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Long-Run Stockholder Consumption Risk and Asset Returns

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ABSTRACT

We focus on the long-run consumption risk of households who actually own stocks and find support for the CCAPM. The long-run consumption growth covariance of stockholders captures cross-sectional variation in average returns, including the size and value premia. To generate a longer time-series we form factor-mimicking portfolios for stockholder consumption growth that perform at least as well as the Fama-French model in asset pricing tests. The stockholder share of aggregate consumption also captures time-variation in stock and bond market returns that mirrors the dynamics of the aggregate consumption-to-wealth ratio. We interpret our findings under a model of recursive preferences and find that risk aversion as low as 5 is sufficient to match both the cross-sectional price of risk and the equity premium for the wealthiest stockholders.

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A cornerstone of asset pricing theory, the consumption CAPM, focuses on consumption risk as the key determinant of equilibrium prices. Recent evidence suggests that the consumption CAPM finds more support using long-run aggregate consumption risk to capture 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 aggregate consumption risk raises several questions. First, does the long-run consumption risk of households who actually own and trade stocks capture variation in returns, perhaps even better than aggregate consumption? Second, why does consumption take so long to adjust to news in stock returns and what are the underlying shocks driving stock returns and consumption?

In this paper, we focus on the first question. We leave aside the issue of why some households do not participate in capital markets and focus instead on the consumption patterns of those participating, for whom the Euler equation should hold. We show that the consumption risk of households who hold financial assets is particularly relevant for asset pricing. Using micro-level household data from the Consumer Expenditure Survey (CEX) for the period 1982 to 2001, we show that covariance of returns with long-run consumption growth rates of households who bear stock market risk captures the cross-sectional variation of average stock returns better than the covariance of returns with long-run aggregate or non-stockholder consumption growth. In particular, small and value stocks earn low returns when the future consumption growth of stockholders is low. A simple explanation for the high average returns observed for small and value stocks is thus that these covariance patterns induce stockholders to require a premium for holding such stocks.

As one of several possible interpretations of these findings, we adopt the theoretical framework of Hansen, Heaton, and Li (2005), who employ a special case of the recursive utility framework of Kreps and Porteus (1978), Epstein and Zin (1989), Weil (1989), and Bansal and Yaron (2004), which allow 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 long-run consumption risk of stockholders is around 8, and as low as 5 for the

[1] Under constant relative risk aversion (CRRA) preferences, investors care only about the covariance with contemporaneous consumption growth. Hence, Parker and Julliard (2004) offer several explanations for the importance of long-run consumption growth that range from measurement error in consumption to costs of or constraints on adjusting consumption as in models of staggered decision making (Lynch (1996) and Gabaix and Laibson (2002)). Recursive preferences provides a preference based role for long-run consumption risk. Other preferences that could give rise to long-run consumption growth mattering for asset prices are habit persistence (Sundaresan (1989), Constantinides (1990), Heaton (1995), Campbell and Cochrane (1999).

wealthiest third of stockholders with the largest holdings of equity. These implied risk aversion estimates are significantly smaller than those obtained from either aggregate or non-stockholder long-run consumption growth.

Since CEX data is limited to the period 1982 to 2001, and long-run risks are difficult to measure (as emphasized by Hansen, Heaton, and Li (2005)), we generate a longer time-series of data by projecting consumption growth on a set of macroeconomic variables and asset returns and use the coefficients to construct proxies for/factor-mimicking portfolios for long-run stockholder consumption growth as far back as 1926 and extending through 2004. The longer time-series serve to reduce estimation error in long-run consumption risk. Using aggregate consumption growth as an instrument, we show that long-run stockholder consumption growth is more sensitive to aggregate growth than non-stockholders, and hence requires a lower risk aversion parameter to match the moments of asset returns. In addition, a consumption growth factor-mimicking (CGF) portfolio formed from a set of size and value portfolios generates a factor that performs at least as well as the Fama and French (1993) three factor model in asset pricing tests. The CGF can explain 88 percent of the variation in average returns across the 25 Fama and French (1996) portfolios, even over the out of sample period prior to CEX data availability.

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

Finally, we examine aggregate stock and bond market returns and find that the risk aversion implied by the aggregate equity premium and its long-run stockholder consumption risk matches the magnitudes we find in the cross-section: around 9 for all stockholders and 6 for the wealthiest third of stockholders. We then examine time-series predictability in aggregate returns and find that the ratio of stockholder consumption to total consumption is negatively correlated with both stock and bond market yields and the future excess return on stocks and long-term government bonds. This result is potentially consistent with limited participation theories (Basak and Cuoco (1998) and Guvenen (2004)), where the equity and term premium are low when the stockholder consumption share (or the consumption of assetholders more generally) is high. 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 savings rates increase with wealth (Dynan, Skinner, and Zeldes (2004), Carroll (2000)) and offers a new interpretation for the empirical success of *cay*.

Our results highlight the importance of consumption risk for those who bear capital market risk in pricing assets. The fact that the Euler equation holds for those households who actually own the asset, and implies fairly reasonable risk aversion estimates, is comforting and alleviates concerns about potential spuriousness in the success of long-run consumption growth in pricing assets.

The rest of the paper is organized as follows. Section I cites the literature on consumption risk in asset pricing and asset pricing with limited participation. Section II outlines a theoretical framework with recursive preferences which links asset prices to long-run consumption risk and which we use to interpret our findings. Section III summarizes the data sources and variable construction. Section IV examines the relation between long-run stockholder consumption risk and the cross-section of returns. Section V examines the equity premium and aggregate stock and bond market return predictability, and links the stockholder consumption share to the aggregate 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 average returns and contemporaneous consumption risk (Kandel and Stambaugh (1990), Mankiw and Shapiro (1996), Breeden, Gibbons, and Litzenberger (1989), Cochrane (1996), and Lettau and Ludvigson (2001b)) and a stronger relation with long-run consumption risk (Bansal, Dittmar, and Lundblad (2004), Parker and Julliard (2004), and Hansen, Heaton, and Li (2005) consider the cross-section of stocks, while Kandel and Stambaugh (1990), 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 (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 of consumption risk and the long-run stockholder consumption risk measures we focus on here.

Our work also complements the existing literature on limited stock market participation by focusing on long-run consumption growth risk in the cross-section of stocks. Focusing on stockholders, Mankiw and Zeldes (1991) document a 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 emphasize, 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. Attanasio and Vissing-Jørgensen (2003) furthermore show (using a different methodology from that used here) 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 by using long-run stockholder consumption growth and our factor-mimicking portfolio for long-run stockholder consumption growth and focusing on the long-run stockholder consumption risk in the cross-section of stocks. Using contemporaneous consumption risk measures, Brav, Constantinides, and Geczy (2002) find that a stochastic discount factor (SDF) based on stockholder consumption is able to reconcile the value premium for low values of risk aversion, but only 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. In addition, we test whether the stockholder consumption share has forecasting power for aggregate returns. These tests are motivated by the theoretical work on limited participation by Basak and Cuoco (1998) and Guvenen (2004).

II. Asset Pricing with Recursive Preferences

Our theoretical setup follows Hansen, Heaton, and Li (2005), who adopt a set of recursive preferences following Kreps and Porteus (1978), Epstein and Zin (1989), and Weil (1989). The innovation of our study is in the empirical implementation of this setup, which provides one potential explanation for the role of long-run consumption growth in asset pricing. We use this model to compute the value of risk aversion implied by our findings. This allows us to evaluate the economic plausibility of our results. As stated previously, the recursive preference setting is just one of several possible filters for the data that allows an economic evaluation of the importance of focusing on the long-run consumpion growth of stockholders rather than that of non-stockholders or of a representative agent.

Each household has recursive 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 the elasticity of substitution equals one, following Hansen, Heaton, and Li (2005) who extend an approach suggested by Kogan and Uppal (2001) [2]. For $\sigma = 1$, the recursion 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[3]

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

[2] 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. 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.

[3] 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. We adopt their moving-average process for consumption growth, which Hansen, Heaton, and Li (2005) claim finds some support in the data. This representation is valid as long as consumption covariances are stationary. This assumption also puts a similar assumption on the dividend growth process. Hansen, Heaton, and Li (2005) also consider an alternative specification for dividend growth with a time trend and find that long-run differential responses to permanent shocks between value and growth portfolios are roughly the same under both specifications.

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

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

(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 $(E_{t+1} - E_t)\,\Sigma_{s=0}^{\infty}\beta^s\,(c_{t+1+s} - c_{t+s})$, 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 $(E_{t+1} - E_t)\,\Sigma_{s=0}^{\infty}\beta^s\,(c_{t+1+s} - c_{t+s})$ can be rewritten as $(1-\beta)\,(E_{t+1} - E_t)\,[\Sigma_{s=0}^{\infty}\beta^s c_{t+s}]$, 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

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

$\left(k_0^i\right)^2$ is the conditional variance of the log asset return, $V_t\left[r_{t+1}^i\right]$. 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, $cov_t\left(r_{t+1}^i, (E_{t+1} - E_t)\,\Sigma_{s=0}^{\infty}\beta^s\,(c_{t+1+s} - c_{t+s})\right)$. For large risk aversion, 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 main asset pricing relation used in our empirical analysis is thus,

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

In this setting, under the special case where the EIS = 1, the Euler equation (2) contains only one term: the discounted future consumption growth of the agent. 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. Hence, one motivation for examining the special case where the EIS $= 1$ is to isolate the importance of long-run consumption growth itself on asset prices, ignoring other hard to measure terms.

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 find more reasonable risk aversion estimates can match the data (in addition to aggregate market returns we also examine the cross-section of returns) when employing the long-run consumption growth of stockholders, 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.

If $\beta = 1$ (no discounting), and if we end the sum at $s = 11$ (in quarterly data) rather than ∞, then $cov_t\left(r^i_{t+1}, (E_{t+1} - E_t)\,\Sigma^{\infty}_{s=0}\beta^s\left(c_{t+1+s} - c_{t+s}\right)\right)$ reduces to $cov_t\left(r^i_{t+1}, \Delta c_{t+1} + ... + \Delta c_{t+12}\right)$, the measure of risk considered by Parker and Julliard (2004). Hence, the recursive framework 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.[4]

In addition to estimating equation (2) using aggregate and stockholder consumption growth, we also use non-stockholder consumption growth for comparison. Since non-stockholders do not hold financial assets and because non-stockholders may have an EIS $\neq 1$, the Euler equation (2) will not hold for these households. Hence, we do not treat risk aversion estimates from equation (2) for non-stockholders as valid, but merely report them for comparison with those obtained for stockholders. In addition, if non-stockholders have an EIS different from one, then the aggregate consumption-to-wealth ratio will vary through time and may be related to the relative wealth shares of stock and non-stockholders; an implication we investigate in Section V.

III. Data

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

[4]Discount rates within the range of 0-10% ($\beta = 0.9$-1) have little effect on the results.

A. Household level consumption: Stockholders and non-stockholders

We calculate separate consumption growth rates for stockholders and non-stockholders using data from the Consumer Expenditure Survey (CEX). This data set covers the period 1980Q1 to 2002Q1.

A.1. CEX consumption measure

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. 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, we compute quarterly growth rates at a monthly frequency. Financial information is gathered in the fifth quarter only. Aside from attrition, with 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. 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.[5] Nominal consumption values are deflated by the BLS deflator for nondurables for urban households. 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 and use the residual from this regression as our basic quarterly consumption growth measure. Appendix A describes the various drop criteria imposed on our CEX sample, including the omission of extreme consumption growth outliers.

[5]Results in the paper are generally similar when using total consumption, thus adding remaining services and durables to our current measure. We report results only for our nondurable and service measure of consumption for ease of comparison with the existing literature.

A.2. Stockholder status

We consider both a simple definition of stockholders based on responses to the CEX indicating positive holdings of "stocks, bonds, mutual funds and other such securities" and a more sophisticated definition of stockholders that supplements the CEX definition with a probit analysis designed to predict the probability that a household owns stocks. The motivation for this more sophisticated definition is to more accurately classify households as stock and non-stockholders. For instance, households holding bonds exclusively will be misclassified as stockholders when household are categorized based on the above category only. Using the Survey of Consumer Finances (SCF) from 1989, 1992, 1995, 1998, and 2001, which contains the entire wealth decomposition of households (including direct and indirect holdings), we estimate a probit model for whether a household owns stock on a set of observable characteristics that also exist in the CEX (age, education, race, income, holdings in checking and savings accounts, dividend income, and year). The estimated coefficients from the probit model in the SCF are then used to predict the probability of stock ownership for households in the CEX who have information on the same observable characteristics and valid responses to CEX checking and savings account questions. The details of this procedure and the probit estimates are described in Appendix A. Under the more sophisticated classification, stockholders are then defined as households who own "stocks, bonds, mutual funds, and other such securities" *and* have a predicted probability of owning stock from the probit analysis of at least 0.50. Non-stockholders are similarly defined as those responding negatively to the CEX question *and* having a predicted probability of owning stock of less than 0.50. These alternative definitions of stock and non-stockholders refine the simple CEX definition to increase confidence that each household actually holds (or does not hold) stocks. Under the simple CEX definition, we classify 77.3% (22.7%) of households as non-stockholders (stockholders). This is too high (low) relative to other sources such as the SCF, probably due to omission of indirect stockholdings in retirement plans by many CEX respondents. Under the alternative definition that includes the probit analysis, we classify 42.8% of households as non-stockholders, 13.5% as stockholders, and exclude the 43.7% of households that cannot be confidently classified. Consistent with data from other sources (such as the SCF), the fraction of stockholders increases over our sample period.

We also compute consumption separately for the wealthiest third of stockholders based on their beginning quarter dollar amount of holdings under both stockholder definitions. The top third of CEX stockholders hold a *minimum* of $26,469 in stocks in 1982 dollars under the simple stockholder definition and \$78,292 under the alternative definition. Since these are minimum values (in 1982 dollars) we are clearly capturing wealthy households.[6]

Appendix A describes the stockholder/non-stockholder classifications in detail. Our final sample contains 167,434 quarterly consumption growth observations across 62,007 households. The average number of consumption observations per month is 162 for stockholders and 551 for non-stockholders under the first definition and 96 and 307 under the alternative defintion for stockholders and non-stockholders, respectively.

A.3. 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. Since households do not exist for more than four quarters in the CEX, we cannot calculate a long-run consumption growth rate for a particular household. Instead we construct a time series of average consumption growth for a particular group of households (e.g., stockholders), and average the (log) consumption growth rates for households in that particular group. Our baseline approach computes the average growth rate for a group as follows:

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

where $c_t^{h,g}$ is the quarterly log consumption of household h in group g for quarter t, $c_{t+1}^{h,g}$ 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. The asset pricing relation from Section II for holders of asset i is then,

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

This equation is simply the asset pricing relation based on the consumption of a particular household, summed cross-sectionally across households in the group. It is important to emphasize that this cross-sectional summation simply exploits the fact that the Euler equation should hold for each stockholder at each point in time. It does not assume a representative stockholder, an assumption that would be violated in an incomplete market setting with uninsurable idiosyncratic consumption shocks.

For robustness, we also consider two other methods for computing consumption growth rates for a particular group. The first makes a representative agent assumption. Thus, rather than aggre-

[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 absolute wealthiest stockholders.

gating correctly by taking the cross-sectional average of the change in household log consumption growth rates, this approach uses the log change in the cross-sectional average of consumption growth,

$$log\left(\frac{1}{H_t^g}\sum_{h=1}^{H_t^g} C_{t+1+s}^{h,g}\right) - log\left(\frac{1}{H_t^g}\sum_{h=1}^{H_t^g} C_{t+s}^{h,g}\right)$$

where $C_t^{h,g}$ is consumption at quarter t for household h in group g. Estimating the Euler equation in (3) using this alternative consumption growth series we can assess how sensitive our results are to the representative agent assumption and hence gauge whether comovements of cross-household inequality with asset returns play a role for our results. In addition, this computation reduces the influence of large positive or negative growth rates for some individual households.

Finally, we also compute an average growth rate for group g by weighting households by their beginning period consumption,

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

where C_t denotes consumption and c_t denotes log of consumption in quarter t. This computation places more weight on households with higher levels of consumption (e.g., wealthier households) and reduces the influence of outliers that more likely come from households with extremely small beginning of period consumption. By comparing the results across all three aggregation methods we can get a sense of how important correct aggregation is and how sensitive results are to weighting wealthier households (who matter more for equilibrium pricing) more.

One concern with computing long-run growth rates for these groups of households is that the composition and attributes of households may change over time. Hence, consumption at the beginning of the period may pertain to a somewhat different set of households than consumption several quarters later. For instance, increased stock market participation over time suggests households at date t will be from a higher part of the wealth distribution than households at date $t + 12$, say. However, since results from our third (consumption-weighted) consumption growth measure are fairly similar to those from the baseline (equal-weighted) consumption growth measure, this should not be a series concern.

B. 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]

C. 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.

D. Aggregate consumption data

We use seasonally-adjusted monthly aggregate consumption of non-durables from NIPA Table 2.8.3 [line 3] available monthly beginning January, 1959 and quarterly beginning 1947Q1. 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].

[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.

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

Table I examines the relation between long-run stockholder consumption risk and the cross-section of expected stock returns by estimating the Euler equation from the model of Section II.

A. 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 plus half their variance (measured quarterly from July, 1926 to December, 2004) against the covariance of returns with long-run discounted consumption growth of 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,

$$E[r_{i,t+1}] - r_f + \frac{\sigma_i^2}{2} = \alpha + (\gamma - 1)cov\left(r_{i,t+1}, \sum_{s=0}^{11} \beta^s \left[\frac{1}{H_t^g} \sum_{h=1}^{H_t^g} \left(c_{t+1+s}^{h,g} - c_{t+s}^{h,g}\right)\right]\right) + e_i \tag{4}$$

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 $t+1$. We also estimate this regression suppressing the constant term to zero so that the model is forced to price the equity premium as well

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

Although CEX consumption data is only available from 1982 to 2002, since mean returns are notoriously difficult to estimate, we employ the entire time-series of returns from July, 1926 to December, 2004 to estimate average returns and variances on the 25 Fama-French portfolios.

We estimate regressions (4) and (5) via GMM. Appendix B derives the moment conditions and shows that the point estimate we obtain from our GMM framework is equivalent to that obtained from a standard OLS approach in which expected returns and covariances are estimated in a first stage and then used to regress estimated expected returns on covariance estimates in a second stage. We then compute standard errors under our GMM framework that account for (a) correlation of error terms across assets at a point in time, (b) estimation error in covariances, (c) serial correlation in the consumption growth series induced by overlapping consumption data, and (d) the different length of the data series used to estimate covariances and average returns.

We can also estimate the Euler equation in reverse by running the reverse regression

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

and estimating γ from the resulting estimate of $\frac{1}{\gamma-1}$. For brevity Appendix B omits the proof that GMM (with the appropriately redefined moment conditions) again is equivalent to OLS, though taking the GMM perspective again allows us to account for the four issues ((a)-(d)) discussed above. Equations (4) (and (5)) and (6) both provide consistent estimates of γ that could differ substantially in sample. Since both estimates are consistent, we report both.

By computing implied risk aversion estimates directly from the model, the 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 that the use of covariances rather than betas of consumption growth reduces the impact of measurement error in consumption since covariances are not scaled by the variance of consumption growth.

Panel A of Table I reports the results from regressions (4), (5), and (6) for each consumption series: stockholders, the top third of stockholders, and non-stockholders. For comparability we report implied risk aversion estimates (γ) and three sets of t-statistics. All three sets of t-statistics account for the different sample length used to estimate means and covariances. The first set furthermore adjusts only for cross-correlated residuals, the second in addition accounts for first-stage estimation error, and finally the third and main one additionally accounts for consumption growth autocorrelation. For regression (6), t-statistics are computed using the delta method.

The first three columns of Table I Panel A report results for stockholder consumption risk. The implied risk aversion from regression (4) is 17 with an R^2 of 0.55, suggesting that long-run stockholder consumption risk captures much of the variation in average returns across the 25 Fama-French portfolios and requires a relatively modest risk aversion coefficient. The decline in t-statistics as we account for more effects indicates the relative importance of cross-correlated errors, first-stage estimation, consumption growth autocorrelation, and different sample lengths used to compute different moments. However, the statistical significance of stockholder consumption risk remains even after accounting for all of these effects. The constant term is insignificant, however, indicating that we cannot reject that the pricing errors are different from zero, which helps support the model.

For regression (5), which forces the model to also price the equity premium by forcing the regression through the origin, we find slightly higher risk aversion of 30 is needed to explain returns. Finally, regression (6), which places consumption risk on the left-hand side of the regression, generates a similar risk aversion estimate and indicates a significant relation between stockholder consumption risk and the cross-section of expected returns.

Using the consumption risk of the top third of stockholders, the cross-sectional fit improves and implied risk aversion estimates fall to 10.0, 16.6, and 14.9 under regressions (4), (5), and (6), respectively. Hence, the covariance with consumption growth of the wealthiest stockholders, who own and trade most of the equity in the market, has an even greater impact on asset prices.

The last three columns of Table I Panel A report results for non-stockholder consumption risk. Although the Euler equation may not hold for these households (since they do not hold stocks and may have an EIS $\neq 1$), it is still interesting to examine the estimates obtained from non-stockholder consumption for comparison, even if we do not think they are valid measures of risk aversion for non-stockholders. There appears to be no significant relation between long-run non-stockholder consumption risk and expected returns. The cross-sectional R^2 is only 9 percent and the coefficients from the regressions are unreliably different from zero. The implied risk aversion coefficients from the model are not only very noisy, but their point estimates are quite large compared to those obtained for stockholders and the top third of stockholders. Under regression (4) γ is 31.3, jumps to 56.0 under regression (5), and is 345.7 under regression (6)! However, standard errors are so large that we cannot reject that these are different from zero. These results indicate much weaker covariation between returns and consumption of non-stockholders.

The results in Panel A are the second-stage regression estimates of average returns on covariances with consumption 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. The risk aversion estimate is inversely related to the dispersion in covariances, giving us a sense of economic significance and we adjust for first-stage estimation error in computing second-stage test statistics to account for the statistical significance of the covariances. For reference, Panel B of Table I reports the first-stage covariance estimates and t-statistics for each of the 25 Fama-French portfolios with long-run stockholder consumption growth using GMM standard errors. These standard errors account for autocorrelation in the long-run consumption growth series (details on this GMM procedure are available upon request). As Panel B indicates, many of the covariance estimates are at least two standard errors from zero. An F-tests testing whether they are all zero is clearly rejected at teh 1% significance level. Moreover, there appears to be wide dispersion in covariances across the portfolios. An F-test for the joint equality of the first-stage covariances across the 25 portfolios is also clearly rejected at the 1% significance level. The F-test for the joint equality of the first-stage covariances is also rejected at the 5% significance level for the top third of stockholders. However, consistent with there being little relation between non-stockholder consumption risk and returns, the first-stage covariance estimates based on non-stockholder consumption are not significantly different from each other.[8]

Panels C and D of Table I report results for the representative agent and consumption-weighted series, respectively. Aggregating households under a representative agent assumption (Panel C) improves the results, generating even lower (and significant) risk aversion coefficients for stockholders and the top stockholders, and higher risk aversion estimates for non-stockholders. For stockholders, risk aversion declines from 17 under the baseline aggregation (and using regression (4)) to 10.8 under the representative agent aggregation and from 10 to 7 for the top third of stockholders. For non-stockholders, risk aversion increases from 30 to 50 under the new aggregation. The baseline (correct) aggregation that averages log consumption growth rates not only captures the mean growth rate, but also heterogeneity in consumption growth across households. Hence, by comparing the results to those under a representative agent aggregation, which captures only the mean effect, we can gauge whether the pricing relations are driven more by mean effects or heterogeneity across households. The improvement in results under the representative agent assumption suggests that it is the covariances of returns with average consumption growth and not covariances with the dispersion in consumption growth across households (Mankiw (1986), Constantinides and Duffie (1996)) that is driving the asset pricing relations.

[8] 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, for example if survey response rates and responses are related to the performance of the market, then the coefficient estimates from these regressions will be inconsistent. 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 the regressions are remarkably similar to those using the full sample of returns (which partially overlaps with the CEX consumption data) and are 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.

One reason the representative agent series might deliver slightly better results is that it mitigates the influence of outliers by taking the log of the average rather than the average of the logs across households. Since extreme consumption growth observations are more likely to occur among households with very small consumption levels, we also employ a series that aggregates households correctly but now weights a households log consumption growth rate by the initial consumption level of the household. As Panel D of Table I shows, results for the consumption-weighted series are better than our baseline series and similar to the representative agent series, consistent with mitigating the influence of large positive or negative growth observations.

Panels E through G of Table I repeat the analysis for our alternative measure of stockholders using information from the CEX and a probit analysis from the SCF to classify stockholders and non-stockholders. The tighter definition of stockholders and the top third of stockholders improves the results. Risk aversion coefficients are lower under the more precise stockholder definitions, which are now 12 in Panel E, regression (4) for stockholders (compared to 17 in Panel A) and about 8 for the top third of stockholders (compared to 10 in Panel A). Non-stockholder consumption growth continues to exhibit a weaker relation with returns and requires much larger risk aversion parameters.

For reference, Panel H of Table I reports results for aggregate consumption from NIPA over the CEX sample period (1982 to 1999) and over the longer period (1947 to 2004) for which the NIPA series is available. Both samples yield similar results: the cross-sectional R^2 is high, but the required risk aversion ranges from 50 to 92, which is substantially higher than those obtained with stockholder (and top stockholder) consumption.

B. Robustness Across Horizons

Although we focus primarily on 12-quarter growth rates in consumption, results are robust to increasing or decreasing the horizon by several quarters. Figure 1 plots implied risk aversion coefficients and their 95% confidence bounds from estimates of regression (4) using the (correctly aggregated) consumption growth of stockholders, the top third of stockholders, non-stockholders, and aggregate consumption over various horizons S from quarter t to $t + S$. Standard errors are computed under the GMM framework of Appendix B.

As the first plot of Figure 1 shows, risk aversion estimates from stockholder consumption growth are unreliable (and negative) at horizons less than 8 quarters out. Examining growth rates at 8 quarters and beyond, risk aversion estimates are reliably different from zero and relatively stable over horizon lengths. Risk aversion of 10.7 is obtained from 8 quarter growth rates and 12.8 from 24 quarter ahead growth rates. Hence, defining long-run growth rates from 8 to 24 quarters out makes little difference. In fact, the 12 quarter consumption growth we primarily use to define long-run growth rates in this paper actually produces the highest risk aversion estimate across horizons. This evidence should alleviate concerns that the horizon length was chosen ex post to maximize fit.

For the top third stockholder consumption growth a similar pattern emerges across horizons, where the magnitude of the risk aversion coefficient declines and its precision increases at longer horizons. Once again, risk aversion is relatively stable, ranging from 5 to 10 over the 8 to 24 quarter ahead horizons, and is lower than those obtained for all stockholder consumption.

The third plot in Figure 1 shows a slightly different pattern for non-stockholder consumption growth, where point estimates of risk aversion are much higher across all horizons and become a little more reliable beyond 4 quarters. Finally, we plot the same figure for aggregate consumption growth for comparison, and find a pattern similar to that for non-stockholder consumption growth, with high risk aversion point estimates.

In sum, Figure 1 shows that there is nothing special about the 12-quarter growth rates used in Table I, but that it is important to examine long-run growth rates going out at least 8 quarters. For brevity, we focus on 12 quarter growth rates for the remainder of the paper, but emphasize that results are robust to other definitions of long-run consumption.

C. Consumption growth factor-mimicking portfolios

The CEX covers only the period 1982 to 2002, from which we try to measure long-run consumption risk. Due to the limited sample and the difficulty in estimating long-run risks, one may question the reliability of such measures and the ability to draw strong inferences. To improve estimation, we construct a longer time-series for stockholder consumption growth by projecting the present value of consumption growth on a set of instruments available over a longer period and use the coefficients to construct a longer time-series of data. Since stock return data is available for much longer periods than the CEX consumption data, we first employ the excess returns of a set of tradable assets as instruments. Specifically, we regress stockholder consumption growth on a constant and the excess returns on a set of portfolios to estimate factor portfolio weights in sample which we then use to project a time-series of returns for the factor portfolio out of sample. The resulting factor portfolio is designed to mimic long-run stockholder consumption growth over a much longer period. The longer time-series should help improve the accuracy of our findings. In addition, if measurement error in consumption is uncorrelated with the asset returns used to construct the factor-mimicking portfolio, then the factor portfolio may 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,

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

where c_t is the average log of consumption for a particular group, 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 long-run stockholder consumption growth in sample. Equation (8) is not limited to the CEX sample period if b is relatively stable over time. One caveat, of course, is that the relation between consumption growth and asset returns may not be constant over time. For instance, 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. To address this concern somewhat we also construct a CGF for the top third of stockholders (who may be more similar to stockholders in the early part of the sample) and present results for a host of CGFs using various sets of instruments for robustness. The tradeoff we face is to use actual consumption growth over a limited time-series versus instrumented consumption growth, which may not capture true consumption growth over time, over a longer time series. By presenting both sets of results, we can compare the tradeoffs.

We create the CGF over the entire period for which reliable returns data is available on the base assets. The first set of instruments or base assets we employ are the value-weighted small growth (intersection of the smallest 40% size, lowest 40% BE/ME stocks, based on NYSE break-points), 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.[9]

Regression (7) is estimated from June, 1982 to May, 1999 for which 12-quarter discounted CEX stockholder consumption growth rates are available. The results from this first-stage regression, with Newey-West t-statistics allowing for autocorrelation up to order 36 (months), are reported at the bottom of Panel A of Table II.[10]o save space Table II only shows results based on our baseline definition of stockholders and non-stockholders. Results based on the more sophisticated definition are similar and are illustrated in Figure 2. Consistent with our earlier findings, long-run consumption growth for stockholders is negatively related to small growth and positively related to small value. These relations are even stronger for the top third of stockholders and are negligible for non-stockholder consumption.

Repeating the regressions from Table I (equations (4) and (6)) by replacing actual CEX consumption growth with the factor-mimicking portfolio CGF over July, 1926 to December, 2004, Panel A of Table II shows a marked improvement in the cross-sectional fit and significance of consumption risk as well as considerably lower implied risk aversion estimates. For stockholders, the cross-sectional R^2 jumps from 55 to 80 percent when using the longer CGF data and implied risk aversion coefficients drop to between 7.5 and 9.2. For the top third of stockholders the cross-sectional R^2 increases to 82 percent and risk aversion estimates are around 5.

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 availability 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 cross-sectional regressions using covariances calculated from 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 during the CEX period are roughly the same as those in Table I that use actual consumption growth. This finding indicates that the

[9]Breeden, Gibbons, and Litzenberger (1989) justify the use of covariances with respect to a portfolio that has maximum correlation with consumption growth in place of covariances with respect to actual consumption growth in a proof where the test assets are the same as the base assets used to construct the portfolio. Technically, therefore, we should construct the CGF from all 25 Fama-French portfolios or a set of base assets that span the mean-variance frontier. Since, with a limited time-series, we cannot obtain 25 reliable coefficients, and since the four size and value portfolios approximately span the mean-variance frontier of the 25 Fama-French portfolios, we use these four portfolios as the base assets for our CGF construction.

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longer time series helps improve the model's fit rather than the use of factor portfolios per se. This improvement may be related to the covariance between asset returns and consumption growth being stronger in the early sample period according to Parker (2001).

We also report results for the CGF of non-stockholder consumption growth, obtaining a poorer fit for cross-sectional average returns and higher risk aversion estimates. The R^2 for the non-stockholder CGF is much lower than it is for stockholders, but it is still 0.52, which is much higher than that obtained from using actual non-stockholder consumption. Also, risk aversion estimates range from 15 to 28, which are lower than those obtained from actual non-stockholder consumption. Nonetheless, risk aversion estimates are still higher than those obtained from the stockholder (and especially the top stockholder) CGF. Furthermore, any significance between the CGF of non-stockholders and cross-sectional returns should be viewed with caution because the first-stage regression used to construct the CGF for non-stockholders exhibits negligible coefficients on the base assets that are not reliably different from zero. The t-statisics reported in the table are those from our GMM procedure in Appendix B, which do not account for first-stage estimation error in the CGF. To account for first-stage estimation error in the construction of the CGFs, we also run a block bootstrap simulation that resamples consumption growth, base asset return data, and test asset return data using 12-quarters blocks and then recomputes the first-stage coefficients for CGF and the risk aversion coefficient from the Euler equation that uses the bootstrapped CGF. Standard errors for γ are then computed from the observed distribution of γ values across replications and used to calculate t-statistics. This block bootstrap procedure is run using 5,000 replications and is conducted only for regression (4) and Panel A for brevity. As the table indicates, the block bootstrap t-statistics confirm our intuition based on the reliability of the first-stage estimates of each CGF. For stockholders and the top third of stockholders, risk aversion coefficients remain highly significant using the t-statistics from the bootstrap, but for non-stockholders the risk aversion estimates are not reliably different from zero. This result is not surprising given that none of the non-stockholder CGF weights are significantly different from zero. Given the noise in first-stage weights for constructing the non-stockholder CGF, the implied risk aversion could have easily been negative. In fact, the bootstrap based on non-stockholder consumption produce a negative implied risk aversion 24% of the time, compared to only 4% and 3% of the time for stockholders and the top stockholders, respectively.

We furthermore report results using a factor-mimicking portfolio constructed for aggregate consumption growth. As the last columns of Table II Panel A show, risk aversion estimates based on the aggregate consumption growth CGF are higher than those for stockholders and about the same as when using the CGF for non-stockholders.

Although the 25 Fama-French portfolios have returns dating back to July, 1926, many of the portfolios are quite thin in the early parts of the sample and contain only a handful of stocks. To check the robustness of our results we rerun the regressions in Panel A of Table II excluding return-months for which any of the 25 Fama-French portfolios contain fewer than five stocks. As Panel A of Table II shows, the cross-sectional fit improves when we exclude these return-months, generating higher R^2s for all consumption series. However, we obtain higher risk aversion estimates as well. For stockholders and the top stockholders, risk aversion estimates increase only slightly. However, for non-stockholders the risk aversion estimate almost doubles, suggesting that the non-stockholder CGF is more sensitive than the CGFs of stockholders. This finding is consistent with the simulation evidence above indicating that the non-stockholder CGF is less reliable.

Figure 2 summarizes the evidence in Table I and Panel A of Table II. The figure shows the positive relation between consumption risk and average returns for stockholders (under both stockholder definitions) and virtually no relation for non-stockholders. It is also evident from the figure that the top third of stockholders generate an even greater spread in consumption risk across the Fama-French portfolios and better fit the cross-section of average returns. The stockholder CGFs provide an even better model fit and bigger spread in covariances that translates into an even lower risk aversion estimate.

C.1. Other instruments and factor-mimicking portfolios

Panel B of Table II reports results for CGFs instrumented with macroeconomic variables in lieu of returns to tradable assets. Since aggregate consumption is available from 1947 to 2004 quarterly, we first instrument stockholder, top stockholder, and non-stockholder long-run consumption growth from the CEX using long-run aggregate consumption growth from NIPA. The sensitivity of stockholder, top stockholder, and non-stockholder long-run consumption growth, i.e. $\frac{1}{H_t^g}\sum_{h=1}^{H_t^g}\left(c_{t+1}^{h,g}-c_t^{h,g}\right)$, to aggregate long-run consumption growth, $c_{t+1}^{aggr.}-c_t^{aggr.}$, is estimated in a first-stage regression over the CEX sample period 1982 to 1999 and the coefficients from these regressions are used to generate CGFs over the longer sample period for which aggregate consumption data exists (1947 to 2004). The column labeled (A) under stockholders, top stockholders, and non-stockholders reports the risk aversion coefficient from the estimation of equation (4) using the

CGFs generated based on aggregate consumption growth. Risk aversion estimates for stockholders, top stockholders, and non-stockholders are 23.7, 13.8, and 54.7, respectively. Consistent with our previous results, risk aversion estimates are lower for stockholders and particularly the top stockholders. For reference, the first-stage estimates of the sensitivity of the actual consumption growth series from the CEX to aggregate consumption from NIPA are reported. As Panel B of Table II indicates, stockholders are more than twice as sensitive to aggregate consumption growth than non-stockholders and the top third of stockholders are more than 3.5 times as sensitive to aggregate consumption growth as non-stockholders. This heightened sensitivity to aggregate growth is precisely why we obtain much lower risk aversion estimates when using the consumption series or CGFs of stock and top stockholders. Figure 3 plots the sensitivities of each group to illustrate that stockholders bear a disproportionate amount of aggregate consumption risk relative to non-stockholders.[11]

The remainder of Panel B of Table II reports results for other CGFs formed from macroeconomic variables. Specification (B) adds the T-bill and inflation rate to aggregate consumption. Risk aversion estimates for stockholders decline to about 10 and for non-stockholders remain high at 42. Specifications (C) and (D) replace aggregate consumption growth with aggregate labor income growth (the sum of compensation of employees [line 2] and proprietors' income [line 9] from NIPA Table 1.12). The risk aversion estimates are very similar to those obtained from aggregate consumption.

Panel C of Table II reports results for CGFs constructed from other tradable assets. Column (E) reports results for CGFs created from the excess return on the market, the default spread (return on AAA-rated bonds minus T-bill rate), and the term spread (return on 30-year government bonds minus T-bill rate). Column (F) reports results for the CGF created from six industry portfolios (nondurables, durables, manufacturing, energy, shops, and utilities from Kenneth French's webpage). Column (G) reports results for the CGF created from momentum (past 12-month return) quartile portfolios. Across all CGFs, the pattern is similar: risk aversion estimates are largest for non-stockholders, smaller for stockholders, and smallest for the top third of stockholders. Furthermore, the magnitude of the risk aversion estimates are remarkably similar across the different CGFs and are similar to those obtained from the previous CGFs using macro variables

[11]Possible explanations for greater stockholder consumption sensitivity to aggregate consumption risk are that stockholders may have more risky income streams, that the tax system transfers aggregate risk from non-stockholders to stockholders, or that the bond market transfers risk to stockholders from non-stockholders, who can only use the bond market to smooth their consumption (Guvenen (2004)).

or the size and value portfolios. Most striking is the ability to create a factor from a set of industry or momentum portfolios (two sets of portfolios that have been shown to exhibit little relation with size and value premia) that can capture substantial variation in the average returns of the 25 Fama-French portfolios. A linear combination of six industry (four momentum) portfolios that is maximally correlated with stockholder consumption growth can explain almost 30 (40) percent of the cross-sectional variation in the 25 Fama-French portfolios. This evidence highlights the robustness of our results to the use of CGFs from a variety of instruments and suggests there is a common component in expected returns related to long-run stockholder consumption growth that is embedded in all these assets.

D. The entire cross-section of individual stocks

So 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 stocks in the economy. In Table III, therefore, we investigate whether long-run stockholder consumption risk helps explain the entire cross-section of all individual stock returns. This analysis helps alleviate the concern that we may be "overfitting" the 25 Fama-French test assets but potentially capturing little of the remaining variation in the cross-section of returns (see Daniel and Titman (2005)).

We estimate the covariance of returns with long-run consumption growth for each individual stock traded on the NYSE, AMEX, and Nasdaq with beginning of month share prices above $5. To reduce noise in individual stock consumption growth covariances, we follow the procedure of Fama and French (1992) and compute portfolio covariance estimates and assign these to each individual stock within the portfolio. In this procedure the individually estimated covariances are used to rank stocks and form portfolios (this is referred to as the pre-ranking step). One then computes return covariances for the constructed portfolios over the full data sample and assigns these to each stock in a particular portfolio (these covarianace are called post-ranking covariances). 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.

Specifically, for each individual stock we estimate the covariance of its returns with long-run 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 100 pre-ranking covariance centiles. We then compute the equal-weighted quarterly log excess returns on these 100 portfolios over the next 12 months, from July to June. This procedure is repeated every year, forming a time-series of returns on these 100 portfolios. We then reestimate covariances for the portfolios formed from the 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 100 portfolios due to its pre-ranking covariance changing, that stock will receive a different post-ranking covariance. This procedure reduces estimation error by shrinking individual covariance estimates to a portfolio average and employing the full sample of data.

We then 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. To assess the marginal impact of consumption risk controlling for other known determinants of returns, we also include market β, the log of market capitalization, and the log of BE/ME as regressors. The time-series average of the monthly coefficient estimates and time-series t-statistics that employ a Newey-West adjustment of 36 lags are reported in the style of Fama and MacBeth (1973).[12]

Table III Panel A indicates that covariance with long-run consumption growth of stockholders (using our baseline stockholder definition) captures significant cross-sectional variation in average returns. The univariate regression is a direct estimate of the Euler equation (4) and yields an implied risk aversion coefficient of 3.7. However, we are cautious in our interpretation of this value due to the noise in individual stock covariance estimates which may bias downward this number. Consistent with this concern, we find (in unreported results) implausibly large risk aversion values when we reverse the regression. Controlling for other known determinants of returns (market, size, and BE/ME) does not eliminate the significance of consumption risk.[13] For the top third of stockholders, results are even stronger and for non-stockholders there is little relation between returns and consumption risk. These results are consistent with those from the 25 Fama-French portfolios and provide additional evidence supporting long-run stockholder consumption risk as a

[12]A correction for first-stage covariance estimation error via Shanken (1987) has little effect on the standard errors.

[13]Unfortunately, consumption risk also does not eliminate the significance of market cap and BE/ME on the cross-section. However, it is probably heroic to expect a highly noisy estimate of individual stock long-run consumption risk to drive out precisely-measured characteristics of firms that could be capturing the same underlying risks.

significant variable explaining the entire cross-section of returns.

Panel B of Table III employs CGFs (constructed based on the small-growth, small-value, large-growth, and large-value portfolios) in place of actual CEX consumption over the longer sample period July, 1926 to December, 2004. Covariance with the CGF of stockholders carries a significant positive risk premium in the entire cross-section of stocks, even in the presence of market, size, and BE/ME premia. The evidence for the non-stockholder CGF is weaker and insignificant.

E. Asset Pricing Tests

Another useful feature of the stockholder and top stockholder CGFs is that we can use them in asset pricing tests to compare its performance against other factor-mimicking portfolios such as those of Fama and French (1993). For brevity, we employ only the CGF constructed from the four size and value portfolios. Equations (7) and (8) used to form a given 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 at least as well as the Fama and French (1993) factors SMB and HML.

Recall, that the stockholder and top stockholder CGFs place more weight on small value and small growth (with opposite sign) than on large value and large growth, and places more absolute weight on small value than small growth (Table II Panel A), compared to the Fama-French factors which are an equal dollar long and short in small and large stocks and value and growth stocks, respectively. The stockholder (and top stockholder) CGF is a linear combination of the size and value portfolios that maximizes the in sample covariance with consumption growth. This new combination, guided by the theory of Section II, has a much stronger aggregate market component than either SMB or HML. For example, the stockholder CGF exhibits a 0.58 correlation with the market, whereas HML exhibits only a 0.09 correlation with the market. Lakonishok, Shleifer, and Vishny (1994) tout the weak cyclicality of HML as evidence favoring a non-risk based explanation for the value premium. However, a slight recombination of size and value portfolios that is consistent with the consumption-based models produces a strongly cyclical factor.

Panel A of Table IV reports the mean returns and t-statistics of the CGFs 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 stockholder CGF is 51 basis points per month, and is highly statistically different from zero. For the top third of stockholders, the premium is 96 basis points per month. These premiums are a bit larger when using the more sophisticated stockholder (and top stockholder) definition. The Sharpe ratios of the CGFs (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 II Panel A. Although we focus on the CGF created from size and value portfolios, the average correlation between all of the stockholder CGFs formed from other instruments in Table II is 0.42.

We next investigate whether the CGF has any additional explanatory power over the Fama and French factors. If both models are capturing the same effects, then the CGF will be subsumed by the Fama and French model and vice versa. If, however, the consumption-based framework is correct, then the CGF may be a better way to form a factor that more closely matches the theory. Panel B of Table IV reports regression results of the CGF on the Fama and French (1996) 4-factor model, which consists of $RMRF$, SMB, HML, and UMD. The intercepts or α's from these time-series regressions are significant and are 21 and 45 basis points per month for the stockholder and top stockholder CGF, respectively. 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 similar and highly statistically significant.

Since the CGF is not fully captured by the 4-factor model, we investigate whether the CGF can capture the Fama-French factors. Panel C of Table IV reports results from time-series tests of SMB and HML on the CGF. 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 CGFs are constructed. The CGF, which itself is not fully captured by SMB and HML, appears to capture the size and value premia associated with these two factors.

Panel D of Table IV performs asset pricing tests that examine 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, a two factor model consisting of $RMRF$ and the CGF, and 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 average absolute α of 21 basis points per month across the portfolios. Adding SMB and HML, which is the Fama and French 3-factor model, improves the fit. 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 CGFs. 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 0.16 to 0.19). In part, failure to reject 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 the intercepts. The average absolute α is only 8 basis points with the stockholder or top stockholder CGF is used in conjunction with $RMRF$. Even when we remove $RMRF$ and examine the CGF by itself, which is what the theory dictates, the average absolute α's remain small (11 to 14 basis points), similar to those under the three factor model of Fama and French (1993).

For robustness, we perform the same time-series tests over the subperiod July, 1926 to December, 1981 that preceeds the availability of CEX consumption data to provide an out of sample test of the consumption growth factors, whose weights are constructed from June, 1982 to May, 1999. As the last four columns of Panel D of Table IV indicate, the value of the CGFs is apparent even in this out of sample period.

[14] As Fama and French (1996) point out, part of the reason their model rejects the F-test is that it captures a large 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.

V. Aggregate Returns

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

A. The equity premium

We first address to what extent a model emphasizing limited stock market participation and long-run consumption growth can help resolve the equity premium puzzle. In particular, we use our factor-mimicking portfolios (based on the small-growth, small-value, large-growth, and large-value portfolios) to construct a longer (estimated) time series of stockholder consumption growth to compare to previous work on limited participation by Parker (2001), Attanasio, Banks, and Tanner (2002), Brav, Constantinides, and Geczy (2002) and Vissing-Jørgensen (2002).

Table V reports implied measures of risk aversion for the equity premium across various measures of long-run consumption growth using aggregate, stockholder, top stockholder, and non-stockholder consumption. 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, obtained from the GMM framework of Appendix B. (Confidence intervals are similar using a block bootstrap method that resamples 10,000 times, 12 quarters at a time and uses the percentile method.)

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 results for the factor-mimicking portfolios over the two shorter time periods. To minimize the effect of estimation uncertainty in the equity premium, all risk aversion coefficients are calculated using the equity premium estimate from the full July, 1926 to December, 2004 period of quarterly excess log stock returns. Differences in risk aversion estimates across periods are therefore driven by differences in the covariance of the excess return on stocks with the long-run consumption growth measures. These covariances are also reported in the table.

As Table V shows, a risk aversion coefficient of about 45 is needed to explain the equity premium when using long-run aggregate consumption growth and estimating the covariance based on actual consumption data over the CEX sample period. Over the longer period of NIPA data availability, a risk aversion of around 60 is needed. Using the factor-mimicking portfolio we obtain slightly higher risk aversion estimates. However, prior to 1959, the covariance between the aggregate consumption growth CGF and market returns is higher, implying a lower estimated risk aversion coefficient of about 27.8 over the full sample from 1926 to 2004 with a confidence interval of 13.3 to 74.5.

Turning to the CEX-based long-run stockholder consumption growth, we obtain much smaller risk aversion estimates. Over the CEX sample period, we obtain a point estimate of 23.8 based on stockholder consumption growth and 44.1 based on non-stockholder consumption growth. For the top stockholders the risk aversion estimate is 19.6. 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 Euler equations being 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). In unreported results we find fairly similar risk aversion coefficients using the CEX data whether we aggregate correctly or assume a representative agent for a given group of households (e.g. stockholders).[15]

Since we do not have CEX data prior to 1982, we employ the factor-mimicking consumption growth factors, CGFs, 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 stockholders, the risk aversion estimate is as low as 5.9 with a 95% confidence interval of 3.2 to 15.1. These risk aversion estimates are similar to those found previously in the cross-section of returns. The results suggest that fairly reasonable risk aversion estimates can be obtained using the full 1926 to 2004 period and focusing on the long-run consumption risk of wealthy stockholders.

B. Stock and bond market predictability

Another implication of our limited participation framework is that variation in stockholder consumption shares should generate time-series predictability in aggregate excess stock returns. 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. This prediction holds even if stockholders and non-stockholders have identical preferences, though the effects are stronger if stockholders are less risk averse than

[15]This is not necessarily inconsistent with Parker's results which uses shorter time-horizons than we do.

non-stockholders.[16]

If non-stockholders have an EIS different from that of stockholders, as suggested by Attanasio and Vissing-Jørgensen (2002), then there may also be a link between stockholder consumption shares and the dynamics of the aggregate consumption-to-wealth ratio. The ability of stockholder consumption shares to predict excess stock returns may therefore be 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. Lettau and Ludvigson (2001a) suggest that the habit-formation framework of Campbell and Cochrane (1999) is consistent with their findings. Julliard (2005) suggests that a representative agent framework with recursive utility can explain these patterns if there is a predictable component of consumption growth (as in the model of Bansal and Yaron (2002)). We offer evidence in favor of an alternative explanation based on limited stock market participation and time-variation in the stockholder consumption share. Micro-level evidence suggests that wealthy households (e.g., stockholders) have higher savings rates (Dynan, Skinner, and Zeldes (2004), Carroll(2000), Bosworth, Burtless, and Sabelhaus (1991)). Hence, when stockholder wealth increases relative to non-stockholder wealth (or participation rises), the aggregate consumption-to-wealth ratio will decline, providing a link between stockholder wealth (or consumption) shares and the aggregate consumption-to-wealth ratio.

Table VI reports results from predictability regressions of future stock market excess returns and stock market excess 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 stockholders and the top third of stockholders, to total aggregate quarterly consumption in the CEX, calculated using CEX survey weights. Results are similar if we employ the ratio of stockholder to non-stockholder consumption instead (available upon request). For robustness, we report results for the alternative definition of stockholders using the probit analysis as well. Regressions are estimated from June, 1982 to February, 2002 for which CEX consumption data exists. We use

[16]Household responses from the Survey of Consumer Finances on questions of risk-taking behavior indicate stockholders have lower risk aversion than 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; (4) Not willing to take any financial risks. Based on their responses, non-stockholders appear more risk averse (average response = 3.5) than stockholders (average response = 2.8), who in turn are more risk averse than the top third of stockholders (average response = 2.5). The differences in responses are statistically significant, with an F-test for equal risk aversion across groups clearly rejected at less than the 1% significance level.

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. Standard errors used to compute t-statistics employ a Newey-West adjustment of 36 lags when consumption ratios are on the right-hand side of the regression, and 3 lags when they are on the left-hand side of the regression.

Panel A of Table VI reports results for predicting 12 quarter ahead stock market returns in excess of the Treasury bill rate. 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 stockholder consumption shares, we find a negative but statistically modest relation between stockholder consumption shares and future market returns. Reversing the regression to mitigate measurement error in consumption, we find a statistically significant negative relation that is stronger for the wealthiest 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 scaling donations by aggregate consumption) as another proxy for the wealth share of stockholders. This ratio also exhibits a significant negative relation with future stock returns.

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 of Table VI. Yields are forward looking and may have less noise than returns. As Panel B shows, stock market excess yields are positively related to *cay* and *caylr* and negatively related to the consumption shares of stockholders.

Panels C and D of Table VI examine the predictability of aggregate bond excess returns and yields (over the T-bill rate) to help alleviate data mining concerns. This is motivated by the likely large overlap between the set of households who hold stocks and the set that holds bonds. As Panel C indicates, *cay* and *caylr* are positively related to bond market excess returns, defined as the difference in returns between long-term (10-year) government bonds and the T-bill rate. Stockholder consumption shares, including donations, are strongly negatively related to aggregate bond excess returns. Panel D shows similar results for bond market excess yields.

Finally, Panel E reports regression results of stockholder consumption shares and donations on *cay* and *caylr*, as well as the correlations between them. The relation between the stockholder consumption share and the consumption-to-wealth ratio is strongly negative. Figure 4 plots the consumption-to-wealth ratio against stockholder consumption shares (and donations), highlighting the strong negative relation between the two. This evidence suggests that the consumption-to-wealth ratio may be linked to stock market participation, possibly providing an economic story for its empirical success in pricing assets.

VI. Conclusion

We find empirical support for consumption-based asset pricing by focusing on the long-run consumption risk of stockholders. In particular, long-run stockholder consumption risk captures the return premia associated with size and value. This evidence supplements recent findings of long-run aggregate consumption risk having explanatory power for asset prices. The fact that the consumption 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. Analysis of aggregate returns furthermore suggests that the stockholder consumption share is associated with time-varying risk premia in aggregate stock and bond markets.

Understanding further why stockholder consumption growth responds slowly to news in asset returns will improve our understanding of what is driving these long-run relations. The fact that stockholders are more sensitive to aggregate consumption movements helps explain why the consumption risk of stockholders delivers lower risk aversion estimates from the cross-section of stock returns. The predictability results suggest there is also substantial variation over time in stockholder consumption shares, possibly induced by preference differences and differences in sensitivity to aggregate shocks. One possible direction for future inquiry is to examine the potential link between long-run consumption growth, asset returns, and macroeconomic shocks. Evidence from the macroeconomic growth and real business cycle literature shows that consumption and production respond slowly to technology shocks with full effects reached at at 10-15 quarter horizon (see Altig, et. al (2005)), which is about the same horizon consumption is responding to asset returns.

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Appendix A: Details of CEX Sample Choice and Stockholder Definitions

We describe our CEX sample criteria and stockholder definitions.

CEX sample choice

For each household we calculate quarterly consumption growth rates based on 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 consumption 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 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 4 months of 1985 and 1995 and the first 9 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 (with a resulting longer period of missing long-run consumption growth rates) 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.

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 for our simplest and baseline definition. 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.[17] [18]

It is known from the Survey of Consumer Finances (SCF), 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. This may indicate that some CEX households with stockholdings in pension plans do not report these, leading them to be miscategorized as non-stockholders. We are unable to address this issue, and our results are thus conservative in that they likely would strengthen with a cleaner separation of stockholders and non-stockholders.

A complementary issue which can be addressed is that households holding bonds or bond funds exclusively will be misclassified as stockholders when household are categorized based on the above category only. We therefore also consider a more sophisticated definition of who likely holds stocks using a probit analysis from another data source, the SCF, to predict the probability that a household owns stocks. Using the SCF from 1989, 1992, 1995, 1998, and 2001, which contains the entire wealth decomposition of households (including direct and indirect holdings), we estimate the following probit model for whether a household owns stocks on a set of observable characteristics that also exist in the CEX: age of household, age squared, an indicator for at least 12 but less

[17]1,530 households in our final sample of 62,007 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. 128 households report an increase in their holdings of stocks, bonds or mutual funds larger than the value of the reported end of period holdings. We classify these as non-stockholders.

[18]For interviews conducted from October 1990 to March 1996, 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 amount of stockholding, we exclude them in estimations that are based on the wealthiest third of stockholders (by stockholdings). Thus, the total number of stockholders used to define the wealthiest third of stockholders is smaller than the total number of stockholders.

than 16 years of education for head of household (*highschool*), an indicator for 16 or more years of education (*college*), an indicator for race not being white/caucasion, year dummies, the log of real total household income before taxes, the log of real dollar amount in checking and savings accounts (set to zero if checking and savings = 0), an indicator for checking and savings accounts = 0, and an indicator for positive dividend income, plus a constant.[19] Data are averaged across SCF imputations and SCF weights are employed in the probit model to avoid the estimates being unduely influenced by the oversampling of high wealth individuals in the SCF. The estimates of the coefficients from the probit model in the SCF are then used to predict the probability of stock ownership for households in the CEX who have information on the same observable characteristics and valid responses to checking and savings account questions. The estimated probit model coefficients and t-statistics are

$$Prob(Stockholder) = \Phi(x'b)$$

$$\begin{array}{lllllllllll}
x'b & = & \begin{array}{l}-7.457\\(-46.35)\end{array} & + & \begin{array}{l}0.0297age\\(7.44)\end{array} & + & \begin{array}{l}-0.0004age^2\\(-9.21)\end{array} & + & \begin{array}{l}0.3102highschool\\(9.86)\end{array} & + & \begin{array}{l}0.5160college\\(14.23)\end{array}\\
 & + & \begin{array}{l}-0.2594nonwhite\\(-9.56)\end{array} & + & \begin{array}{l}0.2508y_{1992}\\(7.11)\end{array} & + & \begin{array}{l}0.3795y_{1995}\\(10.91)\end{array} & + & \begin{array}{l}0.6299y_{1998}\\(18.20)\end{array} & + & \begin{array}{l}0.6575y_{2001}\\(19.19)\end{array}\\
 & + & \begin{array}{l}0.5513log(income)\\(34.35)\end{array} & + & \begin{array}{l}0.0747log(chk+sav)\\(10.20)\end{array} & + & \begin{array}{l}0.1067(chk+sav=0)\\(1.72)\end{array} & + & \begin{array}{l}1.2438(div>0)\\(36.14)\end{array} & &
\end{array}$$

where the pseudo R^2 from the first-stage probit model in the SCF is 0.32.

When calculating a household's predicted probability of stock ownership in the CEX we use the 1992 dummy coefficient for the years 1990-1993, the 1995 dummy coefficient for 1994-1996, the 1998 dummy coefficient for 1997-1999, and the 2001 dummy coefficient from 2000 onward. We then define stockholders under the more sophisticated definition as the intersection of households who respond positively to holdings of "stocks, bonds, mutual funds and other such securities"and have a predicted probability of owning stocks of at least 0.50. Non-stockholders are similarly defined as those responding to negatively to the CEX question *and* having a predicted probability of owning stocks less than 0.50. These alternative definitions of stock and non-stockholders refine the CEX classification.

In addition, the set of stockholders (under both definitions) 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 with those for the top third of stockholders. [20]

Appendix B: GMM estimation

This appendix first shows the equivalence, in terms of the point estimate for risk aversion, between GMM and a cross-sectional OLS regression in our setting. It then uses the GMM framework to

[19]Note that we include checking and savings account holdings and not total financial wealth in the probit because of the suspected underreporting of indirect financial wealth holdings in the CEX discussed above.

[20]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 holdings 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 amount of stockholdings, we exclude them when defining the top third of stockholders.

derive standard errors that account for (a) correlation of error terms across assets in a given time period, (b) estimation error in covariances, (c) autocorrelation in our consumption series due to the use of overlapping consumption growth data, and (d) the different length of the data series used to estimate covariances and average returns.

To save space we show the derivations here for the Euler equation written in the form of equation (4). The derivations follow in a similar manner when the Euler equation is written in the form of equation (6) with GMM moment conditions reorganized correspondingly.

A. The GMM setup

The conditional Euler equation from the Epstein-Zin setting is given by

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

Assume V_t and cov_t do not vary over time. Then the corresponding unconditional Euler equation is

$$\begin{aligned} E\left(r_{t+1}^i - r_{t+1}^f\right) + \frac{1}{2}V\left(r_{t+1}^i\right) &\simeq (\gamma - 1)\, cov\left(r_{t+1}^i, \Sigma_{s=0}^{\infty}\beta^s\left(c_{t+1+s} - c_{t+s}\right)\right) \\ &= (\gamma - 1)\, E\left(r_{t+1}^i \varepsilon_{c,t+1}\right). \end{aligned}$$

where $\varepsilon_{c,t+1} = \Sigma_{s=0}^{\infty}\beta^s\left(c_{t+1+s} - c_{t+s}\right) - \mu_{\Sigma_{s=0}^{\infty}\beta^s(c_{t+1+s}-c_{t+s})}$.

In moment form, adding the moment for $\varepsilon_{c,t+1}$ to allow estimation of $\mu_{\Sigma_{s=0}^{\infty}\beta^s(c_{t+1+s}-c_{t+s})}$,

$$\begin{bmatrix} 0_{Nx1} \\ 0 \end{bmatrix} = E\begin{bmatrix} r_{t+1} - r_{t+1}^f + \frac{1}{2}\sigma^2 - (\gamma - 1)\, r_{t+1}\varepsilon_{c,t+1} \\ \varepsilon_{c,t+1} \end{bmatrix} = E\begin{bmatrix} g_t^r \\ g_t^m \end{bmatrix}.$$

N is the number of stocks (25 in our setting using the 25 FF test assets), g_t^r and g_t^m are defined by the last equality, and

$$r_{t+1} = \begin{bmatrix} r_{t+1}^1 \\ \dots \\ r_{t+1}^N \end{bmatrix}, \quad \sigma^2 = \begin{bmatrix} V\left(r_{t+1}^1\right) \\ \dots \\ V\left(r_{t+1}^N\right) \end{bmatrix}.$$

For simplicity we assume σ^2 is known. Since σ^2 in practice can be estimated with high accuracy any biases resulting from this should be minimal.

In order to use all available data, we use a different sample length T for the two parts of $E\left(g_t^r\right)$. We use a sample of length T_1 for estimating mean excess returns (adjusted for the variance term) $E\left(r_{t+1} - r_{t+1}^f + \frac{1}{2}\sigma^2\right)$ and a subsample of length T_2 for estimating covariances $E\left(r_{t+1}\varepsilon_{c,t+1}\right)$. We also use the subsample of length T_2 for estimating $E\left(g_t^m\right)$. When estimating equation (4) we include an intercept term α, identical for all stocks, in the first set of moment conditions. This is done following Parker and Julliard (2005) to provide estimates of risk aversion that best fit the cross-section of average returns, without the additional constraint of also fitting the level of these returns (and thus the equity premium). Equation (5) omits this intercept term.

The GMM objective function is

$$\min_{\gamma, \mu_{\Sigma_{s=0}^{\infty}\beta^s(c_{t+1+s}-c_{t+s})}, \alpha} g'Wg$$

with

$$g=\left[\begin{array}{c} g^r \\ g^m \end{array}\right]=\left[\begin{array}{c} \frac{1}{T_1}\Sigma_{t=1}^{T_1}\left(r_{t+1}-r_{t+1}^f+\frac{1}{2}\sigma^2\right)-\alpha 1-(\gamma-1)\frac{1}{T_2}\Sigma_{t=1}^{T_2}r_{t+1}\varepsilon_{c,t+1} \\ \frac{1}{T_2}\Sigma_{t=1}^{T_2}\varepsilon_{c,t+1} \end{array}\right].$$

where 1_N is a $Nx1$ vector of ones. The estimation picks three parameters to fit 26 moments as best possible. We use a pre-specified weighting matrix

$$W=\left[\begin{array}{cc} I_{25} & 0 \\ 0 & h \end{array}\right]$$

rather than efficient GMM. This is done to give each of the 25 portfolios equal importance in the estimation as opposed to downweighting portfolios with less precisely estimated average returns. Following Parker and Julliard (2005) we set h to a sufficiently large number that the estimate of $\mu_{\Sigma_{s=0}^{\infty}\beta^s(c_{t+1+s}-c_{t+s})}$ approximately equals $\frac{1}{T_2}\Sigma_{t=1}^{T_2}\Sigma_{s=0}^{\infty}\beta^s\left(c_{t+1+s}-c_{t+s}\right)$ and that further increases in h has only minimal effects on the estimates of γ and α.

B. Equivalence of GMM and OLS

The first order conditions for the GMM minimization are $(\nabla g)' Wg=0$, i.e.

$$\underbrace{\left[\begin{array}{cc} \left(\frac{\partial g^r}{\partial\gamma}\right)' & \frac{\partial g^m}{\partial\gamma} \\ \left(\frac{\partial g^r}{\partial\mu}\right)' & \frac{\partial g^m}{\partial\mu} \\ \left(\frac{\partial g^r}{\partial\alpha}\right)' & \frac{\partial g^m}{\partial\alpha} \end{array}\right]}_{3\times 26}\underbrace{\left[\begin{array}{cc} I_{25} & 0 \\ 0 & h \end{array}\right]}_{26\times 26}\underbrace{\left[\begin{array}{c} g^r \\ g^m \end{array}\right]}_{26\times 1}=0_{3x1}$$

$\Longleftrightarrow$

$$\left[\begin{array}{cc} \left(\frac{\partial g^r}{\partial\gamma}\right)' I_{25} & 0_{1x1} \\ \left(\frac{\partial g^r}{\partial\mu}\right)' I_{25} & -h \\ \left(\frac{\partial g^r}{\partial\alpha}\right)' I_{25} & 0_{1x1} \end{array}\right]\left[\begin{array}{c} g^r \\ g^m \end{array}\right]=0$$

$\Longleftrightarrow$

$$\left[\begin{array}{c} \left(\frac{\partial g^r}{\partial\gamma}\right)' g^r \\ \left(\frac{\partial g^r}{\partial\mu}\right)' g^r-hg^m \\ \left(\frac{\partial g^r}{\partial\alpha}\right)' g^r \end{array}\right]=0$$

$\Longleftrightarrow$

$$\left[\begin{array}{c} \left(-\frac{1}{T_2}\Sigma_{t=1}^{T_2}r_{t+1}\varepsilon_{c,t+1}\right)'\left(\frac{1}{T_1}\Sigma_{t=1}^{T_1}\left(r_{t+1}-r_{t+1}^f+\frac{1}{2}\sigma^2\right)-\alpha 1_N-(\gamma-1)\frac{1}{T_2}\Sigma_{t=1}^{T_2}r_{t+1}\varepsilon_{c,t+1}\right) \\ (\gamma-1)\frac{1}{T_2}\Sigma_{t=1}^{T_2}r_{t+1}'\left(\frac{1}{T_1}\Sigma_{t=1}^{T_1}\left(r_{t+1}-r_{t+1}^f+\frac{1}{2}\sigma^2\right)-\alpha 1_N-(\gamma-1)\frac{1}{T_2}\Sigma_{t=1}^{T_2}r_{t+1}\varepsilon_{c,t+1}\right)-h\frac{1}{T_2}\Sigma_{t=1}^{T_2}\varepsilon_{c,t+1} \\ (-1_N)'\left(\frac{1}{T_1}\Sigma_{t=1}^{T_1}\left(r_{t+1}-r_{t+1}^f+\frac{1}{2}\sigma^2\right)-\alpha 1_N-(\gamma-1)\frac{1}{T_2}\Sigma_{t=1}^{T_2}r_{t+1}\varepsilon_{c,t+1}\right) \end{array}\right]=0.$$

The third row implies

$$\begin{aligned}\widehat{\alpha} &= \frac{1}{N}\Sigma_{i=1}^{N}\left(\frac{1}{T_1}\Sigma_{t=1}^{T_1}\left(r_{t+1}^i-r_{t+1}^f+\frac{1}{2}\sigma_i^2\right)\right)-(\widehat{\gamma}-1)\frac{1}{N}\Sigma_{i=1}^{N}\left(\frac{1}{T_2}\Sigma_{t=1}^{T_2}r_{t+1}^i\widehat{\varepsilon}_{c,t+1}\right) \\ &= \frac{1}{N}\Sigma_{i=1}^{N}\left(y_i-(\widehat{\gamma}-1)x_i\right)\end{aligned}$$

where $y_i = \frac{1}{T_1}\Sigma_{t=1}^{T_1}\left(r_{t+1}^i - r_{t+1}^f + \frac{1}{2}\sigma_i^2\right)$ is the average excess return on asset i (adjusted for the variance term) and $x_i = \frac{1}{T_2}\Sigma_{t=1}^{T_2} r_{t+1}^i \varepsilon_{c,t+1}$ is the covariance of asset i's return with the long-run consumption growth measure.

The first row implies

$$\begin{aligned}
\widehat{\gamma} - 1 &= \frac{\Sigma_{i=1}^N\left(\frac{1}{T_2}\Sigma_{t=1}^{T_2}\left(r_{t+1}^i\widehat{\varepsilon}_{c,t+1}\right)\right)\left(\frac{1}{T_1}\Sigma_{t=1}^{T_1}\left(r_{t+1}^i - r_{t+1}^f + \frac{1}{2}\sigma_i^2 - \widehat{\alpha}\right)\right)}{\Sigma_{i=1}^N\left(\frac{1}{T_2}\Sigma_{t=1}^{T_2}\left(r_{t+1}^i\widehat{\varepsilon}_{c,t+1}\right)\right)^2} \\
&= \frac{\Sigma_{i=1}^N\left(x_i - \overline{x}\right)\left(y_i - \overline{y}\right)}{\Sigma_{i=1}^N\left(x_i - \overline{x}\right)^2}
\end{aligned}$$

where the last equality follows after substituting in the above expression for $\widehat{\alpha}$. Thus the GMM estimates of α and $(\gamma - 1)$ are identical to the OLS estimates these parameters obtained from a cross-sectional OLS regression (with N data points) of average excess returns (adjusted for the variance term) on sample covariances.

C. GMM standard errors

Following Newey and McFadden (1994), Theorem 3.4, the asymptotic distribution of the GMM estimator is (under appropriate regularity conditions) given by

$$\sqrt{T_1}\left(\begin{bmatrix} \widehat{\gamma} \\ \widehat{\mu}_{\Sigma_{s=0}^{\infty}\beta^s(c_{t+1+s}-c_{t+s})} \\ \widehat{\alpha} \end{bmatrix} - \begin{bmatrix} \gamma \\ \mu_{\Sigma_{s=0}^{\infty}\beta^s(c_{t+1+s}-c_{t+s})} \\ \alpha \end{bmatrix}\right) \rightarrow^{\mathrm{d}} N\left(0, \left[G'WG\right]^{-1} G'W\Omega WG\left[G'WG\right]^{-1}\right)$$

where $G = E\left[\nabla g\right]$ and $\sqrt{T_1}\left[g\right] \rightarrow^{\mathrm{d}} N\left(0, \Omega\right)$. Since

$$\nabla g = \begin{bmatrix} \frac{\partial g^r}{\partial \gamma} & \frac{\partial g^r}{\partial \mu} & \frac{\partial g^r}{\partial \alpha} \\ \frac{\partial g^m}{\partial \gamma} & \frac{\partial g^m}{\partial \mu} & \frac{\partial g^m}{\partial \alpha} \end{bmatrix} = \begin{bmatrix} \left(-\frac{1}{T_2}\Sigma_{t=1}^{T_2} r_{t+1}\varepsilon_{c,t+1}\right) & \left((\gamma - 1)\frac{1}{T_2}\Sigma_{t=1}^{T_2} r_{t+1}\right) & -1_N \\ 0 & -1 & 0 \end{bmatrix},$$

G is estimated by

$$\widehat{G} = \begin{bmatrix} \left(-\frac{1}{T_2}\Sigma_{t=1}^{T_2} r_{t+1}\widehat{\varepsilon}_{c,t+1}\right) & \left((\widehat{\gamma} - 1)\frac{1}{T_2}\Sigma_{t=1}^{T_2} r_{t+1}\right) & -1_N \\ 0 & -1 & 0 \end{bmatrix}.$$

Ω is a $26x26$ matrix:

$$\Omega = E\left[T_1 g g'\right] = E\left[T_1\begin{pmatrix} g^r g^{r\prime} & g^r g^{m\prime} \\ g^m g^{r\prime} & g^m g^{m\prime} \end{pmatrix}\right] = \begin{bmatrix} \Omega^{rr} & \Omega^{rm} \\ \Omega^{mr} & \Omega^{mm} \end{bmatrix}.$$

Ω^{mm} can be rewritten as follows

$$\Omega_{1x1}^{mm} = E\left(T_1\frac{1}{T_2}\Sigma_{t=1}^{T_2}\varepsilon_{c,t+1}\frac{1}{T_2}\Sigma_{t=1}^{T_2}\varepsilon_{c,t+1}'\right) = \frac{T_1}{T_2}E\left(\frac{1}{T_2}\Sigma_{t=1}^{T_2}\varepsilon_{c,t+1}^2 + \Sigma_{l=1}^{L}\frac{1}{T_2}\Sigma_{t=l+1}^{T_2}\left[\varepsilon_{c,t+1}\varepsilon_{c,t+1-l} + \varepsilon_{c,t+1-l}\varepsilon_{c,t+1}\right]\right)$$

where L is the highest order of autocorrelation induced by the use of overlapping consumption growth data (L=36 when using 12-quarter discounted consumption growth rates, available at the monthly frequency).

Define now $z_{t+1}=r_{t+1}-r^f_{t+1}+\frac{1}{2}\sigma^2-\alpha$ and $w_{t+1}=r_{t+1}\varepsilon_{c,t+1}$ (where z is assumed i.i.d, while w is autocorrelated). Then we can write Ω^{rm} as

$$\begin{aligned}
\Omega^{rm}_{25x1} &= E\left(T_1\left(\frac{1}{T_1}\Sigma_{t=1}^{T_1}\left(r_{t+1}-r^f_{t+1}+\frac{1}{2}\sigma^2-\alpha_S\right)-(\gamma-1)\frac{1}{T_2}\Sigma_{t=1}^{T_2}r_{t+1}\varepsilon_{c,t+1}\right)\times\frac{1}{T_2}\Sigma_{t=1}^{T_2}\varepsilon_{c,t+1}\right)\\
&= E\left(\frac{1}{T_2}\Sigma_{t=1}^{T_2}z_{t+1}\times\Sigma_{t=1}^{T_2}\varepsilon_{c,t+1}\right)-(\gamma-1)\frac{T_1}{T_2}E\left(\left(\frac{1}{T_2}\Sigma_{t=1}^{T_2}w_{t+1}\right)\times\Sigma_{t=1}^{T_2}\varepsilon_{c,t+1}\right)\\
&= E\left(\frac{1}{T_2}\Sigma_{t=1}^{T_2}z_{t+1}\varepsilon_{c,t+1}+\Sigma_{l=1}^{L}\frac{1}{T_2}\Sigma_{t=l+1}^{T_2}\left[z_{t+1}\varepsilon_{c,t+1-l}+z_{t+1-l}\varepsilon_{c,t+1}\right]\right)\\
&\quad -(\gamma-1)\frac{T_1}{T_2}E\left(\frac{1}{T_2}\Sigma_{t=1}^{T_2}w_{t+1}\varepsilon_{c,t+1}+\Sigma_{l=1}^{L}\frac{1}{T_2}\Sigma_{t=l+1}^{T_2}\left[w_{t+1}\varepsilon_{c,t+1-l}+w_{t+1-l}\varepsilon_{c,t+1}\right]\right).
\end{aligned}$$

where the second equality exploits the fact that values of ε_c are uncorrelated with values of z which are from the part of the long sample (of length T_1) which does not overlap with the short sample (of length T_2).

Finally,

$$\begin{aligned}
\Omega^{rr}_{25x25} &= E\left(T_1\left(\frac{1}{T_1}\Sigma_{t=1}^{T_1}z_{t+1}-(\gamma-1)\frac{1}{T_2}\Sigma_{t=1}^{T_2}w_{t+1}\right)\times\left(\frac{1}{T_1}\Sigma_{t=1}^{T_1}z_{t+1}-(\gamma-1)\frac{1}{T_2}\Sigma_{t=1}^{T_2}w_{t+1}\right)'\right)\\
&= T_1E\left(\frac{1}{T_1}\Sigma_{t=1}^{T_1}\left[z_{t+1}-\mu_{z_1}\right]\right)\left(\frac{1}{T_1}\Sigma_{t=1}^{T_1}\left[z_{t+1}-\mu_z\right]'\right)-T_1\mu_z\mu_{z_1}'\\
&\quad +T_1(\gamma-1)^2E\left(\left(\frac{1}{T_2}\Sigma_{t=1}^{T_2}\left[w_{t+1}-\mu_w\right]\right)\times\left(\frac{1}{T_2}\Sigma_{t=1}^{T_2}\left[w_{t+1}-\mu_w\right]'\right)\right)-T_1(\gamma-1)^2\mu_w\mu_{w_2}'\\
&\quad -(\gamma-1)T_1E\left(\frac{1}{T_1}\Sigma_{t=1}^{T_1}\left[z_{t+1}-\mu_z\right]\right)\times\left(\frac{1}{T_2}\Sigma_{t=1}^{T_2}\left[w_{t+1}-\mu_w\right]'\right)+(\gamma-1)T_1\mu_z\mu_w'\\
&\quad -(\gamma-1)T_1E\left(\frac{1}{T_2}\Sigma_{t=1}^{T_2}\left[w_{t+1}-\mu_w\right]\right)\times\left(\frac{1}{T_1}\Sigma_{t=1}^{T_1}\left[z_{t+1}-\mu_z\right]'\right)+(\gamma-1)T_1\mu_w\mu_z'
\end{aligned}$$

$$\begin{aligned}
= \; & E\left(\frac{1}{T_1}\Sigma_{t=1}^{T_1}\left[z_{t+1}-\mu_z\right]\left[z_{t+1}-\mu_z\right]'\right)-T_1\mu_z\mu_z'\\
&+\frac{T_1}{T_2}(\gamma-1)^2E\left(\begin{array}{c}\frac{1}{T_2}\Sigma_{t=1}^{T_2}\left[w_{t+1}-\mu_w\right]\left[w_{t+1}-\mu_w\right]'\\ +\Sigma_{l=1}^{L}\frac{1}{T_2}\Sigma_{t=l+1}^{T_2}\left(\left[w_{t+1}-\mu_w\right]\left[w_{t+1-l}-\mu_w\right]'+\left[w_{t+1-l}-\mu_w\right]\left[w_{t+1}-\mu_w\right]'\right)\end{array}\right)\\
&-T_1(\gamma-1)^2\mu_w\mu_w'\\
&-(\gamma-1)E\left(\begin{array}{c}\frac{1}{T_2}\Sigma_{t=1}^{T_2}\left[z_{t+1}-\mu_z\right]\left[w_{t+1}-\mu_w\right]'\\ +\Sigma_{l=1}^{L}\frac{1}{T_2}\Sigma_{t=l+1}^{T_2}\left(\left[z_{t+1}-\mu_z\right]\left[w_{t+1-l}-\mu_w\right]'+\left[z_{t+1-l}-\mu_z\right]\left[w_{t+1}-\mu_w\right]'\right)\end{array}\right)\\
&+(\gamma-1)T_1\mu_z\mu_w'\\
&-(\gamma-1)E\left(\begin{array}{c}\frac{1}{T_2}\Sigma_{t=1}^{T_2}\left[w_{t+1}-\mu_w\right]\left[z_{t+1}-\mu_z\right]'\\ +\Sigma_{l=1}^{L}\frac{1}{T_2}\Sigma_{t=l+1}^{T_2}\left(\left[w_{t+1}-\mu_w\right]\left[z_{t+1-l}-\mu_z\right]'+\left[w_{t+1-l}-\mu_w\right]\left[z_{t+1}-\mu_z\right]'\right)\end{array}\right)\\
&+(\gamma-1)T_1\mu_w\mu_z'.
\end{aligned}$$

where μ_z is the mean of z_{t+1} and μ_w is the mean of w_{t+1}. Ω is estimated by removing the E's and using the estimated values of $\varepsilon_{c,t+1}$, α, γ, μ_z, and μ_w. When invertibility problems arise we employ Newey-West weightings up to lag $L=36$ to ensure invertibility. In cases where invertibility is not a problem we confirm that the Newey-West weights deliver similar standard errors.

In Table I we show three sets of t-statistics based on this GMM approach. All three sets of t-statistics account for the different sample length used to estimate means and covariances. The first set furthermore adjusts only for cross-correlated residuals, the second in addition accounts for first-stage estimation error in the covariances, and finally the third and main one additionally accounts for consumption growth autocorrelation. The above variance formulas are those used for the third set of t-statistics. The first two sets of t-statistics set L=0 in these formulas. The first set furthermore exploits only the moments $E(g^r) = 0$ and thus disregards the terms in G and Ω that are due to the moment $E(g^m) = 0$.

Table I:

Euler Equation Estimation for Stockholders and Non-stockholders
Actual Consumption Data from the CEX, 1982-1999

Panel A reports estimates of the Euler equation based on the long-run (discounted) consumption growth of stockholders, the top third of stockholders, and non-stockholders. Long-run consumption growth from quarter t to $t+12$ is calculated using data from the *Consumer Expenditure Survey* over the period June, 1982 to May, 1999, assuming a discount rate of 5% per year (quarterly discount factor $\beta = 0.95^{1/4}$). Results are presented for regressions (4), (5), and (6) of Section IV. Average log excess returns and variances on the 25 Fama-French portfolios are estimated using data from July, 1926 to Dec., 2004. Covariances between log excess returns and long-run consumption risk are estimated over the period of CEX data avaibility. Regression (4) regresses average log excess returns on return covariances, regression (5) repeats regression (4) without a constant (forcing the model to price the equity premium) and, regression (6) reverses regression (4) by interchanging the dependent and independent variables. In the table α denotes the intercept for all regressions. Estimation is performed by OLS which is shown to be equivalent to GMM in Appendix B. Reported are the intercepts and implied risk aversion coefficients (γ) with three sets of t-statistics (in parentheses) computed using the GMM approach in Appendix B. All three sets of t-statistics account for the different sample length used to estimate means and covariances. The first set further adjusts for cross-correlated residuals, the second set also accounts for first-stage estimation error in the covariances, and the third set additionally accounts for consumption growth autocorrelation. For regression (6), t-statistics are computed using the delta method. R^2s from the cross-sectional regressions and an F-test of the equality of the first-stage covariances with long-run consumption growth across the 25 portfolios are also reported. Panel B reports the first-stage covariance estimates and GMM t-statistics of each of the 25 Fama-French portfolios with stockholder consumption growth, and F-test of the joint equality of the first-stage covariances, and an F-test of whether all the covariances equal zero. Panels C through G report the risk aversion coefficient estimates and t-statistics for various alternative weighting schemes and definitions of stockholder consumption growth using a representative agent series, weighting by the beginning of period consumption of each household, and augmenting the definition of stockholders using the predicted probability of being a stockholder from a probit analysis conducted in the SCF. Panel H reports risk aversion estimates based on aggregate consumption from NIPA over the CEX period 1982 to 1999 and over the period 1947 to 2004.

	Stockholders			**Top stockholders**			**Non-stockholders**		
Regression	(4)	(5)	(6)	(4)	(5)	(6)	(4)	(5)	(6)
Panel A: Baseline definition of stockholders from CEX									
α	0.015	0	-0.001	0.014	0	-0.001	0.014	0	0.002
$t^{cross}(\alpha)$	(3.29)	-	(-0.25)	(3.54)	-	(-0.78)	(3.79)	-	(3.61)
$t^{gmm1}(\alpha)$	(1.14)	-	(-0.13)	(1.05)	-	(-0.68)	(1.79)	-	(2.05)
$t^{gmm2}(\alpha)$	(0.89)	-	(-0.01)	(0.92)	-	(-0.10)	(2.12)	-	(2.10)
γ	17.02	30.21	30.02	10.02	16.56	14.88	31.30	56.04	345.68
$t^{cross}(\gamma)$	(8.07)	(7.78)	(5.12)	(9.35)	(7.94)	(6.06)	(1.46)	(1.31)	(0.79)
$t^{gmm1}(\gamma)$	(2.80)	(3.87)	(1.78)	(2.77)	(7.64)	(1.79)	(0.76)	(0.92)	(0.37)
$t^{gmm2}(\gamma)$	(1.97)	(4.37)	(2.88)	(2.43)	(7.14)	(5.44)	(0.27)	(0.36)	(0.36)
R^2	0.55			0.65			0.09		
F-test statistic (p-value) on joint equality of first-stage covariance estimates									
	7.80 (0.003)			4.01 (0.032)			1.77 (0.193)		

Panel B: First-stage covariance estimates for stockholder consumption growth

		growth 1	2	3	4	value 5	avg.	1	2	3	4	5
		Consumption growth covariance $\times 10^4$						t^{gmm2}-statistics				
small	1	4.90	8.36	11.23	11.48	13.96	9.99	0.74	1.48	2.19	2.30	2.26
	2	4.01	10.08	11.89	14.06	14.52	10.91	0.58	2.12	2.61	3.55	2.92
	3	3.74	9.60	13.11	10.08	12.81	9.87	0.59	2.00	2.90	2.16	2.57
	4	1.83	8.83	12.73	8.37	12.35	8.82	0.16	1.92	2.46	1.80	2.91
large	5	5.17	10.79	10.71	10.79	11.92	9.88	0.93	2.27	2.58	1.92	2.97
avg.		3.93	9.53	11.93	10.96	13.11						
F-stat		$H_0: \sigma_{i,c} = 0$	9.48	(0.001)		$H_0: \sigma_{i,c} = \sigma_{j,c}$			7.80	(0.003)		

	Stockholders			Top third stockholders			Non-stockholders		
Regression	(4)	(5)	(6)	(4)	(5)	(6)	(4)	(5)	(6)
Panel C: Baseline definition, representative agent series									
α	0.019	0	-0.001	0.019	0	-0.001	0.001	0	0.001
$t^{gmm2}(\alpha)$	(1.10)	-	(-1.15)	(1.14)	-	(-1.40)	(1.20)	-	(3.64)
γ	10.81	23.97	19.64	7.11	15.11	11.88	50.34	50.48	93.11
$t^{gmm2}(\gamma)$	(2.06)	(3.60)	(2.19)	(2.39)	(3.16)	(2.02)	(1.75)	(1.75)	(1.17)
R^2	0.53			0.56			0.54		
Panel D: Baseline definition, consumption-weighted series									
α	0.015	0	-0.001	0.021	0	-0.001	0.006	0	0.001
$t^{gmm2}(\alpha)$	(0.78)	-	(-0.08)	(4.52)	-	(-1.00)	(0.65)	-	(1.29)
γ	11.34	20.02	20.66	6.74	16.44	13.83	30.01	36.63	46.17
$t^{gmm2}(\gamma)$	(1.99)	(4.28)	(2.62)	(2.12)	(2.91)	(1.46)	(2.94)	(3.38)	(2.70)
R^2	0.53			0.45			0.64		
Panel E: Alternative definition of stockholders: CEX and Probit from SCF									
α	0.018	0	0.001	0.014	0	-0.001	0.016	0	0.001
$t^{gmm2}(\alpha)$	(0.90)	-	(0.21)	(1.53)	-	(-0.57)	(4.26)	-	(3.00)
γ	12.00	25.85	29.58	7.86	12.94	12.09	21.80	44.00	215.95
$t^{gmm2}(\gamma)$	(1.67)	(4.58)	(2.06)	(2.36)	(5.65)	(4.01)	(3.15)	(2.89)	(0.35)
R^2	0.39			0.62			0.11		
Panel F: Alternative definition, representative agent series									
α	0.022	0	-0.001	0.020	0	-0.001	0.014	0	0.001
$t^{gmm2}(\alpha)$	(1.27)	-	(-0.34)	(1.56)	-	(-1.71)	(3.70)	-	(2.51)
γ	8.35	23.62	23.79	6.05	13.20	9.78	20.66	35.64	83.72
$t^{gmm2}(\gamma)$	(1.73)	(3.01)	(1.10)	(2.55)	(2.73)	(1.76)	(4.40)	(3.22)	(0.95)
R^2	0.32			0.58			0.24		
Panel G: Alternative definition, consumption-weighted series									
α	0.018	0	0.001	0.020	0	-0.001	0.014	0	0.001
$t^{gmm2}(\alpha)$	(0.94)	-	(0.01)	(2.05)	-	(-1.72)	(3.49)	-	(1.30)
γ	9.65	20.11	21.62	5.93	13.33	9.80	15.58	26.98	39.56
$t^{gmm2}(\gamma)$	(1.98)	(6.27)	(3.04)	(2.62)	(2.54)	(1.58)	(4.43)	(3.17)	(1.48)
R^2	0.42			0.56			0.38		
Panel H: Aggregate consumption growth from NIPA									
	1982-1999			**1947-2004**					
Regression	(4)	(5)	(6)	(4)	(5)	(6)			
α	0.015	0	-0.001	0.014	0	-0.001			
$t^{gmm2}(\alpha)$	(2.24)	-	(-1.15)	(2.50)	-	(-1.31)			
γ	49.42	92.15	76.42	52.15	87.98	71.80			
$t^{gmm2}(\gamma)$	(3.85)	(2.88)	(2.25)	(2.86)	(3.80)	(2.59)			
R^2	0.64			0.72					

Table II:

Euler Equation Estimation for Stockholders and Non-stockholders Consumption Growth Factor-Mimicking Portfolios, 1926-2004

The table reports estimates of the Euler equation using the long-run consumption growth of stockholders, the top third of stockholders, non-stockholders, and aggregate consumption over longer periods than the CEX sample by constructing consumption growth factor-mimicking (CGF) portfolios. Panel A reports results for a CGF formed by regressing long-run (12-quarter, discounted) consumption growth on a constant and the quarterly log excess returns (over the T-bill rate) 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. The regression is estimated from June, 1982 to May, 1999 and the coefficients are then used to project consumption growth from July, 1926 to Dec., 2004. Panel B reports results from CGFs created from macroeconomic variables including aggregate consumption and labor income growth, and the inflation and Treasury bill rates. Panel C reports results from CGFs created from other return portfolios including the excess return on the market, term and default spread, six industry portfolios, and momentum quartile portfolios. Each panel reports the risk aversion estimate from the Euler equation that employs the CGF in place of actual consumption growth and a t-statistic calculated from the GMM procedure of Appendix B as well as a t-statistic computed from a block bootstrap procedure to account for first-stage estimation error in the CGF (the bootstrap uses 5,000 replications). Also reported are the first-stage coefficient estimates and t-statistics (with a Newey-West adjustment of 36 lags) from the regression used to construct each CGF, as well as the R^2 from that regression.

Panel A: CGF from size and value portfolios, 1926-2004

	Stockholders			**Top stockholders**			**Non-stockholders**			**Aggregate**		
Regression	(4)	(5)	(6)	(4)	(5)	(6)	(4)	(5)	(6)	(4)	(5)	(6)
α	0.004	0	0.001	0.006	0	-0.001	0.002	0	0.001	0.006	0	-0.001
$t^{gmm}(\alpha)$	(0.68)	-	(0.60)	(1.19)	-	(-0.23)	(0.41)	-	(3.14)	(1.62)	-	(-0.13)
γ	7.56	8.43	9.20	4.57	5.43	5.34	15.14	16.32	28.08	20.75	25.64	25.51
$t^{gmm}(\gamma)$	(6.83)	(9.93)	(6.18)	(7.12)	(14.20)	(5.90)	(4.70)	(5.08)	(3.78)	(5.13)	(3.71)	(5.28)
$t^{boot}(\gamma)$	(3.30)			(2.57)			(0.77)			(2.05)		
R^2	0.80			0.82			0.52			0.81		
	Excluding return-months with fewer than 5 stocks per portfolio											
α	0.005	0	0.001	0.008	0	-0.001	-0.004	0	0.001	0.007	0	-0.001
$t^{gmm}(\alpha)$	(0.89)	-	(0.75)	(1.44)	-	(-0.33)	(-0.52)	-	(2.51)	(1.24)	-	(-0.20)
γ	10.50	12.42	11.81	5.86	7.42	6.50	29.83	26.62	43.40	27.18	34.11	31.17
$t^{gmm}(\gamma)$	(6.22)	(24.92)	(5.89)	(6.61)	(6.96)	(5.81)	(5.56)	(6.91)	(5.20)	(5.34)	(12.25)	(5.42)
$t^{boot}(\gamma)$	(4.59)			(3.29)			(1.52)			(2.69)		
R^2	0.88			0.88			0.68			0.87		

First-stage estimates of weights for CGF

	Stockholders	**Top stockholders**	**Non-stockholders**	**Aggregate**
Small,Growth	-0.52	-1.02	-0.08	-0.21
	(-2.36)	(-2.39)	(-0.53)	(-2.68)
Large,Growth	0.11	0.18	0.04	0.07
	(0.59)	(0.51)	(0.22)	(1.15)
Small,Value	0.65	1.48	0.05	0.31
	(1.79)	(2.46)	(0.27)	(2.24)
Large,Value	-0.01	-0.31	0.11	-0.10
	(-0.03)	(-0.77)	(0.80)	(-1.18)
R^2	0.06	0.04	0.03	0.13

Panel B: *CGF* from macroeconomic variables, 1947-2004												
	Stockholders				**Top stockholders**				**Non-stockholders**			
CGF	(A)	(B)	(C)	(D)	(A)	(B)	(C)	(D)	(A)	(B)	(C)	(D)
γ	23.73	9.98	22.77	13.78	15.41	10.21	13.58	17.34	54.69	42.34	64.15	53.17
$t^{gmm}(\gamma)$	(3.02)	(2.74)	(2.49)	(1.97)	(3.09)	(3.29)	(2.57)	(2.10)	(2.95)	(2.30)	(2.42)	(2.53)
R^2	0.69	0.31	0.26	0.14	0.69	0.49	0.26	0.34	0.69	0.48	0.26	0.13
					First-stage estimates of weights for *CGF*							
Agg. C	2.28	2.57			3.60	3.75			0.97	0.97		
	(6.12)	(7.64)			(4.25)	(4.10)			(3.05)	(2.57)		
Labor			1.56	1.54			2.69	2.50			0.54	0.52
			(5.88)	(5.87)			(4.73)	(4.56)			(2.40)	(2.00)
T-bill		8.43		4.57		13.82		7.99		1.07		-0.34
		(4.23)		(2.18)		(1.78)		(1.08)		(0.84)		(-0.23)
Inflation		-0.03		-0.02		-0.08		-0.06		-0.02		-0.02
		(-1.11)		(-0.88)		(-1.87)		(-1.31)		(-2.62)		(-2.18)
R^2	0.37	0.51	0.47	0.51	0.17	0.26	0.27	0.29	0.23	0.25	0.19	0.21

Panel C: *CGF* from other return portfolios, 1926-2004												
	Stockholders			**Top stockholders**			**Non-stockholders**			**Aggregate**		
CGF	(E)	(F)	(G)	(E)	(F)	(G)	(E)	(F)	(G)	(E)	(F)	(G)
γ	11.28	25.92	20.45	5.62	19.43	13.96	24.03	24.00	37.86	33.16	48.58	54.10
$t^{gmm}(\gamma)$	(2.36)	(3.47)	(2.45)	(1.64)	(3.27)	(3.07)	(3.04)	(2.63)	(3.43)	(2.66)	(3.29)	(2.69)
R^2	0.09	0.26	0.39	0.04	0.17	0.39	0.16	0.27	0.27	0.13	0.22	0.41
					First-stage estimates of weights for *CGF*							
RMRF	0.13			0.21			0.07			0.05		
	(1.95)			(1.82)			(2.93)			(2.02)		
DEF	-0.78			-3.06			0.19			-0.02		
	(-1.70)			(-3.20)			(0.89)			(-0.14)		
TERM	0.01			0.03			0.31			0.01		
	(0.02)			(0.07)			(2.65)			(0.19)		
NonDur		0.20			1.57			-0.29			-0.10	
		(0.82)			(1.43)			(-2.28)			(-1.18)	
Dur		0.22			0.68			0.01			0.13	
		(1.54)			(2.14)			(0.13)			(2.54)	
Manuf		-0.09			-0.89			0.25			-0.14	
		(-0.38)			(-0.88)			(1.92)			(-2.44)	
Energy		-0.27			-0.65			-0.13			-0.02	
		(-2.29)			(-2.36)			(-3.15)			(-0.80)	
Shops		-0.57			-1.22			-0.05			0.03	
		(-2.72)			(-2.08)			(-0.60)			(0.45)	
Utils		0.00			0.31			0.28			0.04	
		(-0.04)			(0.79)			(3.32)			(0.82)	
*MOM*1			-0.29			-0.57			-0.05			-0.05
			(-3.39)			(-1.69)			(-0.80)			(-1.81)
*MOM*2			1.15			1.27			0.06			0.30
			(3.46)			(1.20)			(0.42)			(1.90)
*MOM*3			-0.58			0.15			0.04			-0.16
			(-1.79)			(0.15)			(0.20)			(-1.35)
*MOM*4			-0.04			-0.31			0.11			-0.01
			(-0.21)			(-0.80)			(0.82)			(-0.05)
R^2	0.02	0.03	0.03	0.03	0.06	0.02	0.04	0.05	0.02	0.02	0.06	0.03

Table III:

Asset and Stockholder Long-Run 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 (12-quarter, discounted) consumption growth covariances calculated using CEX data from June, 1982 to May, 1999. Covariances are estimated using a procedure similar to Fama and French (1992). Specifically, for each individual stock we estimate the covariance of its returns with long-run 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 100 *pre-ranking* covariance centiles. We then compute the equal-weighted quarterly log excess returns on these 100 portfolios over the next 12 months, from July to June. This procedure is repeated every year, forming a time-series of returns on these 100 portfolios. We then reestimate covariances for the portfolios formed from the 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. Every month the cross-section of stock returns in excess of the 90 day T-bill rate is then 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), and the log of the book-to-market equity ratio (BE/ME). 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), adjusted for autocorrelation. Covariances with long-run consumption are estimated for stockholders, the top third of stockholders, 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 log excess returns (plus one-half the variance of excess returns) on covariances with respect to the stockholder consumption mimicking factor (CGF). Covariances are estimated using the pre- and post-ranking procedure described for Panel A. A Newey-West adjustment for autocorrelation is again used to compute the Fama and MacBeth (1973) t-statistics.

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

$$r_{i,t+1} - r_{f,t+1} = a + \gamma Cov(r_{i,t+1}, \sum_{s=0}^{11} \beta^s (c_{t+s+1} - c_{t+s})) + \lambda'_m X_i + e_i$$

June, 1982–May, 1999	γ	β_{RMRF}	$ln(ME)$	$ln(BE/ME)$
Stockholders	3.7487			
	(2.32)			
	2.7710	-0.0062	0.0017	0.0021
	(2.07)	(-0.71)	(2.03)	(1.23)
Top stockholders	2.3569			
	(2.45)			
	1.7044	-0.0067	0.0017	0.0022
	(2.27)	(-0.76)	(2.03)	(1.28)
Non-stockholders	3.5670			
	(1.04)			
	0.4369	-0.0064	0.0017	0.0022
	(0.20)	(-0.74)	(2.013)	(1.30)

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

$$r_{i,t+1} - r_{f,t+1} = a + \gamma Cov(r_{i,t+1}, CGF_{t+1}) + \lambda'_m X_i + e_i$$

July, 1926–Dec., 2004	γ	β_{RMRF}	$ln(ME)$	$ln(BE/ME)$
Stockholders	5.3700			
	(3.03)			
	2.7008	0.0027	-0.0005	0.0011
	(2.30)	(1.31)	(-1.81)	(1.81)
Top stockholders	3.2978			
	(3.21)			
	1.8611	0.0025	-0.0005	0.0011
	(2.77)	(1.22)	(-1.75)	(1.83)
Non-stockholders	7.9897			
	(1.99)			
	3.4916	0.0024	-0.0005	0.0011
	(1.43)	(1.15)	(-1.83)	(1.86)

Table IV:

Asset Pricing Tests for the 25 Fama-French Portfolios

Panel A reports summary statistics on the consumption growth factor (CGF) mimicking portfolios for stockholder (and top stockholder) long-run consumption growth from Table II Panel A, including their means, time-series t-statistics, percentage of months with negative returns, and their correlation with the Fama and French (1993) factor portfolios $RMRF$, SMB, and HML. For comparison with existing factors which are available at the monthly frequency, the CGF's used in this table are calculated using the coefficients from the first stage reported in Table II combined with monthly (as opposed to quarterly) returns on the regressors from that first stage. Panel B reports results from time-series regressions of the CGF on the four-factor model of Fama and French (1996) containing $RMRF$, SMB, HML, and UMD. Panel C reports the reverse regression of the Fama-French factor portfolios SMB and HML on the CGF. Panel D 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 models with 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. 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 the CGFs.

Panel A: Summary statistics of factor portfolios

	CGF_{stock}	CGF^{top}_{stock}	$RMRF$	SMB	HML	UMD
Mean	0.51	0.96	0.65	0.23	0.40	0.73
t-stat	7.49	7.54	3.64	2.05	3.43	4.84
% < 0	0.38	0.38	0.40	0.49	0.45	0.36
			Correlations			
CGF_{stock}	1.00	0.97	0.58	0.29	0.68	-0.20
CGF^{top}_{stock}		1.00	0.49	0.40	0.61	-0.12
$RMRF$			1.00	0.35	0.09	-0.18
SMB				1.00	0.08	-0.09
HML					1.00	-0.27

	July, 1926 – Dec., 2004				**July, 1926 – Dec., 1981**			
	Stockholders		Top Stockholders		Stockholders		Top Stockholders	
Panel B: Four-factor regressions: $CGF_t = a + bRMRF_t + sSMB_t + hHML_t + mUMD_t + e_t$								
	coeff.	t-stat	coeff.	t-stat	coeff.	t-stat	coeff.	t-stat
a	0.21	5.42	0.45	5.20	0.19	3.75	0.39	3.46
b	0.15	10.37	0.15	4.77	0.15	8.71	0.15	4.03
s	0.07	1.85	0.35	4.20	0.13	3.25	0.48	5.30
h	0.37	12.64	0.64	9.94	0.32	9.70	0.54	7.35
m	0.03	1.77	0.08	2.02	0.04	1.67	0.09	1.89
R^2	0.65		0.54		0.66		0.55	
Panel C: One-factor regressions: $y = \alpha + \beta_{CGF} CGF_t + \epsilon_t$								
	coeff.	t-stat	coeff.	t-stat	coeff.	t-stat	coeff.	t-stat
				$y = SMB_t$				
α	0.01	0.01	-0.03	-0.24	-0.10	-0.97	-0.10	-0.86
β_{CGF}	0.46	9.06	0.68	13.17	0.35	13.46	0.44	16.87
R^2	0.08		0.21		0.16		0.31	
				$y = HML_t$				
α	-0.18	-1.99	-0.13	-1.18	-0.11	-1.19	-0.06	-0.50
β_{CGF}	1.17	28.40	1.10	23.90	0.56	23.28	0.51	18.93
R^2	0.47		0.47		0.37		0.36	
Panel D: Time-series tests $r_{i,t} - r_{f,t} = \alpha_i + \beta'_{i,F} F_t + \epsilon_{i,t}$								
	GRS F	p-value	avg. $\|\alpha\|$	avg. R^2	GRS F	p-value	avg. $\|\alpha\|$	avg. R^2
$RMRF$	3.18	0.000	0.21	0.77	1.99	0.019	0.19	0.79
$RMRF$, SMB, HML	2.82	0.001	0.13	0.90	1.83	0.034	0.13	0.91
$RMRF$, CGF_{stock}	1.40	0.156	0.08	0.79	1.07	0.445	0.09	0.81
$RMRF$, CGF^{top}_{stock}	1.34	0.188	0.08	0.79	1.09	0.419	0.09	0.81
CGF_{stock}	1.58	0.081	0.14	0.34	1.04	0.356	0.18	0.44
CGF^{top}_{stock}	1.56	0.089	0.11	0.24	1.14	0.362	0.11	0.31

Table V:

The Equity Premium and Implied Risk Aversion

Implied measures of risk aversion from the equity premium and the covariance of aggregate stock returns with long-run (12-quarter, discounted) consumption growth are reported separately for the consumption of stockholders, the top third of stockholders, and non-stockholders 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 II Panel A, 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 (monthly) 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 risk aversion estimate γ based on this covariance and the equity premium. Risk aversion estimates are based on the Epstein-Zin Euler equation of Section II

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

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). For the longest time period we also report confidence intervals around the point estimates for γ, calculated using the GMM framework of Appendix B.

	CEX period 198206–199905	Period with aggregate consumption data 195901–200412	Full return period 192607–200412
Aggregate consumption growth (from NIPA)			
Covariance with market	2.59×10^{-4}	1.97×10^{-4}	
Implied γ	83.8	110.00	
$CGF^{aggregate}$			
Covariance with market	2.37×10^{-4}	1.87×10^{-4}	7.50×10^{-4}
Implied γ (95% confidence interval)	91.7	115.7	27.8 [13.3,74.5]
Stockholder consumption growth (from CEX)			
Covariance with market	8.73×10^{-4}		
Implied γ	23.8		
CGF^{stock}			
Covariance with market	6.41×10^{-4}	5.02×10^{-4}	26.49×10^{-4}
Implied γ (95% confidence interval)	25.6	32.2	8.9 [4.7,20.8]
Top third stockholder consumption growth (from CEX)			
Covariance with market	13.94×10^{-4}		
Implied γ	19.6		
CGF^{top}			
Covariance with market	9.15×10^{-4}	6.64×10^{-4}	41.18×10^{-4}
Implied γ (95% confidence interval)	15.4	21.2	5.9 [3.2,15.1]
Non-stockholder consumption growth (from CEX)			
Covariance with market	4.98×10^{-4}		
Implied γ	44.1		
$CGF^{non-stock}$			
Covariance with market	5.04×10^{-4}	4.90×10^{-4}	14.34×10^{-4}
Implied γ (95% confidence interval)	43.2	43.0	15.7 [7.7,41.8]

Table VI:
Aggregate Return Predictability

The table reports the results from predicting the excess return on the stock market over the T-bill rate (Panel A) and the excess return on 10-year bonds over the T-bill rate (Panel C) over the next 36 monthts, using (in separate regressions) Lettau and Ludvigson's (2002) consumption-to-wealth ratio *cay*, Julliard's (2005) augmented *caylr* variable, and the ratio of the quarterly consumption of stockholders (and the top third of stockholders) to total CEX consumption. Results are shown based on both our baseline definition of stockholders (and top stockholders) and the alternative definition of stockholders from Table I and Appendix A (labeled *stock2*), as well as based on annual charitable donations as a fraction of personal disposable income from NIPA from 1962 to 2002 as another proxy for the consumption share of wealthy stockholders. Standard errors employ a Newey-West adjustment of 36 lags when using data at the monthly frequency, 12 lags when using data at the quarterly frequency (*cay* and *caylr*), and 3 lags when using data at the annual frequency (donations). *t*-statistics are reported in parentheses. Due to measurement error in consumption, we run regressions both ways by interchanging the dependent and independent variables to reduce attenuation bias. Panel B and D supplements the predictability regressions with regressions showing the contemporaneous relations between stock market excess yields (Panel B) and bond market excess yields (Panel D) and the same set of regressors. Panel E reports the results of regressing each of our stockholder consumption share measures on *cay* and *caylr*. Also reported are the correlations between the stockholder consumption share measures and *cay* and *caylr*.

$x =$	*cay*	*caylr*	$\frac{c_t^{stock2}}{c_t^{agg}}$	$\frac{c_t^{top2}}{c_t^{agg}}$	$\frac{c_t^{stock}}{c_t^{agg}}$	$\frac{c_t^{top}}{c_t^{agg}}$	$\frac{donate_t}{y_t}$
Panel A: Stock market excess returns, $y = \Sigma_{\tau=t+1}^{t+36} (ln(1+r_{mkt,\tau}) - ln(1+r_{f,\tau}))$							
y on x	14.112	12.010	-1.679	-2.887	-1.265	-2.774	-0.588
	(6.72)	(4.98)	(-1.10)	(-1.09)	(-0.88)	(-1.14)	(-2.47)
x on y	0.033	0.029	-0.046	-0.027	-0.033	-0.028	-0.429
	(13.04)	(10.22)	(-2.82)	(-2.90)	(-2.18)	(-3.07)	(-3.58)
Panel B: Stock market excess yields, $y = (E/P)_t - r_{f,t}$							
y on x	0.423	0.371	-0.096	-0.167	-0.052	-0.147	-0.003
	(2.83)	(2.51)	(-5.14)	(-5.19)	(-2.62)	(-4.75)	(-0.21)
x on y	0.089	0.080	-1.067	-0.620	-0.545	-0.603	-0.369
	(2.83)	(2.51)	(-3.68)	(-3.75)	(-1.89)	(-3.67)	(-0.21)
Panel C: Bond market excess returns, $y = \Sigma_{\tau=t+1}^{t+36} (ln(1+r_{10,\tau}) - ln(1+r_{f,\tau}))$							
y on x	1.992	1.304	-0.987	-1.479	-1.086	-1.399	-0.203
	(1.22)	(0.86)	(-2.40)	(-2.30)	(-2.94)	(-2.29)	(-1.89)
x on y	0.019	0.013	-0.181	-0.092	-0.188	-0.094	-0.490
	(2.78)	(1.82)	(-4.53)	(-4.30)	(-5.59)	(-4.34)	(-2.02)
Panel D: Bond market excess yields, $y = y_{10y,t} - y_{1mo,t}$							
y on x	0.260	0.324	-0.106	-0.180	-0.065	-0.151	-0.014
	(3.58)	(4.62)	(-5.79)	(-5.63)	(-3.27)	(-4.92)	(-1.31)
x on y	0.226	0.292	-1.184	-0.666	-0.675	-0.621	-2.919
	(3.58)	(4.62)	(-3.82)	(-4.10)	(-2.24)	(-3.71)	(-1.31)
Panel E: Regression of consumption share measures on *cay* and *caylr*							
cay			-1.556	-0.949	-1.556	-1.030	-9.561
			(-4.87)	(-4.89)	(-4.78)	(-4.68)	(-4.35)
cay, including trend in regression			-0.617	-0.472	-0.635	-0.507	-8.76
			(-3.01)	(-3.03)	(-2.85)	(-2.79)	(-4.19)
caylr			-1.440	-0.875	-1.444	-0.967	-10.868
			(-4.32)	(-4.30)	(-4.27)	(-4.20)	(-4.71)
caylr, including trend in regression			-0.422	-0.358	-0.450	-0.400	-9.877
			(-1.95)	(-2.16)	(-1.92)	(-2.08)	(-4.44)
correlation with *cay*			-0.49	-0.49	-0.48	-0.48	-0.58
correlation with *caylr*			-0.45	-0.44	-0.44	-0.44	-0.50

Figure 1. Implied Risk Aversion Coefficients Across Consumption Growth Horizons

Estimates of the Euler equation in equation (4) based on the long-run consumption growth of stockholders, the top third of stockholders, non-stockholders, and aggregate consumption over various horizons S from quarter t to $t+S$ are estimated and the implied risk aversion coefficients (and 95% confidence intervals) are plotted. The long-run consumption growth rates are discounted using a quarterly discount factor of $\beta = 0.95^{1/4}$. The average log excess returns on the 25 Fama-French portfolios (estimated from July, 1926 to Dec., 2004) are regressed on covariances between log returns and long-run consumption risk calculated over the period of the CEX sample. Standard errors are computed using the GMM approach detailed in Appendix B which accounts for cross-asset correlations in returns, first-stage estimation error in covariances, autocorrelation in consumption growth, and the differing length of sample periods used to estimate means and covariances.

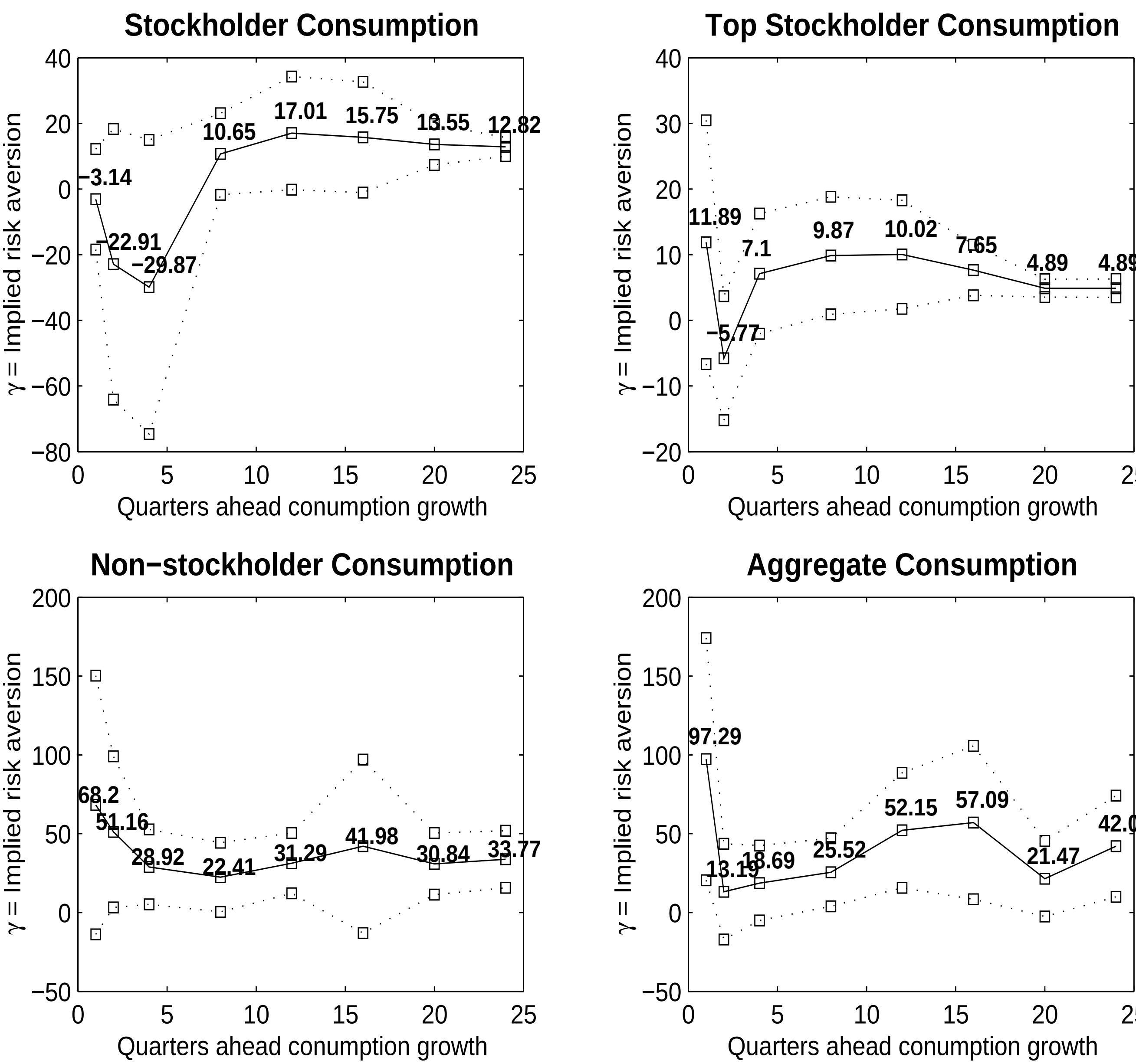
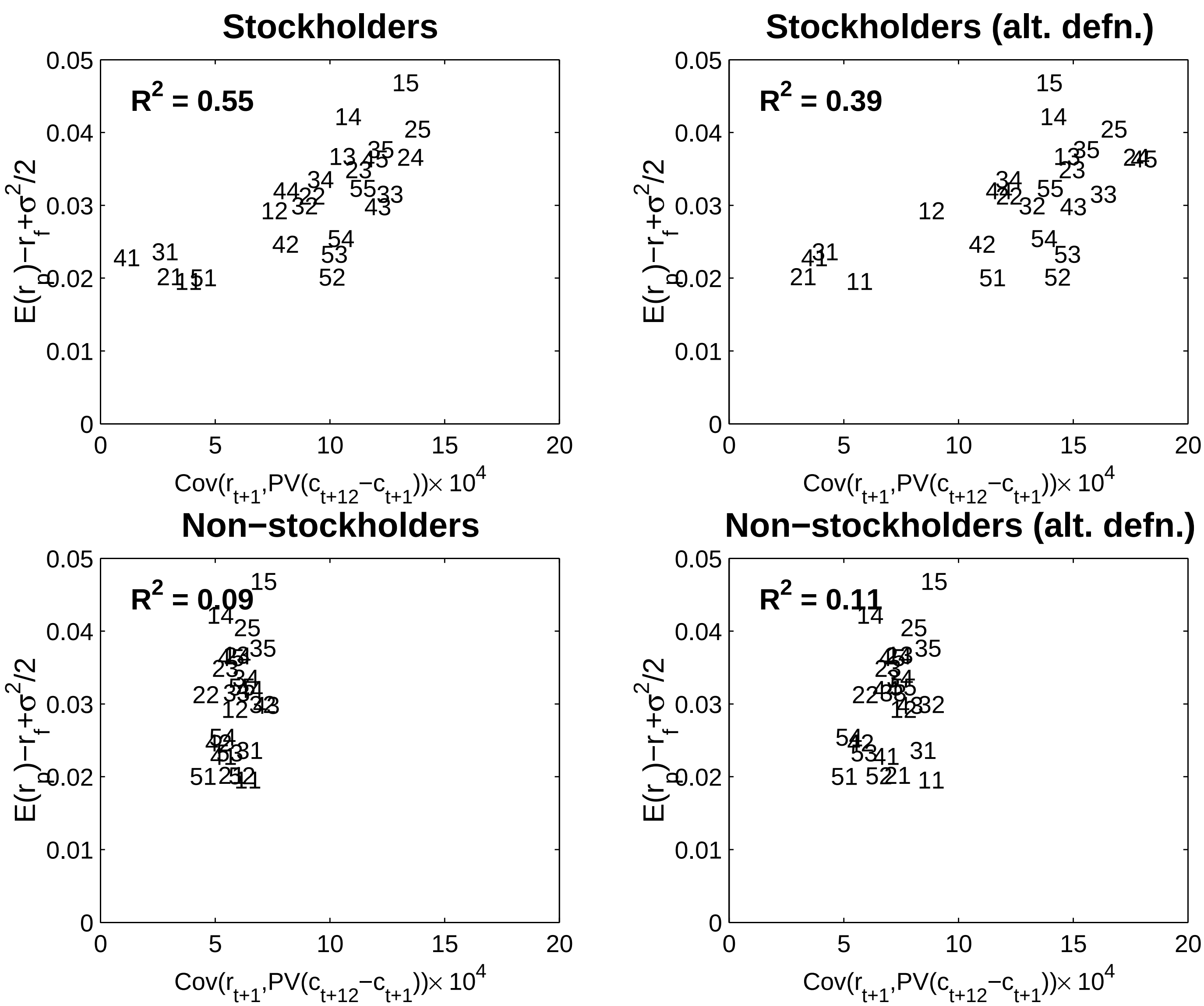


Figure 2. Stockholder Long-Run Consumption Risk and Expected Returns on the 25 Fama-French Portfolios

The graphs plot the average returns on the 25 Fama-French portfolios against the covariance of returns with long-run (12-quarter, discounted) consumption growth for stockholders, the top third of stockholders, and non-stockholders from Table I Panels A (baseline stockholder definition) and B (alternative stockholder definition). The entire time-series of returns from July, 1926 to Dec., 2004 is used to estimate mean returns which covariances are calculated over the CEX sample (June, 1982, to May, 1999). Also plotted are the implied risk aversion coefficients from employing the consumption growth factor (CGF) portfolios for stockholders and the top third of stockholders. The construction of the factor-mimicking portfolios for consumption growth are described in Table II Panel A. CGFs are available from July, 1926 to Dec., 2004. Also reported are the R^2s 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.



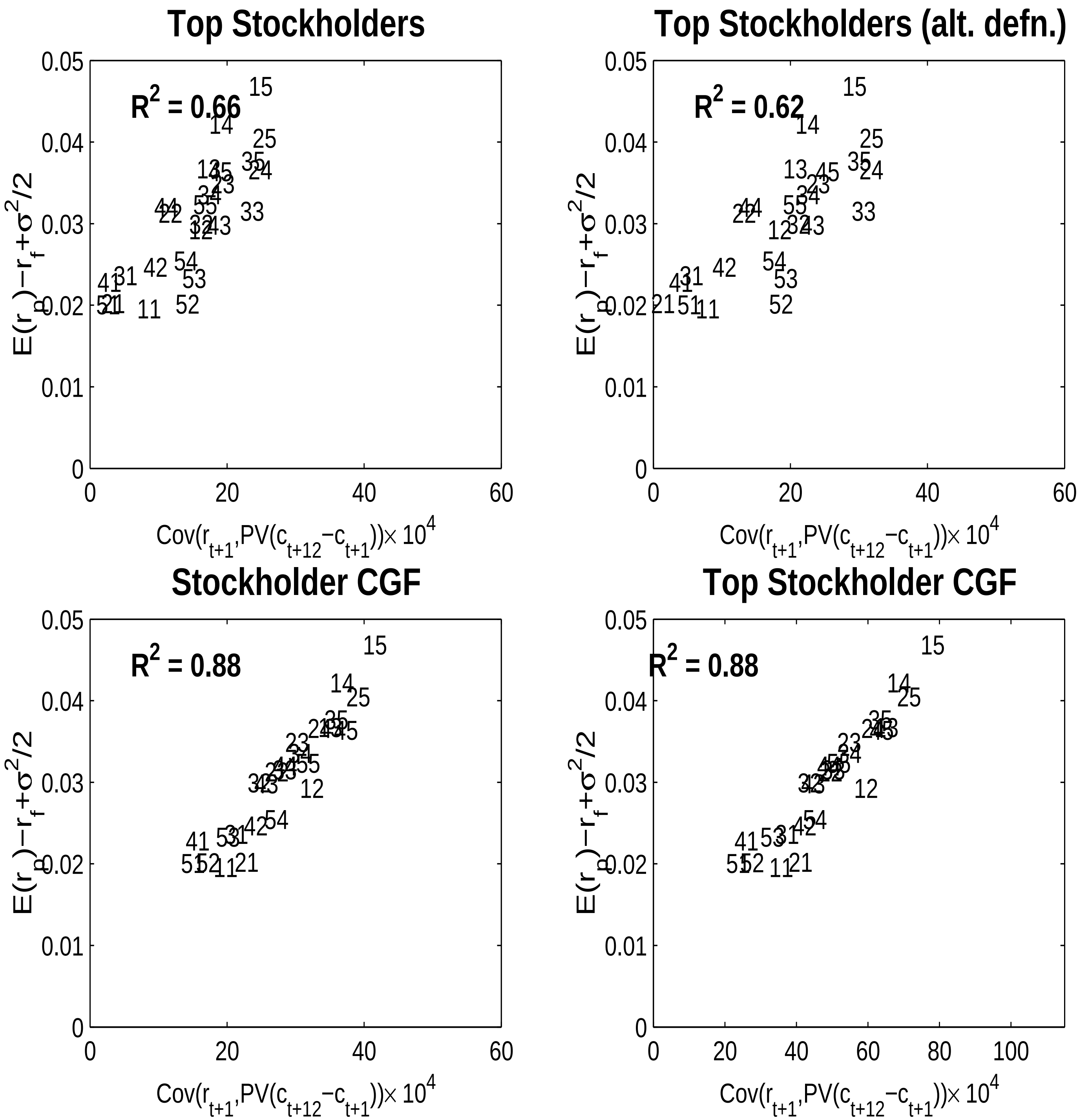

Top Stockholders
R2 = 0.66
Top Stockholders (alt. defn.)
R2 = 0.62
Stockholder CGF
R2 = 0.88
Top Stockholder CGF
R2 = 0.88
E(r_p)−r_f+σ^2/2
Cov(r_t+1,PV(c_t+12−c_t+1))× 10^4

Figure 3. Plots of the Sensitivity of Stockholder, Top Stockholder, and Non-stockholder Long-Run Consumption Growth to Aggregate Long-Run Consumption Growth

This figure plots the sensitivity of stockholder, top stockholder, and non-stockholder long-run (12-quarter, discounted) consumption growth from the CEX to aggregate long-run consumption growth from NIPA.

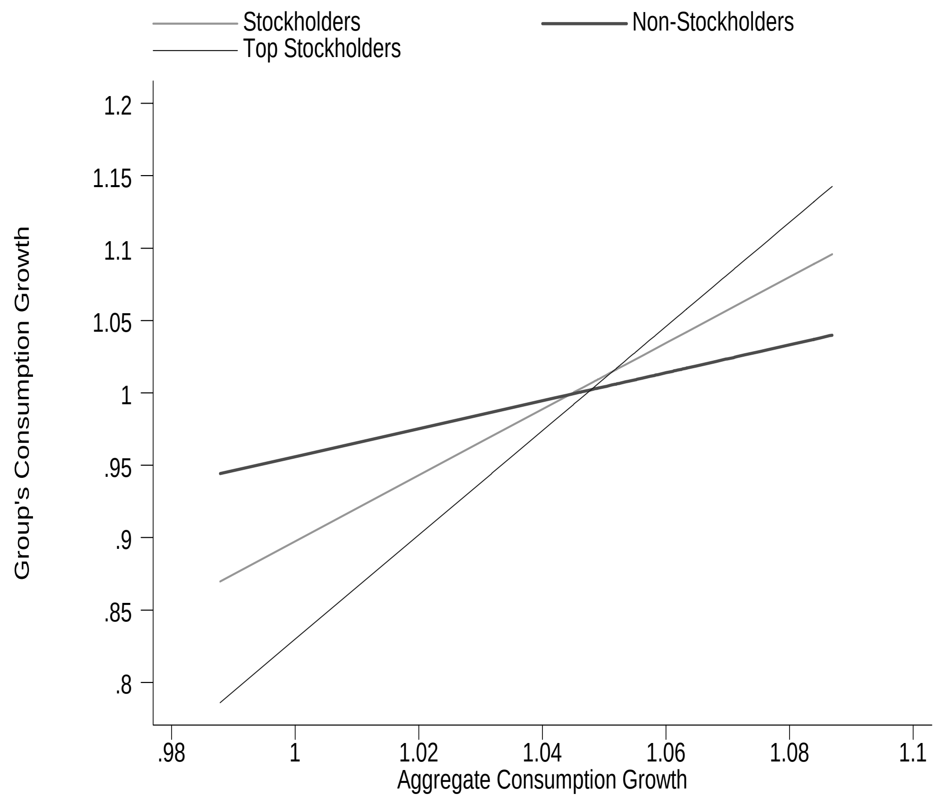


Figure 4. Plots of the Consumption-to-Wealth ratio *cay* and the Stockholder Consumption Share and Donations/Disposable Personal Income

The figure plots the consumption-to-wealth ratio of Lettau and Ludvigson (2002) along with the ratio of the quarterly consumption of stockholders (using our baseline stockholder definition) to total quarterly CEX consumption (top graph) or annual charitable donations as a fraction of annual NIPA disposable personal income (bottom graph). For readability, each time series is standardized by subtracting its mean and dividing by its standard deviation.