

Long-Run Stockholder Consumption Risk and Asset Returns

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ABSTRACT

We provide new evidence on the success of long-run risks in asset pricing by focusing on the risks borne by *stockholders*. Exploiting microlevel household consumption data, we show that long-run stockholder consumption risk better captures cross-sectional variation in average asset returns than aggregate or nonstockholder consumption risk, and implies more plausible risk aversion estimates. We find that risk aversion around 10 can match observed risk premia for the wealthiest stockholders across sets of test assets that include the 25 Fama and French portfolios, the market portfolio, bond portfolios, and the entire cross-section of stocks.

A CORNERSTONE OF ASSET PRICING THEORY, the consumption Capital Asset Pricing Model (CCAPM) focuses on consumption risk as the key determinant of equilibrium asset prices. Recent studies find success using long-run aggregate consumption risk to capture cross-sectional and aggregate stock returns (Parker (2001), Bansal and Yaron (2004), Parker and Julliard (2005), Bansal, Dittmar, and Lundblad (2005), Hansen, Heaton, and Li (2008)).¹ The empirical success of long-run consumption risk raises several questions. First,

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¹ Prior research finds a weak asset pricing role for contemporaneous consumption risk (Mankiw and Shapiro (1986), Breeden, Gibbons, and Litzenberger (1989), Kandel and Stambaugh (1990), Cochrane (1996), and Lettau and Ludvigson (2001b)). 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 sufficient dispersion in risks across assets needed to explain observed return premia (Lewellen and Nagel (2006)).

how well does aggregate consumption risk reflect the risks faced by capital market participants who actually own and trade stocks? Second, why does future consumption respond to news in stock returns and what are the underlying economic shocks driving these patterns? This article focuses on the first question.

We show that the long-run consumption risk of households who hold financial assets is particularly relevant for asset pricing. Examining disaggregated measures of long-run consumption risks across stockholders and nonstockholders separately, we provide new evidence on the long-run properties of consumption growth and its importance for asset pricing. Our work intersects the recent long-run risk literature and the literature on limited stock market participation and consumption (Mankiw and Zeldes (1991), Parker (2001), Vissing-Jørgensen (2002), Brav, Constantinides, and Geczy (2002), Attanasio, Banks, and Tanner (2002), and Gomes and Michaelides (2008)). Using microlevel household data from the Consumer Expenditure Survey (CEX) for the period 1982–2004, we show that the covariance of returns with long-run consumption growth rates of households who own stocks provides a better fit and more plausible risk aversion magnitudes when capturing the cross-sectional variation in average returns than long-run aggregate or nonstockholder consumption growth. Using the 25 Fama and French stock portfolios, the market portfolio, eight Treasury bond portfolios, and the entire cross-section of individual stocks, we show that when small stocks, value stocks, the overall stock market, and long-term bonds do poorly, the future consumption growth of stockholders will also be low. The high average returns observed for those asset classes may therefore reflect a premium stockholders require to bear long-run consumption risk.

The recent empirical success of long-run consumption measures has prompted asset pricing models that feature long-run risks. One set of models includes variations of the recursive utility framework of Kreps and Porteus (1978) and Restoy and Weil (1998), which allows for the separation of the elasticity of intertemporal substitution (EIS) from risk aversion. Epstein and Zin (1989), Weil (1989), Bansal and Yaron (2004), and Hansen et al. (2008), for example, introduce models that feature a role for future consumption. While other preferences, such as habit persistence (Sundaresan (1989), Constantinides (1990), Heaton (1995), and Campbell and Cochrane (1999)) or constraints on adjusting consumption (Lynch (1996) and Gabaix and Laibson (2002)), can also generate a pricing role for long-run consumption growth, we adopt the recursive preference framework to interpret our empirical results, consistent with recent work on long-run consumption risk.²

² From a microeconomic perspective, a framework with habit formation may be less attractive because support for such preferences is mixed (see Dynan (2000) and Brunnermeier and Nagel (2008) for evidence against habit formation based on consumption and portfolio choice data and Ravina (2005) for a dissenting view based on credit card charges). In contrast, estimates of the EIS in Vissing-Jørgensen (2002) and the results in the present paper on risk aversion suggest that consumption and asset returns may be consistent with a framework where stockholders have an EIS around one and a risk aversion higher than the reciprocal of the EIS, a structure that is possible within the recursive preference framework.

We use the structural framework of Hansen et al. (2008), focusing mainly on the special case where the EIS equals one. The pricing kernel under this special case simplifies to an expression that depends only on the present value of long-horizon consumption growth rates, substantially simplifying estimation and interpretation of our findings. For robustness, we also provide estimations where the EIS differs from one.³ Altering the EIS parameter makes little difference for the point estimate of the risk aversion parameter because we focus on the *cross-sectional* price of risk. Studies that argue for the importance of an EIS greater than one (e.g., Bansal and Yaron (2004)) focus on fitting the low level and low volatility of the riskless rate.

Our main result can be summarized as follows. Using the consumption of stockholders, as opposed to nonstockholders or aggregate consumption, and using long-run measures of consumption growth both serve to make our consumption series more sensitive to aggregate consumption shocks and more correlated with asset returns. Consequently, our structural framework requires a much lower risk aversion parameter to match the cross-sectional variation of average returns. For the 25 Fama–French portfolios, our structural estimate of the risk aversion coefficient implied by the premium for long-run consumption risk of stockholders is around 15, and for the wealthiest third of stockholders with the largest holdings of equity is around 10. These implied risk aversion estimates are significantly smaller than those obtained from either aggregate or nonstockholder long-run consumption growth (ranging from about 50 to 100).

The rest of the paper provides a battery of robustness tests that verify our basic findings. Our findings are robust to using various estimation methods, alternative definitions and aggregation of stockholders, and values of the EIS different from one. Furthermore, because microlevel consumption data are limited to the period 1982–2004, and long-run risks are difficult to measure, we also construct factor-mimicking portfolios for long-run stockholder consumption growth over a much longer period to reduce the impact of measurement error. Reestimating the Euler equations using the longer series of consumption growth factor-mimicking (*CGF*) portfolio returns, we find even lower risk aversion estimates of around 6–9 for stockholders and 5–7 for the wealthiest stockholders.

To further explore the scope of our findings, we also estimate our model for the aggregate equity premium, the cross-section of eight Treasury bond portfolios, and the entire cross-section of U.S. individual stock returns traded on the New York Stock Exchange (NYSE), Amex, and NASDAQ. The magnitudes of the risk aversion parameters required to explain returns for these three separate sets of test assets are remarkably similar to those obtained for the 25 Fama–French portfolios, highlighting the robustness of our results.

³ Attanasio and Vissing-Jørgensen (2003) estimate conditional Euler equations for stockholders in the CEX data using after-tax returns and find a value for the EIS 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. This evidence suggests that an EIS of one is not unreasonable.

Our findings provide additional evidence for long-run consumption risks as an economic mechanism for explaining asset prices. We highlight the importance of focusing on the long-run consumption risk of those households who directly bear capital market risk and for whom the estimated Euler equations should hold. The more plausible risk aversion estimates obtained for this set of households further supports a long-run consumption asset pricing framework.

The paper is organized as follows. Section I outlines a structural model with recursive preferences linking asset prices to long-run consumption risk. Section II summarizes the data sources and variable construction and provides preliminary measures of long-run risks. Section III examines the relation between long-run stockholder and nonstockholder consumption risk and the cross-section of returns on the 25 Fama–French portfolios. Section IV constructs factor-mimicking portfolios for stockholder and nonstockholder consumption growth. Section V examines other test assets, including the equity premium, the cross-section of Treasury bond portfolios, and the entire cross-section of individual stocks. Section VI concludes.

I. A Structural Asset Pricing Model with Recursive Preferences

Our theoretical setup follows Hansen et al. (2008) (hereafter HHL), who adopt a set of recursive preferences following Kreps and Porteus (1978), Epstein and Zin (1989), Weil (1989), and Bansal and Yaron (2004). The innovation of our study is in the empirical implementation of this setup. We use this model to compute the value of risk aversion implied by our findings to evaluate the economic plausibility of our results. Under this model each household has recursive preferences of the form

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

where C_t is consumption, σ is the elasticity of intertemporal substitution, γ is relative risk aversion, and β is the discount factor.

Following HHL, we assume that consumption growth is a linear function of the state of the economy x_t , which in turn evolves according to a first-order VAR:

$$c_{t+1} - c_t = \mu_c + U_c x_t + \lambda_0 w_{t+1} \quad (2)$$

$$x_{t+1} = Gx_t + Hw_{t+1}. \quad (3)$$

In this setup, c_t is log consumption, x_t is a vector of state variables, G has eigenvalues with absolute values that are strictly less than one, and the sequence $\{w_{t+1} : t = 0, 1, \dots\}$ consists of vectors of normal random variables that are independently and identically distributed with mean zero and covariance matrix Σ . The only departure from HHL is that for much of our analysis, in the case of an EIS equal to one, we do not need to assume that shocks are uncorrelated ($\Sigma = I$). In the case where the EIS is not equal to one we follow

HHL and assume $\Sigma = I$. Repeated substitution implies a stationary MA(∞) process for consumption growth,

$$\begin{aligned} c_t - c_{t-1} &= \mu_c + \lambda(L)w_t \\ &= \mu_c + \left(\sum_{s=0}^{\infty} \lambda_s L^s\right) w_t = \mu_c + \sum_{s=0}^{\infty} \lambda_s w_{t-s}. \end{aligned} \quad (4)$$

A. Euler Equations for an Elasticity of Intertemporal Substitution = 1

We focus on the special case where the EIS (for stockholders) equals one. The stochastic discount factor when the EIS = 1 simplifies to an expression that only depends on the present value of expectations about future consumption growth rates, making empirical estimation and interpretation of the Euler equation substantially easier. Specifically, following equation (3) and HHL, the expression for the log stochastic discount factor when the EIS = 1 is given by

$$\begin{aligned} s_{t+1} &= \ln \beta - [\mu_c + \lambda(L)w_{t+1}] + (1 - \gamma)\lambda(\beta)w_{t+1} - \frac{1}{2}(1 - \gamma)^2\lambda(\beta)^2 \\ &= \ln \beta - [c_{t+1} - c_t] + (1 - \gamma)\left(\sum_{s=0}^{\infty} \lambda_s \beta^s\right)w_{t+1} - \frac{1}{2}(1 - \gamma)^2\left(\sum_{s=0}^{\infty} \lambda_s \beta^s\right)^2 \\ &\simeq \ln \beta + (1 - \gamma)\left[(E_{t+1} - E_t)\sum_{s=0}^{\infty} \beta^s(c_{t+1+s} - c_{t+s})\right] - \frac{1}{2}(1 - \gamma)^2\left(\sum_{s=0}^{\infty} \lambda_s \beta^s\right)^2. \end{aligned} \quad (5)$$

In the last line of the above expression, we drop the term $[c_{t+1} - c_t]$. This term does not materially affect the results because one-period consumption growth is known to be poorly correlated with one-period excess returns. Written in terms of consumption growth rates, the term $(\sum_{s=0}^{\infty} \lambda_s \beta^s)w_{t+1}$ equals $(E_{t+1} - E_t)\sum_{s=0}^{\infty} \beta^s(c_{t+1+s} - c_{t+s})$, representing the innovation in expectations about the present value of consumption growth rates, and the term $(\sum_{s=0}^{\infty} \lambda_s \beta^s)^2$ is the variance of this innovation.

For any valid stochastic discount factor S_{t+1} , the unconditional Euler equation for asset i states that $E(S_{t+1}R_{t+1}^i) = 1$. Assuming joint (unconditional) log-normality of the stochastic discount factor, the return on asset i , and the return on the riskless asset, and using the expression for the log stochastic discount factor in (5), the log-linearized unconditional Euler equation for the excess log return on asset i over the riskless rate is

$$\begin{aligned} E(r_{t+1}^i - r_{t+1}^f) + \frac{1}{2}V(r_{t+1}^i) - \frac{1}{2}V(r_{t+1}^f) \\ &= -\text{cov}(s_{t+1}, r_{t+1}^i - r_{t+1}^f) \\ &\simeq (\gamma - 1)\text{cov}(E_{t+1}\sum_{s=0}^{\infty} \beta^s(c_{t+1+s} - c_{t+s}) - E_t\sum_{s=0}^{\infty} \beta^s(c_{t+1+s} - c_{t+s}), r_{t+1}^i - r_{t+1}^f) \\ &= (\gamma - 1)\text{cov}(\sum_{s=0}^{\infty} \beta^s(c_{t+1+s} - c_{t+s}), r_{t+1}^i - r_{t+1}^f) \\ &\quad - (\gamma - 1)\text{cov}(E_t(\sum_{s=0}^{\infty} \beta^s(c_{t+1+s} - c_{t+s})), E_t(r_{t+1}^i - r_{t+1}^f)). \end{aligned} \quad (6)$$

We assume a discount factor of 5% per annum, implying $\beta = 0.95^{1/4}$ quarterly.⁴

In addition to estimating the Euler equation using stockholder consumption growth, we also use data for nonstockholders and aggregate U.S. per capita consumption. Because nonstockholders do not hold financial assets and because evidence (e.g., Vissing-Jørgensen (2002)) suggests that nonstockholders likely have an EIS lower than one, the Euler equations may not hold for nonparticipating households. Hence, we do not treat risk aversion parameters from these estimations as valid for nonstockholders, but merely report them for comparison.⁵

A.1. Unconditional Covariances

We estimate two versions of the unconditional Euler equation (6). The first approach is to estimate equation (6) using only the unconditional covariance term, $\text{cov}(\sum_{s=0}^{\infty} \beta^s (c_{t+1+s} - c_{t+s}), r_{t+1}^i - r_{t+1}^f)$, and ignore the covariance of the conditional expectation of discounted consumption growth rates and the conditional expectation of the excess asset return. This approach leads to a consistent estimate of risk aversion γ if expected excess returns are constant over time, or if the covariance between $E_t(\sum_{s=0}^{\infty} \beta^s (c_{t+1+s} - c_{t+s}))$ and $E_t(r_{t+1}^i - r_{t+1}^f)$ is identical across the set of test assets, implying that expected excess returns on all assets move up or down in parallel as the conditional expectation of discounted consumption growth rates changes.⁶ The main advantage of this first approach is that it does not require the estimation of conditional expectations, a difficult task in the CEX data, which that span only 23 years.

A.2. Covariance of Conditional Expectations

A second approach is to estimate equation (6) by estimating the expression $\text{cov}(E_{t+1} \sum_{s=0}^{\infty} \beta^s (c_{t+1+s} - c_{t+s}) - E_t \sum_{s=0}^{\infty} \beta^s (c_{t+1+s} - c_{t+s}), r_{t+1}^i - r_{t+1}^f)$. This estimation is more consistent with the theory because it incorporates conditional expectations and calculates sums to infinity. Estimation of $E_{t+1} \sum_{s=0}^{\infty} \beta^s (c_{t+1+s} - c_{t+s}) - E_t \sum_{s=0}^{\infty} \beta^s (c_{t+1+s} - c_{t+s})$ based on a VAR model to capture consumption

⁴ Discount rates within the range of 0% to 10% ($\beta = 0.9^{1/4}$ to 1) have little effect on our results.

⁵ If all households have an EIS equal to one, then the aggregate consumption-to-wealth ratio is constant, which contradicts evidence from aggregate U.S. data that this ratio varies over time (Lettau and Ludvigson (2001a, 2001b)). However, as long as nonstockholders have an EIS different from one, a general equilibrium model of limited stock market participation also delivers a time-varying aggregate consumption-to-wealth ratio.

⁶ In our estimations we include a constant term to ensure consistent estimates of γ in this scenario, where the constant estimates the term $-(\gamma - 1)\text{cov}(E_t(\sum_{s=0}^{\infty} \beta^j (c_{t+1+s} - c_{t+s})), E_t(r_{t+1}^i - r_{t+1}^f))$. Further, although we cannot continue the sum $\sum_{s=0}^{\infty} \beta^s (c_{t+1+s} - c_{t+s})$ to the upper limit of infinity, it is likely that consumption growth rates very far out in the future will not matter substantially for the resulting estimate of $\text{cov}(\sum_{s=0}^{\infty} \beta^s (c_{t+1+s} - c_{t+s}), r_{t+1}^i - r_{t+1}^f)$ because they are unlikely to be correlated with excess asset returns in the current period. We consider several possible upper limits in the sum and show that beyond 16 quarters, additional quarters do not alter the results much.

dynamics would introduce additional and substantial estimation error given the relatively short sample period of CEX consumption data. Alternatively, we can rewrite the above covariance using return innovations rather than innovations to the present value of consumption growth rates:

$$\begin{aligned} & \text{cov}(E_{t+1}\Sigma_{s=0}^{\infty}\beta^s(c_{t+1+s}-c_{t+s})-E_t\Sigma_{s=0}^{\infty}\beta^s(c_{t+1+s}-c_{t+s}), r_{t+1}^i-r_{t+1}^f) \\ &= \text{cov}(\Sigma_{s=0}^{\infty}\beta^s(c_{t+1+s}-c_{t+s}), r_{t+1}^i-r_{t+1}^f-E_t(r_{t+1}^i-r_{t+1}^f)), \end{aligned} \quad (7)$$

because $\Sigma_{s=0}^{\infty}\beta^s(c_{t+1+s}-c_{t+s})-E_t\Sigma_{s=0}^{\infty}\beta^s(c_{t+1+s}-c_{t+s})$ by definition is uncorrelated with $E_t(r_{t+1}^i-r_{t+1}^f)$ and $E_t\Sigma_{s=0}^{\infty}\beta^s(c_{t+1+s}-c_{t+s})$ is uncorrelated with $r_{t+1}^i-r_{t+1}^f-E_t(r_{t+1}^i-r_{t+1}^f)$.

Because we have return data going back to 1926, we can more precisely estimate the dynamics of returns than consumption. To estimate return innovations we use lags of the test asset returns themselves (consistent with the literature on momentum and reversals), along with the log real T-bill return and the log price-dividend ratio on the market (proxied by the S&P 500).⁷

B. Euler Equations for an Elasticity of Intertemporal Substitution $\neq 1$

When the EIS (of stockholders) differs from one, additional terms enter the expression for the stochastic discount factor and thus the Euler equation. Hansen et al. (2008), building on earlier work by Kogan and Uppal (2002), provide the following first-order approximation of the log stochastic discount factor that is valid for values of the EIS different from one:

$$s_{t+1} \simeq s_{t+1}^1 + \left(\frac{1}{\sigma} - 1\right) Ds_{t+1}^1,$$

where s_{t+1}^1 is the log stochastic discount factor for an EIS of one and Ds_{t+1}^1 is the derivative of the log stochastic discount factor with respect to the inverse of the EIS, evaluated at an EIS of one. Hansen et al. (2007) show that, omitting constants that do not affect our subsequent analysis,

$$s_{t+1} \simeq (1-\gamma)\lambda(\beta)w_{t+1} + \left(\frac{1}{\sigma} - 1\right) \left(\frac{1}{2}w'_{t+1}\Theta_0w_{t+1} + w'_{t+1}\Theta_1x_t + \vartheta_1x_t + \vartheta_2w_{t+1}\right), \quad (8)$$

where $\Theta_0, \Theta_1, \vartheta_1$, and ϑ_2 are functions of the statistical model parameters μ_c, U_c, λ_0, G , and H and the preference parameters β and γ . Appendix A details the expressions of equation (8) following the statistical model of HHL and the derivations of Hansen et al. (2007), and describes how we estimate the Euler equation when the EIS differs from one.

⁷ We also tried using the same predictors for all test assets (lags of SMB, lags of HML, lags of the excess return on the market, the log real T-bill return, and the log price-dividend ratio on the S&P 500) and find very similar results.

C. What Is the Likely Impact of the EIS on Risk Aversion Estimates?

We argue that the value of the EIS has little effect on the risk aversion estimate when examining the cross-sectional price of risk. The higher the EIS, the lower the level of expected asset returns that is needed in equilibrium to induce households to consume a growing consumption stream (for a given time discount factor β), and the smaller the fluctuations in expected asset returns generated by fluctuations in expected consumption growth rates over time. Thus, a higher EIS keeps the level of expected asset returns low and the volatility of expected asset returns low, which helps fit the moments of the riskless rate.

In terms of the equity premium, in the recursive framework the covariance between the excess stock market return and the stochastic discount factor is driven by the sensitivity of each to shocks to the long-run component of dividend growth and consumption growth (see Bansal and Yaron (2004)). When the EIS is high, a positive shock to the long-run component leads to a large positive change in the stock market price-dividend ratio and thus a large positive stock return. The shock leads to high dividends but generates only a small increase in the discount rate when the EIS is high. The sensitivity of the stochastic discount factor to shocks to the long-run component does not depend strongly on the EIS. As a result, the covariance of the excess stock market return and the stochastic discount factor is higher for a higher EIS.

What are the implications of the EIS for the cross-section of returns? If a higher EIS increases covariances of all stock returns with the log stochastic discount factor by a roughly similar amount, then the risk aversion coefficient identified purely from the cross-section of returns will not be substantially affected by the value of the EIS. We confirm this intuition empirically in Section III. This property of the model suggests that using cross-sectional inference to estimate risk aversion may be particularly useful when there is lack of agreement about the value of the EIS.

II. Data and Preliminaries

We briefly describe the data sources and variables used in the study and provide some preliminary measures of long-run risks of stockholder and non-stockholder consumption growth.

A. Asset Returns

Our initial tests focus on the returns of the 25 size and book-to-market equity sorted portfolios of Fama and French (1996) obtained from Kenneth French's website from July 1926 to November 2004. The ending date of November 2004 is chosen to match the end of the consumption data we use. We also employ the returns on eight Treasury bond portfolios with average maturities of 3 months, 1 year, 2 years, 5 years, 7 years, 10 years, 20 years, and 30 years, obtained from the Center for Research in Security Prices (CRSP) from March 1982 to November 2004. We extract the full cross-section of NYSE,

Amex, and NASDAQ individual stock returns with beginning-of-month share prices above \$5 from CRSP from July 1926 to November 2004. We acquire all available stock-level market capitalization and book equity figures from CRSP and Compustat and obtain stock-level book values of NYSE firms prior to June 1962 from Kenneth French's website. We use the 30-day T-bill rate from CRSP over the period July 1926–November 2004 as the riskless rate of interest.

B. Consumption Growth

We calculate separate quarterly consumption growth rates for stockholders, the wealthiest third of stockholders, and nonstockholders using data from the CEX for the period March 1982 (end month of the first quarterly growth rate used) to November 2004. We also calculate aggregate per capita nondurable and service consumption growth rates from the National Income and Product Accounts (NIPA) from June 1959 to November 2004.

B.1. Household-Level Consumption from the CEX

We begin by describing the disaggregated CEX household-level data and then discuss how we compute average growth rates for stockholders, nonstockholders, and top stockholders. CEX data are available from the start of 1980 to the first quarter of 2005. Before 1999 the CEX contains interviews of about 4,500 households 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 subsequent four interviews are conducted 3 months apart and households are asked to report consumption for the previous 3 months. While each household is interviewed at 3-month intervals, interviews across households are spread out over the quarter. This means that there will be households interviewed in each month of the sample, enabling us to compute quarterly growth rates at a monthly frequency. Financial information is gathered in the fifth quarter only. Aside from attrition, with about 60% of households making it through all five quarters, the sample is representative of the U.S. population.

The consumption definition and sample selection criteria follow Vissing-Jørgensen (2002) and are described in Appendix B. Nominal consumption values are deflated by the Bureau of Labor Statistics (BLS) deflator for non-durables for urban households. To control for consumption changes driven by changes in family size and for seasonal consumption changes, we regress the change in log consumption on the change in log family size at the household level over the same period plus a set of seasonal dummies and use the residual as our quarterly consumption growth measure.

B.2. Identifying Stockholders

We use both a simple definition of stockholders, following Vissing-Jørgensen (2002), based on responses to the CEX indicating positive holdings of “stocks, bonds, mutual funds and other such securities,” and a more sophisticated

definition that seeks to mitigate response error by supplementing the CEX definition with a probit analysis designed to predict the probability that a household owns stocks. Using the Survey of Consumer Finances (SCF) from 1989, 1992, 1995, 1998, and 2001, which contain the entire wealth decomposition of households (including direct and indirect stock holdings), we estimate a probit model for whether a household owns stock on a set of observable characteristics that are also available in the CEX (age, education, race, income, checking and savings accounts, dividend income, and year). The estimated coefficients from the probit model in the SCF data are then used to predict the probability of stock ownership for households in the CEX data who have information on the same observable characteristics. The details of this procedure and the probit estimates are described in Appendix B. Under the more sophisticated classification, stockholders are then defined as the intersection of households who state that they own “stocks, bonds, mutual funds, and other such securities” *and* have a predicted probability of owning stock from the probit analysis of greater than 0.50. Nonstockholders are similarly defined as the intersection of those responding negatively to the CEX question *and* having a predicted probability of owning stock of less than 0.50. Households whose responses are inconsistent with their predicted probabilities are excluded.

The alternative definition of stockholders refines 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 nonstockholders (stockholders). This percentage 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 40.3% of households as nonstockholders and 13.7% as stockholders, excluding the 46.0% of households that cannot be confidently classified. The fraction of stockholders increases over our sample period as found in other sources (such as the SCF).

We also compute consumption separately for the wealthiest third of stockholders based on their beginning-of-quarter dollar amount of holdings. We define the wealthiest third of stockholders under the first definition and also examine the intersection of those top third stockholders with the more refined second definition of stockholders. The cutoffs for being in the top third are defined by year. Under the simple stockholder definition, the average cutoff for being in the top third of CEX stockholders is stockholdings of \$18,955 (in 1982 dollars) in the first half of our sample (1982–1992) and \$39,813 in the second half (1993–2004). Under the more refined stockholder definition the corresponding average cutoffs are \$18,978 for the first half of our sample and \$39,813 for the second half. The similarity in cutoff values indicates that there is little discrepancy in identifying stockholders under either definition when looking at the wealthiest third of households. Because these are minimum values (in 1982 dollars) we are clearly capturing wealthy households.⁸

⁸ For robustness, we also tried identifying the wealthiest stockholders using a fixed cutoff in real terms, or using the top 5% of the population in terms of stockholdings. Results are similar

Appendix B describes the stockholder and nonstockholder classifications in detail. Our final sample contains 206,067 quarterly consumption growth observations across 76,568 households. The median number of consumption observations per month under the first definition is 172 for stockholders and 586 for nonstockholders. Under the alternative definition, these medians are 103 and 314, respectively.

B.3. Aggregation of Household Consumption Growth Rates

The panel dimension for each household in the CEX allows us to calculate consumption growth rates at the household level. However, because households do not appear in the CEX for more than four quarters, 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), by averaging the (log) consumption growth rates for households in that group in each period. We compute the average growth rate from t to $t + 1$ for a group as follows:

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

where $c_t^{h,g}$ is the quarterly log consumption of household h in group g for quarter t and H_t^g is the number of households in group g in quarter t . We then substitute the average growth rate of the group into the asset pricing relations from Section I. For example, for the special case where the EIS = 1, for holders of asset i (and the riskless asset) the Euler equation we estimate using unconditional covariances is

$$\begin{aligned} E(r_{t+1}^i - r_{t+1}^f) + \frac{1}{2}V(r_{t+1}^i) - \frac{1}{2}V(r_{t+1}^f) \\ \simeq (\gamma - 1)\text{cov}\left(\sum_{s=0}^{\infty}\beta^s\left[\frac{1}{H_{t+s}^g}\sum_{h=1}^{H_{t+s}^g}(c_{t+1+s}^{h,g} - c_{t+s}^{h,g})\right], r_{t+1}^i - r_{t+1}^f\right). \end{aligned} \quad (9)$$

This equation is the asset pricing relation based on the consumption of a particular household, summed cross-sectionally across households in the group. The cross-sectional summation 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 another method for computing average consumption growth rates for a particular group by assuming a representative

across these various definitions. In addition, the CEX tends to underweight the super wealthy (see Bosworth, Burtless, and Sabelhaus (1991)). Thus, our results likely understate the importance of focusing on stockholders, particularly the wealthiest stockholders.

agent aggregation. This approach, rather than aggregating properly by taking the cross-sectional average of the change in household log consumption growth rates, takes the log change in the cross-sectional average of consumption growth,

$$\text{Representative agent} \Rightarrow \log \left(\frac{1}{H_{t+s}^g} \sum_{h=1}^{H_{t+s}^g} C_{t+1+s}^{h,g} \right) - \log \left(\frac{1}{H_{t+s}^g} \sum_{h=1}^{H_{t+s}^g} C_{t+s}^{h,g} \right), \quad (10)$$

where $C_t^{h,g}$ is consumption in quarter t for household h in group g . Estimating the Euler equation in (9) using this alternative consumption growth series addresses how sensitive our results are to the representative agent assumption and hence can gauge whether comovements of asset returns with cross-household inequality play a significant role. In addition, this computation reduces the influence of large positive or negative growth rates for some individual households.

B.4. Aggregate Consumption Data

We also compute aggregate consumption growth rates by using real seasonally adjusted monthly aggregate consumption of nondurables from NIPA Table 2.8.3 available starting in January 1959. We use data up to November 2004 and calculate quarterly consumption growth rates at the monthly frequency to coincide with the CEX data. Real per capita growth rates are calculated by subtracting the log population growth rate over the quarter, using monthly population from NIPA Table 2.6.

As is common practice in the literature, we assume in our analysis that the decision interval of the agent and the sampling interval of the data observed by the econometrician (quarterly) coincide.⁹

C. Long-Run Consumption Risks across Groups

Before proceeding to the Euler equation estimation, the main result of our paper is summarized very simply in Table I, which reports the sensitivity of stockholder, top stockholder, and nonstockholder consumption growth to aggregate consumption growth across various horizons ($S = 1, 2, 4, 8, 12, 16, 20$, and 24 quarters). As Table I shows, stockholder consumption is more sensitive to aggregate consumption shocks than nonstockholder consumption growth, and the top stockholders' consumption is even more sensitive to aggregate shocks. The differences in consumption sensitivity across the groups are exaggerated when

⁹ See Bansal, Kiku, and Yaron (2006) for evidence that misalignment of the sampling frequency with the agent's decision interval (time averaging) can lead to an upward bias in estimates of risk aversion in long-run risk models. Also see Grossman, Melino, and Shiller (1987) and Kandel and Stambaugh (1990). For evidence on the impact of delayed consumption adjustment, see Lynch (1996), Gabaix and Laibson (2002), and Jagannathan and Wang (2007).

Table I
Sensitivity of Stockholder, Top Stockholder, and Nonstockholder
Consumption Growth to Aggregate Consumption Growth
Across Horizons

The sensitivity of stockholder, top stockholder, and nonstockholder consumption growth to aggregate consumption growth from NIPA is reported over horizons of $S = 1, 2, 4, 8, 12, 16, 20$, and 24 quarters. The sensitivity of each group's consumption growth to aggregate consumption growth is computed as the regression coefficient from regressing a group's discounted consumption growth over horizon S on aggregate discounted consumption growth over the same horizon. Standard errors (in parentheses) on the regression sensitivity measure are computed using a Newey–West estimator that allows for autocorrelation of up to $S \times 3 - 1$ month lags. Group consumption growth rates are calculated using data from the Consumer Expenditure Survey over the period March 1982 to November 2004.

$S =$	1	2	4	8	12	16	20	24
Stockholder (s.e.)	0.68 (0.35)	0.93 (0.37)	1.21 (0.32)	1.57 (0.36)	2.12 (0.39)	2.68 (0.49)	2.68 (0.49)	2.42 (0.41)
Top stockholder (s.e.)	0.70 (0.90)	1.01 (0.77)	1.56 (0.62)	2.14 (0.49)	2.88 (0.53)	3.94 (0.67)	3.91 (0.73)	3.48 (0.63)
Nonstockholder (s.e.)	0.51 (0.23)	0.41 (0.27)	0.59 (0.26)	0.84 (0.27)	0.96 (0.29)	1.01 (0.26)	0.95 (0.24)	0.79 (0.27)

we examine long-run consumption growth rates. For example, at 16-quarter consumption growth rates, nonstockholder consumption growth has a β of about one with aggregate consumption, revealing why we obtain similar results using nonstockholder consumption as we do with aggregate consumption in our subsequent analysis.

However, at 16-quarter growth rates stockholder consumption is about three times more sensitive to aggregate consumption shocks and top stockholders' consumption is four times more sensitive. These differences are muted at short horizons. It is the combination of stockholder consumption *and* long-run risks that is crucial. Clearly, stockholders bear a disproportionate amount of aggregate consumption risk relative to nonstockholders and this burden increases in the long run, shedding light on why long-run stockholder (and top stockholder) consumption risk is able to match asset return premia at much lower values of risk aversion.¹⁰

¹⁰ An open question is why stockholders bear more consumption risk, especially in the long-run. Possible explanations for greater stockholder consumption sensitivity to aggregate consumption risk are (1) stockholders may have different preferences leading them to accumulate more wealth and take on more capital market risk, (2) stockholders may have different labor income or entrepreneurial income, (3) the tax system may transfer risk from nonstockholders to stockholders, or (4) the bond market may transfer risk to stockholders from nonstockholders, because the latter only use the bond market to smooth their consumption. Gomes and Michaelides (2008) focus on mechanism (1), while Guvenen (2007) analyzes mechanism (4). But to our knowledge no one has focused on why shareholders bear more of this risk in the long run.

III. Stockholder Consumption Risk and the 25 Fama and French Portfolios

Table II examines the relation between stockholder, top stockholder, non-stockholder, and aggregate consumption risk over various horizons and the cross-section of expected returns on the 25 Fama and French portfolios.

A. Unconditional Euler Equation Estimation for an $EIS = 1$

We first estimate the unconditional Euler equation (6) for the special case where the $EIS = 1$, using only the unconditional covariances of the test assets with consumption growth. Following equation (9), we run cross-sectional regressions of the average quarterly log excess returns on the 25 Fama–French portfolios plus a variance adjustment (measured quarterly from July 1926 to November 2004) against the covariance of returns with long-run discounted consumption growth. The relation is estimated separately for stockholders, the top third of stockholders, and nonstockholders. Although CEX consumption data are available only from 1982 to 2004, we employ the entire time series of returns dating back to July 1926 to estimate sample average returns and variances on the 25 Fama–French portfolios, to reduce estimation error, particularly for mean returns, which are notoriously difficult to estimate. In the next subsection we confirm that results are similar when estimating means and variances of returns over the shorter CEX sample period. Specifically, we run the cross-sectional regression

$$\hat{E}[r_{t+1}^i - r_{t+1}^f] + \frac{\hat{\sigma}_i^2}{2} - \frac{\hat{\sigma}_f^2}{2} = \alpha + (\gamma - 1)\hat{\sigma}_{ic} + e_i$$

$$\hat{\sigma}_{ic} = c\hat{v} \left(\sum_{s=0}^{S-1} \beta^s \left[\frac{1}{H_{t+s}^g} \sum_{h=1}^{H_{t+s}^g} (c_{t+1+s}^{h,g} - c_{t+s}^{h,g}) \right], r_{t+1}^i - r_{t+1}^f \right), \quad (11)$$

where $\beta = 0.95^{1/4}$ and γ is the implied risk aversion coefficient from the model. We estimate regression (11) via GMM. Appendix C details the moment conditions. The point estimate for risk aversion we obtain from our GMM framework is equivalent to that obtained from OLS. We compute standard errors under our GMM framework that account for correlation of error terms across assets at a point in time, estimation error in covariances, serial correlation in the consumption growth series induced by overlapping consumption data, and the different lengths of the data series used to estimate covariances and average returns. The details are provided in Appendix C.

We can also obtain consistent estimates of γ by estimating the Euler equation in reverse:

$$\hat{\sigma}_{ic} = \alpha + \frac{1}{\gamma - 1} \left(\hat{E}[r_{t+1}^i - r_{t+1}^f] + \frac{\hat{\sigma}_i^2}{2} - \frac{\hat{\sigma}_f^2}{2} \right) + u_i. \quad (12)$$

Table II
Euler Equation Estimation for Stockholder, Top Stockholder,
Nonstockholder, and Aggregate Consumption Growth
across Horizons

Panel A reports estimates of the Euler equation based on discounted consumption growth of stockholders over horizons of $S = 1, 2, 4, 8, 12, 16, 20$, and 24 quarters. Results are presented for regressions of the average log excess return plus half of the excess variance of the 25 Fama–French portfolios (estimated using quarterly data from September 1926 to November 2004) on a constant term and the covariances of log excess returns on the portfolios and discounted consumption growth over horizon S (estimated using data from 1982 to 2004). Results are also reported for the reverse regression that regresses covariances on average log excess returns plus half the excess variance and a constant term. Estimation is performed by OLS, which is equivalent to the GMM estimation setup in Appendix C. Reported are the intercepts (α) and implied risk aversion coefficients (γ) with t -statistics (in parentheses) computed using the GMM approach in Appendix C, which accounts for the different sample lengths used to estimate means and covariances, cross-correlated residuals, first-stage estimation error in the covariances, and consumption growth autocorrelation. For the reverse regressions, t -statistics are computed using the Delta method. R^2 's from the cross-sectional regressions are also reported at the bottom of the panel along with p -values of a test of overidentifying restrictions and model performance using the Hansen and Jagannathan (1997) distance. Panel B reports the same set of regression results using the consumption growth of the top third of stockholders and Panel C reports results using the consumption growth of nonstockholders. Group consumption growth rates are calculated using data from the Consumer Expenditure Survey over the period March 1982–November 2004. Panel D reports regression results and risk aversion estimates based on aggregate consumption from NIPA over both the CEX period (1982–2004) and the longer period 1959–2004 for which NIPA data are available.

$S =$	1	2	4	8	12	16	20	24
Panel A: Stockholder Consumption Growth over Horizon S								
Regression: $E[r_{t+1}^i - r_{t+1}^f] + \frac{\sigma_i^2}{2} - \frac{\sigma_f^2}{2} = \alpha + (\gamma - 1)\sigma_{i,cS} + e_i$								
α	0.03 (4.97)	0.06 (3.93)	0.06 (3.38)	0.02 (3.50)	0.02 (1.95)	0.01 (1.40)	0.01 (1.34)	0.02 (4.04)
γ	-20.27 (-1.23)	-43.50 (-2.10)	-44.53 (-2.13)	12.10 (1.53)	18.83 (2.61)	17.02 (3.44)	13.92 (5.03)	13.68 (7.91)
Reverse regression: $\sigma_{i,cS} = \alpha + \frac{1}{(\gamma-1)}(E[r_{t+1}^i - r_{t+1}^f] + \frac{\sigma_i^2}{2} - \frac{\sigma_f^2}{2}) + u_i$								
$\alpha \times 10^2$	0.03 (1.48)	0.07 (7.50)	0.09 (10.46)	0.05 (1.33)	-0.01 (-0.39)	-0.02 (-0.90)	-0.01 (-0.24)	-0.04 (-1.16)
γ	-390.02 (-0.29)	-345.61 (-0.39)	-134.71 (-0.73)	136.96 (0.39)	40.25 (1.53)	26.56 (2.40)	20.10 (3.65)	20.52 (5.80)
R^2	0.05	0.13	0.34	0.08	0.45	0.63	0.68	0.65
HJ p -value	0.000	0.000	0.001	0.000	0.001	0.008	0.008	0.001
Panel B: Top Stockholder Consumption Growth over Horizon S								
Regression: $E[r_{t+1}^i - r_{t+1}^f] + \frac{\sigma_i^2}{2} - \frac{\sigma_f^2}{2} = \alpha + (\gamma - 1)\sigma_{i,cS} + e_i$								
α	0.03 (6.33)	0.02 (1.94)	0.04 (5.63)	0.02 (1.09)	0.02 (2.82)	0.01 (0.87)	0.01 (1.78)	0.02 (3.81)
γ	3.47 (0.81)	13.03 (2.59)	-3.28 (-0.86)	10.37 (1.60)	7.62 (2.16)	10.99 (2.87)	8.79 (4.06)	8.23 (5.12)

(continued)

Table II—Continued

<i>S</i> =	1	2	4	8	12	16	20	24
Panel B: Top Stockholder Consumption Growth over Horizon <i>S</i>								
Reverse regression: $\sigma_{i,cS} = \alpha + \frac{1}{(\gamma-1)}(E[r_{t+1}^i - r_{t+1}^f] + \frac{\sigma_i^2}{2} - \frac{\sigma_f^2}{2}) + u_i$								
$\alpha \times 10^2$	0.01 (0.05)	0.10 (3.22)	0.13 (1.41)	0.07 (1.70)	−0.04 (−0.76)	−0.01 (−0.27)	−0.03 (−0.51)	−0.04 (−0.67)
γ	409.91 (0.05)	192.03 (0.35)	−149.62 (−0.19)	50.76 (0.53)	22.68 (0.99)	16.07 (1.92)	12.67 (2.87)	14.99 (3.50)
R^2	0.01	0.06	0.03	0.19	0.31	0.66	0.67	0.52
HJ <i>p</i> -value	0.000	0.000	0.000	0.000	0.000	0.007	0.001	0.000
Panel C: Nonstockholder Consumption Growth over Horizon <i>S</i>								
Regression: $E[r_{t+1}^i - r_{t+1}^f] + \frac{\sigma_i^2}{2} - \frac{\sigma_f^2}{2} = \alpha + (\gamma - 1)\sigma_{i,cS} + e_i$								
α	0.03 (3.08)	0.02 (1.27)	0.02 (4.29)	0.04 (4.56)	0.05 (6.75)	0.01 (0.40)	0.01 (2.09)	0.02 (0.82)
γ	86.31 (1.91)	74.39 (1.96)	24.64 (1.78)	−21.11 (−1.38)	−30.92 (−1.80)	48.00 (3.41)	28.77 (4.24)	30.26 (6.01)
Reverse regression: $\sigma_{i,cS} = \alpha + \frac{1}{(\gamma-1)}(E[r_{t+1}^i - r_{t+1}^f] + \frac{\sigma_i^2}{2} - \frac{\sigma_f^2}{2}) + u_i$								
$\alpha \times 10^2$	−0.01 (−1.60)	0.00 (−0.77)	0.03 (1.80)	0.05 (2.69)	0.07 (5.47)	0.03 (3.25)	0.00 (−0.20)	−0.02 (−1.51)
γ	694.00 (0.37)	218.10 (0.97)	304.00 (0.49)	−382.14 (−0.27)	−170.82 (−0.57)	148.76 (1.09)	44.04 (3.58)	43.46 (5.31)
R^2	0.12	0.34	0.08	0.06	0.19	0.32	0.65	0.69
HJ <i>p</i> -value	0.000	0.000	0.000	0.000	0.000	0.001	0.001	0.001
Panel D: Aggregate Consumption Growth over Horizon <i>S</i>								
CEX sample period 1982–2004								
Regression: $E[r_{t+1}^i - r_{t+1}^f] + \frac{\sigma_i^2}{2} - \frac{\sigma_f^2}{2} = \alpha + (\gamma - 1)\sigma_{i,cS} + e_i$								
α	0.03 (7.94)	0.03 (8.81)	0.03 (7.28)	0.02 (0.80)	0.02 (1.04)	0.01 (0.56)	0.02 (2.39)	0.02 (1.21)
γ	27.05 (0.90)	−17.18 (−0.97)	21.66 (1.62)	74.66 (2.75)	53.44 (4.31)	46.15 (6.11)	17.82 (3.92)	32.07 (4.17)
Reverse regression: $\sigma_{i,cS} = \alpha + \frac{1}{(\gamma-1)}(E[r_{t+1}^i - r_{t+1}^f] + \frac{\sigma_i^2}{2} - \frac{\sigma_f^2}{2}) + u_i$								
$\alpha \times 10^2$	0.01 (0.34)	0.02 (0.89)	0.01 (0.55)	0.01 (−0.31)	−0.01 (−2.00)	0.01 (0.63)	0.02 (0.72)	0.01 (−0.40)
γ	2,103.56 (0.12)	−810.34 (−0.18)	466.74 (0.40)	147.83 (1.87)	78.7 (1.66)	89.53 (3.30)	125.51 (4.20)	75.48 (1.59)
R^2	0.01	0.02	0.04	0.50	0.67	0.51	0.14	0.42
HJ <i>p</i> -value	0.000	0.000	0.000	0.001	0.002	0.000	0.000	0.000
NIPA sample period 1959–2004								
Regression: $E[r_{t+1}^i - r_{t+1}^f] + \frac{\sigma_i^2}{2} - \frac{\sigma_f^2}{2} = \alpha + (\gamma - 1)\sigma_{i,cS} + e_i$								
α	0.02 (6.23)	0.03 (7.22)	0.02 (5.69)	0.02 (2.81)	0.01 (2.44)	0.01 (0.76)	0.02 (2.82)	0.01 (1.38)
γ	97.01 (2.48)	13.06 (0.85)	18.22 (1.56)	25.14 (2.36)	51.85 (4.40)	53.98 (3.57)	21.37 (1.88)	39.79 (2.87)

(continued)

Table II—Continued

$S =$	1	2	4	8	12	16	20	24
Panel D: Aggregate Consumption Growth over Horizon S								
	Reverse regression: $\sigma_{i,cS} = \alpha + \frac{1}{(\gamma-1)}(E[r_{t+1}^i - r_{t+1}^f] + \frac{\sigma_i^2}{2} - \frac{\sigma_f^2}{2}) + u_i$							
$\alpha \times 10^2$	0.01 (1.16)	0.02 (0.74)	0.03 (1.22)	0.03 (2.10)	0.01 (-0.05)	0.02 (2.07)	0.03 (1.72)	0.02 (2.51)
γ	540.68 (1.33)	1,036.79 (0.11)	363.16 (0.41)	155.18 (1.11)	79.94 (3.83)	133.49 (2.16)	190.97 (0.73)	147.39 (1.58)
R^2	0.18	0.01	0.05	0.16	0.64	0.40	0.11	0.26
HJ p -value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000

The intercept, α , will obviously differ from that of equation (11). Here, we use the delta method to compute standard errors on γ from the GMM framework. Because equations (11) and (12) both provide consistent estimates of γ and because these estimates could differ in sample, we report results from both specifications.

Our focus is primarily on the structural estimate of risk aversion, γ . Implied risk aversion estimates provide a direct economic measure of the plausibility of the model. If the covariance between consumption growth and returns is too small to capture cross-sectional return premia (e.g., Hansen and Singleton (1983), Hansen and Jagannathan (1991), and Lewellen and Nagel (2006)), the regression will produce an implausibly large risk aversion estimate. While we also report cross-sectional R^2 s and pricing errors, we show that these diagnostics can be misleading. Lewellen, Nagel, and Shanken (2006) demonstrate how cross-sectional R^2 s and pricing errors can lead to a false impression of the model's success, particularly when the test assets are highly correlated with each other and contain a strong factor structure as the 25 Fama–French portfolios do (see also Daniel and Titman (2006)). One of the prescriptions for improving asset pricing tests is to impose the theoretical restrictions of the model and to take seriously the parameter value estimates. By backing out the implied risk aversion value from our Euler equation estimation, we evaluate the plausibility of the economic magnitudes of this parameter value as a metric for the success of the model. Another benefit from the structural approach is that the use of covariances, rather than betas, of returns with consumption growth reduces the impact of measurement error in consumption growth because covariances are not scaled by the variance of consumption growth.

B. Results across Horizons and Groups

Table II reports the results from regressions (11) and (12) for each consumption series: stockholders, the top third of stockholders, nonstockholders, and aggregate from quarter t to $t + S$ for $S = 1, 2, 4, 8, 12, 16, 20$, and 24. We report the estimated intercept, α , and implied risk aversion coefficient, γ , from the Euler equations with GMM t -statistics as described in Appendix C.

Panel A of Table II reports results for stockholder consumption risk. Risk aversion estimates from stockholder consumption growth are unreliable at horizons less than 12 quarters. In contrast, examining growth rates up to 12–24 quarters, risk aversion estimates are reliably positive and relatively stable. Risk aversion of 19, 17, 14, and 14 are obtained from 12-, 16-, 20-, and 24-quarter growth rates, respectively. The cross-sectional R^2 is between 0.63 and 0.68 for 16- to 24-quarter growth rates, indicating that long-run stockholder consumption risk captures substantial variation in average returns across the 25 Fama–French portfolios. While cross-sectional R^2 s alone are not strong tests of the model, a high R^2 combined with a relatively modest risk aversion coefficient lends reasonable support for the model. In addition, the constant term is insignificant at several of these horizons, indicating that the average pricing errors are insignificantly different from 0. We also report in the last row of Panel A of Table II the p -value of a test of overidentifying restrictions and model performance using the Hansen–Jagannathan (1997) distance. We calculate the p -values using the methodology of Parker and Julliard (2005). As is typically the case with these tests, we reject these restrictions.

Results are also reported for the reverse regression (12), which places consumption risk on the left-hand side of the regression equation. The implied risk aversion estimates from the reverse regression are mostly noise for less than 12 quarters' consumption growth and are between 20 and 27, and statistically significant, from 16 to 24 quarters out. These values are quite similar in magnitude to those obtained above from the forward regressions.

Panel B of Table II reports results for the consumption risk of the top third of stockholders. The implied risk aversion estimates fall to between 8 and 11 for 16- to 24-quarter growth rates when estimating the regression forward, and to between 13 and 16 when estimating the equation in reverse. Once again, while short-horizon growth rates, even for the wealthiest stockholders, do not pick up variation in average returns, long-run growth rates do.

Comparing the results for top stockholders (Panel B) to those for all stockholders (Panel A) illustrates the dangers of focusing on R^2 s or pricing errors to judge model success. The R^2 s are no higher (and sometimes lower) for top stockholder consumption risk, yet the implied risk aversion estimates are more reasonable. Taking the economic magnitudes of the model seriously is a more stringent test of the theory. For roughly the same R^2 , we obtain lower, more plausible risk aversion estimates when using the consumption risk of top stockholders, who own and trade a large part of the equity market.

Panel C of Table II reports results for nonstockholder consumption risk. Because the Euler equation may not hold for these households, because they do not hold stocks and, moreover, because nonstockholders may have an $EIS < 1$, the risk aversion parameter γ for nonstockholders may not be a valid measure of their true risk aversion. Nevertheless, it is still interesting to compare the results from nonstockholder consumption risk to those from stockholders to emphasize the importance of focusing on capital market participants. As Panel C of Table II shows, there is no significant relation between nonstockholder consumption risk and expected returns for horizons less than 16 quarters. The

risk aversion coefficients are not reliably different from 0 and the point estimates from the forward and reverse regressions are highly unstable. For 16- to 24-quarter growth rates, nonstockholder consumption risk captures some of the variation in average returns, which is not surprising because stockholder and nonstockholder consumption is likely cointegrated at long horizons. More importantly, however, the implied risk aversion coefficients from the model are quite large compared to those obtained for stockholders and the top third of stockholders, ranging from 29 to 149. This result illustrates, again, the danger of focusing exclusively on R^2 s. For 20- and 24-quarter growth rates, nonstockholder consumption risk produces R^2 s of 65 and 69%, respectively, which are roughly similar to those for stockholders. However, the risk aversion estimates obtained from the model are far less plausible, indicating much weaker covariation between returns and the consumption of nonstockholders. Put another way, high cross-sectional R^2 s, particularly on the 25 Fama–French portfolios, are a relatively low hurdle to clear for a model. Using theory to interpret the economic magnitude of the coefficients is a much tougher test.

Finally, Panel D of Table II reports results for aggregate NIPA consumption growth based on covariances for the sample period for which we have CEX microlevel data (1982–2004) and for the longer period for which aggregate consumption data are available from NIPA (1959–2004). Average returns are estimated from the full 1926 to 2004 sample in both cases. Both sets of covariances yield similar results. Risk aversion required to reconcile aggregate consumption risk with average asset returns ranges from 18 to 191 at long horizons and is sensitive to whether the regression is run forward or in reverse. These estimates are substantially higher than those for stockholder and top stockholder consumption growth.

Overall, Table II indicates that the long-run consumption risk of stockholders better captures asset return variation. Because stockholders are the households for which the Euler equation should hold and because the household consumption data we employ come from a completely separate source from aggregate consumption data, these results help alleviate general concerns about the empirical importance of long-run consumption risk being spurious.

Because Table II shows that our results are fairly consistent at long horizons of 16–24 quarters, we focus the remainder of the paper on 16-quarter (discounted) growth rates for brevity, except for the VAR-based results presented for the $EIS \neq 1$ case. We emphasize that there is nothing special per se about 16-quarter growth rates and that our results are robust to other definitions of long-run consumption.

C. Dispersion in Consumption Risks

To illustrate what drives the lower risk aversion estimates we obtain by using stockholder and top stockholder consumption growth, Table III reports the dispersion in long-run consumption growth covariances at 16-quarter horizons for the 25 Fama–French portfolios along with t -statistics from a GMM estimator that accounts for autocorrelation induced by overlapping consumption data.

Table III
Dispersion in Consumption Growth Covariances across
the 25 Fama–French Portfolios

Panel A reports the first-stage covariance estimates and GMM *t*-statistics allowing for autocorrelation up to 47 monthly lags of each of the 25 Fama–French portfolios with stockholder discounted consumption growth over a 16-quarter horizon. An *F*-test on the joint equality of the covariances is reported (with the *p*-value in parentheses). Panels B, C, and D report the covariance estimates for top stockholder, nonstockholder, and aggregate consumption growth, respectively.

	Growth				Value		Growth				Value
	1	2	3	4	5	Avg.	1	2	3	4	5
Panel A: Stockholder Consumption Growth											
	Consumption growth covariance $\times 10^{-4}$						GMM t -statistics				
1 (small)	4.80	8.14	11.37	12.17	15.45	10.38	0.90	1.56	1.98	2.48	2.37
2	3.18	9.19	10.27	14.34	14.83	10.36	0.90	2.20	2.23	3.49	2.94
3	4.61	9.82	13.37	9.68	13.75	10.25	1.12	2.11	3.34	2.47	3.21
4	0.51	6.90	11.79	7.98	11.40	7.72	0.62	2.50	3.08	2.76	3.32
5 (large)	6.22	9.59	9.13	10.98	12.54	9.69	1.58	3.12	3.67	3.49	4.52
Avg.	3.86	8.73	11.19	11.03	13.59		F -stat = 3.42 (p -value = 0.050)				
Panel B: Top Stockholder Consumption Growth											
	Consumption growth covariance $\times 10^{-4}$						GMM t -statistics				
1 (small)	13.14	19.26	20.66	24.06	31.12	21.65	1.03	1.87	2.08	2.73	3.16
2	11.12	15.70	20.44	27.16	27.88	20.46	1.10	2.13	2.82	4.16	4.13
3	12.93	19.90	25.94	19.79	26.81	21.08	1.39	2.37	3.94	3.40	3.69
4	4.99	14.23	21.57	14.53	18.45	14.75	0.74	2.44	3.24	2.70	3.47
5 (large)	10.86	18.01	16.29	17.50	23.80	17.29	1.44	2.50	3.09	3.24	4.99
Avg.	10.61	17.42	20.98	20.61	25.61		F -stat = 3.54 (p -value = 0.046)				
Panel C: Nonstockholder Consumption Growth											
	Consumption growth covariance $\times 10^{-4}$						GMM t -statistics				
1 (small)	4.61	4.58	5.02	4.86	6.33	5.08	1.72	2.64	2.89	3.09	3.30
2	3.96	3.85	4.60	6.14	5.63	4.84	1.59	2.30	3.24	5.18	3.28
3	5.06	6.32	5.63	4.25	5.76	5.40	1.87	3.32	4.02	3.70	4.47
4	3.70	3.71	5.39	4.20	3.91	4.18	1.58	2.44	3.59	3.38	2.52
5 (large)	3.13	4.93	4.53	3.22	4.62	4.08	1.53	2.60	2.79	3.31	4.89
Avg.	4.09	4.68	5.04	4.53	5.25		F -stat = 0.23 (p -value = 0.794)				
Panel D: Aggregate Consumption Growth, NIPA Sample Period											
	Consumption growth covariance $\times 10^{-4}$						GMM t -statistics				
1 (small)	5.06	3.81	4.99	4.32	5.50	4.74	1.77	1.51	2.25	2.16	2.74
2	2.51	3.13	3.66	4.26	4.69	3.65	1.05	1.58	2.16	2.47	2.89
3	1.76	3.50	4.07	3.69	4.44	3.49	0.99	2.00	2.59	2.56	2.93
4	2.15	3.29	4.27	4.45	3.96	3.62	1.24	2.03	3.44	3.51	3.15
5 (large)	3.29	3.21	2.66	4.05	4.09	3.46	2.10	2.30	2.87	3.56	3.83
Avg.	2.95	3.39	3.93	4.15	4.54		F -stat = 1.50 (p -value = 0.245)				

A typical concern with the estimates presented in Table II is that consumption covariances may not differ substantially across the test assets and may be imprecise. However, our results in Table II adjust for covariance estimation error in computing risk aversion estimates and the value of the risk aversion coefficient γ is identified off of the economic magnitudes of the differences in covariances across the assets.

Panel A of Table III highlights the dispersion in long-run stockholder consumption covariance estimates (and their t -statistics) for each of the 25 Fama–French portfolios. Many of the individual covariance estimates are at least two standard errors from 0, but, more importantly, there is wide economic dispersion in covariances across the portfolios. An F -test for the joint equality of the first-stage covariances across the 25 portfolios is rejected at the 5% significance level. More dispersed covariances across the test assets implies that a smaller risk aversion coefficient is required to match return data. If dispersion in consumption risk were negligible across the test assets, then a very large γ would be required to explain the cross-section of average returns. In addition, the significant dispersion in covariances lines up with average returns as small value portfolios have the highest covariances with long-run stockholder consumption growth.

For the top third of stockholders in Panel B of Table III, there is even more dispersion in consumption covariances across the 25 assets, which generates the even lower risk aversion estimate we obtain in Table II. The first-stage covariance estimates for nonstockholder (Panel C) and aggregate (Panel D) consumption, however, exhibit little dispersion and are not significantly different from each other across the test assets, explaining why a very large and unreliable risk aversion coefficient is needed to capture observed return premia.

Figure 1 summarizes the evidence in Tables II and III by plotting average log excess returns (plus half the excess variance) of the 25 Fama–French portfolios against their covariances with long-run consumption growth for each consumption group in Panel A, and against the fitted returns from the model (including the constant) in Panel B. To highlight the importance of economic magnitudes, the four graphs within each panel are plotted on the same scale. The plots in Panel A make apparent that there is not much dispersion in nonstockholder or aggregate consumption covariances, implying that high risk aversion values are needed to match average return dispersion across the assets. On the other hand, stockholder, and particularly top stockholder, consumption covariances are widely dispersed, with the result that much more plausible risk aversion values are sufficient to explain asset returns. Panel B reports the pricing errors of the model using each consumption series. Pricing errors are smaller for stockholder and top stockholder consumption compared to nonstockholder or aggregate consumption. However, the differences between the different consumption series are less striking here because we allow the model to pick α and γ freely without regard to what is reasonable for these parameters (i.e., not taking the economic magnitudes of the model seriously). Allowing these parameters to be chosen freely illustrates the problem of focusing only on model fit rather than on both fit and plausibility of preference parameter estimates.

Panel A: Mean returns versus consumption covariances

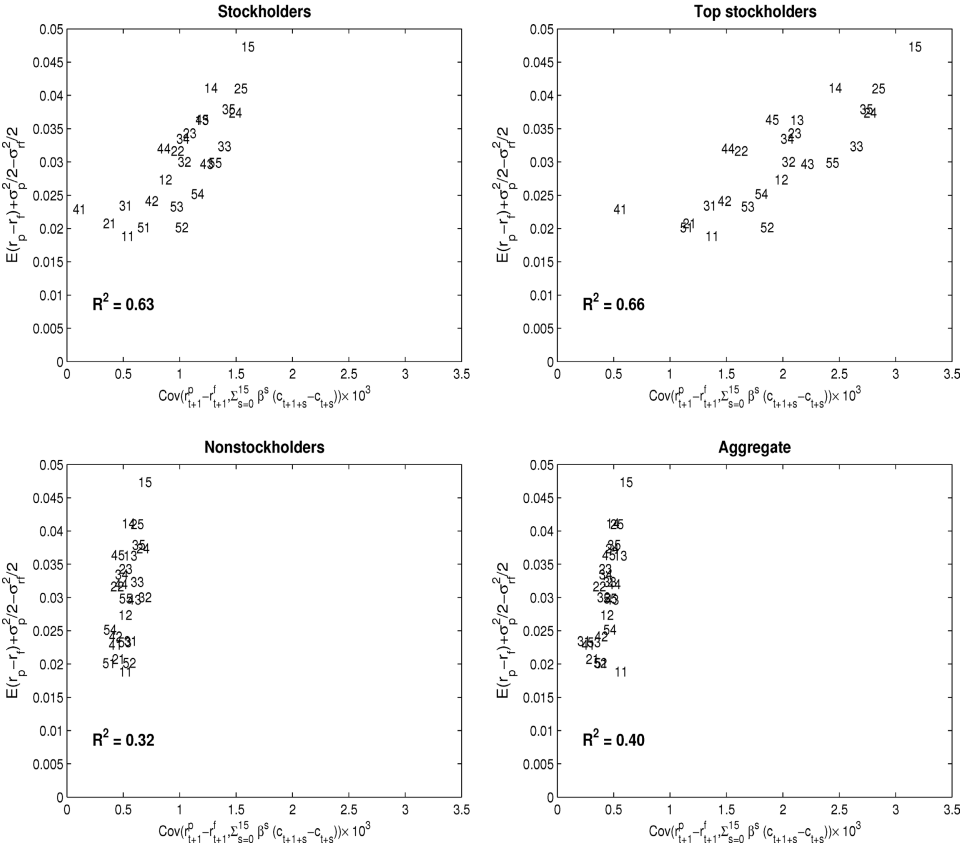


Figure 1. Consumption risk and expected returns on the 25 Fama-French portfolios. Panel A plots the average log excess returns plus half the excess variance of the 25 Fama-French portfolios against the covariance of the log excess returns with long-run (16-quarter, discounted) consumption growth for stockholders, the top third of stockholders, nonstockholders, and aggregate NIPA consumption data. Panel B plots the pricing errors of the 25 Fama-French portfolios with respect to long-run consumption growth for each group. A 45-degree line is added to highlight the pricing errors (vertical distances to the 45-degree line). The entire time-series of returns from July 1926 to November 2004 is used to estimate mean returns, while covariances are calculated over the CEX sample (March 1982–November 2004) except for the aggregate consumption data where the covariances are estimated using data for 1959–2004. Also reported in both panels are the R^2 s from the cross-sectional regressions. Panel B also reports the estimated intercept, α , and risk aversion coefficient, γ , from the cross-sectional regression. The 25 Fama-French portfolios are labeled from 1 to 5 based on growth to value and 1–5 from small to large (e.g., small, growth = 1,1; large, growth = 5,1; small, value = 1,5; and large, value = 5,5).

D. Robustness

In this section, we demonstrate that our results are robust to different estimation methods, various definitions of stockholder consumption, and the aggregation method used to compute each group's average consumption growth rates.

Panel B: Pricing errors

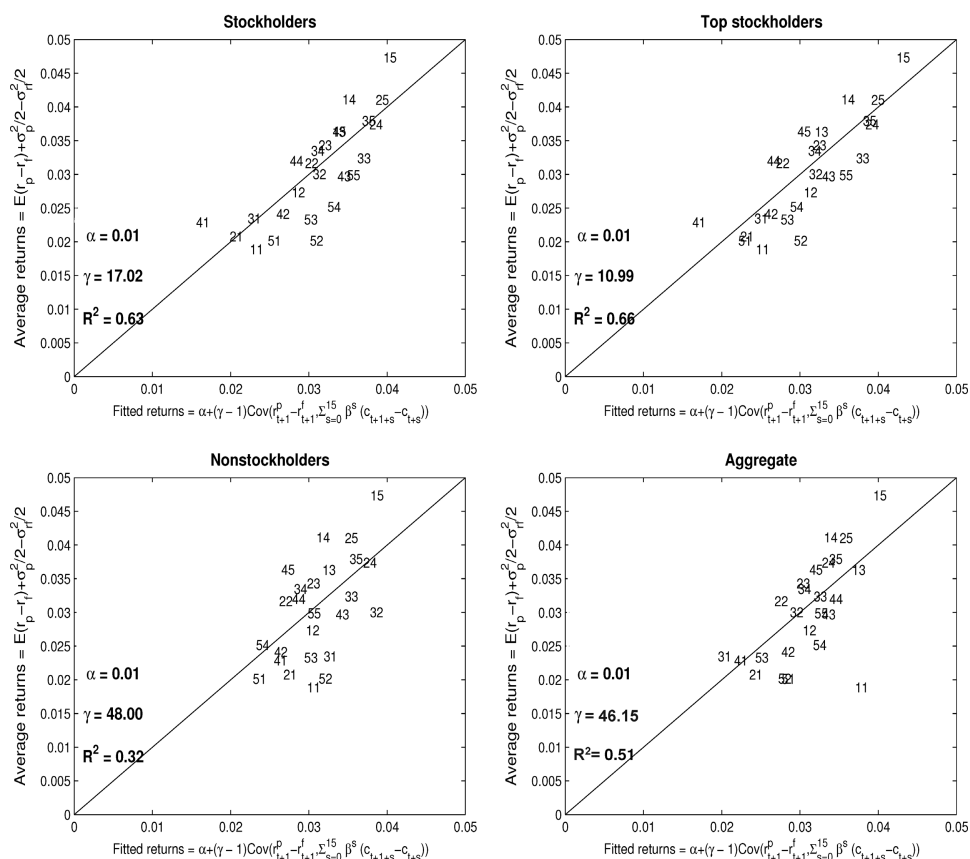


Figure 1. Continued

D.1. Alternative Estimation Methods

Panel A of Table IV reports results from the Euler equation estimation when we force the constant term to be 0, thereby forcing the model to also price the average level of the 25 assets' excess returns (i.e., approximately, the equity premium). This specification imposes the theoretical restriction of the model that the intercept should be zero and hence provides a more stringent test of the model. As Panel A of Table IV shows, risk aversion estimates are higher under this specification, but still quite similar to those obtained when the intercept is a free parameter. For stockholders, risk aversion estimates are between 30 and 32, for the top stockholders they are about 16, and for nonstockholders they are between 65 and 67. Hence, results are roughly similar when imposing this additional theoretical restriction.

Panel B of Table IV reports results from the Euler equation estimation when we estimate the mean and variance of returns over the same sample period as the consumption growth covariance estimates (1982–2004). The disadvantage

Table IV

Robustness of Estimation, Stockholder Definition, and Aggregation

Estimates of the Euler equation based on the long-run (discounted) consumption growth of stockholders, the top third of stockholders, and nonstockholders at 16-quarter horizons are reported across different estimation methods, stockholder definition, and aggregation method. Long-run consumption growth from quarter t to $t + 16$ is calculated using data from the Consumer Expenditure Survey over the period March 1982–November 2004, assuming a discount rate of 5% per year (quarterly discount factor $\beta = 0.95^{1/4}$). Results are presented for regressions of average log excess returns plus half the excess variance on consumption covariances (“Forward” regressions) and for regressions of consumption covariances on average log excess returns plus half the excess variance (“Reverse” regressions). In the table α denotes the intercept from the regressions and γ is the implied risk aversion coefficient calculated from the regression estimates. Intercept values from the reverse regressions are reported $\times 100$ for readability. Estimation is performed by OLS, which is equivalent to GMM with t -statistics (in parentheses) computed using the GMM approach in Appendix C, which accounts for the different sample lengths used to estimate means and covariances, cross-correlated residuals, first-stage estimation error in the covariances, and consumption growth autocorrelation. R^2 s from the cross-sectional regressions are also reported. Panel A reports results for regressions that force the intercept to be zero. Panel B reports results where the mean and variance of asset returns are estimated over the same time period as the CEX sample. Panel C reports results from an alternative definition of stockholders using the predicted probability of being a stockholder from a probit analysis conducted in the SCF. Panel D reports estimates using a representative agent aggregation assumption to compute stockholder and nonstockholder consumption growth.

Regression:	Stockholders		Top Stockholders		Nonstockholders	
	Forward	Reverse	Forward	Reverse	Forward	Reverse
Panel A: No Intercept						
α	0	0	0	0	0	0
	—	—	—	—	—	—
γ	30.35	31.89	16.35	16.85	64.61	67.35
	(2.96)	(1.89)	(3.12)	(3.45)	(3.40)	(3.81)
Panel B: Mean Returns Estimated over CEX Sample Period (1982–2004)						
α	0.02	0.01	0.01	0.06	0.02	0.04
	(1.51)	(0.12)	(1.88)	(0.70)	(1.00)	(1.46)
γ	13.40	30.57	7.54	21.64	22.59	289.11
	(3.96)	(4.77)	(3.25)	(1.53)	(0.91)	(0.15)
R^2	0.42		0.32		0.07	
Panel C: Alternative Definition of Stockholders: CEX and Probit from SCF						
α	0.02	−0.02	0.01	−0.04	0.01	0.03
	(1.30)	(−0.80)	(1.06)	(−0.85)	(2.57)	(2.08)
γ	13.49	25.07	8.38	11.80	29.56	102.84
	(2.59)	(1.54)	(3.07)	(2.08)	(1.96)	(0.95)
R^2	0.52		0.68		0.28	
Panel D: Representative Agent Aggregation						
α	0.02	−0.03	0.01	0.00	0.01	0.01
	(1.56)	(−0.96)	(0.98)	(0.02)	(0.79)	(0.80)
γ	12.11	19.99	9.48	14.13	41.00	62.46
	(3.50)	(2.24)	(3.29)	(2.12)	(3.99)	(2.56)
R^2	0.59		0.65		0.65	

of this approach is that, because mean returns are difficult to estimate, by shortening the sample considerably we introduce additional noise into our estimates. As Panel B of Table IV shows, the results are quite similar to those obtained in Table II, where the entire time series of returns is used to estimate the mean and variance of returns.

D.2. Alternative Definition of Stockholder Status

Panel C of Table IV reports results using the alternative definition of stockholders, which combines information from the CEX with a probit analysis from the SCF to classify households as stockholders and nonstockholders. The tighter definition of stockholders and the top third of stockholders improves the results slightly. Risk aversion coefficients are a little lower under the more precise stockholder definition. Nonstockholder consumption growth, however, continues to exhibit a weaker relation with returns and requires much larger risk aversion to match returns.

D.3. Aggregation Method

Panel D of Table IV reports results when the representative agent aggregation from equation (10) is used to compute the average consumption growth series of each group. Aggregating households under a representative agent assumption improves the results, generating even lower risk aversion coefficients for stockholders. Risk aversion estimates for stockholders decline from 17.0 under the baseline aggregation to 12.1 under the representative agent aggregation, when running the regression forward, and from 26.6 to 20.0 when running the regression in reverse. For the top stockholders, risk aversion declines from 11.0 to 9.5 (16.1 to 14.1) when running the forward (reverse) regression. For nonstockholders, risk aversion is still large at 41–62 under the new aggregation. Hence, results are robust to the aggregation method employed.

The baseline, and theoretically consistent, aggregation that averages log consumption growth rates not only captures the growth rate in average consumption for the group, but also heterogeneity in consumption growth across households in the group. By comparing our baseline results to those under a representative agent aggregation, which captures only the average consumption effect, we can gauge the extent to which our baseline results are affected by consumption dispersion within a given group (e.g., stockholders). The finding that stockholder risk aversion estimates are slightly lower using the representative agent (representative stockholder) approach suggests that our baseline results are driven by the covariances of returns with average stockholder consumption growth and not the covariances of returns with the dispersion in consumption growth across households (Mankiw (1986), Constantinides and Duffie (1996)).

E. Covariance of Conditional Expectations

So far we have ignored the covariance of the conditional expectation of discounted consumption growth and the conditional expectation of the excess asset

Table V
Euler Equation Estimation with Time-Varying Conditional Expectations and EIS Values Different from One

The table reports estimates of the Euler equation based on the long-run (discounted) consumption growth of stockholders, the top third of stockholders, and nonstockholders, and on aggregate consumption from NIPA. Estimates take into account time-varying conditional expectations and allow for values of the EIS different from one. Estimation of both the intercept, α , and implied risk aversion parameter, γ , are reported. Panel A reports results for the case where $EIS = 1$ using return innovations to compute conditional covariances instead of a VAR for consumption dynamics. Return innovations are defined with respect to 16 quarterly lags of the excess return on the asset itself, one lag of the real 3-month Treasury bill rate, and one lag of the log price-dividend ratio on the S&P 500 index. Panel B reports results for the same $EIS = 1$ case using a VAR for consumption dynamics that places restrictions on the coefficient estimates as described in Appendix A. Panel C reports results using the same VAR for the case where the $EIS = 1.5$ and Panel D for the case where $EIS = 0.5$. The VARs are estimated for each portfolio individually using 16 quarterly lags of the excess return on the asset itself, one lag of the real 3-month Treasury bill rate, and one lag of the log price-dividend ratio on the S&P 500 index. Standard errors used to compute t -statistics (reported in parentheses) are computed using a block bootstrap approach that resamples in 16-quarter blocks 5,000 times.

	Stockholders	Top Stockholders	Nonstockholders	Aggregate
Panel A: Return Innovations ($EIS = 1$)				
α	0.02 (2.37)	0.02 (1.71)	0.01 (0.74)	0.01 (0.54)
γ	13.41 (3.92)	9.29 (3.31)	42.15 (2.88)	48.45 (1.53)
Panel B: VAR ($EIS = 1$)				
α	0.01 (2.50)	0.02 (2.97)	0.02 (2.91)	0.02 (5.70)
γ	12.54 (2.89)	6.87 (2.06)	30.21 (2.64)	41.01 (2.01)
Panel C: VAR ($EIS = 1.5$)				
α	0.02 (2.71)	0.02 (2.46)	0.02 (2.88)	0.02 (3.49)
γ	10.84 (2.69)	6.16 (2.21)	27.93 (2.51)	37.35 (3.15)
Panel D: VAR ($EIS = 0.5$)				
α	0.01 (2.77)	0.02 (3.66)	0.02 (3.47)	0.02 (3.26)
γ	14.10 (2.39)	7.15 (2.27)	24.45 (2.79)	43.47 (3.11)

returns in our estimation. Estimating the conditional moments of consumption is difficult in our short sample of CEX data. To reduce estimation error, we use return innovations as in equation (7) to estimate conditional covariances for the $EIS = 1$ case because returns data go back to 1926.

Panel A of Table V reports the results from using return innovations and an $EIS = 1$ to estimate the Euler equation. For brevity, we focus on the results

from the forward regression (equation (11)). Standard errors used to compute t -statistics are calculated using a block bootstrap that resamples the data in 16-quarter blocks 5,000 times. For stockholder consumption growth, the estimated risk aversion needed to match the cross-section of returns is 13.4, which is smaller than the 17.0 point estimate we obtain in Table II with 16-quarter consumption growth rates under our baseline estimation, which does not account for covariance of conditional expectations. Similarly, the required risk aversion for top stockholder consumption growth drops from 11.0 under our baseline approach to 9.3 when accounting for covariance of conditional expectations. For nonstockholders, however, we still need large risk aversion parameters to be able to match return moments.

F. Euler Equation Estimation When the EIS $\neq 1$

When the EIS $\neq 1$, it is necessary to estimate consumption dynamics using a VAR model in order to construct the additional term in the stochastic discount factor following HHL. Given the short history of CEX consumption data, including all test asset returns and their lags in the VAR is not feasible. Instead, we estimate a separate VAR for each asset as outlined in Appendix A.

To first gauge whether the asset-by-asset VAR (and imposing six identifying restrictions outlined in Appendix A) significantly affects the results, we estimate this series of VARs for the EIS = 1 case, where we can compare it to our other estimates. Panel B of Table V reports results from using a separate VAR for each test asset when the EIS = 1 and adding the six restrictions on the dynamics of the VAR. The same block bootstrap procedure is used to calculate standard errors as above. As Panel B of Table V indicates, the separate VAR and the additional restrictions do not matter for the point estimates of risk aversion, which are similar to those obtained previously for each consumption series.

Panel C of Table V reports results from our implementation of HHL's stochastic discount factor for an EIS = 1.5. As Section I details, when the EIS differs from one, an additional term enters the Euler equation and this could alter our estimate of γ . However, the test asset returns may not differ significantly in their covariance with this additional term, so it may only have a small effect on estimated risk aversion when pricing the cross-section of returns. Confirming this intuition, Panel C of Table V reports that for an EIS of 1.5 (the value used by Bansal and Yaron (2004)), the point estimates of γ are very similar to those obtained when the EIS = 1 for each consumption series.

Likewise, Panel D of Table V, which reports results for an EIS = 0.5, shows that the results are nearly identical to those for an EIS = 1 or 1.5. When estimating the cross-sectional price of risk, the value for the EIS makes little difference for the resulting risk aversion estimates.

IV. Consumption Growth Factor-Mimicking Portfolios

Because measuring long-run risks in the short CEX sample is challenging, another way to assess the importance of long-run stockholder consumption growth

for asset pricing is to construct factor-mimicking portfolios that allow a longer time series of data to be constructed to improve estimation of long-run risks. Factor-mimicking portfolios are tradeable assets designed to be maximally correlated with long-run stockholder consumption growth. These tradeable portfolios can generate a longer time series of data and contain market prices. In addition to the benefits of more data, if measurement error in consumption is uncorrelated with the asset returns used to construct the factor-mimicking portfolio, then the factor portfolio may also contain less measurement error than actual consumption growth. To construct factor-mimicking portfolios, we project the discounted consumption growth series on a set of instruments (available over a longer period) and use the estimated coefficients to construct a longer time series of instrumented stockholder consumption growth.

A. Constant Factor Loadings

Following Breeden et al. (1989) and Lamont (2001), we create the consumption growth factor portfolio, *CGF*, by estimating the regression

$$\sum_{s=0}^{15} \beta^s \left[\frac{1}{H_{t+s}^g} \sum_{h=1}^{H_{t+s}^g} (c_{t+1+s}^{h,g} - c_{t+s}^{h,g}) \right] = a + b' R_{t+1} + \eta_{t+1}, \quad (13)$$

where $c_t^{h,g}$ is the log of consumption of household h in group g for quarter t , H_t^g is the number of households in group g in quarter t , R_{t+1} are the quarterly excess returns over the riskless rate on the base assets (instruments), and $\beta = 0.95^{1/4}$. We use returns, as opposed to log returns, in this regression so that the coefficients b are easily interpretable as the weights in a zero-cost portfolio. The return on the portfolio *CGF* is

$$CGF_{t+1} = b' R_{t+1}, \quad (14)$$

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. Moreover, equation (14) is not limited to the CEX sample period if the vector b is relatively stable over time, an assumption that we discuss below.

Regression (13) is estimated using data from March 1982 to November 2004, using the 16-quarter discounted CEX stockholder, top stockholder, and non-stockholder consumption growth rates. The first set of instruments we use are four portfolios that span the size and value dimensions of the 25 Fama and French portfolios. Specifically, we use a small growth portfolio (average of the two smallest size, two lowest BE/ME portfolios from the 25 Fama–French portfolios), a large growth portfolio (average of the two largest size, two lowest BE/ME portfolios), a small value portfolio (average of the two smallest size, two highest BE/ME portfolios), and a large value portfolio (average of the two largest size, two highest BE/ME portfolios) as instruments. The excess returns

on these portfolios are available from July 1926 to November 2004.¹¹ The results from this first-stage regression, with Newey and West (1987) *t*-statistics allowing for autocorrelation up to order 47 (months), are reported in Panel A of Table VI. 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 stronger for the top third of stockholders and are much weaker (in terms of the factor loadings) for nonstockholder consumption growth.

Using the coefficient estimates from the first-stage regression equation (13) to construct the *CGF* in equation (14), we generate a factor-mimicking portfolio for consumption growth over the longer time period from July 1926 to November 2004. The sample average return of each *CGF* and its *t*-statistic are reported across the different consumption series. The stockholder *CGF* produces an average 145 basis points per quarter by essentially going long small value and short small growth. The top stockholder *CGF* almost doubles its exposure to small value and small growth and hence produces 260 basis points per quarter, whereas the nonstockholder and aggregate consumption *CGFs* do not load heavily on these portfolios and hence only produce 47 and 34 basis points per quarter, respectively. These results further highlight the strength of the relation between stockholder consumption and asset prices—assets with high comovement with stockholder consumption should have high average returns in equilibrium.

We then estimate the Euler equation for the 25 Fama–French portfolios in the second stage by replacing actual consumption growth with its *CGF* returns. Standard errors used to compute *t*-statistics are calculated via a block bootstrap procedure that resamples consumption growth, base asset return data, and test asset return data using 16-quarter blocks and then recomputes the first-stage coefficients for the *CGF* and the risk aversion coefficient from the second-stage Euler equation that uses the bootstrapped *CGF*. This procedure, therefore, also accounts for first-stage estimation error in the *CGF*. The block bootstrap procedure is run using 5,000 replications.

The second-stage Euler equation estimates based on covariances for the CEX period 1982–2004 are shown in the middle of Panel A of Table VI. The estimated risk aversion coefficients are very similar to those obtained using actual consumption growth data in Table II. The reliability of the estimates is also similar despite the additional estimation error introduced from the first stage. This additional error may be mitigated by a reduction in measurement error from using returns instead of actual consumption growth.

¹¹ Breeden et al. (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. Because, with a limited time series, we cannot obtain 25 reliable coefficients, and because the four size and value portfolios we use 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.

Table VI
Consumption Growth Factor-Mimicking Portfolios

Estimates of the Euler equation using the long-run (discounted) consumption growth of stockholders, the top third of stockholders, nonstockholders, and aggregate consumption using consumption growth factor-mimicking (*CGF*) portfolios are reported. Panel A reports results for a *CGF* formed by regressing 16-quarter (discounted) consumption growth on a constant and the quarterly excess returns (over the T-bill rate) of a small growth (average of the two smallest size, two lowest BE/ME portfolios from the 25 Fama–French portfolios), a large growth (average of the two largest size, two lowest BE/ME portfolios), a small value (average of the two smallest size, two highest BE/ME portfolios), and a large value (average of the two largest size, two highest BE/ME portfolios) portfolio. The regression is estimated over the CEX sample period (March 1982–November 2004) and the coefficients are then used to construct a *CGF* portfolio for July 1926 to November 2004. Coefficient estimates (factor loadings) from the first stage are assumed to be constant over time. Also reported in Panel A are the mean returns (and *t*-statistics) on the *CGF* portfolios. Panel B reports results allowing for time-variation in factor loadings by interacting the four size and value portfolios with Lettau and Ludvigson’s (2001a) consumption-to-wealth ratio *cay* (linearly interpolated between quarters to produce monthly values). The consumption-to-wealth ratio proxies for the fraction of consumption consumed by stockholders and is available over a longer time period than CEX consumption data. We model factor loadings as having a constant component and a time-varying component that is a function of *cay*. Panel C reports results from a *CGF* created by projecting stockholder and nonstockholder consumption growth on aggregate consumption growth. Each panel also reports the first-stage regression parameters that determine the factor loadings in constructing each *CGF*, including Newey and West (1987) *t*-statistics with an adjustment of 47 lags and the R^2 from the regression. Each panel also reports the intercept (α) and the risk aversion estimate (γ) from the Euler equation in the second stage that employs the *CGF* in place of actual consumption growth. Intercept values for the reverse regressions are reported $\times 100$ for readability. Standard errors used to compute *t*-statistics are calculated from a block bootstrap procedure that resamples 16-quarter blocks 5,000 times and repeats both the first-stage *CGF* construction and second-stage Euler equation estimation in each bootstrap simulation.

	Stockholders	Top Stockholders	Nonstockholders	Aggregate				
Panel A: <i>CGF</i> from Size and Value Portfolios with Constant Factor Loadings								
	First-stage estimates of weights for <i>CGF</i>							
Small, growth	−0.32 (−1.71)	−0.58 (−1.91)	−0.14 (−1.69)	−0.06 (−0.87)				
Large, growth	0.02 (0.11)	0.07 (0.25)	0.11 (1.08)	0.01 (0.09)				
Small, value	0.56 (1.42)	1.08 (2.02)	0.20 (1.68)	0.08 (0.76)				
Large, value	−0.07 (−0.40)	−0.25 (−0.69)	−0.10 (−1.05)	0.05 (0.47)				
R^2	0.05	0.04	0.02	0.04				
Mean return (1926-2004)	1.45 (6.45)	2.60 (6.56)	0.47 (7.84)	0.34 (5.84)				
Second-Stage Euler Equation Estimates Using <i>CGF</i>								
Regression:	Forward	Reverse	Forward	Reverse	Forward	Reverse	Forward	Reverse
<i>CEX sample period March 1982–November 2004</i>								
α	0.01 (1.28)	−0.01 (−0.95)	0.01 (1.41)	−0.01 (−0.76)	−0.01 (−0.16)	0.01 (4.27)	0.02 (2.71)	0.00 (0.95)
γ	17.02 (3.19)	26.05 (2.21)	11.01 (4.29)	15.21 (3.02)	68.76 (6.40)	136.98 (3.33)	46.78 (3.75)	133.91 (3.12)
R^2	0.64		0.70		0.50		0.34	

(continued)

Table VI—Continued

Panel A: <i>CGF</i> from Size and Value Portfolios with Constant Factor Loadings								
Second-Stage Euler Equation Estimates Using <i>CGF</i>								
Regression:	Forward	Reverse	Forward	Reverse	Forward	Reverse	Forward	Reverse
<i>Entire sample period July 1926–November 2004</i>								
α	0.01 (1.50)	0.11 (1.23)	0.01 (1.79)	0.17 (0.95)	0.01 (0.27)	0.04 (2.47)	0.00 (1.99)	0.05 (1.35)
γ	6.30 (6.33)	9.39 (5.39)	3.99 (7.17)	5.79 (5.13)	25.87 (4.99)	40.27 (4.80)	21.20 (6.34)	36.07 (5.55)
R^2	0.63		0.62		0.63		0.58	
			Stockholders	Top Stockholders	Nonstockholders		Aggregate	
Panel B: <i>CGF</i> from Size and Value Portfolios with Time-Varying Factor Loadings								
First-stage estimates of weights for <i>CGF</i>								
Small, growth	−0.53 (−2.60)		−0.76 (−2.27)		−0.21 (−2.14)		−0.13 (−1.43)	
$SG \times cay$	−0.18 (−1.44)		−0.37 (−1.71)		−0.06 (−0.86)		−0.07 (−1.12)	
Large, growth	0.07 (0.34)		0.02 (0.06)		0.10 (0.78)		0.02 (0.17)	
$LG \times cay$	−0.13 (−0.59)		0.00 (0.00)		−0.10 (−1.12)		−0.02 (−0.26)	
Small, value	0.73 (1.87)		1.23 (2.42)		0.26 (2.01)		0.13 (1.06)	
$SV \times cay$	0.25 (1.06)		0.53 (1.42)		0.12 (1.18)		0.08 (0.92)	
Large, value	−0.04 (−0.21)		−0.15 (−0.40)		−0.05 (−0.45)		0.08 (0.66)	
$LV \times cay$	0.01 (0.07)		−0.25 (−0.86)		0.04 (0.51)		0.04 (0.49)	
cay	−0.02 (−1.37)		0.00 (0.21)		−0.00 (−1.03)		−0.01 (−0.89)	
R^2	0.08		0.05		0.05		0.07	
Second-Stage Euler Equation Estimates Using <i>CGF</i>								
Regression:	Forward		Forward		Forward		Forward	
α	0.00 (1.03)		0.00 (0.49)		−0.02 (−2.03)		−0.01 (−2.71)	
γ	22.84 (6.78)		12.03 (6.30)		108.25 (5.32)		95.53 (5.34)	
R^2	0.88		0.76		0.70		0.75	
Panel C: <i>CGF</i> from Aggregate Consumption Growth								
First-stage estimates of weights for <i>CGF</i>								
Aggregate consumption	2.68 (5.41)		3.94 (5.87)		1.01 (3.94)		—	
R^2	0.50		0.24		0.27		—	

(continued)

Table VI—Continued

Panel C: CGF from Aggregate Consumption Growth				
Regression:	Forward	Forward	Forward	
Second-stage Euler equation estimates using CGF				
α	0.01 (0.40)	0.01 (0.40)	0.01 (0.40)	—
γ	20.76 (3.69)	14.45 (3.69)	53.32 (3.69)	—
R^2	0.40	0.40	0.40	—

The second-stage Euler equation estimates based on covariances for the full 1926–2004 period at the bottom of Panel A of Table VI show considerably lower implied risk aversion estimates. For the stockholder CGF, implied risk aversion based on the forward regression drops from 17.0 using actual consumption growth (Table II) to 6.3 using the CGF. For the top third of stockholders, implied risk aversion drops from 11.0 using actual consumption growth (Table II) to 4.0 using the CGF. We also report results for the CGF of nonstockholder and aggregate consumption growth, obtaining risk aversion estimates of 25.9 and 21.2, respectively. These estimates are also lower than those obtained with actual consumption growth, but they are still much higher than those obtained for stockholders.

B. Time-Varying Factor Loadings

General equilibrium models of limited participation (e.g., Basak and Cuoco (1998), Guvenen (2007), and Gomes and Michaelides (2008)) imply that time variation in stock market participation rates or the fraction of wealth owned by stockholders predict time variation in stockholder consumption–return covariances and in expected excess returns.

We assess empirically the effect of allowing time-varying factor loadings by using a specification of the following form to construct the consumption growth factor–mimicking portfolios:

$$\sum_{s=0}^{15} \beta^s \left[\frac{1}{H_{t+s}^g} \sum_{h=1}^{H_{t+s}^g} (c_{t+1+s}^{h,g} - c_{t+s}^{h,g}) \right] = a + dz_t + b'R_{t+1} + c'z_t R_{t+1} + \eta_{t+1}, \quad (15)$$

where z_t is a variable capturing time variation in factor loadings. The ideal variable to use as z would be a measure of the fraction of wealth owned by, or the fraction of consumption consumed by, stockholders. This variable can only be estimated with CEX data, thus limiting the sample. However, we argue that the aggregate consumption-to-wealth ratio cay of Lettau and Ludvigson (2001a) may be quite closely related to the stockholder consumption share.

First, microlevel evidence suggests that wealthy households, who tend to be stockholders, have higher savings rates than poorer households (Bosworth et al. (1991), Carroll (2000), Dynan, Skinner, and Zeldes (2004)). Hence, when stockholder wealth increases relative to nonstockholder wealth, the aggregate consumption-to-wealth ratio will decline, providing a link between stockholder wealth (or consumption) shares and the aggregate consumption-to-wealth ratio. Higher savings rates for stockholders and nonstockholders could be driven by stockholders having a higher value of EIS, consistent with the findings of Vissing-Jørgensen (2002).

Second, we can confirm that *cay* and the stockholder consumption share are quite strongly negatively correlated over the CEX period for which we have data on both. Figure 2 plots the aggregate consumption-to-wealth ratio *cay* against the ratio of quarterly consumption of stockholders to aggregate quarterly consumption in the CEX (the stockholder consumption share), calculated using CEX survey weights. The figure highlights that the stockholder consumption share varies over time in a manner that mirrors the dynamics of the consumption-to-wealth ratio, generating a correlation of -0.44 . This evidence suggests that the consumption-to-wealth ratio may be linked to stock market participation and stockholder consumption, possibly providing an economic story for its empirical success in pricing assets (Lettau and Ludvigson (2001a, 2001b)).

Based on this evidence, we employ *cay*, linearly interpolated between quarters to produce monthly values, as our variable z_t to capture time variation in factor loadings. We estimate equation (15) over the CEX sample period March 1982–November 2004 and apply the coefficients (d , b , and c) to return and *cay* data that are available from December 1951 to November 2004 to form a longer time series for the *CGF*. We then employ this *CGF*, whose factor loadings on the four stock portfolios vary over time as determined by the dynamics of *cay*, in the Euler equation estimation.

The first column of Panel B of Table VI reports the first-stage estimation results from equation (15) for stockholder consumption growth. The bottom of Panel B of Table VI reports the subsequent second-stage Euler equation (forward regression) estimates for the time-varying *CGF* for stockholders. The coefficient of risk aversion for the stockholder *CGF* is 22.8, which is roughly in line with our earlier estimates using actual consumption data. To assess whether time variation in the factor loadings or the different sample periods used is contributing to the results, we also compute risk aversion from the constant-loading *CGF* over the same December 1951–November 2004 sample period to coincide with the availability of *cay*. The risk aversion estimate we obtain is almost the same as that generated from the time-varying *CGF* over the same period.

The remaining columns of Panel B of Table VI report results for the time-varying *CGFs* for top stockholder, nonstockholder, and aggregate consumption growth. The findings are similar: For each group, risk aversion estimates are similar in magnitude to actual consumption growth and the constant loading *CGF* over the same sample period. This evidence indicates that accounting for

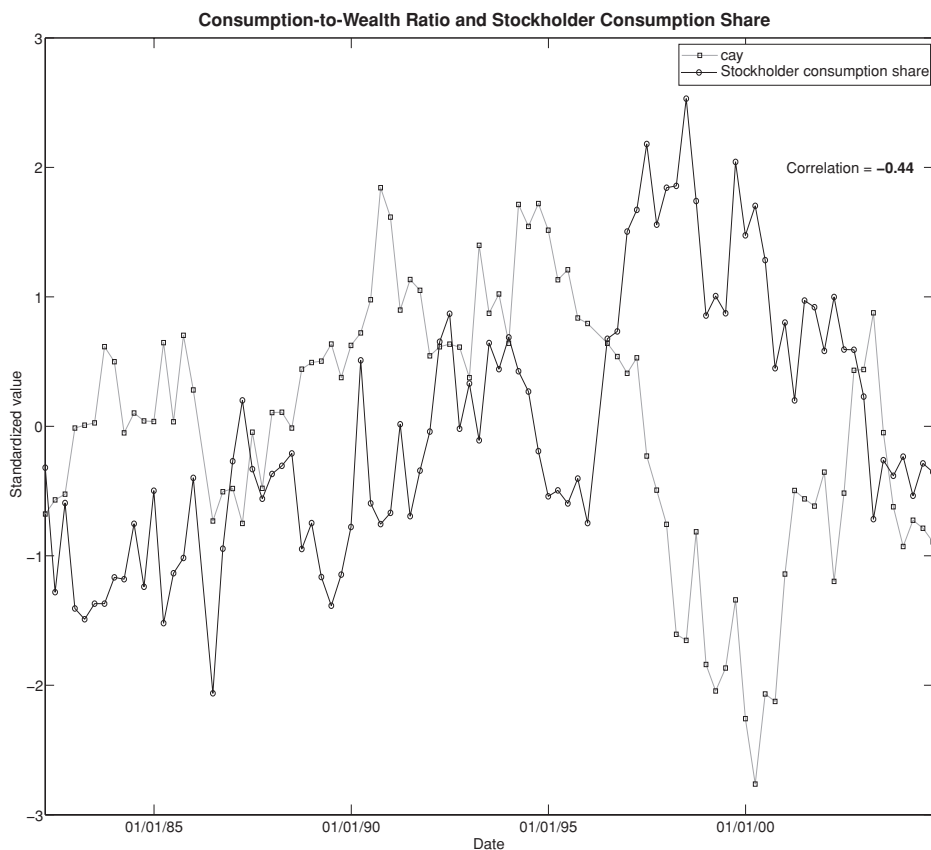


Figure 2. Plot of the consumption-to-wealth ratio *cay* and the stockholder consumption share. The figure plots the consumption-to-wealth ratio of Lettau and Ludvigson (2001a) along with the ratio of the quarterly consumption of stockholders (using our baseline stockholder definition) to total quarterly CEX consumption. For readability, each time series is standardized by subtracting its mean and dividing by its standard deviation calculated over the CEX sample period.

time variation in factor loadings from 1951 to 2004 does not seem to alter our estimates.

The results in Panel B of Table VI suggest that the lower risk aversion estimates obtained based on our constant-loading *CGF* in Panel A over the longer time period (1926–2004) are not primarily driven by bias due to assuming constant factor loadings, but rather by covariances between returns and consumption being larger in the earlier period. Fama and French (2007) find that the covariance of value (growth) stocks with the market portfolio is much larger (smaller) in the pre-war data and Parker (2001) finds that asset returns covary more strongly with aggregate consumption in the pre-CEX period. Investigating why the covariance of asset returns with macroeconomic variables appears larger in the first half of the 20th century is beyond the scope of this paper.

C. Another Instrument

Panel C of Table VI reports results for *CGFs* instrumented with aggregate consumption instead of the size and value portfolios. Risk aversion estimates from the second-stage Euler equation (forward regression) estimation for stockholders, top stockholders, and nonstockholders are 20.8, 14.5, and 53.3, respectively, using the *CGF* associated with aggregate consumption from 1959 to 2004. These magnitudes are consistent with our previous results. As shown in Table I and in the first-stage estimates here, stockholders are almost three times as sensitive to aggregate consumption growth as nonstockholders and the top third of stockholders are almost four times as sensitive to aggregate consumption growth as nonstockholders. The heightened sensitivity of stockholder consumption to aggregate shocks is precisely why we obtain much lower risk aversion estimates. This particular set of results also highlights why R^2 s may not be a good metric to judge model success. In this case, the second-stage R^2 s are identical across the groups because the *CGFs* are constructed from one identical variable. Hence, the only comparison to make across groups is the difference in required risk aversion to explain returns.

V. Other Test Assets

We estimate the Euler equation for stockholder and nonstockholder consumption growth using three different sets of other test assets instead of the 25 Fama–French portfolios to examine whether our results are sensitive to the specific cross-section of average returns chosen as the testing ground. This exercise provides robustness for the empirical success of our model of long-run stockholder consumption risk and addresses some of the criticism of the recent consumption asset pricing literature (e.g., Lewellen et al. (2006), Daniel and Titman (2006)). Using the aggregate market equity return, a cross-section of eight Treasury bond portfolios, and even the entire cross-section of individual stock returns, we find remarkably consistent estimates of the risk aversion parameter across these various testing grounds. Not only do the qualitative results hold, namely, that risk aversion estimates are lowest for the wealthiest stockholders and highest for nonstockholders, but, more notably, the quantitative values of risk aversion are quite similar.

A. The Equity Premium

Panel A of Table VII reports implied measures of risk aversion using actual stockholder, top stockholder, nonstockholder, and aggregate consumption growth. Implied risk aversion coefficients are computed from equation (11) with the intercept set to zero. The first row of Panel A of Table VII uses the (quarterly) excess log return on the CRSP value-weighted index as the single test asset. We report t -statistics for the risk aversion estimate (in parentheses) that are calculated from a block bootstrap procedure that resamples 5,000 times, using 16-quarter blocks. To minimize the effect of estimation uncertainty in the equity premium, all risk aversion coefficients are calculated using the equity

Table VII
Other Test Assets: The Equity Premium, Bond Returns,
and Individual Stocks

The estimated risk aversion from the Euler equation for an $EIS = 1$ that ignores the covariance between the conditional expectations of returns and consumption growth is reported for different sets of test assets. Panel A reports implied measures of risk aversion for the covariance of asset returns with long-run (16-quarter, discounted) consumption growth for stockholders, top stockholders, and nonstockholders from the CEX over the period 1982–2004, as well as aggregate consumption from NIPA over the CEX period 1982–2004. The test assets used are the equity premium (excess return on the CRSP value-weighted index) and eight Treasury bond portfolios with maturities of 3 months, 1 year, 2 years, 5 years, 7 years, 10 years, 20 years, and 30 years. Panel B reports risk aversion estimates from the Euler equation estimated using the constant-loading *CGF*s in place of actual consumption growth. The test assets used are the equity premium and the entire cross-section of individual stocks. The cross-section of individual stocks Euler equation is estimated in the style of Fama and MacBeth (1973), where each month the log excess returns plus one half the excess variance of all NYSE, Amex, and NASDAQ stocks with share prices above \$5 is regressed on the covariance of the log excess return with the consumption growth factor—mimicking portfolio *CGF* for either stockholder, top stockholder, nonstockholder, or aggregate consumption growth and the time-series average of the monthly coefficient estimates and their time-series *t*-statistics are reported. Covariances are estimated using a procedure similar to the pre- and postranking beta procedure of Fama and French (1992). The equity premium and individual stock estimates cover the full return period from July 1926 to November 2004. Except for the individual stock estimates, the *t*-statistics (in parentheses) of the risk aversion estimates are calculated using a block bootstrap method that resamples 5,000 times 16 quarters at a time.

Test Assets	Stockholder	Top Stockholder	Nonstockholder	Aggregate
Panel A: Actual Consumption Data				
Equity premium	25.67 (2.41)	15.10 (2.85)	49.34 (2.15)	69.36 (5.57)
Bond portfolios	13.36 (2.39)	6.90 (2.40)	31.43 (1.98)	81.27 (1.52)
Panel B: <i>CGF</i> over Full Return Period				
Equity premium 1926–2004	8.16 (5.15)	5.20 (5.58)	26.71 (5.89)	24.15 (4.80)
Individual stocks 1926–2004	9.94 (3.37)	6.02 (3.57)	40.42 (3.51)	29.85 (3.27)

premium estimate from the full September 1926 to November 2004 period of quarterly excess log stock returns.

Our estimates for the equity premium yield risk aversion coefficients of 25.7, 15.1, 49.3, and 69.4 for stockholder, top stockholder, nonstockholder, and aggregate consumption, respectively.¹² These risk aversion values estimated

¹² For robustness, we also calculated Euler equation estimates and implied risk aversion values for $EIS = 1.5$, 1, and 0.5 using the VAR we created from Hansen et al.'s (2008) statistical model. The results across the EIS values are similar and are essentially the same as those from our baseline approach. Hence, while the value of the EIS matters more for the equity premium than it does for the cross-sectional risk premium (as we argued earlier) it still does not have a big effect. Bansal and Yaron (2004) find that the EIS matters significantly for the value of risk aversion in their calibration

from the equity premium are in line with those obtained for the 25 Fama–French portfolios.

The first row of Panel B of Table VII reports results using covariances based on the constant-loading *CGFs* over the full return period of 1926–2004. Required risk aversion of only 8.2 is obtained for the stockholder *CGF* and only 5.2 for the top stockholder *CGF*. For nonstockholder and aggregate consumption *CGFs*, required risk aversion is 26.7 and 24.2, respectively. These risk aversion estimates are similar to those found previously for the cross-section of returns on the 25 Fama–French portfolios over the same sample period. The results suggest that fairly reasonable risk aversion estimates can be obtained using the full sample period and focusing on the long-run consumption risk of stockholders, particularly wealthy stockholders.

B. Cross-section of Bond Returns

The second row of Table VII Panel A reports risk aversion estimates from equation (11), including a constant term, using a cross-section of excess returns (over the return on 1-month T-bills) on eight Treasury bond portfolios with average maturities of 3 months, 1 year, 2 years, 5 years, 7 years, 10 years, 20 years, and 30 years over the period for which CEX data are available (March 1982–November 2004). We use sample mean returns over the CEX period because expected (excess) bond returns are thought to move more than expected (excess) equity returns over time. For the same concern of time variation in expected returns, we also do not use a *CGF* for bond returns that assumes constant factor loadings. Consistent with our findings for the 25 Fama–French portfolios and the equity premium, the estimates of risk aversion are roughly similar to those obtained previously from the other test assets.

Figure 3 plots the average excess bond returns (plus a variance adjustment) against their consumption covariances measured via actual CEX consumption data. The plot highlights the spread in covariance risk across the bond portfolios as well as the high cross-sectional fit.

C. The Entire Cross-section of Individual Stocks

As an additional testing ground for the model, we employ the entire cross-section of all individual stock returns in the last row of Table VII. We estimate

of the equity premium. A key difference between Bansal and Yaron’s (2004) calibration and our estimation is that in their calibration when the EIS changes, *both* the stochastic discount factor and return process change endogenously. Obviously, when we estimate the risk aversion parameter from the data, we take the returns as given. Hence, when the EIS changes, only the stochastic discount factor changes in our estimation procedure because returns data are exogenous inputs in the estimation. In terms of their model, when the EIS changes, both $\beta_{m,e}$ and $\lambda_{m,e}$ change in their equation (7), with most of the action coming via $\beta_{m,e}$ (we checked this through calibration exercises). In our estimation using return data, we are essentially only changing $\lambda_{m,e}$ and therefore conclude that the EIS does not matter much for the risk aversion estimates we obtain from the equity premium. This difference helps reconcile why Bansal and Yaron (2004) reach different conclusions about the importance of the EIS for the risk aversion parameter needed to fit the equity premium.

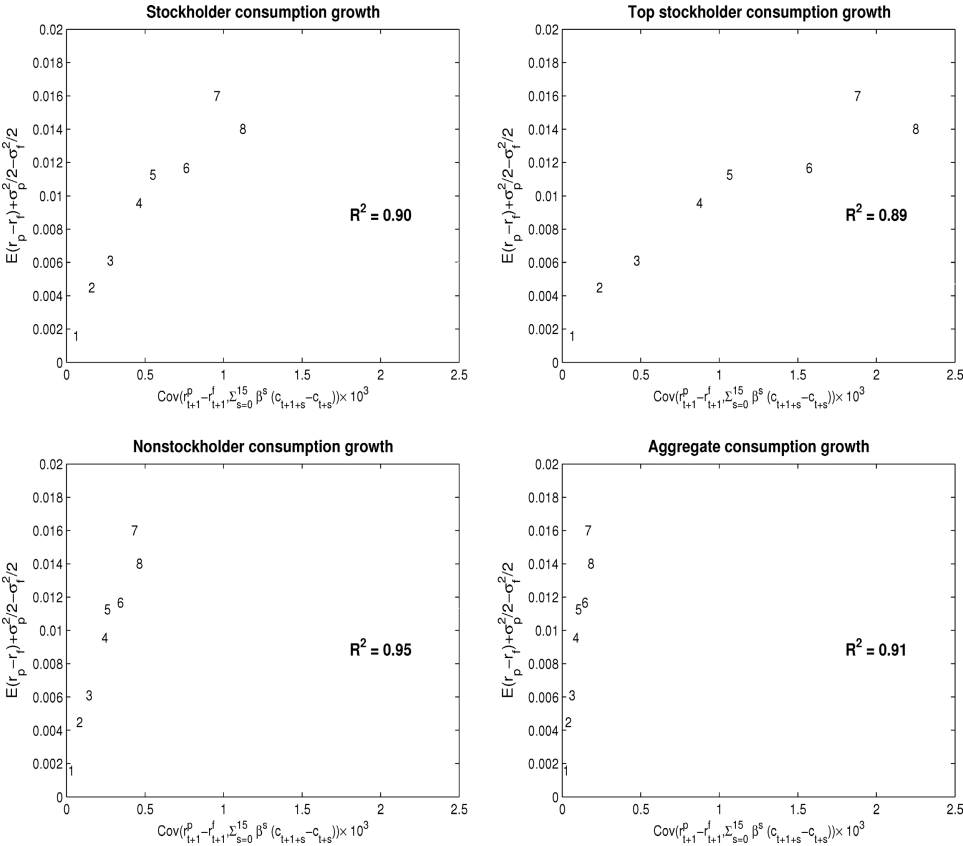


Figure 3. Consumption risk and expected returns on eight Treasury bond portfolios. The graphs plot the average log excess returns (over the return on 1-month T-bills) plus half the excess variance of eight Treasury bond portfolios ranging in maturity from 3 months, 1 year, 2 years, 5 years, 7 years, 10 years, 20 years, and 30 years against the covariance of log excess returns with long-run (16-quarter, discounted) consumption growth for stockholders, the top third of stockholders, nonstockholders, and in aggregate over the period March 1982–November 2004. Also reported are the R^2 s from the cross-sectional regressions. The Treasury bond portfolios are labeled from 1 to 8 in ascending order of maturity.

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 them 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 preranking step). We then compute return covariances for the constructed portfolios over the full data sample and assign these estimates to each stock in a particular portfolio (these covariances are called postranking

covariances).¹³ Fama and French (1992) follow a similar procedure by using size and pre-ranking beta sorted portfolios to form postranking betas in testing the unconditional CAPM. In addition, because of the noise in individual stock returns and difficulty in estimating precise covariances for individual stocks, we only employ the *CGFs* to measure consumption risk because they are available over a much longer sample period than actual CEX consumption data and because the *CGFs* contain only the component of consumption growth correlated with returns, mitigating the effects of measurement error.

We run Fama and MacBeth (1973) month-by-month cross-sectional regressions of the entire cross-section of log excess stock returns on their covariance with long-run consumption growth. 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 using the Newey and West (1987) procedure.¹⁴

The last row of Table VII shows that the covariance of returns with long-run consumption growth of stockholders captures significant cross-sectional variation in average returns. The regression is a direct estimate of the Euler equation (11) (including a constant term) and yields an implied risk aversion coefficient of 9.9, which is roughly similar to the magnitude obtained from the other test assets: 25 Fama–French portfolios, the equity premium, and 8 bond portfolios.¹⁵ For the top third of stockholders, the results are even stronger, requiring a risk aversion of only 6, while for nonstockholders risk aversion is 40 and for aggregate consumption it is 30. Overall, the results and point estimates of risk aversion are reassuringly consistent across the various sets of test assets.

VI. Conclusion

We find empirical support for consumption-based asset pricing by focusing on the long-run consumption risk of stockholders. Long-run stockholder

¹³ Specifically, for each individual stock we estimate the covariance of its returns with *CGFs* using the past 24 to 60 months (as available) of monthly log excess returns before July of year *t*. Stocks are sorted at the end of June into 100 preranking 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 preranking sorts using the full sample of returns from July 1926 to 2004 to obtain *postranking* covariances. The postranking covariance estimate for a given group is then assigned to each stock in the group, with group assignments updated at the end of June of each year. Even though the postranking covariances themselves do not change over time, as an individual stock moves into and out of one of the 100 portfolios due to its preranking covariance changing, that stock will receive a different postranking covariance. This procedure reduces estimation error by shrinking individual covariance estimates to a portfolio average and employing the full sample of data.

¹⁴ A correction for first-stage covariance estimation error via Shanken (1992) has little effect on the standard errors.

¹⁵ Moreover, controlling for other known determinants of returns (market β , log size, and log BE/ME) does not eliminate the significance of consumption risk.

consumption risk captures the return premia associated with size and value portfolios, the aggregate stock market, bond portfolios, and the entire cross-section of stocks, requiring a modest risk aversion coefficient of about 10 for the wealthiest stockholders to match return premia. The stronger link between asset prices and the consumption of households that actually own financial assets is consistent with theory and suggests that recent evidence on the success of long-run consumption risk is unlikely to be due to chance.

The fact that stockholders are more sensitive to aggregate consumption movements helps explain why the consumption risk of stockholders delivers lower risk aversion estimates. Understanding why consumption growth, particularly that of stockholders, responds slowly to news in asset returns will improve our knowledge of what is driving these long-run relations. How do these long-run patterns emerge as an equilibrium outcome? One potential avenue for research is to examine the potential link between long-run stockholder consumption growth, asset returns, and macroeconomic shocks. Evidence from the macroeconomic growth and real business cycle literature highlights that consumption and production respond slowly to technology shocks, peaking at 3- to 4-year horizons (see Altig et al. (2005)), which is, perhaps not coincidentally, about the same horizon over which consumption responds fully to asset returns. Theory also needs to determine why stockholders, in equilibrium, take on more of this aggregate risk. As a first step, Gomes and Michaelides (2008) show that a setting with fixed costs of stock market participation combined with preference heterogeneity can generate equilibrium differences in stockholder and nonstockholder consumption patterns.

Appendix A: Expression and Estimation for Stochastic Discount Factor When the EIS Differs from One

Hansen et al. (2007) (henceforth HHLR) use the same data-generating process as Hansen et al. (2008), but use slightly different notation. HHLR set up the model with the following notation:

$$c_{t+1} - c_t = \mu_c + G'x_t + H'w_{t+1} \quad (\text{A1})$$

$$x_{t+1} = Ax_t + Bw_{t+1}. \quad (\text{A2})$$

From Section 5.1.3, Section 5.1.4, and Appendix A.1 of HHLR it follows that, omitting constants that do not affect our subsequent analysis, the first-order expansion of the logarithm of the stochastic discount factor for the Epstein–Zin utility setting is

$$s_{t+1} \simeq (1 - \gamma)\lambda(\beta)w_{t+1} + \left(\frac{1}{\sigma} - 1\right)\left(\frac{1}{2}w'_{t+1}\Theta_0w_{t+1} + w'_{t+1}\Theta_1x_t + \vartheta_1x_t + \vartheta_2w_{t+1}\right), \quad (\text{A3})$$

where

$$\begin{aligned}\lambda(\beta) &= H' + \beta G'(1 - \beta A)^{-1} B \\ \Theta_0 &= (\gamma - 1) B' \Omega B \\ \Theta_1 &= (\gamma - 1) B' \Omega A \\ \vartheta_1 &= -G' + (\gamma - 1)^2 [H' + \beta G'(I - \beta A)^{-1} B] B' \Omega A \\ \vartheta_2 &= -(1 - \gamma) \omega' B + U_v' B\end{aligned}$$

and

$$\begin{aligned}\mu_v &= \frac{\beta}{1 - \beta} \left[\mu_c + \frac{1 - \gamma}{2} |H' + \beta G'(I - \beta A)^{-1} B|^2 \right] \\ U_v &= \beta (I - \beta A')^{-1} G \\ \Omega &= \frac{1 - \beta}{\beta} U_v U_v' + \beta A' \Omega A \\ \omega &= (I - \beta A')^{-1} \left(\frac{1 - \beta}{\beta} \mu_v U_v + \beta (1 - \gamma) A' \Omega B (H + B' U_v) \right).\end{aligned}$$

The first term on the right-hand side of equation (A3), $(1 - \gamma)\lambda(\beta)w_{t+1}$, is the expression for the stochastic discount factor in the $EIS = 1$ case (aside from omitted constants and the $c_{t+1} - c_t$ term). This term equals $E_{t+1} \sum_{s=0}^{\infty} \beta^s (c_{t+1+s} - c_{t+s}) - E_t \sum_{s=0}^{\infty} \beta^s (c_{t+1+s} - c_{t+s})$. For consistency across all terms of the stochastic discount factor in the $EIS \neq 1$ case, we estimate this term as $(1 - \gamma)\lambda(\beta)w_{t+1}$ based on the VAR model described below. In order to calculate the additional term in the log stochastic discount factor in equation (A3), $(\frac{1}{\sigma} - 1)[\frac{1}{2}w_{t+1}'\Theta_0w_{t+1} + w_{t+1}'\Theta_1x_t + \vartheta_1x_t + \vartheta_2w_{t+1}]$, we need to estimate μ_c , G , H , A , B , and w_{t+1} in addition to making assumptions about x_t and the values of β and σ (the EIS). In order to estimate these elements, however, we need to estimate the dynamics of consumption growth through a VAR and cannot adopt the simple approaches (of using unconditional covariances or return innovations) employed for the case where the EIS equals one.

When choosing the elements of the state vector x_t in the VAR it is essential to capture the ability of each test asset return to predict future consumption growth. Otherwise, the estimated shocks w_{t+1} , and thus the stochastic discount factor, could be spuriously correlated with the test asset returns. However, given the relatively short time series of CEX consumption data and the large number of test assets (25 in the case of the Fama–French portfolios we study), it is not feasible to include all test asset returns and their lags in x_t . To overcome this problem we estimate a separate VAR for consumption dynamics for the calculation of $\text{cov}(s_{t+1}, r_{t+1}^i - r_{t+1}^f)$ for each test asset. This approach is motivated by different test assets having different ability to predict future consumption growth and by a desire to reduce the number of parameter estimates in the short CEX sample. The VAR ensures that the ability of a particular test asset to predict future consumption growth is preserved using the following

state vector:

$$x_t = \left[r_t^f, \log(P/D)_t^{SP500}, r_t^i - r_t^f, r_{t-1}^i - r_{t-1}^f, \dots, r_{t-16}^i - r_{t-16}^f \right]'$$

The system in equation (A1) and equation (A2) then contains four shocks. Denote these by $w_{t+1}^c, w_{t+1}^{r_f}, w_{t+1}^{\log P/D}, w_{t+1}^{r_i-r_f}$.¹⁶ Denote the error terms from equation (A1) by ε_{t+1}^c and the error terms from the first three rows of equation (A2) by $\varepsilon_{t+1}^{r_f}, \varepsilon_{t+1}^{\log P/D}$, and $\varepsilon_{t+1}^{r_i-r_f}$, and stack them to get

$$\begin{bmatrix} \varepsilon_{t+1}^c \\ \varepsilon_{t+1}^{r_f} \\ \varepsilon_{t+1}^{\log P/D} \\ \varepsilon_{t+1}^{r_i-r_f} \end{bmatrix} = \begin{bmatrix} H^c & H^{r_f} & H^{\log P/D} & H^{r_i-r_f} \\ B_{11} & B_{12} & B_{13} & B_{14} \\ B_{21} & B_{22} & B_{23} & B_{24} \\ B_{31} & B_{32} & B_{33} & B_{34} \end{bmatrix} \begin{bmatrix} w_{t+1}^c \\ w_{t+1}^{r_f} \\ w_{t+1}^{\log P/D} \\ w_{t+1}^{r_i-r_f} \end{bmatrix} \iff \varepsilon^* = M^* w^*,$$

where ε^*, M^* , and w^* are defined by the equation. Estimating equations (A1) and (A2) equation by equation (using OLS) results in estimates of μ_c, G, A , and the residuals in ε^* . To recover H, B , and $w_{t+1} = (M^*)^{-1} \varepsilon^*$ it is necessary to estimate the 16 elements of M^* . Assuming $V(w^*) = I$ (following HHL), the covariance matrix for ε^* is given by $\text{cov}(\varepsilon^*, \varepsilon^*) = M^* M^{*'}.$ Because $\text{cov}(\varepsilon^*, \varepsilon^*)$ has 10 (nonidentical) values there are 10 equations in 16 unknowns. We impose six restrictions on M^* , $H^{r_f} = B_{22} = B_{32} = 0$, and $H^{\log P/D} = B_{23} = B_{33} = 0$, and solve for the remaining 10 elements of M^* . These restrictions imply that shocks to r_f and $\log(P/D)$ have no immediate effect on the other variables, but are allowed to affect them with a lag through the VAR dynamics.

Compared to the $EIS = 1$ case, our approach for an $EIS \neq 1$ imposes additional structure. First, while no VAR was necessary in the $EIS = 1$ case we now need to estimate a VAR. For data reasons (the limited time-series dimension of the CEX data) the VAR differs across test assets to avoid spurious correlation between a given test asset and the stochastic discount factor. The drawback of this approach is that the conditional expectations for consumption growth are different depending on the excess return of interest so that we do not get a consistent model for consumption dynamics across the test assets. There is no obvious bias in this procedure, but the varying consumption dynamics across test assets are somewhat unappealing. To address the impact of estimating a separate VAR for each test asset, we repeat the approach above and estimate a VAR—a separate VAR for each test asset—for the $EIS = 1$ case and compare the results to those for the $EIS \neq 1$ case that did not rely on a VAR (specifically, the results for the $EIS = 1$ case that are based on using return innovations). As we show in Table V, the results are extremely similar under both approaches.

Second, for the $EIS \neq 1$ case we need to impose six restrictions on the estimated VAR in order to estimate the second term of the stochastic discount factor in equation (A3). For the $EIS = 1$ case, which relies only on the first

¹⁶ In equation (A2), the error terms in the equations for $r_t^i - r_t^f, \dots, r_{t-(16)}^i - r_{t-(16)}^f$ will be zero. See Hamilton (1994) p. 259 for a similar VAR setting.

term of equation (A3), a VAR with or without these restrictions is feasible. In order to determine if these additional restrictions are likely to materially affect the results, we compare the results from a VAR with and without the six restrictions for the $EIS = 1$ case. The results are not materially affected by these restrictions in the sense that the risk aversion estimates are very similar.

To gauge whether the long-run component of consumption growth is estimated in a reasonable manner under our VAR, we report the time-series standard deviation of expected consumption growth $G'x_t$ and the time-series standard deviation of the main element driving the stochastic discount factor when the $EIS = 1$, namely, the change in expectations about discounted consumption growth $(\sum_{s=0}^{\infty} \lambda_s \beta^s) w_{t+1}$ (which equals $(E_{t+1} - E_t) \sum_{s=0}^{\infty} \beta^s (c_{t+1+s} - c_{t+s})$). Because we run a separate VAR for each asset, we average the time-series volatilities across the 25 VARs corresponding to the 25 Fama–French portfolios used as test assets.

The standard deviation of the estimated long-run component of consumption, $G'x_t$, from equation (A1) of the VAR is 0.009 for stockholders, 0.021 for top stockholders, 0.005 for nonstockholders, and 0.003 for aggregate consumption. The volatility of the predictable component of consumption growth is thus twice as large for stockholders as for nonstockholders, and the predictable component of the consumption of top stockholders has about four times the volatility of that of nonstockholders. Likewise, the volatility of the change in expected discounted consumption growth is 0.034 for stockholders, 0.064 for the top stockholders, and only 0.020 and 0.007 for nonstockholder and aggregate consumption. The differences in the volatility of expected consumption growth and the change in expected discounted consumption growth are consistent with the decline in risk aversion estimates from nonstockholders to stockholders to top stockholders.

Appendix B: CEX Sample Choice and Stockholder Definitions

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 because these costs have substantial durable components. Attanasio and Weber (1995) use a similar definition of consumption.¹⁷

CEX Sample Choice

For each household we calculate quarterly consumption growth rates based on reported monthly consumption values. We drop household-quarters in

¹⁷ In leaving out durables, it is implicitly assumed that utility is separable in durables and nondurables/services. Results in the paper are generally similar when using total consumption, which adds 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.

which a household reports nonzero consumption for more than 3 or less than 3 months or where consumption is negative.

Extreme outliers are dropped because these may reflect reporting or coding errors. Specifically, we drop observations for which the consumption growth ratio (C_{t+1}^h/C_t^h) is less than 0.2 or above 5.0. In addition, nonurban 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 that 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 because 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 adjust for this change without substantial loss of data. See Battistin (2003) for details on the questions asked. Furthermore, we obtain small sample sizes in the last 3 months of the sample because all households used in those months must be in their last interview in order for financial information to be available. We therefore drop these 3 months of data.

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 because 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 that 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 those 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 nonstockholders for our simplest and baseline definition. The Euler equation involving consumption in periods t and $t + 1$ should hold for those that 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 periods 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.¹⁸

As discussed in the data section we also define a group consisting of the wealthiest third of stockholders. This is done based on the value of beginning-of-period t holdings of “stock, bonds, mutual funds and other such securities” calculated using the end-of-period value minus the change in holdings over the period. Two data issues arise. First, the CEX started using computers to conduct the interviews (rather than paper) beginning in April 2003. This change appears to imply that the dollar asset change question (the variable COMPSECX) is not asked if the respondent already replied having the same asset holdings as a year ago (in the data this shows up as “valid blank” responses for the variable COMPSECX). This problem is easily fixed by simply setting the asset change to zero in these cases. Second, for interviews conducted from October 1990 to September 1997, about 5% of households reported holdings of stocks, bonds, and mutual funds of \$1. We contacted the BLS to determine if this is a coding error, but they were not sure how to interpret the \$1 answers. Because all of the households reporting \$1 asset 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. However, because 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).

It is known from the SCF that many households hold stocks or bonds only in their pension plans. 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 in the CEX is smaller than in other sources. This fact may indicate that some CEX households with stockholdings in pension plans do not report these, leading them to be miscategorized as nonstockholders. We are unable to

¹⁸ Note that 1,834 households in our final sample of 76,568 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. In addition, 154 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 nonstockholders.

address this issue. However, our results may be conservative and would likely strengthen with a cleaner separation of stockholders and nonstockholders.

A complementary issue that can be addressed is that households holding bonds or bond funds exclusively will be misclassified as stockholders when households are categorized based on the definitions above. 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 contain the entire wealth decomposition of households (including direct and indirect holdings), we estimate a 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 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/Caucasian, 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, an indicator for positive dividend income, and a constant.¹⁹ Data are averaged across SCF imputations and SCF weights are employed in the probit model to ensure that the estimates are not unduly 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 that 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

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

and the pseudo- R^2 from the probit model 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 that

¹⁹ 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.

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. Nonstockholders are similarly defined as those responding negatively to the CEX question *and* having a predicted probability of owning stocks less than 0.50. These alternative definitions of stockholders and nonstockholders refine the CEX classification.

Appendix C: GMM Estimation

This appendix uses the GMM framework to 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 (11). The derivations follow in a similar manner when the Euler equation is written in the form of the reverse regression equation (12) with GMM moment conditions reorganized correspondingly.

A. The GMM Setup

The Euler equation for asset i is

$$\begin{aligned} E(r_{t+1}^i - r_{t+1}^f) + \frac{1}{2}V(r_{t+1}^i) - \frac{1}{2}V(r_{t+1}^f) \\ \simeq (\gamma - 1)\text{cov}(r_{t+1}^i - r_{t+1}^f, \Sigma_{s=0}^{\infty}\beta^s(c_{t+1+s} - c_{t+s})) \\ = (\gamma - 1)E((r_{t+1}^i - r_{t+1}^f)\varepsilon_{c,t+1}), \end{aligned} \quad (\text{C1})$$

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

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

$$\begin{aligned} \begin{bmatrix} 0_{N \times 1} \\ 0 \end{bmatrix} &= E \begin{bmatrix} r_{t+1} - r_{t+1}^f \mathbf{1}_N + \frac{1}{2}\sigma^2 - \frac{1}{2}\sigma_f^2 - (\gamma - 1)(r_{t+1} - r_{t+1}^f \mathbf{1}_N)\varepsilon_{c,t+1} \\ \varepsilon_{c,t+1} \end{bmatrix} \\ &= E \begin{bmatrix} g_{t+1}^r \\ g_{t+1}^m \end{bmatrix}, \end{aligned}$$

where N is the number of assets (25 in our setting using the 25 Fama–French test assets), $\mathbf{1}_N$ is a $N \times 1$ vector of ones, g_{t+1}^r and g_{t+1}^m are defined by the last equality, and

$$r_{t+1} = \begin{bmatrix} r_{t+1}^1 \\ \vdots \\ r_{t+1}^N \end{bmatrix}, \quad \sigma^2 = \begin{bmatrix} V(r_{t+1}^1) \\ \vdots \\ V(r_{t+1}^N) \end{bmatrix}, \quad \sigma_f^2 = \begin{bmatrix} V(r_{t+1}^f) \\ \vdots \\ V(r_{t+1}^f) \end{bmatrix}.$$

For simplicity we assume σ^2 and σ_f^2 are known.

In order to use all available data, we use a different sample length T for the two parts of $E(g_{t+1}^r)$. We use a sample of length T_1 for estimating mean excess returns (adjusted for the variance terms) $E(r_{t+1} - r_{t+1}^f 1_N + \frac{1}{2}\sigma^2 - \frac{1}{2}\sigma_f^2)$ and a subsample of length T_2 for estimating covariances $E((r_{t+1} - r_{t+1}^f 1_N)\varepsilon_{c,t+1})$. We also use the subsample of length T_2 for estimating $E(g_{t+1}^m)$. When estimating equation (C1) we include an intercept term α , identical for all stocks, in the first set of moment conditions 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). We also estimate the equation omitting the intercept term.

The GMM objective function is

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

with

$$g = \begin{bmatrix} g^r \\ g^m \end{bmatrix} = \begin{bmatrix} \frac{1}{T_1} \sum_{t=0}^{T_1-1} \left(r_{t+1} - r_{t+1}^f 1_N + \frac{1}{2}\sigma^2 - \frac{1}{2}\sigma_f^2 \right) - \alpha 1_N - (\gamma - 1) \frac{1}{T_2} \sum_{t=0}^{T_2-1} (r_{t+1} - r_{t+1}^f 1_N) \varepsilon_{c,t+1} \\ \frac{1}{T_2} \sum_{t=0}^{T_2-1} \varepsilon_{c,t+1} \end{bmatrix}.$$

The estimation picks three parameters to fit 26 moments as best possible. We use a pre-specified weighting matrix

$$W = \begin{bmatrix} I_{25} & 0 \\ 0 & h \end{bmatrix}$$

rather than efficient GMM to give each of the 25 portfolios equal importance in the estimation. Following Parker and Julliard (2005) we set h to a sufficiently large number that the estimate of $\mu_{\sum_{s=0}^{\infty} \beta^s(c_{t+1+s} - c_{t+s})}$ approximately equals $\frac{1}{T_2} \sum_{t=0}^{T_2-1} \sum_{s=0}^{\infty} \beta^s(c_{t+1+s} - c_{t+s})$ and that further increases in h have only minimal effects on the estimates of γ and α .

B. 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

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

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

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

G is estimated by

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

The matrix Ω is 26×26 and given by:

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

We can rewrite Ω^{mm} as follows

$$\begin{aligned} \Omega_{1 \times 1}^{mm} &= E \left(T_1 \frac{1}{T_2} \sum_{t=0}^{T_2-1} \varepsilon_{c,t+1} \frac{1}{T_2} \sum_{t=0}^{T_2-1} \varepsilon'_{c,t+1} \right) \\ &= \frac{T_1}{T_2} E \left(\frac{1}{T_2} \sum_{t=0}^{T_2-1} \varepsilon_{c,t+1}^2 + \sum_{l=1}^L \frac{1}{T_2} \sum_{t=l}^{T_2-1} [\varepsilon_{c,t+1} \varepsilon_{c,t+1-l} + \varepsilon_{c,t+1-l} \varepsilon_{c,t+1}] \right), \end{aligned}$$

where L is the highest order of autocorrelation induced by the use of overlapping consumption growth data ($L = 47$ when using 16-quarter discounted consumption growth rates, available at the monthly frequency).

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

$$\begin{aligned} \Omega_{25 \times 1}^{rm} &= E \left(T_1 \left(\frac{1}{T_1} \sum_{t=0}^{T_1-1} z_{t+1} - (\gamma - 1) \frac{1}{T_2} \sum_{t=0}^{T_2-1} w_{t+1} \right) \times \frac{1}{T_2} \sum_{t=0}^{T_2-1} \varepsilon_{c,t+1} \right) \\ &= E \left(\frac{1}{T_2} \sum_{t=0}^{T_2-1} z_{t+1} \times \sum_{t=0}^{T_2-1} \varepsilon_{c,t+1} \right) \\ &\quad - (\gamma - 1) \frac{T_1}{T_2} E \left(\left(\frac{1}{T_2} \sum_{t=0}^{T_2-1} w_{t+1} \right) \times \sum_{t=0}^{T_2-1} \varepsilon_{c,t+1} \right) \\ &= E \left(\frac{1}{T_2} \sum_{t=0}^{T_2-1} z_{t+1} \varepsilon_{c,t+1} + \sum_{l=1}^L \frac{1}{T_2} \sum_{t=l}^{T_2-1} [z_{t+1} \varepsilon_{c,t+1-l} + z_{t+1-l} \varepsilon_{c,t+1}] \right) \\ &\quad - (\gamma - 1) \frac{T_1}{T_2} E \left(\frac{1}{T_2} \sum_{t=0}^{T_2-1} w_{t+1} \varepsilon_{c,t+1} + \sum_{l=1}^L \frac{1}{T_2} \sum_{t=l}^{T_2-1} \right. \\ &\quad \left. \times [w_{t+1} \varepsilon_{c,t+1-l} + w_{t+1-l} \varepsilon_{c,t+1}] \right), \end{aligned}$$

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

Finally,

$$\begin{aligned}
 \Omega_{25 \times 25}^{rr} &= E \left(T_1 \left(\frac{1}{T_1} \sum_{t=0}^{T_1-1} z_{t+1} - (\gamma - 1) \frac{1}{T_2} \sum_{t=0}^{T_2-1} w_{t+1} \right) \right. \\
 &\quad \times \left. \left(\frac{1}{T_1} \sum_{t=0}^{T_1-1} z_{t+1} - (\gamma - 1) \frac{1}{T_2} \sum_{t=0}^{T_2-1} w_{t+1} \right)' \right) \\
 &= T_1 E \left(\frac{1}{T_1} \sum_{t=0}^{T_1-1} [z_{t+1} - \mu_z] \right) \left(\frac{1}{T_1} \sum_{t=0}^{T_1-1} [z_{t+1} - \mu_z]' \right) - T_1 \mu_z \mu_z' \\
 &\quad + T_1 (\gamma - 1)^2 E \left(\left(\frac{1}{T_2} \sum_{t=0}^{T_2-1} [w_{t+1} - \mu_w] \right) \times \left(\frac{1}{T_2} \sum_{t=0}^{T_2-1} [w_{t+1} - \mu_w]' \right) \right) \\
 &\quad - T_1 (\gamma - 1)^2 \mu_w \mu_w' - (\gamma - 1) T_1 E \left(\frac{1}{T_1} \sum_{t=0}^{T_1-1} [z_{t+1} - \mu_z] \right) \\
 &\quad \times \left(\frac{1}{T_2} \sum_{t=0}^{T_2-1} [w_{t+1} - \mu_w]' \right) + (\gamma - 1) T_1 \mu_z \mu_w' \\
 &\quad - (\gamma - 1) T_1 E \left(\frac{1}{T_2} \sum_{t=0}^{T_2-1} [w_{t+1} - \mu_w] \right) \times \left(\frac{1}{T_1} \sum_{t=0}^{T_1-1} [z_{t+1} - \mu_z]' \right) \\
 &\quad + (\gamma - 1) T_1 \mu_w \mu_z' \\
 &= E \left(\frac{1}{T_1} \sum_{t=0}^{T_1-1} [z_{t+1} - \mu_z] [z_{t+1} - \mu_z]' \right) - T_1 \mu_z \mu_z' \\
 &\quad + \frac{T_1}{T_2} (\gamma - 1)^2 E \left(\begin{aligned} &\frac{1}{T_2} \sum_{t=0}^{T_2-1} [w_{t+1} - \mu_w] [w_{t+1} - \mu_w]' \\ &+ \sum_{l=1}^L \frac{1}{T_2} \sum_{t=l}^{T_2-1} ([w_{t+1} - \mu_w] [w_{t+1-l} - \mu_w]' + [w_{t+1-l} - \mu_w] [w_{t+1} - \mu_w]') \end{aligned} \right) \\
 &\quad - T_1 (\gamma - 1)^2 \mu_w \mu_w' \\
 &\quad - (\gamma - 1) E \left(\begin{aligned} &\frac{1}{T_2} \sum_{t=0}^{T_2-1} [z_{t+1} - \mu_z] [w_{t+1} - \mu_w]' \\ &+ \sum_{l=1}^L \frac{1}{T_2} \sum_{t=l}^{T_2-1} ([z_{t+1} - \mu_z] [w_{t+1-l} - \mu_w]' + [z_{t+1-l} - \mu_z] [w_{t+1} - \mu_w]') \end{aligned} \right) \\
 &\quad + (\gamma - 1) T_1 \mu_z \mu_w' \\
 &\quad - (\gamma - 1) E \left(\begin{aligned} &\frac{1}{T_2} \sum_{t=0}^{T_2-1} [w_{t+1} - \mu_w] [z_{t+1} - \mu_z]' \\ &+ \sum_{l=1}^L \frac{1}{T_2} \sum_{t=l}^{T_2-1} ([w_{t+1} - \mu_w] [z_{t+1-l} - \mu_z]' + [w_{t+1-l} - \mu_w] [z_{t+1} - \mu_z]') \end{aligned} \right) \\
 &\quad + (\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} . To estimate Ω , remove the E s and use the estimated values of $\varepsilon_{c,t+1}$, α , γ , μ_z , and μ_w .²⁰

²⁰ When invertibility problems arise we employ Newey and West (1987) weightings up to lag $L = 47$ to ensure invertibility. In cases where invertibility is not a problem we confirm that the Newey and West (1987) weights deliver similar standard errors.

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