are most interested in the wealthiest stockholders and supplement our results for stockholders and assetholders with results for the top third of households within each of these groups.

The top third of stockholders consists of those reporting initial holdings above $25,974 in real 1982-1984 dollars, using the CPI for total consumption of urban households to deflate nominal values.[17] The top third of assetholders are classified based on the sum of the holdings of stock, bonds, and mutual funds, and U.S. savings bonds. The top third held more than $13,614 combined in these asset categories.

Consumption measure and sample choice

The consumption measure used is nondurables and some services aggregated from the disaggregate CEX consumption categories to match the definitions of nondurables and services in NIPA. We use consumption as reported in the Interview Survey part of the CEX. The service categories excluded are housing expenses, but not costs of household operations, medical care costs, and education costs, since these three costs have substantial durable components. Attanasio and Weber (1995) use a similar definition of consumption. In leaving out durables, it is implicitly assumed that utility is separable in durables and nondurables/services. Nominal consumption values are deflated by the BLS deflator for nondurables for urban households.

For each household we calculate quarterly consumption growth rates based on the reported monthly consumption values. We drop household-quarters in which a household reports non-zero consumption for more than three or less than three months or where our consumption measure is negative.

Extreme outliers are dropped since these may reflect reporting or coding errors. Specifically, we drop observations for which the consumption growth ratio is less than 0.2 or above 5. In addition, non-urban households (missing for part of the sample) and households residing in student housing are dropped, as are households with incomplete income responses. Furthermore, we drop households who report a change in age of household head between any two interviews different from 0 or 1 year. These exclusions are standard. We also drop all consumption observations for households interviewed in 1980 and 1981, since the CEX food question was changed in 1982 leading to a drop in reported food consumption. The question was changed back to the initial question in 1988, but there is no obvious way to address this without substantial loss of data. See Battistin (2003) for details on the questions asked.

Finally, because financial information is reported in interview five, and because we wish to calculate consumption growth values by household, households must be matched across quarters. Therefore, we drop households for which any of interviews two to five are missing. Matching households across interviews creates problems around the beginning of 1986 and the beginning of

[17]For interviews conducted from 1991 to 1997, about 5 percent of households report holdings of stocks, bonds and mutual funds of $1. We contacted the Bureau of Labor Statistics to determine if this was a coding error, but they were not sure how to interpret the $1 answers. Since all of the households reporting $1 assetholdings answer the question comparing current holdings to holdings a year ago it is likely that they are holding such assets. We therefore include them as stockholders when doing the stockholder-nonstockholder classification. However, since the $1 households cannot be classified by layer of stockholding, we exclude them in estimations based on layers of stockholders. Thus, the total number of stockholders used to define wealth layers is smaller than the total number of stockholders.

1996 since sample design and household identification numbers were changed, with no records being kept of which new household identification numbers correspond to which old ones. We therefore exclude households who did not finish their interviews before the ID change, implying that fewer observations are available for the last two months of 1985 and 1995 and the first 6 months of 1986 and 1996 around the ID changes. Furthermore, no households were interviewed in April, 1986 and April, 1996. To avoid a missing value in our time series we set quarterly consumption growth for March 1986 and March 1996 equal to aggregate quarterly real nondurable per capita consumption growth for this month. Results are similar if we simply drop these months.

While each household is interviewed three months apart, the interviews are spread out over the quarter implying that there will be households interviewed in each month of the sample. Thus, the data frequency is monthly. Our final sample consists of 145,815 quarterly consumption growth observations. The average number of consumption growth observations for stockholders per month is 141, while the average number for non-stockholders is 479. There are on average 195 assetholder and 426 non-assetholder consumption growth observations per month.

We also control for consumption changes driven by changes in family size, by regressing the change in log consumption on the change in log family size at the household level within each group of households using quarterly growth rates. We then use the residual from this regression as our basic quarterly consumption growth measure.

Table I:

Ultimate Labor Income Risk and Expected Returns

Cross-sectional regressions of the average log excess returns on the 25 Fama-French portfolios against the covariance of returns with future labor, national, and capital income growth from quarter t to $t+S$ are reported. Data are obtained from NIPA Table 1.12, and are seasonally adjusted, quarterly from 1947 to 2004, in real per capita dollars using the consumer price index for urban consumers from the Bureau of Labor Statistics and population numbers from NIPA Table 2.1. Panel A reports results for labor income = compensation of employees + proprietor's income. Panel B reports results for national income and Panel C for capital income = corporate profits + rental income + net interest. Panel D reports results for aggregate nondurable consumption for comparison. The present value of the future growth rate in income (or consumption) is calculated by discounting at an annual rate of 5% (quarterly discount factor $\beta = 0.95^{1/4}$). The entire time-series of returns from July, 1926 to Dec., 2004 is employed to estimate the mean returns on the 25 Fama-French portfolios and t-statistics are reported in parentheses below the coefficient estimates under both OLS and correcting standard errors for both cross-correlation in the residuals and for first-stage estimation error in the covariances. Panel E reports the first-stage covariance estimates and t-statistics (with Newey-West adjusted standard errors employing 12 quarter lags) of each of the 25 Fama-French portfolios with ultimate labor income growth from quarter t to $t+12$. An F-test on the equality of the first-stage covariances across the 25 portfolios is also reported.

Cross-sectional regressions on the 25 Fama-French portfolios

$$E[r_{i,t}] - r_f + \sigma_i^2/2 = c + \lambda Cov(r_{i,t+1}, \sum_{s=0}^{S-1} \beta^s (y_{t+s+1} - y_{t+s})) + e_i$$

$S =$	1	2	4	8	12	16	20	24
Panel A: Labor income								
λ	74.45	47.39	24.06	16.35	45.99	37.68	33.99	36.42
$t(\lambda)$	(1.70)	(1.68)	(1.87)	(1.29)	(4.04)	(2.67)	(4.25)	(4.14)
$t^*(\lambda)$	(1.64)	(0.75)	(0.64)	(0.84)	(3.16)	(1.99)	(3.00)	(2.76)
R^2	0.11	0.11	0.13	0.07	0.42	0.24	0.44	0.43
Panel B: National income								
λ	39.79	34.75	15.20	15.06	52.31	39.08	39.95	23.43
$t(\lambda)$	(1.39)	(1.88)	(1.45)	(1.12)	(5.45)	(2.22)	(4.11)	(2.34)
$t^*(\lambda)$	(1.03)	(0.84)	(0.50)	(0.73)	(4.26)	(1.65)	(2.90)	(1.56)
R^2	0.08	0.13	0.08	0.05	0.56	0.18	0.42	0.19
Panel C: Capital income								
λ	12.45	11.80	7.03	7.43	18.15	7.36	9.88	0.49
$t(\lambda)$	(1.24)	(1.99)	(1.40)	(0.93)	(4.17)	(0.88)	(1.50)	(0.08)
$t^*(\lambda)$	(0.92)	(0.89)	(0.48)	(0.61)	(3.26)	(0.65)	(1.06)	(0.05)
R^2	0.06	0.15	0.08	0.04	0.43	0.03	0.09	0.00
Panel D: Aggregate Nondurables Consumption								
λ	86.43	7.37	13.90	21.70	48.95	52.79	17.13	38.31
$t(\lambda)$	(2.08)	(0.33)	(0.87)	(1.90)	(7.12)	(4.17)	(1.42)	(2.73)
$t^*(\lambda)$	(2.01)	(0.15)	(0.30)	(1.24)	(2.76)	(3.11)	(1.00)	(1.82)
R^2	0.16	0.00	0.03	0.14	0.67	0.43	0.08	0.24

Panel E: First-stage covariance estimates for 12-quarter labor income growth

		growth 1	2	3	4	value 5	average
small	1	8.71	8.02	8.92	8.60	10.28	8.91
	2	7.97	7.60	8.27	9.10	8.82	8.35
	3	6.95	8.21	8.46	8.04	8.45	8.02
	4	6.28	7.23	8.29	8.77	10.15	8.14
large	5	8.66	6.02	6.32	8.13	10.09	7.84
average		7.71	7.42	8.05	8.53	9.56	
				t-statistics			F-stat
small	1	0.91	1.00	1.41	1.62	1.92	7.33
	2	0.84	1.13	1.85	1.99	1.96	
	3	0.88	1.59	1.98	2.16	1.66	p-value
	4	1.00	1.46	2.19	2.91	2.52	0.003
large	5	2.59	1.65	1.98	2.67	3.47	

*Standard errors corrected for cross-correlation of the residuals and for first-stage estimation error.

Table II:

Sensitivity of Stock and Non-stockholder Consumption to Aggregate Consumption and Labor Income Shocks

Results from regressions of long-run growth in stockholder, non-stockholder, and the wealthiest top third of stockholders' consumption growth from quarter t to $t+12$, estimated from the CEX using the broader definition for assetholders, on labor income growth (Panel A) and aggregate consumption growth (Panel B) from NIPA, are reported from June, 1982 to May, 1999. Details of the series construction are in Appendix A. The present value of future labor income and consumption growth is calculated using an annual discount rate of 5% (quarterly discount factor $\beta = 0.95^{1/4}$). Stock, non-stock, and the wealthiest third of stockholders' consumption sensitivity to aggregate consumption and labor income shocks are estimated in a linear and piecewise linear model allowing the coefficients to vary across three even categories. Standard errors are calculated under Newey-West allowing for 12 quarter lags, where t-statistics are reported in parentheses.

	Stockholders (1)	Non-stockholders (2)	Top third (3)	(1)−(2)	(3)−(2)
Panel A: Sensitivity to aggregate labor income shocks					
	$\sum_{s=0}^{11} \beta^s (c_{t+s+1} - c_{t+s}) = a + \delta \sum_{s=0}^{11} \beta^s (y_{t+s+1} - y_{t+s}) + e_t$				
δ	1.26	0.49	1.94	0.77	1.45
	(9.46)	(2.74)	(6.77)	(3.61)	(6.95)
R^2	0.48	0.17	0.27		
	Piecewise linear regression: allowing δ to vary across three even categories				
δ_1	2.35	3.61	0.42		
	(5.18)	(3.01)	(1.53)		
δ_2	-2.46	-4.61	-1.14		
	(-2.85)	(-2.61)	(-1.73)		
δ_3	1.78	2.28	-0.896		
	(3.28)	(1.27)	(-2.84)		
R^2	0.53	0.30	0.29		
F-test $\delta_2 = \delta_3 = 0$	24.86	35.42	46.52		
p-value	< 0.001	< 0.001	< 0.001		
Panel B: Sensitivity to aggregate consumption shocks					
	$\sum_{s=0}^{11} \beta^s (c_{t+s+1} - c_{t+s}) = a + \delta \sum_{s=0}^{11} \beta^s (C_{t+s+1}^{Agg.} - C_{t+s}^{Agg.}) + e_t$				
δ	2.14	0.90	2.92	1.24	2.02
	(9.10)	(3.09)	(5.98)	(3.04)	(4.97)
R^2	0.45	0.19	0.20		
	Piecewise linear regression: allowing δ to vary across three even categories				
δ_1	4.67	5.51	0.74		
	(4.12)	(1.57)	(0.91)		
δ_2	-3.96	-5.89	-1.96		
	(-1.72)	(-1.01)	(-1.28)		
δ_3	3.56	3.16	-0.80		
	(2.88)	(0.73)	(-0.97)		
R^2	0.50	0.21	0.27		
F-test $\delta_2 = \delta_3 = 0$	28.30	13.75	68.43		
p-value	< 0.001	< 0.001	< 0.001		

Table III:

Statistics on Labor Income and Stockholdings for Stockholders from the Survey of Consumer Finances 1989–2001

The table uses data from the 1989, 1992, 1995, 1998, and 2001 *Survey of Consumer Finances* to document labor income to total income (Panel A), labor and business income to total income (Panel B), fraction of total stock market owned (Panel C), and stockholdings to financial wealth (Panel D). This is done separately for the full set of stockholders, the top third of stockholders defined based on dollar stockholdings, and the top one percent of stockholders defined based on dollar stockholdings. Stockholdings include both direct and indirect holdings. Labor income refers to income from wages and salaries plus income from unemployment benefits and worker's compensation. Business income refers to income or losses from a professional practice, business or farm. The other income categories in total income are income from non-taxable investments (e.g., municipal bonds), other interest income, dividend income, income or losses from net gains or losses from the sale of stocks, bonds, or real estate, income from rent, trusts, or royalties, income from child support or alimony, welfare payments, income from social security and other pensions, annuities or disability or retirement programs, and other income (including settlements, gambling winnings, education scholarships or grants, inheritances, and gifts). All income measures are pre-tax. Panel E reports measures of risk aversion for stock and non-stockholders. The SCF poses the following question to each respondent: "Which of the statements on this page comes closest to the amount of financial risk that you are willing to take when you save or make investments?" (1) TAKE SUBSTANTIAL FINANCIAL RISKS EXPECTING TO EARN SUBSTANTIAL RETURNS; (2) TAKE ABOVE AVERAGE FINANCIAL RISKS EXPECTING TO EARN ABOVE AVERAGE RETURNS; (3) TAKE AVERAGE FINANCIAL RISKS EXPECTING TO EARN AVERAGE RETURNS; or (4) NOT WILLING TO TAKE ANY FINANCIAL RISKS. Panel E summarizes the responses to this question across stock and non-stockholders separately. The p-value of an F-test that the responses are the same across the groups is also reported. All numbers are calculated using SCF weights and based on data that are averaged across SCF imputations.

Survey year	1989	1992	1995	1998	2001
Panel A: Mean (median) of labor income to total income					
All stockholders	0.79 (0.96)	0.81 (0.97)	0.81 (0.98)	0.82 (0.98)	0.79 (0.98)
Top third	0.71 (0.75)	0.69 (0.84)	0.72 (0.91)	0.69 (0.86)	0.64 (0.83)
Top 1 percent	0.30 (0.03)	0.35 (0.19)	0.43 (0.11)	0.43 (0.23)	0.42 (0.20)
Panel B: Mean (median) of labor and business income to total income					
All stockholders	0.86 (0.97)	0.92 (0.98)	0.86 (1.00)	0.88 (0.99)	0.84 (0.99)
Top third	0.85 (0.84)	0.90 (0.92)	0.81 (0.96)	0.76 (0.92)	0.71 (0.89)
Top 1 percent	0.45 (0.21)	0.50 (0.51)	0.77 (0.31)	0.63 (0.46)	0.57 (0.41)
Panel C: Fraction of total stock market owned					
All stockholders	1	1	1	1	1
Top third	0.93	0.93	0.93	0.93	0.93
Top 1 percent	0.32	0.33	0.37	0.36	0.33
Panel D: Mean (median) of stocks to financial wealth					
All stockholders	0.31 (0.26)	0.36 (0.31)	0.43 (0.39)	0.48 (0.47)	0.52 (0.50)
Top third	0.47 (0.44)	0.52 (0.50)	0.58 (0.57)	0.65 (0.67)	0.67 (0.70)
Top 1 percent	0.69 (0.76)	0.67 (0.70)	0.77 (0.85)	0.78 (0.82)	0.76 (0.84)
Panel E: Mean risk aversion category (1=lowest, 4=highest)					
Non-stockholders	3.45	3.48	3.45	3.44	3.51
All stockholders	3.03	3.07	2.96	2.79	2.79
Top third	2.94	2.80	2.73	2.61	2.60
Top 1 percent	2.91	2.59	2.79	2.49	2.54
F-test p-value	< 0.001	< 0.001	< 0.001	< 0.001	< 0.001

Table IV:

Stockholder and Non-stockholder Ultimate Consumption Risk and Expected Returns on the 25 Fama-French Portfolios

Panel A reports estimates of the Euler equation for the ultimate consumption growth of stockholders, the top third of stockholders, non-stockholders, financial assetholders, the top third of assetholders, and non-assetholders. Ultimate consumption growth from quarter t to $t+12$ is defined as in Table II from CEX data over the period June, 1982 to May, 1999, assuming a discount rate of 5% per year (quarterly discount factor $\beta = 0.95^{1/4}$). Reported are the regression coefficients, t-statistics (in parentheses) under both OLS and with standard errors corrected for residual cross-correlation and first-stage estimation error, and R^2 from the cross-section of average log excess returns on the 25 Fama-French portfolios (estimated from July, 1926 to Dec., 2004) against covariation between log returns and ultimate consumption risk. We convert the coefficient from this regression into the implied risk aversion coefficient (γ) from equation (3) and use the delta method to compute standard errors on γ. An F-test on the equality of the first-stage covariances with ultimate consumption growth across the 25 portfolios is reported. Regressions are run twice by interchanging the dependent and independent variable, in order to reduce attenuation bias resulting from measurement error in consumption:

$$(1) \quad E[r_{i,t+1} - r_{f,t+1}] + \sigma_i^2/2 = c + (\gamma - 1)Cov(r_{i,t+1}, \textstyle\sum_{s=0}^{11} \beta^s (c_{t+s+1} - c_{t+s})) + e_i$$

$$(2) \quad Cov(r_{i,t+1}, \textstyle\sum_{s=0}^{11} \beta^s (c_{t+s+1} - c_{t+s})) = c + \frac{1}{\gamma-1}(E[r_{i,t+1} - r_{f,t+1}] + \sigma_i^2/2) + u_i.$$

We also construct factor-mimicking portfolios for ultimate asset and stockholder consumption risk by projecting ultimate consumption growth on a constant and the excess returns of small growth (intersection of the smallest 40% size, lowest 40% BE/ME stocks, based on NYSE breakpoints), large growth (largest 40% size, lowest 40% BE/ME stocks), small value (smallest 40% size, highest 40% BE/ME stocks), and large value (largest 40% size, highest 40% BE/ME stocks) portfolios. This regression is estimated from June, 1982 to May, 1999 and the estimated coefficients are used to project consumption growth from July, 1926 to Dec., 2004. Regressions (1) and (2) are rerun using the projected consumption growth series over the longer period July, 1926 to Dec., 2004. Panel B reports the coefficient estimates and t-statistics (with Newey-West standard errors assuming 36 monthly lags) from the first-stage regression used to construct the consumption growth factor portfolio, CGF. Panel C reports results for aggregate nondurable per capita consumption from NIPA over the same period as the CEX sample, the entire NIPA sample period 1947 to 2004, and using an aggregate consumption CGF from 1926 to 2004.

Panel A: Euler equation estimation for stockholders and non-stockholders

	CEX 1982-1999		CGF 1926-2004		CEX 1982-1999		CGF 1926-2004		CEX 1982-1999		CGF 1926-2004	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
	Stockholders				**Top third stockholders**				**Non-stockholders**			
γ	20.26	33.59	10.39	12.67	11.00	19.68	6.40	7.52	23.11	413.09	18.22	36.50
$t(\gamma)$	(6.62)	(5.15)	(8.86)	(8.78)	(6.87)	(4.94)	(10.70)	(9.75)	(1.35)	(1.50)	(5.03)	(4.14)
$t^*(\gamma)$	(4.82)	(4.43)	(5.06)	(5.01)	(4.97)	(4.25)	(5.48)	(4.99)	(1.16)	(1.29)	(3.84)	(3.16)
R^2	0.59		0.80		0.54		0.83		0.05		0.48	
F-stat (p-value) on joint equality of first-stage covariance estimates												
	7.89 (0.003)		3.71 (0.040)		2.87 (0.077)		4.18 (0.028)		0.41 (0.667)		0.94 (0.404)	
	Assetholders				**Top third assetholders**				**Non-assetholders**			
γ	19.82	47.65	12.17	15.00	13.75	23.56	7.42	8.60	20.33	479.30	17.51	39.11
$t(\gamma)$	(4.80)	(3.54)	(8.15)	(9.89)	(6.33)	(5.39)	(11.11)	(11.11)	(1.14)	(1.20)	(4.77)	(3.33)
$t^*(\gamma)$	(4.13)	(3.05)	(5.28)	(6.41)	(4.85)	(4.64)	(5.50)	(5.50)	(0.98)	(1.03)	(3.45)	(2.41)
R^2	0.40		0.80		0.56		0.85		0.05		0.43	
F-stat (p-value) on joint equality of first-stage covariance estimates												
	7.27 (0.004)		7.16 (0.004)		2.64 (0.093)		4.96 (0.016)		0.84 (0.446)		1.39 (0.269)	

*Standard errors corrected for cross-correlation of the residuals and for first-stage estimation error.

Panel B: First-stage estimates of weights for *CGF*

	Assetholders	Top third asset	Non-assetholders	Stockholders	Top third stock	Non-stockholders	Aggregate
Small,growth	-0.51	-0.81	-0.09	-0.46	-0.89	-0.14	-0.09
	(-2.36)	(-2.69)	(-0.60)	(-2.30)	(-1.99)	(-0.95)	(-1.47)
Large,growth	0.22	0.24	0.04	0.08	0.22	0.08	-0.01
	(1.64)	(0.82)	(0.22)	(0.52)	(0.55)	(0.50)	(-0.19)
Small,value	0.53	1.05	0.06	0.58	1.19	0.09	0.15
	(1.54)	(2.57)	(0.33)	(1.87)	(2.11)	(0.46)	(1.67)
Large,value	0.02	-0.15	0.13	0.01	-0.15	0.13	0.02
	(0.08)	(-0.44)	(0.89)	(0.05)	(-0.32)	(0.87)	(0.24)

Panel C: Euler equation estimation for aggregate nondurable consumption

	NIPA, 1982-1999		NIPA, 1947-2004		*CGF*, 1926-2004	
	(1)	(2)	(1)	(2)	(1)	(2)
γ	52.42	76.76	48.95	55.71	22.77	38.08
$t(\gamma)$	(7.79)	(5.77)	(7.12)	(6.97)	(6.31)	(4.13)
$t^*(\gamma)$	(6.71)	(4.96)	(2.76)	(2.70)	(3.16)	(2.07)
R^2	0.68		0.68		0.59	
F-stat (p-value) on joint equality of first-stage covariance estimates						
	6.16 (0.007)		3.66 (0.042)		2.43 (0.110)	

Table V:

Asset and Stockholder Ultimate Consumption Risk and Expected Returns on the Entire Cross-Section of Individual Stocks

Panel A reports results from Fama and MacBeth (1973) cross-sectional regressions of the returns of all NYSE, AMEX, and Nasdaq stocks with share prices above \$5 on long-run consumption growth covariances, using CEX data, from June, 1982 to May, 1999 are reported. Covariances are estimated using a procedure similar to Fama and French (1992). We estimate covariances of returns with consumption growth using the past 120 months of quarterly overlapping log excess returns before July of year t for each stock. Stocks are then sorted in June into 3 size deciles, 3 book-to-market deciles, and 10 pre-ranking covariance deciles, and are assigned to one of 90 groups based on these rankings. The equal-weighted quarterly log excess returns on the portfolios for the next 12 months are computed from July to June and post-ranking covariance estimates are calculated using the full sample of post-ranking quarterly log excess returns of these portfolios on the full sample of consumption growth. Post-ranking covariance estimates are then assigned to each stock each June for those stocks that enter into each group. Every month the cross-section of stock returns in excess of the 90 day T-bill rate is regressed on a constant (not reported), COV_{cg}, the covariance with consumption growth, COV_{RMRF} the covariance with the excess return on the CRSP value-weighted index, the log of market capitalization (ME), the log of the book-to-market equity ratio (BE/ME), and the past year return of the stock, $ret_{-12:-2}$, skipping the most recent month. The time-series average of the monthly coefficient estimates and their time-series t-statistics are reported in the style of Fama and MacBeth (1973). Covariances with long-run consumption are estimated for assetholders, the top third of assetholders, stockholders, the top third of stockholders, and non-asset and non-stockholders from the CEX. Panel B reports results from cross-sectional month-by-month Fama and MacBeth (1973) regressions of the cross-section of excess returns on betas with respect to CGF, orthogonalized with respect to SMB, HML, and UMD. Betas are estimated similar to Fama and French (1992) using the same pre- and post-ranking procedure used to compute individual stock covariances.

Panel A: Fama-MacBeth (1973) cross-sectional regressions using CEX consumption data

	$r_{i,t} - r_{f,t} = a + \lambda_c Cov(r_{i,t+1}, \sum_{s=0}^{11} \beta^s (c_{t+s+1} - c_{t+s})) + \lambda' X_i + e_i$				
June, 1982–May, 1999	COV_{cg}	COV_{RMRF}	$ln(ME)$	$ln(BE/ME)$	$ret_{-12:-2}$
Assetholders	3.96	1.45	0.002	0.003	0.020
	(2.49)	(1.17)	(2.38)	(2.42)	(7.46)
Top asset	3.09	1.38	0.002	0.003	0.020
	(2.76)	(1.10)	(2.52)	(2.56)	(7.44)
Non-assetholders	0.90	1.42	0.002	0.003	0.019
	(0.32)	(1.14)	(2.47)	(2.47)	(7.09)
Stockholders	5.45	1.39	0.002	0.003	0.020
	(2.70)	(1.12)	(2.41)	(2.41)	(7.60)
Top stock	1.51	1.35	0.002	0.003	0.020
	(1.73)	(1.09)	(2.58)	(2.61)	(7.47)
Non-stockholders	1.44	1.40	0.002	0.003	0.018
	(0.63)	(1.12)	(2.48)	(2.50)	(6.96)

Panel B: Fama-MacBeth (1973) cross-sectional regressions using CGF portfolios

	$r_{i,t} - r_{f,t} = a + \lambda_c \beta_{i,CGF} + \lambda' X_i + e_i$				
July, 1926–Dec., 2004	β_{CGF}	β_{RMRF}	$ln(ME)$	$ln(BE/ME)$	$ret_{-12:-2}$
Assetholders	0.22	0.16	-0.06	0.28	1.01
	(2.39)	(0.89)	(-2.42)	(5.67)	(8.16)
Stockholders	0.17	0.16	-0.06	0.27	1.01
	(2.20)	(0.85)	(-2.42)	(5.58)	(8.15)
Top assetholders	0.23	0.17	-0.06	0.27	1.01
	(2.25)	(0.91)	(-2.43)	(5.60)	(8.13)
Top stockholders	0.23	0.17	-0.06	0.27	1.01
	(2.25)	(0.91)	(-2.43)	(5.60)	(8.13)

Table VI:

Factor-Mimicking Portfolios for Stockholder Ultimate Consumption: Asset Pricing Tests for the 25 Fama-French Portfolios

Panel A reports summary statistics on the consumption growth factor (CGF) mimicking portfolios for asset and stockholder ultimate consumption growth from Table IV, including their means, time-series t-statistics, percentage of months with negative returns, and their correlation with the Fama and French (1993) factor-mimicking portfolios $RMRF$, SMB, and HML, as well as the momentum factor-mimicking portfolio UMD. Panel B reports results from time-series regressions of each CGF on the four-factor model of Fama and French (1996) containing $RMRF$, SMB, HML, and UMD. Panel C reports results from time-series tests of the excess returns on the 25 Fama-French portfolios on various factor models, including the CAPM, Fama and French (1993), and consumption growth factors (CGF). The Gibbons, Ross, and Shanken (1989) F-statistic on the joint significance of the intercepts (α) from these time-series regressions are reported along with the average absolute α and R^2. Panel C also reports results from time-series tests of the Fama-French factor portfolios SMB and HML on CGF. Results are reported for the full sample of returns from July, 1926 to December, 2004 and from the period July, 1926 to December, 1981, prior to the existence of the CEX household consumption series used to construct CGF.

Panel A: Summary statistics of factor portfolios

	CGF^{asset}	CGF^{asset}_{top}	CGF^{stock}	CGF^{stock}_{top}	$RMRF$	SMB	HML	UMD
Mean	0.46	0.77	0.47	0.87	0.65	0.23	0.40	0.35
t-stat	7.30	7.43	7.17	7.36	3.64	2.05	3.43	1.56
$\% < 0$	0.38	0.37	0.38	0.37	0.40	0.49	0.45	0.36
				Correlations				
CGF^{asset}	1.00	0.97	0.97	0.96	0.59	0.13	0.61	-0.22
CGF^{asset}_{top}		1.00	0.99	0.99	0.52	0.33	0.62	-0.21
CGF^{stock}			1.00	0.99	0.55	0.30	0.70	-0.24
CGF^{stock}_{top}				1.00	0.52	0.35	0.64	-0.22
$RMRF$					1.00	0.32	0.20	-0.24
SMB						1.00	0.08	-0.09
HML							1.00	-0.27

Panel B: Four-factor regressions: $CGF_t = a + bRMRF_t + sSMB_t + hHML_t + mUMD_t + e_t$

	CGF^{asset}		CGF^{asset}_{top}		CGF^{stock}		CGF^{stock}_{top}	
	coeff.	t-stat	coeff.	t-stat	coeff.	t-stat	coeff.	t-stat
			July, 1926 – Dec., 2004					
a	0.24	6.42	0.40	6.32	0.22	6.35	0.44	6.31
b	0.18	12.12	0.21	8.17	0.14	10.59	0.24	8.39
s	-0.04	-1.40	0.16	2.80	0.07	2.31	0.21	3.22
h	0.29	9.32	0.48	8.88	0.35	12.34	0.57	9.65
m	0.01	0.92	0.02	0.91	0.01	0.83	0.02	0.88
R^2		0.61		0.57		0.68		0.60
			July, 1926 – Dec., 1981					
a	0.22	4.83	0.37	4.68	0.21	4.79	0.41	4.69
b	0.18	10.55	0.21	7.20	0.14	9.12	0.23	7.36
s	0.01	0.11	0.24	3.82	0.12	3.45	0.30	4.20
h	0.24	6.79	0.41	6.39	0.31	9.45	0.49	7.07
m	0.01	0.56	0.01	0.53	0.00	0.49	0.01	0.52
R^2		0.60		0.58		0.69		0.61

Panel C: Time-series tests $r_{i,t} - r_{f,t} = \alpha_i + \beta'_{i,F} F_t + \epsilon_{i,t}$								
	July, 1926 – Dec., 2004				**July, 1926 – Dec., 1981**			
	GRS F	p-value	avg. $\lvert\alpha\rvert$	avg. R^2	GRS F	p-value	avg. $\lvert\alpha\rvert$	avg. R^2
$RMRF$	3.18	0.000	0.21	0.77	1.99	0.019	0.19	0.79
$RMRF$, SMB	3.14	0.001	0.19	0.85	1.93	0.024	0.15	0.87
$RMRF$, HML	2.86	0.001	0.13	0.82	1.88	0.029	0.14	0.84
$RMRF$, SMB, HML	2.82	0.001	0.13	0.90	1.83	0.034	0.13	0.91
$RMRF$, CGF^{asset}	1.55	0.093	0.09	0.79	1.15	0.349	0.10	0.81
$RMRF$, CGF^{asset}_{top}	1.49	0.112	0.11	0.80	1.15	0.347	0.11	0.82
$RMRF$, CGF^{stock}	1.64	0.066	0.13	0.81	1.19	0.306	0.12	0.82
$RMRF$, CGF^{stock}_{top}	1.53	0.096	0.13	0.80	1.17	0.330	0.12	0.82
CGF^{asset}	1.57	0.085	0.13	0.34	1.19	0.307	0.17	0.42
CGF^{asset}_{top}	1.49	0.111	0.14	0.32	1.17	0.334	0.17	0.39
CGF^{stock}	1.65	0.065	0.17	0.36	1.22	0.284	0.22	0.45
CGF^{stock}_{top}	1.54	0.095	0.16	0.33	1.18	0.318	0.19	0.41

$$SMB_t, HML_t = \alpha + \beta_{CGF} CGF_t + \epsilon_t$$

	Assetholders		**Top asset**		**Stockholders**		**Top stock**	
$y =$	SMB	HML	SMB	HML	SMB	HML	SMB	HML
				July, 1926 – Dec., 2004				
α	0.12	-0.12	-0.04	-0.14	-0.01	-0.19	-0.06	-0.15
	(1.06)	(-1.30)	(-0.40)	(-1.43)	(-0.09)	(-2.18)	(-0.53)	(-1.60)
β_{CGF}	0.23	1.13	0.35	0.69	0.50	1.25	0.32	0.63
	(4.14)	(23.85)	(10.66)	(24.05)	(9.58)	(30.13)	(11.41)	(25.42)
R^2	0.02	0.38	0.11	0.38	0.09	0.49	0.12	0.41
				July, 1926 – Dec., 1981				
α	0.07	-0.09	-0.05	-0.10	-0.03	-0.15	-0.06	-0.11
	(0.58)	(-0.80)	(-0.40)	(-0.85)	(-0.23)	(-1.47)	(-0.50)	(-1.00)
β_{CGF}	0.49	1.09	0.45	0.65	0.69	1.18	0.41	0.59
	(8.27)	(20.48)	(13.90)	(20.48)	(13.24)	(25.92)	(14.63)	(21.77)
R^2	0.09	0.39	0.23	0.39	0.21	0.50	0.24	0.42

Table VII:

Aggregate Return Predictability

Predictability regression results of future stock market returns over the next 12 quarters (Panel A), earnings to price ratio (Panel B), returns on 30 year Treasury bonds (Panel C), and the yield on 30 year treasuries (Panel D) all in excess of the T-bill rate or yield, on Lettau and Ludvigson's (2002) consumption-to-wealth ratio *cay*, Julliard's (2005) expected future labor income growth rate *lr* combined with *cay*, and the ratio of ultimate consumption of assetholders, the top third of assetholders, stockholders, and the top third of stockholders, estimated from the CEX, to total aggregate consumption of nondurables and services from NIPA are reported. Regressions are estimated from June, 1982 to May, 1999. In addition, we employ annual charitable donations as a fraction of personal disposable income from NIPA from 1962 to 2002 as another proxy for wealthy consumption. Due to measurement error in consumption, we run regressions both ways by interchanging the dependent and independent variables to reduce attenuation bias. Panel E reports regression results of the log of each of the consumption series and donations on *cay* and *caylr*, as well as the correlations between them. Standard errors employ a Newey-West adjustment of 36 lags for monthly data and 12 lags for quarterly data (*cay* and *caylr*), with t-statistics reported in parentheses.

$x =$	cay	$caylr$	$\frac{c_t^{asset}}{c_t^{agg}}$	$\frac{c_t^{top}}{c_t^{agg}}$	$\frac{c_t^{stock}}{c_t^{agg}}$	$\frac{c_t^{top}}{c_t^{agg}}$	$\frac{donate_t}{y_t}$
Panel A: Stock market excess returns, $y = RMRF_{t+1:t+36}$							
y on x	14.130	0.029	0.237	-3.515	-1.292	-3.173	-0.640
	(6.66)	(10.63)	(0.25)	(-1.55)	(-0.86)	(-1.25)	(-2.32)
x on y	0.033	12.709	0.005	-0.039	-0.033	-0.030	-0.431
	(12.91)	(5.17)	(0.44)	(-4.29)	(-2.31)	(-3.46)	(-3.75)
Panel B: Stock market yields, $y = E/P_t - r_{f,t}$							
y on x	0.437	-0.001	-0.200	-0.429	-0.283	-0.465	-0.042
	(3.28)	(-3.48)	(-6.70)	(-13.08)	(-12.26)	(-12.85)	(-2.41)
x on y	0.114	-132.054	-0.806	-0.986	-1.385	-0.892	-3.060
	(2.12)	(-3.48)	(-5.22)	(-9.02)	(-9.16)	(-9.95)	(-2.41)
Panel C: Bond market excess returns, $y = TERM_{t+1:t+36}$							
y on x	1.992	0.012	-1.089	-1.382	-1.038	-1.424	-0.203
	(1.22)	(1.80)	(-3.02)	(-2.60)	(-2.83)	(-2.41)	(-1.89)
x on y	0.019	1.327	-0.152	-0.104	-0.179	-0.092	-0.490
	(2.78)	(0.85)	(-5.68)	(-4.29)	(-5.61)	(-4.31)	(-2.02)
Panel D: Bond market yields, $y = y_{10y,t} - y_{1mo,t}$							
y on x	0.282	0.159	-0.221	-0.413	-0.286	-0.448	-0.053
	(2.13)	(1.19)	(-7.70)	(-12.57)	(-12.78)	(-12.39)	(-4.26)
x on y	0.076	0.044	-0.917	-0.978	-1.441	-0.885	-5.989
	(1.80)	(1.19)	(-5.76)	(-9.07)	(-10.01)	(-9.63)	(-4.26)
Panel E: Regression of consumption share measures on cay **and** $caylr$							
cay			-1.103	-8.790	-4.936	-12.415	-4.601
			(-1.47)	(-3.28)	(-4.45)	(-3.47)	(-4.51)
correlation with cay			-0.17	-0.35	-0.46	-0.37	-0.59
$caylr$			-0.968	-8.273	-4.591	-11.377	-5.222
			(-1.26)	(-2.99)	(-3.95)	(-3.06)	(-4.88)
correlation with $caylr$			-0.14	-0.33	-0.42	-0.33	-0.53

Table VIII:

The Equity Premium and Implied Risk Aversion

Implied measures of risk aversion from the equity premium and the covariance of aggregate stock returns with ultimate consumption growth (12 quarter ahead growth discounted at a 5% annual rate, $\beta = 0.95^{1/4}$) are reported separately for the consumption of financial assetholders, the top third of assetholders, and non-assetholders from the CEX, as well as aggregate nondurable and service consumption from NIPA over the period Jan., 1959 to Dec., 2004. Since the CEX data only cover the period June, 1982 to May, 1999, we also employ the consumption growth factor-mimicking portfolios, CGF, from Table V, which cover the longer period July, 1926 to Dec., 2004. Results are reported for three periods: the period of CEX data June, 1982 to May, 1999, the period of NIPA consumption data Jan., 1959 to Dec., 2004, and the entire return period July, 1926 to Dec., 2004. Reported are the covariances of long-run consumption growth for each series with the log excess return on the market portfolio (CRSP value-weighted index of all publicly traded stocks), as well as the implied risk aversion γ for the equity premium which is calculated under the Epstein-Zin preferences of Section II as follows,

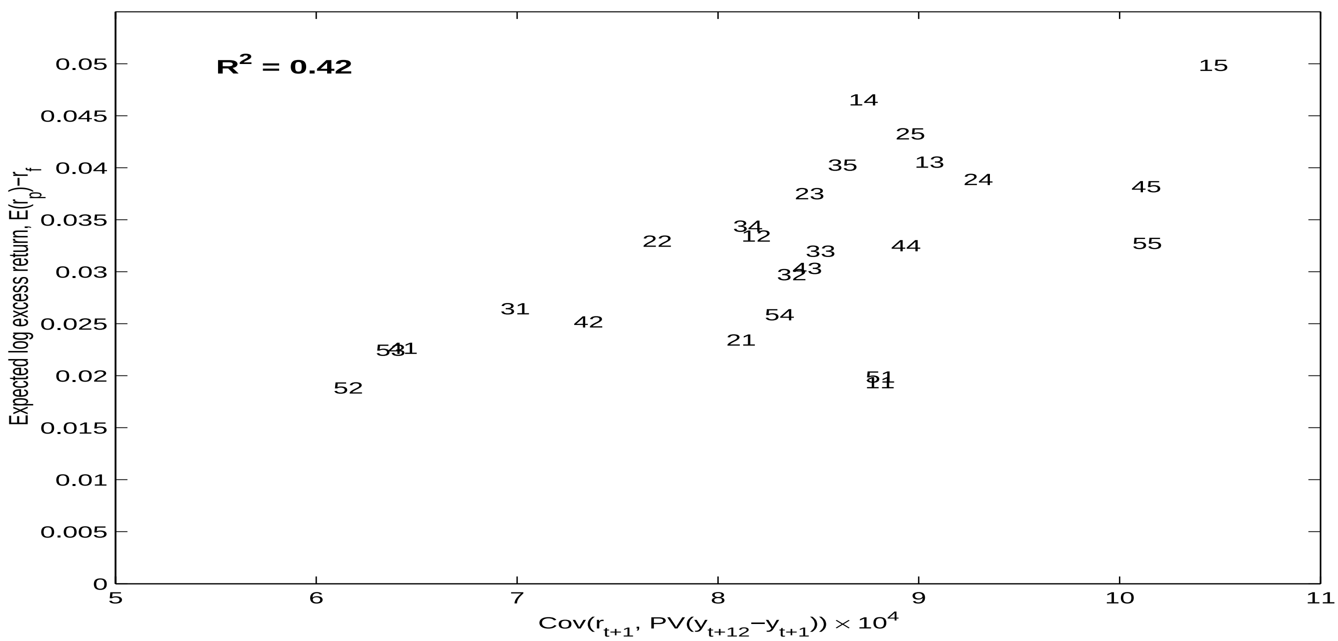
$$E[r_{mkt,t} - r_{f,t}] + \frac{\sigma^2_{mkt}}{2} = (\gamma - 1)Cov(r_{mkt,t} - r_{f,t}, \sum_{s=0}^{11} \beta^s (c_{t+s+1} - c_{t+s}))$$

where all returns and consumption are in logs, σ^2_{mkt} is the variance of the log excess return on the market, and the mean and variance of the log excess return on the market are estimated over the full period for which we have returns (July, 1926 to Dec., 2004). The table reports the risk aversion coefficient γ and a confidence interval around the point estimate for γ which is calculated via a block bootstrap method that resamples 10,000 times 12 quarters at a time constructed using the percentile method.

	CEX period	Period with aggregate consumption data	Full return period
	198206–199905	195901–200412	192607–200412
Aggregate consumption growth (from NIPA)			
Covariance with market	2.54×10^{-4}	2.56×10^{-4}	
Implied γ	78.9	78.3	
$CGF^{aggregate}$			
Covariance with market	2.61×10^{-4}	2.52×10^{-4}	9.14×10^{-4}
Implied γ (95% confidence interval)	76.8	79.8	22.7 (8.2,59.2)
Assetholder consumption growth (from CEX)			
Covariance with market	8.48×10^{-4}		
Implied γ	24.4		
CGF^{asset}			
Covariance with market	7.46×10^{-4}	6.06×10^{-4}	23.23×10^{-4}
Implied γ (95% confidence interval)	27.6	33.7	9.5 (2.7,21.8)
Top third assetholder consumption growth (from CEX)			
Covariance with market	11.83×10^{-4}		
Implied γ	17.8		
CGF^{top}			
Covariance with market	9.90×10^{-4}	8.01×10^{-4}	35.30×10^{-4}
Implied γ (95% confidence interval)	21.0	25.7	6.6 (1.5,15.3)
Non-assetholder consumption growth (from CEX)			
Covariance with market	5.42×10^{-4}		
Implied γ	37.7		
$CGF^{non-asset}$			
Covariance with market	5.46×10^{-4}	5.35×10^{-4}	15.82×10^{-4}
Implied γ (95% confidence interval)	37.3	38.1	13.5 (4.4,31.4)

Figure 1. Ultimate Labor Income Risk and Expected Returns on the 25 Fama-French Portfolios

A plot of the cross-sectional regression results of the average returns on the 25 Fama-French portfolios against the covariance of returns with the present value of 12 and 24-quarter ahead future labor income growth from Table I. The figure plots the average excess returns of the 25 portfolios (estimated from July, 1926 to Dec., 2004) against covariances of returns with the present value of future labor income growth from quarter t to $t+12$ (estimated from quarterly data from 1947 to 2004). Also reported is the R^2 from the cross-sectional regressions. The 25 Fama-French portfolios are labelled as small, growth = 1,1 ... large, growth = 5,1 ... small, value = 1,5 and large, value = 5,5.



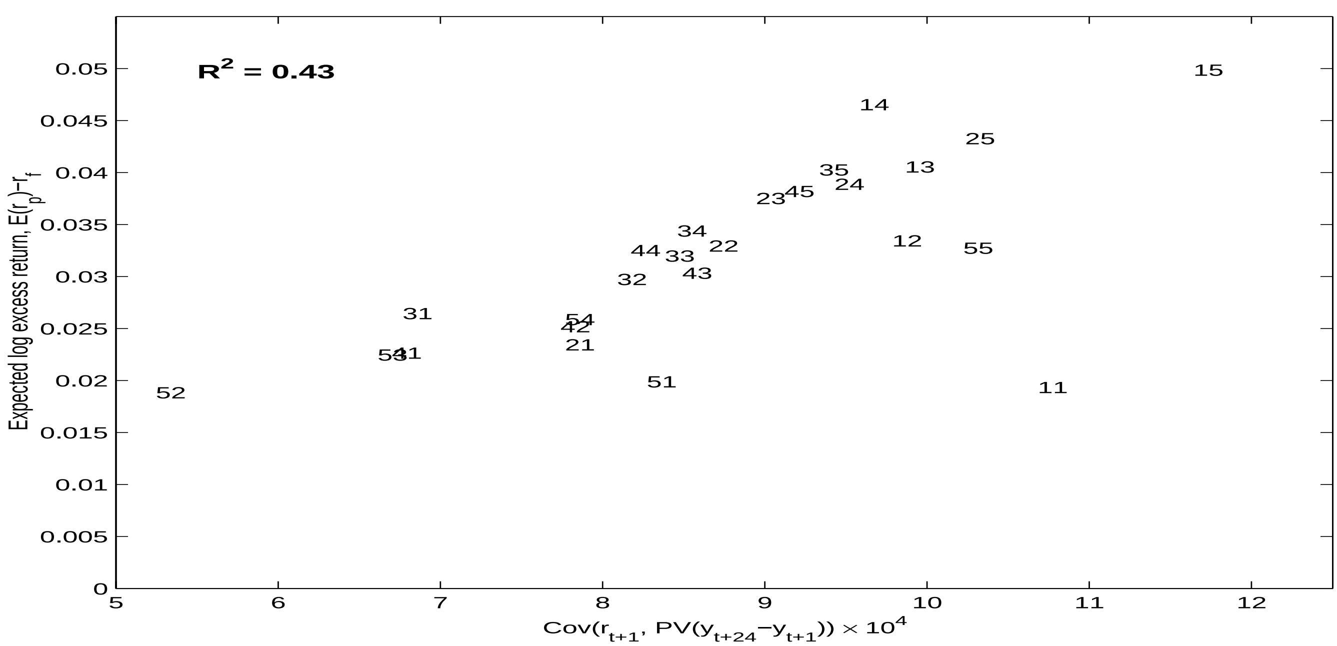


Figure 2. Plots of the Sensitivity of Asset and Non-assetholder Consumption to Aggregate Consumption and Labor Income Shocks

The plots summarize the results of Table II. The first two plots summarize the results from the linear specification of asset, non-asset, and the top third wealthiest assetholders' consumption sensitivity to labor income and aggregate consumption shocks, and the bottom two plots summarize the results from the piecewise linear specification that allows the coefficients to vary across three even categories. The predicted values of asset, non-asset, and the top third wealthiest assetholder consumption growth from both specifications are plotted against aggregate labor income and consumption shocks.

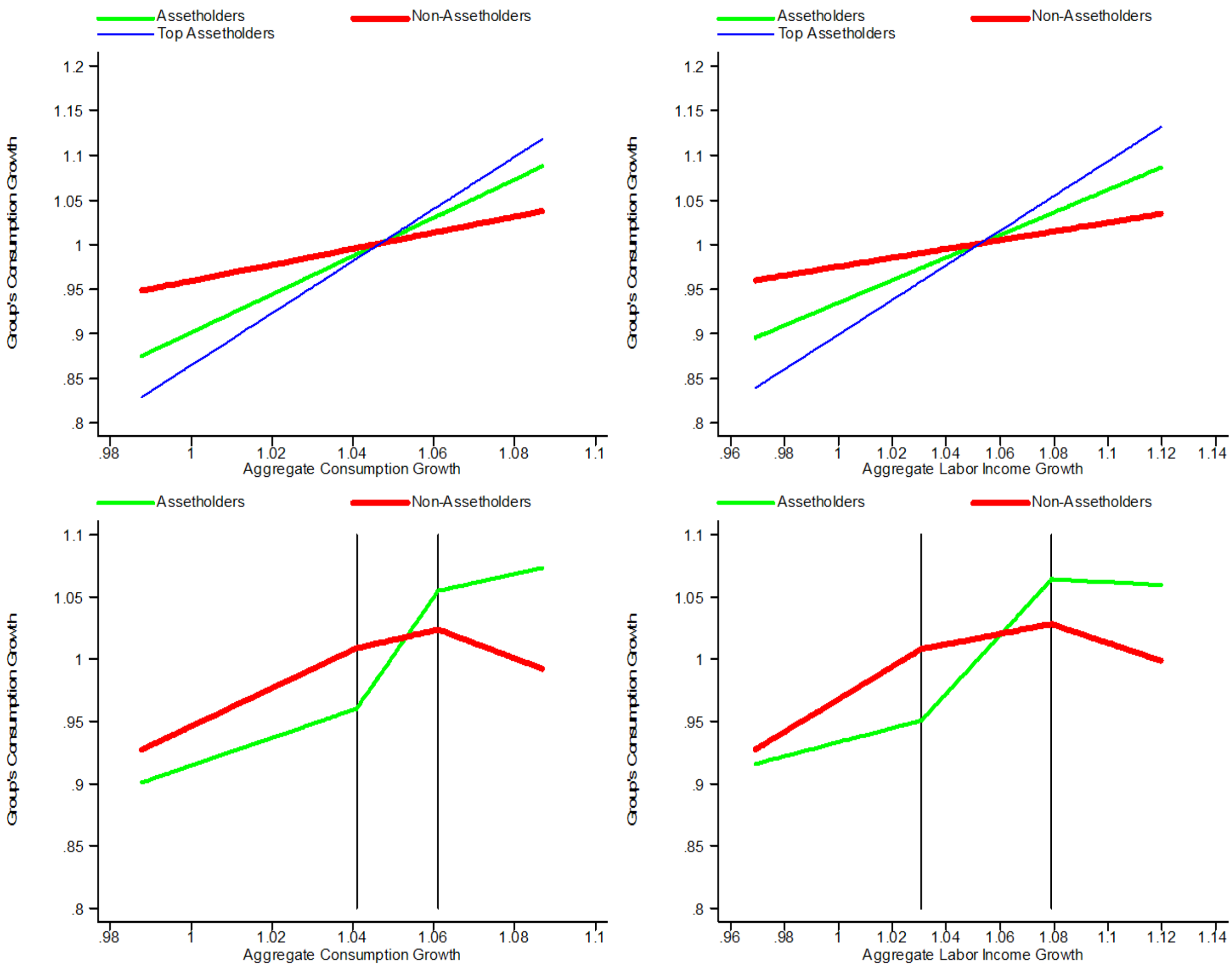


Figure 3. Asset and Stockholder Ultimate Consumption Risk and Expected Returns on the 25 Fama-French Portfolios

Plots of the average returns on the 25 Fama-French portfolios against the covariance of returns with ultimate consumption growth for assetholders, stockholders, the top third of assetholders, the top third of stockholders, and non-asset and non-stockholders from Table IV. The present value of ultimate consumption growth from quarter t to $t+12$ is defined as in Table IV over the period June, 1982 to May, 1999. The entire time-series of returns from July, 1926 to Dec., 2004 is used to estimate mean returns. Also reported is the R^2 from the cross-sectional regression. The 25 Fama-French portfolios are labelled as small, growth = 1,1 ... large, growth = 5,1 ... small, value = 1,5 and large, value = 5,5.

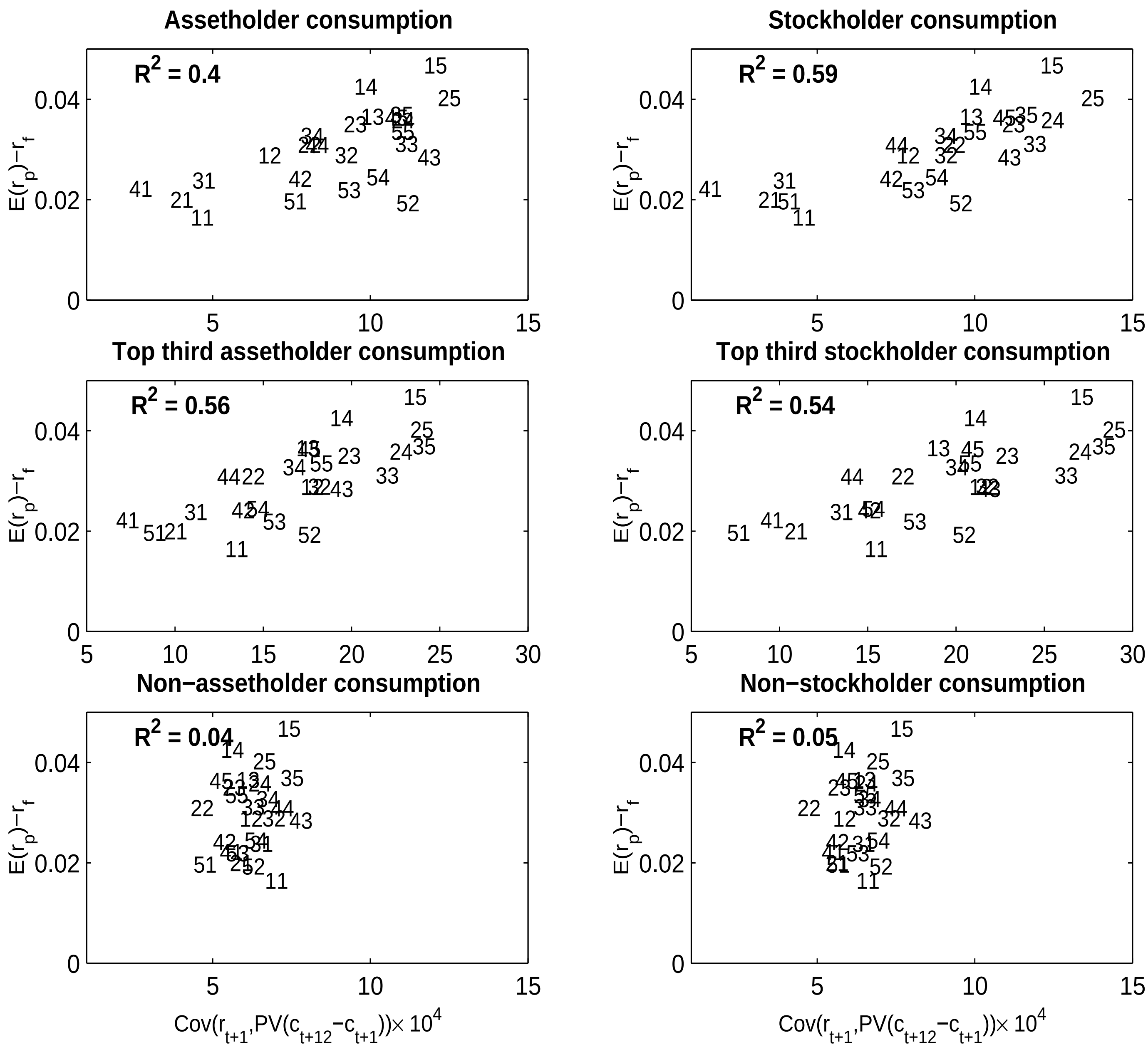


Figure 4. Consumption Growth Factor-Mimicking Portfolios and the 25 Fama-French Portfolios

Plots of the cross-sectional regression results of the average excess returns on the 25 Fama-French portfolios against the covariance of excess returns with the consumption growth factor (CGF) for assetholders, the wealthiest top third of assetholders, stockholders, and the wealthiest top third of stockholders are reported below. The construction of the factor-mimicking portfolios for consumption growth are described in Table V and provide monthly returns from July, 1926 to Dec., 2004. Also reported is the R^2 from the cross-sectional regression. The 25 Fama-French portfolios are labelled as small, growth = 1,1 ... large, growth = 5,1 ... small, value = 1,5 and large, value = 5,5.

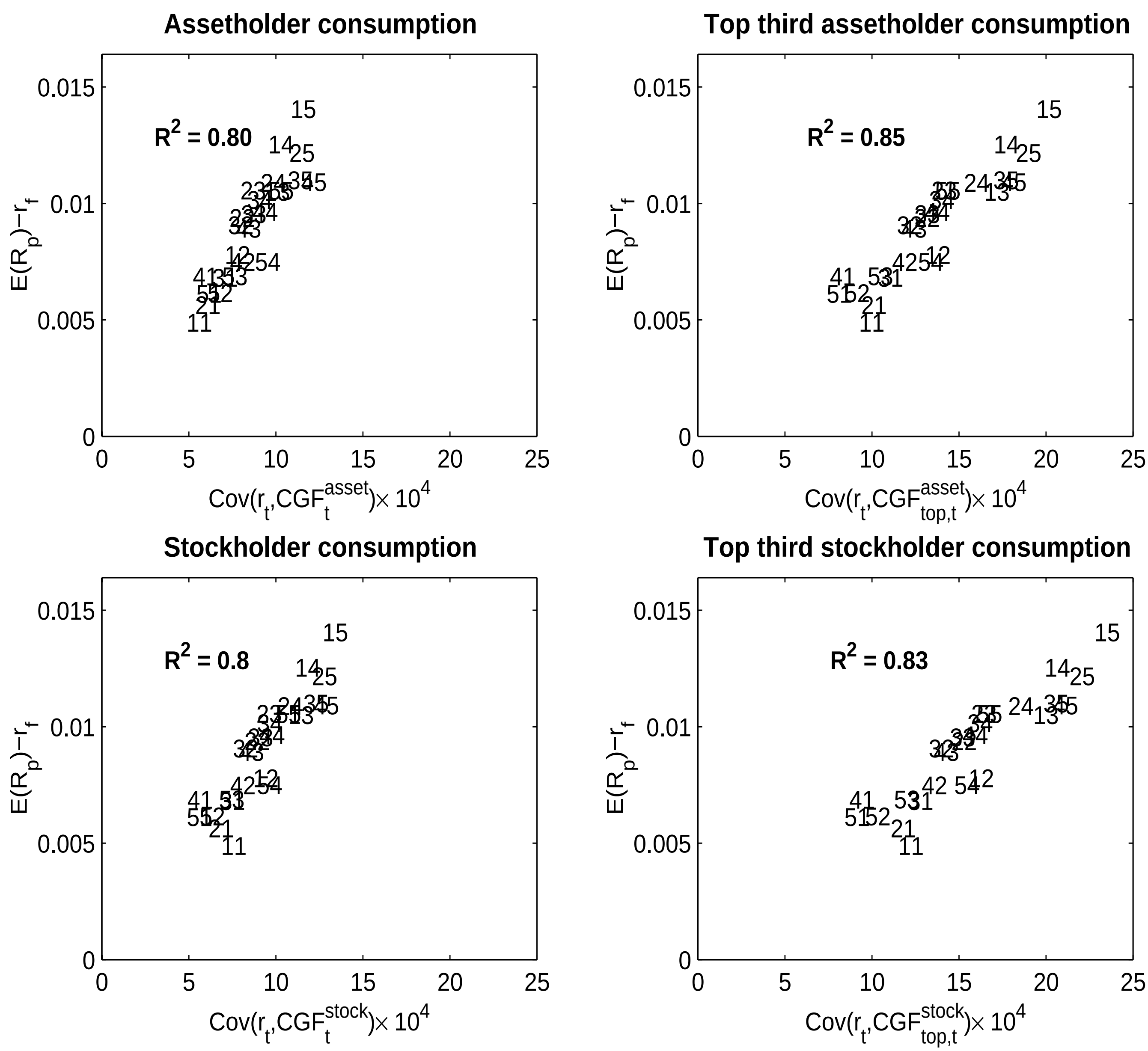


Figure 5. Plots of Consumption-to-Wealth ratio *cay* and Stockholder Consumption Share and Donations

The figure plots the consumption-to-wealth ratio of Lettau and Ludvigson (2002) against the ratio of ultimate consumption of stockholders, estimated from the CEX and described in Table II, to ultimate consumption of non-stockholders and against annual charitable donations as a fraction of personal disposable income from NIPA from 1962 to 2002 as another proxy for wealthy consumption.