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Habit persistence: Explaining cross sectional variation in returns and time-varying expected returns

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Habit persistence: Explaining cross sectional variation in returns and time-varying expected returns*

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Abstract

This paper finds empirical support for the habit persistence model of Campbell and Cochrane (1999) along both cross sectional and time-series dimensions of the US stock market. GMM estimations show that the model is able to explain a substantial part of the cross sectional variation of returns on the 25 Fama and French value and size portfolios over the period 1932-2003, although it has difficulties in fully explaining the value premium, and some of the implied risk free rates are strongly negative. In addition, the model accounts for time-varying expected returns on stocks. Forecasting regressions show that the estimated surplus consumption ratio has strong forecasting power for future real stock returns and holds additional explanatory power relative to traditional financial forecasting variables such as the dividend yield. We also document that the Campbell-Cochrane model is particularly successful up to 1991. Including data from the 1990s reduces somewhat the fit of the model.

Keywords: Campbell-Cochrane model, 25 Fama-French portfolios, GMM, return predictability by surplus-consumption ratio.
JEL codes: C32, G12.

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1 Introduction

By adding the surplus consumption ratio to the standard C-CAPM with power utility, Campbell and Cochrane (1999) show by calibration that their habit formation model accounts for a number of stylized facts on the US stock market, including time-varying expected returns on stocks. The model implies that the price of risk is time-varying and counter-cyclical: when consumption is well above the habit in cyclical upswings, the price of risk is low leading to low expected returns and high asset prices. By contrast, when consumption is close to habit, the price of risk is high leading to high expected returns and low asset prices. This is in accordance with a growing body of literature documenting time-varying expected stock returns. Financial variables such as the dividend yield, the term premium on bonds, and the relative interest rate have been documented as forecasters of stock returns, cf. Campbell and Shiller (1988), Fama and French (1989), Campbell (1991), and Hodrick (1992). Fama and French (1989) link the financial forecasting variables to the business cycle and suggest that investors require a higher expected return at a business cycle trough than they do at a business cycle peak. As an extension to these financial forecasting variables, Lettau and Ludvigson (2001a) introduce the consumption wealth ratio: a macroeconomic variable that forecasts stock returns. Similarly, the surplus consumption ratio in the Campbell-Cochrane model is also a macroeconomic variable, which provides a direct linkage between the business cycle and expected stock returns.

Relatively few studies have estimated the Campbell-Cochrane model. Campbell and Cochrane (1999) themselves do not estimate the model econometrically, but instead they calibrate the model. Tallarini and Zhang (2005) estimate the model on US data and find evidence of counter-cyclical behavior of expected returns, but they statistically reject the model. Fillat and Garduno (2005) find insignificant values of the utility curvature parameter, and they also statistically reject the model. Furthermore the model does not perform well in explaining a cross section of US asset returns. Garcia et al. (2005) do not statistically reject the model on US data, but their iterated GMM approach does not lead to convergence with positive values of the utility curvature parameter. Engsted et al. (2006) test the model outside the US and find that the model does not perform better than the simple power utility model in explaining Danish asset returns.[1] This paper tests the Campbell-Cochrane model on US data over the period 1932-2003 using the GMM estimation procedure outlined in Engsted et al. (2006). The model is estimated in a cross sectional setting using the Fama and French 25 value and size portfolios, which has not been tried yet, cf. Cochrane (2006). The cross sectional estimation documents that the model i) explains a substantial part of the variation in the 25 Fama and French portfolios, but to some extant does not fully account for the value premium; ii) generates a high and time-varying relative risk aversion; iii) has the ability to produce a constant risk free rate at a reasonable low level, but the risk free rate is highly sensitive to the parameter estimates.

[1] Hyde and Sherif (2005) and Hyde et al. (2005) also test the model outside the US and find supporting evidence of the model for the UK by adopting the calibrated values from the US.

In contrast to the previous studies, this paper also formally tests the forecasting power of the surplus consumption ratio. Forecasting regressions show that the surplus consumption ratio tracks a substantial amount of the variation in future real stock returns. Consistent with the Campbell-Cochrane model, the surplus consumption ratio is significantly negatively related to future real stock returns, so that low surplus consumption ratios in recession times predict high real stock returns. Furthermore, the surplus consumption ratio has additional explanatory power for future real stock returns relative to traditional financial forecasting variables such as the dividend yield.

Li (2001, 2005) also tests the forecasting power of the surplus consumption ratio and finds that the surplus consumption ratio provides some information about future stock returns. However, Li (2001, 2005) does not estimate the parameters of the Campbell-Cochrane model, but takes the parameters as given. In contrast to Li (2001, 2005), this paper estimates the parameters of the Campbell-Cochrane model. Furthermore, small sample bias is taken into account by using Lewellen's (2004) correction method and bootstrap.

Subsample analyses reveal that the Campbell-Cochrane model performs particularly well up to the beginning of the 1990s. However, including the 1990s stock market boom reduces somewhat the performance of the model.

The paper is organized as follows. Section 2 introduces the Campbell-Cochrane model, section 3 describes the empirical methodology, section 4 reports the cross sectional evidence, section 5 reports the evidence of time-varying expected returns, section 6 reports robustness checks for statistical pitfalls in forecasting regressions, and section 7 concludes.

2 The Campbell-Cochrane model

The utility function of the representative investor is:

$$E_t \sum_{j=0}^{\infty} \delta^j \frac{(C_{t+j} - X_{t+j})^{1-\gamma} - 1}{1-\gamma}. \tag{1}$$

C_t is consumption, X_t is the external habit level, δ is the impatience parameter, and γ is the utility curvature parameter. Campbell and Cochrane capture the relation between consumption and habit through the surplus consumption ratio:

$$S_t \equiv \frac{C_t - X_t}{C_t}, \tag{2}$$

and specify the logarithm of the surplus consumption ratio $s_t = \log(S_t)$ as a stationary first-order autoregressive process:

$$s_{t+1} = (1-\phi)\,\bar{s} + \phi s_t + \lambda(s_t)\,v_{t+1}, \tag{3}$$

where $0 < \phi < 1$ is the habit persistence parameter, $\bar{s}$ is the steady state level of s_t, and $\lambda(s_t)$ is the sensitivity function that determines how innovations in consumption growth v_{t+1} influence s_{t+1}. The consumption growth process is given by:

$$\triangle c_{t+1} = g + v_{t+1}, \qquad v_{t+1} \sim niid\left(0, \sigma_c^2\right), \tag{4}$$

where $c_t = \log(C_t)$, and g is the mean consumption growth. The sensitivity function $\lambda(s_t)$ is specified as follows:

$$\lambda(s_t) = \left\{ \begin{array}{cc} \frac{1}{\overline{S}}\sqrt{1 - 2(s_t - \bar{s})} & \text{if } s_t \le s_{\max} \\ 0 & \text{else} \end{array} \right\}, \tag{5}$$

where

$$\overline{S} = \sigma_c \sqrt{\frac{\gamma}{1-\phi}}, \qquad s_{\max} \equiv \bar{s} + \frac{1}{2}(1 - \overline{S}^2), \qquad \bar{s} = \log(\overline{S})$$

Specifying $\lambda(s_t)$ in this way implies a constant risk free rate over time. From the Euler equation,

$$E_t\left[M_{t+1} R_{i,t+1}\right] = 1, \tag{6}$$

where $M_{t+1} = \delta\left(\frac{S_{t+1}}{S_t}\frac{C_{t+1}}{C_t}\right)^{-\gamma}$ is the stochastic discount factor, and $R_{i,t+1}$ is the gross return on any asset i, the constant log real risk free rate is:

$$\begin{aligned} r_{f,t+1} &= -\log\left(E_t\left[M_{t+1}\right]\right) = -E_t\left[m_{t+1}\right] - \frac{1}{2}Var\left[m_{t+1}\right] \\ &= -\log(\delta) + \gamma g - \frac{\gamma}{2}(1-\phi), \end{aligned} \tag{7}$$

where $m_{t+1} = \log(M_{t+1})$. From the Euler equation (6), the expected real stock return can be stated as:

$$E_t\left(r_{i,t+1}\right) + \frac{1}{2}\sigma_{i,t}^2 = r_f + \gamma\left[1 + \lambda(s_t)\right]\sigma_{ic,t}, \tag{8}$$

where $\frac{1}{2}\sigma_{i,t}^2$ is a Jensen's inequality term. (8) shows that the expected real stock return is given by the constant risk free rate, r_f, plus the state-dependent price of risk, $\gamma\left[1 + \lambda(s_t)\right]$, times the amount of risk, $\sigma_{ic,t}$ (the conditional covariance between the return on asset i and the consumption growth). Li (2001) finds that $\sigma_{ic,t}$ is close to being constant through time. This lack of time-variation in the amount of risk suggests that time-varying expected returns are generated entirely by time-variation in the price of risk, which is a non-linear function of the surplus consumption ratio. Following Li (2005), I examine the linear relationship between the surplus consumption ratio and future stock market returns. This allows direct comparison with other forecasting variables such as the dividend yield.

3 Empirical methodology

The Euler equation (6) implies the following unconditional moment conditions testable by Hansen's (1982) GMM:[2]

$$E\left[\delta\left(\frac{S_{t+1}}{S_t}\frac{C_{t+1}}{C_t}\right)^{-\gamma}R_{t+1}-1\right]=0, \tag{9}$$

where R_{t+1} contains real gross returns on the 25 Fama and French portfolios. To estimate the model parameters, the sample counterpart of (9) is taken:

$$g_T\left(\theta\right)=\frac{1}{T}\sum_{t=1}^{T}\left[\delta\left(\frac{S_{t+1}}{S_t}\frac{C_{t+1}}{C_t}\right)^{-\gamma}R_{t+1}-1\right]=0, \tag{10}$$

where $\theta=(\delta\quad\gamma)'$, and T is the sample size. Then, GMM estimates θ by minimizing the quadratic form:

$$g_T\left(\theta\right)'Wg_T\left(\theta\right), \tag{11}$$

where W is a positive definite weighting matrix. I use the identity matrix to give equal weight to the 25 Fama and French portfolios.

The estimation of the Campbell-Cochrane model is complicated by the fact that the surplus consumption ratio, S_t, is not observable in the same way as returns, R_t, and consumption, C_t, are directly observable. To overcome this obstacle, I follow the suggested estimation procedure outlined in Engsted et al. (2006):

Step 1: Following Campbell and Cochrane (1999) and Garcia et al. (2005), the persistence parameter, ϕ, is estimated as the first-order autocorrelation parameter for the log dividend yield:

$$d_t-p_t=\alpha+\phi(d_{t-1}-p_{t-1})+\varepsilon_t. \tag{12}$$

This is feasible since in the Campbell-Cochrane model the surplus consumption ratio is the only state variable, whereby the log dividend yield, d_t-p_t, will inherit its dynamic properties from the log surplus consumption ratio, s_t. Recent work by Boudoukh et al. (2007), however, provides evidence of a structural break in the dividend yield due to the rise of share repurchases in the 1980s and 1990s. Consequently, I also estimate ϕ using the log net payout yield of Boudoukh et al. (2007), which I denote npo_t-p_t.

Step 2: g and σ_c are estimated from (4), and the implied process for v_t is obtained.

[2] Due to a small sample of annual observations and a large cross section of portfolios, the use of instrument variables would result in an unmanageable large number of moment conditions relatively to the number of sample observations.

Step 3: An initial value of γ is chosen to obtain $\overline{S}$ and $\overline{s}$. Then by setting $s_t = \overline{s}$ at $t = 0$, the s_t process is obtained from (3).

Step 4: Given the observed time-series for the returns on the 25 Fama and French portfolios and the consumption growth, and given the time-series for the surplus consumption ratio generated in step 3, the model is estimated by minimizing (11). This gives GMM estimates of δ and γ, and step 3 and 4 are repeated until convergence of δ and γ.

Since the chosen weighting matrix is not the efficient Hansen (1982) matrix but the identity matrix I, the formula for the covariance matrix of the parameter vector is (c.f. Cochrane (2005), chpt. 11):

$$Var(\widehat{\theta}) = \frac{1}{T}(d'Id)^{-1}d'ISId(d'Id)^{-1}, \tag{13}$$

where $d' = \partial g_T(\theta)/\partial\theta$, and the spectral density matrix $S = \sum_{j=-\infty}^{\infty} E[\ g_T(\theta)g_{T-j}(\theta)']$ is computed with the usual Newey and West (1987) estimator with a lag truncation. Similarly, the J-test of overidentifying restrictions is computed based on the general formula (c.f. Cochrane (2005) chpt. 11):

$$J_T = Tg_T(\widehat{\theta})' \left[(I - d(d'Id)^{-1}d'I)S(I - Id(d'Id)^{-1}d')\right]^{-1} g_T(\widehat{\theta}). \tag{14}$$

J_T has an asymptotic χ^2 distribution with degrees of freedom equal to the number of overidentifying restrictions. (14) involves the covariance matrix $Var(g_T(\widehat{\theta})) = \frac{1}{T}(I - d(d'Id)^{-1}d'I)S(I - Id(d'Id)^{-1}d')$, which is singular, so it is inverted using the Moore-Penrose pseudo-inversion.

I also calculate the cross sectional R^2, which is a goodness of fit measure used in the literature:[3]

$$R^2 = \frac{Var\left(\bar{R}_i\right) - Var\left(\bar{R}_i - E\left[R_{i,t+1}\right]\right)}{Var\left(\bar{R}_i\right)}, \tag{15}$$

where $\bar{R}_i$ is the average time-series return on portfolio i, and $E\left[R_{i,t+1}\right]$ is the model expected return on portfolio i:

$$E\left[R_{i,t+1}\right] = \frac{1 - Cov\left[R_{i,t+1}, \delta\left(\frac{S_{t+1}}{S_t}\frac{C_{t+1}}{C_t}\right)^{-\gamma}\right]}{E\left[\delta\left(\frac{S_{t+1}}{S_t}\frac{C_{t+1}}{C_t}\right)^{-\gamma}\right]}.$$

[3] As Cochrane (2006) points out the cross sectional R^2 depends on the estimation teqnique, which means that the R^2 in (15) is not directly comparable to applications of linear factor models, estimating the R^2 from the cross section of average returns on a constant and betas.

Hence, $Var\left(\bar{R}_i\right)$ is the cross sectional variance of average returns on the 25 Fama and French portfolios, and $Var\left(\bar{R}_i - E\left[R_{i,t+1}\right]\right)$ is the cross sectional variance of the model residuals.

After estimating the parameters of the Campbell-Cochrane model, I test the forecasting power of the estimated surplus consumption ratio, the dividend yield, and the net payout yield. This is done by forecasting regressions:

$$r_{t+1\to k} = \sum_{i=1}^{k} r_{t+i} = c + \beta x_t + e_{t+1\to k}, \tag{16}$$

where $r_{t+1\to k}$ is the k-period log real stock return, and x_t is the forecasting variable.[4]

4 Habit persistence and the cross section of stock returns

The Campbell-Cochrane model is estimated on annual data using two sample periods: 1932-1991 and 1932-2003. The 1932-1991 sample excludes the 10 year long business cycle expansion from November 1991 to November 2001 as measured by NBER. Campbell and Cochrane (1999) find that during the first part of the 1990s - the end of Campbell and Cochrane's sample - the model has its worst performance. Hence, it is interesting to compare the model performance with and without the cyclical upswing and stock market boom of the 1990s.

Consumption is measured as expenditures on non-durables and services obtained from the National Income and Product Accounts (NIPA) table 2.3.5.[5] I use the Campbell (2003) beginning of period timing assumption that consumption during year t takes place at the beginning of year t. Nominal consumption is converted to real units using the consumption deflator from NIPA table 2.3.4. Real per capita consumption is obtained by using the population numbers in NIPA table 2.1.

The model is estimated on the 25 Fama and French (1993) portfolios sorted on book-to-market and size. To check the robustness of the estimation, the model is estimated on other portfolios as well. Due to potential small sample problems, the model is estimated in a small cross section of the 6 Fama and French portfolios underlying the Fama and French (1993) factors SMB and HML. In addition, the model is estimated on 10 decile portfolios sorted on size, 10 decile portfolios sorted on book-to-market, and 10 portfolios sorted on industry. As a further robustness check following Lewellen et. al (2006), the model is estimated on the 6 Fama and French portfolios together with the 10 industry

[4] I use real stock returns and not excess returns in (16), since the risk free rate is a constant, cf. equations (7) and (8).

[5] The NIPA consumption data starts in 1929, but I start the sample in 1932 in order to exclude the Great Depression.

portfolios, since the model should be able to price these portfolio sets at the same time in a joint cross section. Moreover, Cochrane (2006) points out that the 3 Fama and French factors - the excess return on the market, the return on the SMB portfolio, and the return on the HML portfolio - basically contain all the information in the 25 Fama and French portfolios, so one can use the 3 instead of the 25. With only 3 test portfolios, I use instrument variables without having too many moment conditions relative to the number of sample observations. I test the model conditionally using either the dividend yield or the net payout yield as instrument variable (together with a constant), which involves 6 moment conditions. All portfolio sets are taken from Kenneth French's web site, where details on the construction of the portfolios also are available. Real returns are obtained using the consumption deflator.

Table 1 shows average real gross returns on the 25 Fama and French portfolios with standard deviations in parentheses. Value firms with high book-to-market ratios have higher returns than growth firms with low book-to-market ratios, and small firms have higher average returns than large firms. Can the Campbell-Cochrane model explain these value and size premiums? This is now examined using the above described estimation procedure. First, the annual consumption growth rate, g, the innovations variance, σ_c^2, and the persistence parameter, ϕ, are estimated from equations (4) and (12). Table 2 shows the parameter estimates of ϕ, g, and σ_c.[6] The estimates of ϕ are much lower using the net payout yield instead of the dividend yield. The first-order autocorrelation parameter of the dividend yield is 0.94 in the 1932-2003 sample and 0.80 in the 1932-1991 sample compared to 0.73 and 0.66, respectively, for the net payout yield.[7] Next, based on the estimates in table 2, the Campbell-Cochrane model is estimated using GMM by repeating steps 3 and 4 in section 3 until convergence.

Table 3 reports results of cross sectional GMM estimations of the Campbell-Cochrane model on different portfolio sets for the 1932-2003 sample. The table shows that the estimated value of the persistence parameter, ϕ, has strong impact on the estimated value of the impatience parameter, δ. If $\phi = 0.94$, then δ is estimated precisely around 0.96 to 0.97, but if $\phi = 0.73$, the estimate of δ ranges from 0.66 to 0.89, implying a rather low time preference rate. Similarly, the degree of habit persistence also influences the estimate of the utility curvature parameter, γ. With a high value of $\phi = 0.94$ and a smooth surplus consumption ratio, it generally takes a high value of γ for the stochastic discount factor to match the volatility of the portfolio returns. Overall, the estimate of γ ranges from 0.52 to 17.38 and is significantly positive in almost all cases. The model does not resolve the equity premium puzzle, since the average degree of risk aversion ranges from 21 to 84. Moreover, in recession times with low surplus consumption ratios, the relative risk aversion is even higher.

Table 3 reports mixed evidence about the ability of the Campbell-Cochrane model to resolve the risk free rate puzzle. By inserting the estimated values of δ, γ, g, and ϕ into (7), the model typically implies counterfactually negative risk free rates, but in some

[6] Campbell and Cochrane's parameter choices of ϕ, g, and σ_c are 0.87, 0.019, and 0.015.

[7] The dividend yield is derived from CRSP value weighted returns with and without dividend capitalization, and the net payout yield is taken from the homepage of Michael Roberts.

cases the risk free rate is positive at a reasonable low level.[8] As an illustration of the latter, I look at the 6 Fama and French portfolios estimated jointly with the 10 industry portfolios. For this joint cross section, the estimated values of δ, γ, g, and ϕ of 0.96, 5.41, 0.03, and 0.94 imply a constant risk free rate of 0.41%, even though the average relative risk aversion is 45. Correspondingly, the Campbell-Cochrane model has the important ability to generate a high time-varying relative risk aversion and at the same time keep a low risk free rate.

Now I move on to the pricing errors of the model as measured by the J-test of overidentifying restrictions and the cross sectional R^2. As seen in table 3, the R^2 of the 25 Fama and French portfolios is around 50%. To illustrate the cross sectional fit further, figure 1 plots the realized average real gross returns against the model expected returns (for the case with $\phi = 0.94$). A perfect fit requires that all portfolios lie along the 45 degree line. The figure illustrates that the model has difficulty in pricing the S1BM1 portfolio correctly, but otherwise the pricing errors seem to be fairly low.[9] However, the J-test implies that the pricing errors are significantly different from zero. Possible explanations are the mispricing of the S1BM1 portfolio, or small sample bias, since the number of cross sectional observations of 25 might be too high relative to the number of sample observations of 71. Looking at the results for the small cross section of the 6 Fama and French portfolios, the cross sectional R^2 increases to around 63% and the J-test does not reject at the 1% significance level, but it still rejects at the 5% significance level.

To examine whether the mispricing is along the size dimension or the value dimension, the model is estimated separately on 10 size portfolios and 10 book-to-market portfolios. The results for the 10 size portfolios are convincing: the R^2 is close to 100% and the J-test does not reject the model. For the 10 book-to-market portfolios, the R^2 is around 72%, but J-test rejects the model at the 5% significance level due to high pricing errors on the low book-to-market portfolios. Consequently, it seems that the Campbell-Cochrane model to some extant misprices the low book-to-market portfolios and hence does not fully account for the value premium.

As a further robustness check, the model is estimated on the 6 Fama and French portfolios jointly with the 10 industry portfolios. The 10 industry portfolios have a return structure very different from the size and value portfolios, so it is interesting to examine whether the model can explain the variation in both cross sections at the same time. As it was the case for the 6 Fama and French portfolios estimated separately, the J-test rejects the model when the 10 industry portfolios are included in the joint cross section. On the other hand, the R^2 stays around the same level as the R^2 for the 10 industry portfolios estimated separately. Hence, the Campbell-Cochrane model does not fail the robustness check of expanding the portfolio set beyond the size and value

[8] I have tried adding a short T-bill rate to the moment conditions, but it does not change the value of the model implied risk free rate much.

[9] The mispricing of the S1BM1 portfolio is a common problem of asset pricing models. For instance, the long run C-CAPM of Parker and Julliard (2005), and the conditional C-CAPM with the consumption-wealth ratio as a conditioning variable of Lettau and Ludvigson (2001b) also misprice the S1BM1 portfolio.

portfolios.

How does the model perform in the 1932-1991 subsample, excluding the cyclical upswing and stock market boom of 1990s? Table 4 reports the results. As in the full sample, the estimated value of γ is generally significantly positive, but the implied risk aversion is too high to resolve the equity premium puzzle. The implied risk free rate ranges from -30.35% to 7.84%, giving mixed evidence about the risk free rate puzzle. Comparing the cross sectional fit, the R^2 is generally higher in the 1932-1991 sample than in the 1932-2003 sample. For instance, the R^2 for the 25 Fama and French portfolios increases to around 60%. It seems that the model performs better in the 1932-1991 sample, excluding the cyclical upswing and stock market boom of 1990s.

Finally, the model is estimated conditionally using the 3 Fama and French factors. Table 5 shows that the conditional estimation of the model on the 3 Fama and French factors provides similar results as the unconditional estimation of the model on the 25 Fama and French portfolios in the sense that the estimate of γ is significant and of the same magnitude.[10] As an important difference, the J-test rejects the unconditional model on 25 Fama and French portfolios, but not the conditional model on the 3 Fama and French factors.

5 Habit persistence and time-varying expected returns

Now I turn to formally testing the forecasting power of the surplus consumption ratio based on the estimated model parameters. The forecasting power of the surplus consumption ratio is not sensitive to the estimated value of γ for γ ranging in the intervals reported in tables 3 and 4. The results presented in the following are obtained by using the γ estimates for the Fama and French 25 portfolios.

The log surplus consumption ratio is used to forecast the log real stock return on the value weighted CRSP index including NYSE, AMEX and NASDAQ firms. Real stock returns are obtained using the consumption deflator. I compare the forecasting power of the log surplus consumption ratio to the forecasting power of the log dividend yield and the log net payout yield. Table 6 presents summary statistics of the log real stock return, r_t, the log dividend yield, $d_t - p_t$, the log net payout yield, $npo_t - p_t$, and the log surplus consumption ratio, s_t. The persistence parameter, ϕ, takes on the values given in table 2: $\phi = 0.66, 0.73, 0.80, 0.94$. The Campbell-Cochrane model predicts a perfect relationship between s_t and its proxy variable, but unfortunately s_t does not have much correlation with neither $d_t - p_t$ nor $npo_t - p_t$, which suggests that s_t represents independent information.

[10] The impatience parameter is not estimated due to the use of excess returns: the market return in excess of the risk free rate, the return on small portfolios in excess of the return on big portfolios (SMB), and the return on value portfolios in excess of the return on growth portfolios (HML).

Figure 2 illustrates the surplus consumption ratio, S_t, and the corresponding risk aversion, γ/S_t. The surplus consumption captures the general business cycles trends: increases in high growth periods such as the 1960s and the 1980s, and falls in slow growth periods such as the 1970s. With $\phi = 0.94$, however, the surplus consumption ratio stays below its steady state value during the 10 year long business cycle expansion from November 1991 to November 2001 as measured by NBER. Campbell and Cochrane (1999) themselves also find that during the first part of the 1990s - the end of Campbell and Cochrane's sample - the model has its worst performance, since the calibrated price-dividend ratio moves in the complete opposite direction of the actual price-dividend ratio. This model weakness could reduce the forecasting power of the surplus consumption ratio.

Table 7 reports slope estimates, Newey and West (1987) corrected t-statistics, and adjusted R^2-statistics for 1-5 year ahead forecasting regressions of real stock returns. The forecasting variable is either the dividend yield, the net payout yield, or the surplus consumption ratio. Since the surplus consumption ratio is estimated using the beginning of period consumption timing convention, it is lagged twice in the forecasting regressions.

With $\phi = 0.73$, the surplus consumption ratio tracks a substantial amount of the variation in future real stock returns in the full 1932-2003 sample (row 4 of table 7). The adjusted R^2-statistic is 9.91% at the 1 year horizon and increases to 29.91% at the 5 year horizon. The slope estimates are significantly negative at all horizons. Just like the Campbell-Cochrane model implies: low surplus consumption ratios predict high real stock returns.

If the persistence parameter is increased to $\phi = 0.94$, the surplus consumption ratio still has a significantly negative relationship with future real stock returns, but the t-statistics and adjusted R^2-statistics fall (row 3). Since s_t ($\phi = 0.94$) stays below its steady state value during the 1990s (see figure 2), it cannot convincingly track the cyclical upswing and stock market boom of the 1990s, which seems to reduce its forecasting power.

The benchmarks - the dividend yield and the net payout yield - occasionally have significant coefficients, but the surplus consumption ratio - in particular with $\phi = 0.73$ - has much higher t-statistics and adjusted R^2-statistics.[11] This evidence suggests that the surplus consumption ratio has additional explanatory power for future real stock returns relative to the dividend yield and the net payout yield. Controlling for the dividend yield in a multiple regression confirms that the surplus consumption is significant at all horizons (row 5). Whereas the dividend yield does not show much forecasting power in a univariate regression, it is a significant forecaster in a multiple regression joint with the surplus consumption ratio.

Moving on to the 1932-1991 subsample results, the surplus consumption ratio once again has a significant relationship with future real stock returns and explains around 7% of the 1 year ahead real stock return and around 30% of the 5 year ahead real stock return (non sensitive to the degree of habit persistence). The dividend yield is also a

[11] Surprisingly, the net payout yield looses much of its forecasting power due to the fact that the sample starts in 1932 and not in 1926 as in Boudoukh et al. (2007).

strong forecaster of future real stock returns, excluding the stock market boom of the 1990s. The dividend yield captures around 7% of the 1 year ahead real stock return and around 33% of the 5 year ahead real stock return (row 6). In a multiple regression, the surplus consumption ratio and the dividend yield are significant at all horizons, and the slope estimates are almost identical to the ones from the univariate regressions (row 10).

6 Statistical pitfalls in forecasting regressions

There are two statistical problems with the analysis in section 5. The first is the use of the Newey-West estimator to account for time-overlapping returns when $k > 1$ in (16), which causes serial correlation in the errors and leads to biased t-statistics and R^2-statistics. It is well-known that the Newey-West estimator does not produce reliable inference in a finite sample when the degree of time-overlap is large. Hodrick's (1992) implied statistics from low-order VAR models have been shown to have better finite-sample properties than Newey-West based statistics. Second, the estimates of β in (16) are subject to small sample bias, which arises using a persistent forecasting variable with innovations correlated with the innovations in (16), cf. Stambaugh (1999). To assess these statistical pitfalls, I apply a number of different robustness checks in the following.

Vector autoregressions
The Hodrick (1992) implied R^2-statistic does not rely on overlapping returns and is used as a robustness check of the R^2-statistic of the forecasting regressions. To calculate the implied R^2-statistic, I use first order VARs of the form:

$$Z_{t+1} = AZ_t + u_{t+1}, \tag{17}$$

where $Z_t = [r_t - E(r_t), x_t - E(x_t)]'$ contains the demeaned log real stock return and the demeaned forecasting variable, A contains the VAR coefficients, and u_{t+1} contains the residuals of the VAR. The implied k-period R^2-statistic of the VAR is given by:

$$R^2(k) = 1 - \frac{e1'W_k e1}{e1'V_k e1}, \tag{18}$$

where $e1 = [1\ 0]'$ is a selection vector, $e1'W_k e1$ is the innovation variance of the sum of k consecutive log real stock returns, and $e1'V_k e1$ is the variance of the sum of k consecutive Z_t's. See Hodrick (1992) for further details. Table 8 reports the Hodrick (1992) implied R^2-statistics. The table shows that the Hodrick (1992) implied R^2-statistics are higher, but in most cases not very different from the R^2-statistics of the forecasting regressions. The results confirm that the surplus consumption ratio is able to explain a substantial part of the variation of future stock returns.

The Lewellen (2004) bias correction method
Specifying the forecasting variable x_t as a stationary first-order autoregressive process,

Stambaugh (1999) sets up the following forecasting model:

$$r_{t+1} = c + \beta x_t + e_{t+1}, \qquad e_{t+1} \sim niid\left(0, \sigma_e^2\right) \tag{19}$$

$$x_{t+1} = d + \rho x_t + u_{t+1}, \qquad u_{t+1} \sim niid\left(0, \sigma_u^2\right) \tag{20}$$

where $|\rho| < 1$. Stambaugh (1999) shows that unless the covariance between the innovations in (19) and (20), σ_{eu}, is zero, the OLS estimate of β is biased. Stambaugh (1999) derives the small sample bias in $\hat{\beta}$ as:

$$E\left[\hat{\beta} - \beta\right] = \frac{\hat{\sigma}_{eu}}{\hat{\sigma}_u^2} E\left[\hat{\rho} - \rho\right], \tag{21}$$

and estimates the bias in $\hat{\rho}$ by $-\frac{1+3\hat{\rho}}{T}$, which leads to the following bias correction of $\hat{\beta}$:

$$\hat{\beta}_S = \hat{\beta} + \frac{\hat{\sigma}_{eu}}{\hat{\sigma}_u^2}\left[\frac{1+3\hat{\rho}}{T}\right]. \tag{22}$$

Lewellen (2004) improves the Stambaugh (1999) method by using information about the persistence of the forecasting variable to estimate the bias in $\hat{\rho}$. Assuming that $|\rho|<1$, Lewellen (2004) uses the insight that the bias in $\hat{\beta}$ is maximized when $|\rho|$ is close to 1. Hence, Lewellen (2004) estimates the bias in $\hat{\rho}$ by $\hat{\rho} - 0.9999$, which leads to the most conservative bias correction of $\hat{\beta}$:

$$\hat{\beta}_L = \hat{\beta} - \frac{\hat{\sigma}_{eu}}{\hat{\sigma}_u^2}\left[\hat{\rho} - 0.9999\right]. \tag{23}$$

If $\hat{\beta}$ is significant using the bias correction in (23), then $\hat{\beta}$ is also significant if the true value of ρ is less than 0.9999. To evaluate the small sample bias in the 1 year ahead non-overlapping forecasting regressions, I implement the Lewellen (2004) correction method in the following way:

Step 1) Estimate (19) and (20) by OLS.

Step 2) Estimate the bias corrected $\hat{\beta}$ as in (23).

Step 3) Regress $\hat{e}_{t+1}$ on $\hat{u}_{t+1}$, saving the residuals $\hat{v}_{t+1}$. Then estimate the standard error of $\hat{\beta}_L$ by the 2-2 element of $\sqrt{\hat{\sigma}_v^2\left(X'X\right)^{-1}}$, where X contains a vector of ones and x_t.

Table 9 reports that the bias corrected estimate of β and the corresponding t-statistic are very similar to the unadjusted estimate and t-statistic. Since the surplus consumption ratio is a macroeconomic variable, its innovations have very low correlation with the innovations of returns, which almost eliminates the small sample bias.

Bootstrap

A variety of studies also perform bootstrap analysis to assess the small sample bias in $\hat{\beta}$, see e.g. Nelson and Kim (1993) and Kothari and Shanken (1997). As a supplement to the Lewellen (2004) method, I apply the following bootstrap procedure:

Step 1) Estimate (19) and (20) by OLS.

Step 2) Construct 100,000 bootstrap samples by resampling the residuals from (19) and (20). I set the initial values of r_t and x_t equal to their sample averages and throw away the first 1,000 observations to avoid any effects from using the sample averages as starting values.

Step 3) Estimate β from each bootstrap sample and then estimate the bias in $\hat{\beta}$ by $\overline{\hat{\beta}}^* - \hat{\beta}$, where $\overline{\hat{\beta}}^* = \frac{1}{100{,}000}\sum_{j=1}^{100{,}000} \hat{\beta}_j^*$.

Step 4) Calculate 95% confidence intervals by using the lower 2.5% percentile and the upper 97.5% percentile of the 100,000 bootstrap samples.

The common practise is to bootstrap under the null of no predictability, cf. Nelson and Kim (1993), Kothari and Shanken (1997), Lettau and Ludvigson (2005), and others, but then the bootstrap is only correct if $\beta = 0$. Hence, I do not impose the constraint that $\beta = 0$. Table 10 shows the results of the bootstrap. Consistent with the Lewellen (2004) correction method, the bias in $\hat{\beta}$, measured as $\overline{\hat{\beta}}^* - \hat{\beta}$, is close to zero in all four cases. With $\phi = 0.94$, however, the bootstrap confidence interval implies that the surplus consumption ratio does not forecast real stocks returns. On the other hand, with $\phi = 0.66, 0.73, 0.80$, the bootstrap confidence intervals are similar to the OLS confidence intervals. Hence, the bootstrap analysis confirms the results of the forecasting regressions: with $\phi = 0.66, 0.73, 0.80$, the surplus consumption ratio forecasts real stock returns, but with $\phi = 0.94$, the surplus consumption is too persistent (it stays below its steady value during the cyclical upswing of the 1990s) and is not a strong forecaster of real stock returns.

7 Conclusions

This paper finds that the habit persistence model of Campbell and Cochrane (1999) is able to explain a substantial part of the cross sectional variation of returns on the 25 Fama and French value and size portfolios. In addition, the model accounts for time-varying expected returns on stocks. The Campbell-Cochrane model is particularly successful up to 1991. Including data from the 1990s reduces somewhat the fit of the model. The key feature of the model is a time-varying and counter-cyclical price of risk, which leads to increasing expected returns in recession times where consumption falls towards habit. The model needs an overall high level of risk aversion to fit the high equity premium on

US stocks and does not resolve the equity premium puzzle. Despite the high level of risk aversion, the model has the ability to produce a low risk free rate. But the model does not avoid the risk free rate puzzle in a convincing way, since the risk free rate is highly sensitive to the estimated model parameters. However, it is important to stress that this paper considers the basic version of the Campbell and Cochrane (1999) model, where the risk free rate does not show cyclical variation, that is the risk free rate is a constant. Wachter (2006) shows that the extended version of the model with cyclical variation in the risk free rate does in fact explain the term structure of interest rates for the US.

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9 Tables and figures

	BM1	BM2	BM3	BM4	BM5
S1	1.072 (35.7)	1.130 (33.0)	1.163 (31.5)	1.196 (35.2)	1.216 (38.1)
S2	1.105 (30.1)	1.147 (30.2)	1.162 (29.5)	1.173 (31.4)	1.184 (31.1)
S3	1.113 (29.7)	1.138 (26.2)	1.144 (25.5)	1.156 (25.8)	1.171 (30.5)
S4	1.099 (22.3)	1.111 (23.8)	1.140 (24.6)	1.142 (25.8)	1.162 (33.3)
S5	1.090 (19.9)	1.088 (18.0)	1.109 (19.6)	1.115 (23.3)	1.125 (26.6)

Table 1. Value and growth portfolios. This table reports average real gross returns on the 25 value weighted Fama and French portfolios formed on size and book-to-market (BM). Standard deviations in percent are in parentheses. The sample period is 1932-2003.

	$\phi\,(d_t - p_t)$	$\phi\,(npo_t - p_t)$	g	σ_c
1932-2003	0.936	0.729	0.025	0.016
	(0.044)	(0.077)	(0.002)	
1932-1991	0.802	0.659	0.025	0.017
	(0.075)	(0.091)	(0.003)	

Table 2. Parameter estimates. This table reports estimates of ϕ, g, and σ_c with standard errors in parentheses. The parameters are estimated on two samples: 1932-1991 and 1932-2003.

		δ	γ	γ/S	J_T	R^2	r_f
FF25	$\phi = 0.94$	0.97	15.25	60.74	123.80	50.57%	−8.39%
		(0.12)	(5.74)		(0.00)		
	$\phi = 0.73$	0.69	5.84	77.13	126.25	50.10%	−27.95%
		(0.20)	(2.49)		(0.00)		
FF6	$\phi = 0.94$	0.97	17.37	64.98	11.81	63.37%	−9.61%
		(0.15)	(6.94)		(0.02)		
	$\phi = 0.73$	0.66	6.41	81.04	10.81	63.49%	−30.03%
		(0.26)	(3.07)		(0.03)		
Size10	$\phi = 0.94$	0.97	17.38	65.01	4.10	99.16%	−10.17%
		(0.16)	(7.33)		(0.85)		
	$\phi = 0.73$	0.65	6.92	83.59	5.28	98.44%	−33.44%
		(0.27)	(3.34)		(0.73)		
BM10	$\phi = 0.94$	0.96	17.28	64.80	18.42	72.28%	−9.23%
		(0.13)	(6.88)		(0.02)		
	$\phi = 0.73$	0.67	6.28	79.43	15.93	72.12%	−30.10%
		(0.24)	(3.03)		(0.04)		
Industry10	$\phi = 0.94$	0.96	2.12	20.96	9.55	43.88%	4.54%
		(0.03)	(1.94)		(0.30)		
	$\phi = 0.73$	0.89	0.52	21.30	6.56	53.63%	5.62%
		(0.05)	(0.44)		(0.58)		
FF6Industry10	$\phi = 0.94$	0.96	5.41	34.92	62.79	44.99%	0.41%
		(0.05)	(2.74)		(0.00)		
	$\phi = 0.73$	0.84	1.49	41.52	62.18	47.71%	0.92%
		(0.09)	(0.71)		(0.00)		

Table 3. GMM estimation of the Campbell-Cochrane model, 1932-2003. This table reports GMM estimates of the impatience parameter, δ, and the utility curvature parameter, γ, with standard errors in parentheses. S in γ/S is the average value of the surplus consumption ratio over the sample. J_T is the Hansen's (1982) test of overidentifying restrictions calculated as in (14). P-values are in parentheses. R^2 is a goodness of fit measure calculated as in (15). r_f is the risk free rate calculated as in (7). The first column of the table refers to the cross section of portfolios: FF25 is the 25 Fama and French portfolios sorted on size and book-to-market. FF6 is the 6 Fama and French portfolios underlying the Fama and French (1993) factors SMB and HML. Size10 is 10 portfolios sorted on size. BM10 is 10 portfolios sorted on book-to-market. Industry10 is 10 portfolios sorted on industry. FF6I10 is FF6 joint with I10.

		δ	γ	γ/S	J_T	R^2	r_f
FF25	$\phi = 0.80$	0.76	5.92	60.71	172.26	60.99%	−16.85%
		(0.17)	(2.74)		(0.00)		
	$\phi = 0.66$	0.70	4.03	64.87	164.75	59.59%	−23.69%
		(0.20)	(1.99)		(0.00)		
FF6	$\phi = 0.80$	0.74	6.62	64.43	5.97	79.20%	−18.74%
		(0.22)	(3.31)		(0.20)		
	$\phi = 0.66$	0.67	4.61	69.85	6.29	79.60%	−27.36%
		(0.25)	(2.46)		(0.18)		
Size10	$\phi = 0.80$	0.72	7.26	67.69	9.13	98.93%	−21.00%
		(0.25)	(3.81)		(0.33)		
	$\phi = 0.66$	0.64	5.13	74.21	6.90	98.57%	−30.35%
		(0.28)	(2.80)		(0.55)		
BM10	$\phi = 0.80$	0.72	7.13	67.05	21.67	80.47%	−20.11%
		(0.21)	(3.45)		(0.01)		
	$\phi = 0.66$	0.66	4.68	70.46	18.07	79.87%	−27.59%
		(0.24)	(2.44)		(0.02)		
Industry10	$\phi = 0.80$	0.90	0.58	17.88	7.12	50.28%	6.79%
		(0.04)	(0.51)		(0.52)		
	$\phi = 0.66$	0.88	0.31	14.98	6.94	51.79%	7.84%
		(0.05)	(0.27)		(0.54)		
FF6Industry10	$\phi = 0.80$	0.86	2.36	36.43	56.84	53.00%	−3.03%
		(0.07)	(1.23)		(0.00)		
	$\phi = 0.66$	0.82	1.09	35.89	53.20	55.07%	3.75%
		(0.09)	(0.53)		(0.00)		

Table 4. GMM estimation of the Campbell-Cochrane model, 1932-1991. See notes to table 3.

	γ	γ/S	J_T
Sample: 1932-2003			
$\phi = 0.94$	15.62	61.49	8.95
	(5.71)		(0.11)
$\phi = 0.73$	6.25	62.48	8.95
	(3.00)		(0.10)
Sample: 1932-1991			
$\phi = 0.80$	4.91	70.07	10.09
	(2.39)		(0.07)
$\phi = 0.66$	4.41	68.01	7.96
	(2.27)		(0.16)

Table 5. Conditional GMM estimation of the Campbell-Cochrane model. This table reports GMM estimates of the utility curvature parameter, γ, with standard errors in parentheses. S in γ/S is the average value of the surplus consumption ratio over the sample. J_T is the Hansen's (1982) test of overidentifying restrictions calculated as in (14). P-values are in parentheses. The test portfolios are the 3 Fama and French factors. With $\phi = 0.94$ and 0.73, the dividend yield is used as instrument variable. With $\phi = 0.80$ and 0.66, the net payout yield is used as instrument variable. The sample periods are 1932-2003 and 1932-1991.

	r_t	$d_t - p_t$	$npo_t - p_t$	$s_t\,(\phi = 0.94)$	$s_t\,(\phi = 0.73)$
Sample: 1932-2003					
Means	0.076	-3.293	-2.029	-1.388	-2.619
Standard deviations	0.182	0.420	0.107	0.113	0.282
Correlations					
r_t	1.000	-0.030	0.003	-0.018	0.089
$d_t - p_t$		1.000	0.773	0.377	-0.030
$npo_t - p_t$			1.000	0.427	0.083
$s_t\,(\phi = 0.94)$				1.000	0.826
$s_t\,(\phi = 0.73)$					1.000
	r_t	$d_t - p_t$	$npo_t - p_t$	$s_t\,(\phi = 0.80)$	$s_t\,(\phi = 0.66)$
Sample: 1932-1991					
Means	0.075	-3.149	-2.009	-2.361	-2.738
Standard deviations	0.183	0.263	0.105	0.263	0.300
Correlations					
r_t	1.000	-0.095	-0.022	0.049	0.163
$d_t - p_t$		1.000	0.813	-0.086	-0.181
$npo_t - p_t$			1.000	0.101	0.011
$s_t\,(\phi = 0.80)$				1.000	0.959
$s_t\,(\phi = 0.66)$					1.000

Table 6. Summary statistics. This table reports summary statistics for the log real stock return, r_t, the log dividend yield, $d_t - p_t$, the log net payout yield, $npo_t - p_t$, and the log surplus consumption ratio, s_t. The sample periods are 1932-2003 and 1932-1991.

	Forecasting variable		Forecasting horizon in years 1	2	3	4	5
Sample: 1932-2003							
1	$d_t - p_t$	$\hat{\beta}$	0.08	0.18	0.24	0.27	0.33
		t_{NW}	1.52	1.89	1.95	1.73	1.51
		$\bar{R}^2$	2.25	7.07	9.69	9.51	9.15
2	$npo_t - p_t$	$\hat{\beta}$	0.40	0.67	0.75	0.78	1.01
		t_{NW}	2.12	2.12	1.99	1.60	1.47
		$\bar{R}^2$	4.12	6.96	6.82	5.89	8.15
3	$s_{t-1}\,(\phi = 0.94)$	$\hat{\beta}$	−0.40	−0.60	−0.75	−0.85	−0.97
		t_{NW}	−2.62	−2.77	−2.75	−2.40	−2.30
		$\bar{R}^2$	5.18	6.74	8.87	10.14	10.32
4	$s_{t-1}\,(\phi = 0.73)$	$\hat{\beta}$	−0.21	−0.33	−0.44	−0.53	−0.63
		t_{NW}	−3.22	−3.63	−4.49	−4.11	−3.85
		$\bar{R}^2$	9.91	14.56	21.18	26.88	29.92
5	$d_t - p_t$	$\hat{\beta}$	0.09	0.19	0.26	0.33	0.42
		t_{NW}	2.12	2.88	3.44	3.73	3.64
	$s_{t-1}\,(\phi = 0.73)$	$\hat{\beta}$	−0.21	−0.35	−0.47	−0.58	−0.70
		t_{NW}	−3.30	−3.57	−4.18	−4.17	−4.23
		$\bar{R}^2$	13.06	23.27	33.83	41.98	46.62
Sample: 1932-1991							
6	$d_t - p_t$	$\hat{\beta}$	0.20	0.36	0.47	0.55	0.72
		t_{NW}	2.43	2.60	3.02	3.01	3.00
		$\bar{R}^2$	6.85	13.82	21.73	25.95	33.41
7	$npo_t - p_t$	$\hat{\beta}$	0.43	0.68	0.80	0.92	1.28
		t_{NW}	2.38	2.23	2.31	1.98	1.92
		$\bar{R}^2$	4.42	7.11	9.42	10.81	16.30
8	$s_{t-1}\,(\phi = 0.80)$	$\hat{\beta}$	−0.20	−0.31	−0.42	−0.53	−0.67
		t_{NW}	−2.71	−3.21	−3.91	−3.55	−3.39
		$\bar{R}^2$	7.12	10.36	18.41	26.52	31.62
9	$s_{t-1}\,(\phi = 0.66)$	$\hat{\beta}$	−0.18	−0.25	−0.33	−0.42	−0.58
		t_{NW}	−2.63	−2.93	−4.39	−3.75	−3.70
		$\bar{R}^2$	7.71	8.27	14.01	21.12	29.94
10	$d_t - p_t$	$\hat{\beta}$	0.20	0.36	0.47	0.54	0.72
		t_{NW}	2.75	3.17	4.19	5.57	7.42
	$s_{t-1}\,(\phi = 0.80)$	$\hat{\beta}$	−0.20	−0.30	−0.42	−0.53	−0.67
		t_{NW}	−2.93	−3.90	−5.90	−6.10	−6.06
		$\bar{R}^2$	14.10	24.42	40.75	53.39	66.63

Table 7. Forecasting real stock returns. This table reports slope estimates, Newey West (1987) corrected t-statistics with $k+1$ lags, and adjusted R^2-statistics for k-period forecasting regressions: $r_{t+1\to k} = \sum_{i=1}^{k} r_{t+i} = c + \beta x_t + e_{t+1\to k}$. $r_{t+1\to k}$ is the k-period log real stock return, and x_t is one of the forecasting variables: the log dividend yield, $d_t - p_t$, the log net payout yield, $npo_t - p_t$, or the surplus consumption ratio, s_{t-1}. The sample periods are 1932-2003 and 1932-1991.

Forecasting variable		Forecasting horizon in years 1	2	3	4	5
Sample: 1932-2003						
$s_{t-1}\,(\phi = 0.94)$	Implied R^2	8.49	14.62	19.84	24.10	27.53
	$\bar{R}^2$	5.18	6.74	8.87	10.14	10.32
$s_{t-1}\,(\phi = 0.73)$	Implied R^2	13.89	22.13	27.13	29.87	31.08
	$\bar{R}^2$	10.04	14.78	21.50	27.36	30.56
Sample: 1932-1991						
$s_{t-1}\,(\phi = 0.80)$	Implied R^2	13.76	21.90	26.85	29.56	30.75
	$\bar{R}^2$	7.12	10.36	18.41	26.52	31.62
$s_{t-1}\,(\phi = 0.66)$	Implied R^2	13.29	19.46	22.49	23.54	23.44
	$\bar{R}^2$	7.71	8.27	14.01	21.12	29.94

Table 8. Hodrick (1992) implied R^2-statistics. Implied R^2 is the Hodrick implied R^2-statistic and $\bar{R}^2$ is the adjusted R^2-statistic of the forecasting regressions. The sample periods are 1932-2003 and 1932-1991.

	$\hat{\beta}$	$\hat{\beta}_L$
Sample: 1932-2003		
$s_{t-1}\,(\phi = 0.94)$	-0.40	-0.40
	(-2.18)	(-2.17)
$s_{t-1}\,(\phi = 0.73)$	-0.21	-0.20
	(-2.93)	(-2.86)
Sample: 1932-1991		
$s_{t-1}\,(\phi = 0.80)$	-0.20	-0.20
	(-2.32)	(-2.28)
$s_{t-1}\,(\phi = 0.66)$	-0.18	-0.19
	(-2.40)	(-2.42)

Table 9. Small sample bias: The Lewellen (2004) correction method. $\hat{\beta}$ is the standard OLS slope and $\hat{\beta}_L$ is the Lewellen (2004) bias corrected estimate. t-statistics are in parentheses. The sample periods are 1932-2003 and 1932-1991

	$\hat{\beta}$	$\overline{\hat{\beta}}^*$	95% CI	95% bootstrap CI
Sample: 1932-2003				
$s_{t-1}\,(\phi = 0.94)$	-0.40	-0.40	$(-0.75, -0.04)$	$(-0.88, +0.10)$
$s_{t-1}\,(\phi = 0.73)$	-0.21	-0.21	$(-0.35, -0.07)$	$(-0.37, -0.05)$
Sample: 1932-1991				
$s_{t-1}\,(\phi = 0.80)$	-0.20	-0.20	$(-0.37, -0.03)$	$(-0.40, -0.00)$
$s_{t-1}\,(\phi = 0.66)$	-0.18	-0.18	$(-0.33, -0.03)$	$(-0.34, -0.02)$

Table 10. Small sample bias: Bootstrap. $\hat{\beta}$ and 95% CI are standard OLS slope estimates and 95% confidence intervals, whereas $\overline{\hat{\beta}}^*$ and 95% bootstrap CI are the bootstrap counterparts. The sample periods are 1932-2003 and 1932-1991.

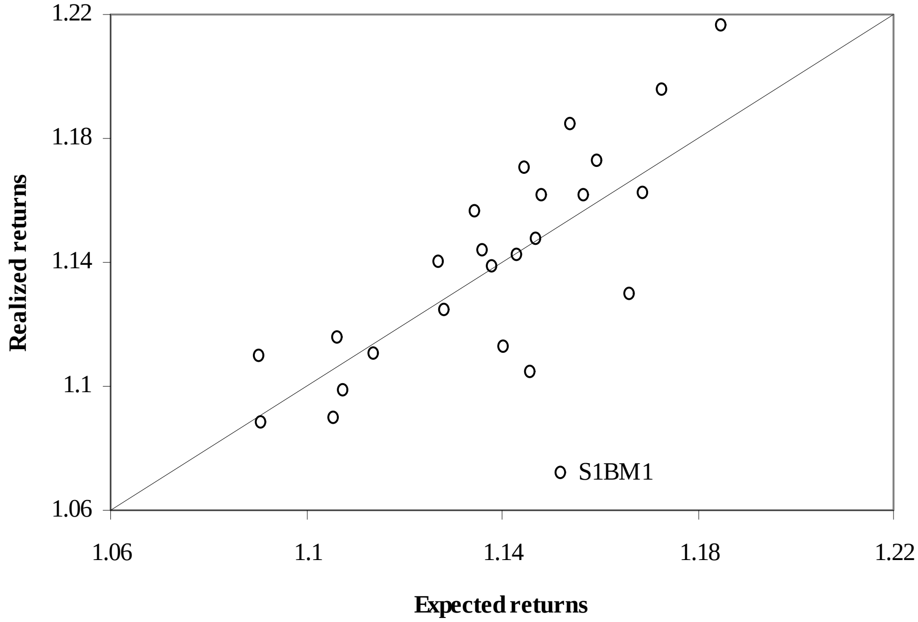

Figure 1. Realized returns vs. expected returns. The figure plots the realized average real gross returns against the model expected returns on the 25 Fama French portfolios. The samples period is 1932-2003.

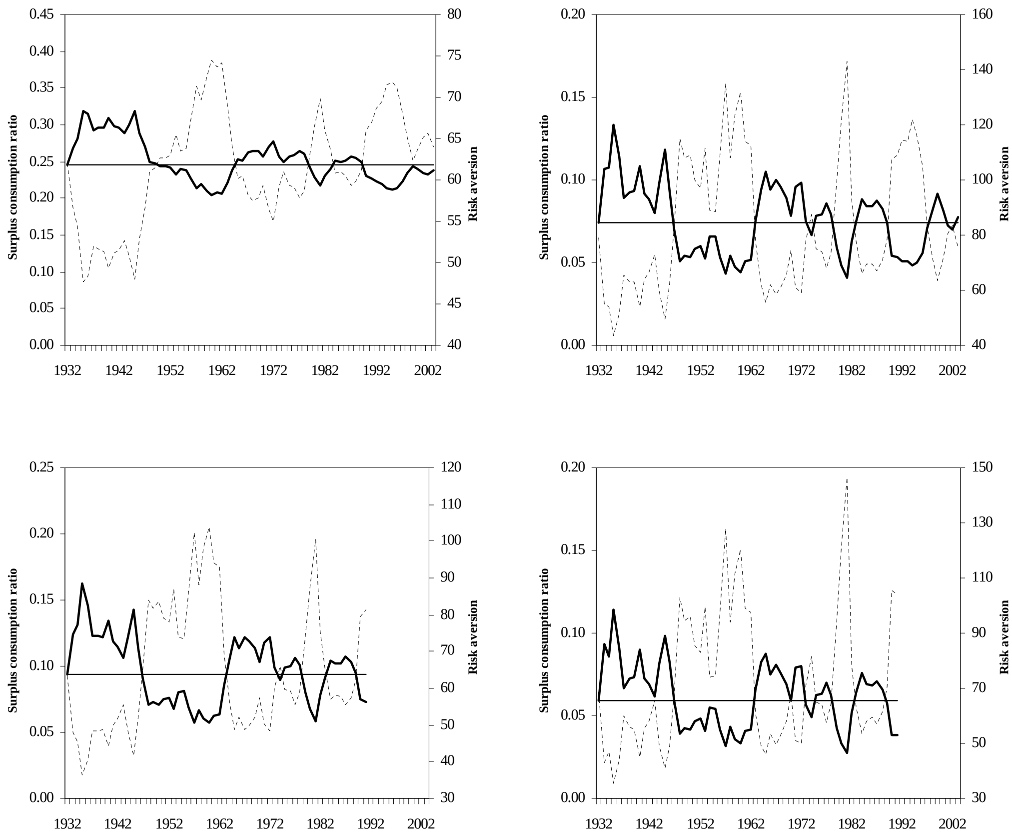

Figure 2. Surplus consumption ratio and relative risk aversion. The figure shows the surplus consumption ratio (solid line), the steady state value (thin line), and the relative risk aversion (dashed line). Top left panel shows the 1932-2003 sample where $\phi = 0.94$. Top right panel shows the 1932-2003 sample where $\phi = 0.73$. Bottom left panel shows the 1932-1991 sample where $\phi = 0.80$. Bottom right panel shows the 1932-1991 sample where $\phi = 0.66$.

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