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Abstract

Based on CAMPBELL and COCHRANE [1999] Consumption-Based Asset Pricing Model (C)CAPM with habit formation, this paper provides empirical evidence in favor of the importance of habit persistence in asset pricing. Using U.S data, we show that the surplus consumption ratio is a strong predictor of excess returns at long-horizons and that it captures a component of expected returns, not explained by the consumption-wealth ratio. Moreover, this paper shows that the (C)CAPM with habit formation performs far better than the standard (C)CAPM in accounting for the cross-sectional variations in average excess returns on the 25 FAMA-FRENCH portfolios sorted by size and book-to-market value.

Keywords: Habit formation, Surplus consumption ratio, Expected returns, Time series predictability, Cross section returns.
JEL classification: G12, E21

Résumé

En se basant sur le modèle d'évaluation des actifs financiers basé sur la consommation (C)CAPM de CAMPBELL et COCHRANE [1999], ce papier met en évidence l'importance de la persistance des habitudes dans la prédiction des rendements futurs. Nous montrons que le ratio de surplus de consommation est un indicateur à fort pouvoir prédictif des rendements excédentaires à long terme et qu'il capture une composante des rendements futurs non expliquée par le ratio de consommation sur richesse. De plus, le papier montre que le (C)CAPM avec formation des habitudes est plus performant que le (C)CAPM standard dans l'explication des variations en coupes transversales des rendements moyens des 25 portefeuilles de FAMA-FRENCH triés suivant les critères de la taille et du ratio de la valeur de marché par rapport à la valeur comptable.

Mots-clés : Formation des habitudes, Ratio de surplus de consommation, Rendements anticipés, Prédiction des séries temporelles, Variation des rendements moyens en coupes transversales.
Code JEL : G21, E21

I. Introduction

One of the motivations of financial studies is to understand the linkage between macroeconomics and financial markets. On the theoretical side, a large body of the literature that tries to deal with this point is based on Consumption–Based Asset Pricing Models (C)CAPM. For instance, in their seminal paper, CAMPBELL and COCHRANE [1999] proposed a consumption–based model with nonlinear habit formation. They show that a low surplus consumption ratio or, equivalently, a low consumption to habit stock ratio, predicts high expected returns in the next period. WACHTER [2006] extended the model to develop a consumption–based term structure model. On the empirical side, extensive research has been devoted to providing evidence that some financial and macroeconomic indicators have predictive power for stock returns (See, among others, FAMA and FRENCH [1988 and 1989]; CAMPBELL and SHILLER [1988] and HODRICK [1992]) and may explain the cross–sectional variations in average stock returns. For instance, LETTAU and LUDVIGSON[2001a and 2001b] investigated the linkage between excess returns and the consumption—wealth ratio. They showed that the consumption–wealth ratio is a good predictor of excess returns at short and intermediate horizons. In addition, used as a conditioning variable, the consumption–wealth ratio improves the performance of the standard Consumption-Based Asset Pricing Model (C)CAPM in explaining the cross section of expected returns. LI [2005] empirically tested the long-horizon predictive power of the actual surplus consumption ratio and the price–dividend ratio on excess stock returns. However, the empirical literature has not investigated the ability of the surplus consumption ratio to explain the cross section of expected returns. This paper is an attempt to fill this gap.

Our main findings can be summarized as follows. First, we show empirically that the surplus consumption ratio is a good predictor for excess stock returns at long horizons, and it captures a component of expected returns which is not explained by the consumption–wealth ratio. Second, we show empirically that the CAMPBELL and COCHRANE [1999]'s model with habit formation successfully captures the cross–sectional variation in average returns on portfolios sorted by size and book-to-market value. The key risk factor is the lagged surplus consumption ratio, as it predicts the price of risk. Therefore, this paper gives empirical evidence in favour of the importance of habit persistence in asset pricing. In the language of macroeconomics, the surplus consumption ratio is the leading business cycle variable in the (C)CAPM models with habit formation. The consumer develops habits for higher or lower past consumptions and therefore, the stock of habit captures her standard of living or the consumption trend. The persistence of habit implies that the standard of living, depending on past consumption, has an impact on how the consumer feels about more consumption today. The time nonseparable property of the utility function, generated by the introduction of habit persistence, implies that after periods of low consumption growth, the volatility of the investors' marginal utility rises, increasing their demand for larger premia on risky assets. Moreover, in recession (expansion) periods, consumption decreases (increases) relative to the reference level, implying a decrease (increase) in the surplus consumption ratio. Thus, the time-varying and pro-cyclical surplus consumption ratio enables the (C)CAPM models with habit formation to replicate the time–varying and counter–cyclical equity premium. In the language of finance, equilibrium asset pricing models imply that time-variation in the equity premium must be explained by time variation in the price and/or the quantity of risk. In the CAMPBELL and COCHRANE [1999]'s (C)CAPM model with habit formation, the quantity of risk is measured by the covariance of stock returns with current consumption growth and surplus consumption ratio. The price of risk is measured by the coefficient of risk aversion, which is negatively linked to the surplus consumption ratio. Therefore, both the quantity and price of risk increase during periods of recession, implying an increase in expected equity premium. Equivalently, the time-varying surplus consumption ratio drives the time–varying and counter–cyclical equity premium.

To empirically investigate the predictive power of the surplus consumption ratio and the dividend-to-price ratio, LI [2005] used VAR estimation as proposed by HODRICK [1992], to mitigate the finite sample bias that may rise when studying long-horizon returns. However, this econometric methodology does not take into account the high persistence of the explanatory variables. This paper proposes a Monte Carlo experiment accounting for the biased coefficient estimators and the distorted distribution of test statistics due to (*i*) the feedback effect, (*ii*) the highly persistent explanatory variables and (*iii*) the overlapping data. Using annual data, we find that the surplus consumption ratio is indeed a strong predictor of excess returns at long horizons, as in LI [2005]. For instance, the surplus consumption ratio explains 35% of the variability of excess returns at the 5–year horizon. Moreover, empirical findings suggest that the surplus consumption ratio predicts a component of expected excess returns which is not captured by the proxy for the consumption–wealth ratio, *cay*, proposed by LETTAU and LUDVIGSON [2001a,b and 2005]. In contrast with LI [2005], the dividend–price ratio fails to predict excess returns at any horizon. The main scope of this paper is the cross–sectional analysis of average stock returns. We show that the habit formation models perform far better than the standard (C)CAPM model in accounting for the cross–sectional variations in average excess returns on the 25 FAMA–FRENCH portfolios sorted by size and book–to–market value. The leading risk factor is the lagged surplus consumption ratio. Indeed, it explains about 42% of the variation in average excess returns and seems to mimic the risk factors related to the size effect. Additional experiments are run to study the robustness of our empirical findings.

The paper is structured as follows. Section II presents the CAMPBELL and COCHRANE [1999]'s consumption-based model (C)CAPM with habit formation, which forms the basis of our empirical work. Section III confronts the theoretical implications of the CAMPBELL and COCHRANE [1999] model with the actual U.S data. First, we investigate the long–horizon predictability, then we explore the ability of the surplus consumption ratio to explain the cross–sectional variations in average returns. The final section provides the conclusion.

II. Theoretical Framework

This section presents, from CAMPBELL and COCHRANE [1999]'s Consumption–Based Asset Pricing Model (C)CAPM with external habit formation of CAMPBELL and COCHRANE [1999], the theoretical framework linking the surplus consumption ratio with expected stock returns.

We consider an endowment economy with complete markets and a representative consumer. The preferences of the representative agent are represented by the following inter–temporal utility function:

$$\mathbb{E}_t \sum_{i=0}^{\infty} \beta^i u_{t+i}$$

where $\beta > 0$ is the subjective discount factor and u_t denotes the instantaneous utility function. Expectations are conditional on information available at the beginning of period t. The preferences of the agent are assumed to be time non–separable. The agent derives her instantaneous utility for period t from her individual current consumption C_t as well as a reference level X_t:

$$u_t = U(C_t, X_t)$$

The reference level X_t is assumed to capture the influence of the history of aggregate consumption choices $\{\overline{C}_{t-\tau}, \tau \succeq 0\}$ on current individual choices. Therefore, X_t is also called external habit stock. In their seminal paper, CAMPBELL and COCHRANE [1999] specify their instantaneous utility function in difference:

$$U(C_t, X_t) = \frac{(C_t - X_t)^{1-\theta} - 1}{1-\theta} \tag{1}$$

where $\theta > 0$ denotes the utility curvature parameter.

Let $S_t = \frac{C_t - X_t}{C_t}$ denote the surplus consumption ratio. It is worth noting that the specification of the utility function in difference generates time–varying risk aversion, as the coefficient of relative risk aversion is equal to $\frac{\theta}{S_t}$.

Following CAMPBELL and COCHRANE [1999], the (log) surplus consumption ratio[1] s_t is assumed to evolve as:

$$s_t = (1-\phi)\overline{s} + \phi s_{t-1} + \lambda(s_{t-1})(\Delta\overline{c}_t - g) \tag{2}$$

where $\Delta\overline{c}_t$ is aggregate consumption growth and g denotes average aggregate consumption growth. The sensitivity function $\lambda(s_t)$ is defined as follows:

$$\lambda(s_t) = \begin{cases} \frac{1}{\overline{S}}\sqrt{1-2(s_t-\overline{s})} - 1 & \text{if } s_t \leqslant s_{max} \\ 0 & \text{otherwise} \end{cases}$$

[1] Throughout, lowercase letters are used for variables in logarithms.

where $\overline{s} = \log \overline{S} = \log \left(\sigma \sqrt{\frac{\theta}{1-\phi}} \right)$ and $s_{max} = \log(S_{max}) = \left(\overline{s} + \frac{1}{2}(1 - \overline{S}^2) \right)$. The parameter σ denotes the standard deviation of consumption growth.

It is worth noting that the benchmark model, i.e. the CAMPBELL and COCHRANE [1999] model, presents two key ingredients. First, the utility function is specified in difference. This implies a time varying coefficient of risk aversion. The second ingredient is that the surplus consumption ratio is nonlinear and moves slowly in response to consumption. The nonlinearity is essential in keeping habit always below consumption, and therefore in guaranteing positive and finite marginal utility.

For any asset j, the first order condition of the agent maximization program yields the following asset pricing equation:

$$1 = \mathbb{E}_t \left[M_{t,t+1} R_{j,t+1} \right] \tag{3}$$

where $R_{j,t+1}$ denotes the gross return on asset j and $M_{t,t+1}$ the inter–temporal marginal rate of substitution between t and $t+1$ or, equivalently, the stochastic discount factor between t and $t+1$.

For these preferences, the inter–temporal rate of substitution is rewritten:

$$M_{t,t+1} = \beta \left(\frac{C_{t+1}}{C_t} \right)^{-\theta} \left(\frac{S_{t+1}}{S_t} \right)^{-\theta} \tag{4}$$

Plugging expression (4) into Euler equation (3), it is useful to consider the following approximation[2] of Equation (3) for the gross return on the stock market portfolio R_{t+1}:

$$\mathbb{E}_t \Delta c_{t+1} \simeq \frac{1}{\theta} \mathbb{E}_t \left(r_{t+1} - \delta \right) - \mathbb{E}_t s_{t+1} + s_t, \tag{5}$$

where $\delta \equiv (1 - \beta) / \beta$. Iterating forward Equation (5), the surplus consumption ratio is given by:

$$s_t \simeq \mathbb{E}_t \sum_{i=1}^{\infty} \left(\triangle c_{t+i} - \frac{1}{\theta} \left(r_{t+i} - \delta \right) \right) + \lim_{i \to \infty} \mathbb{E}_t s_{t+i}. \tag{6}$$

Equation (6) confirms the well-established fact that the surplus consumption ratio s_t is a good candidate to predict stock returns or consumption growth at long horizons. Furthermore, it indicates that the surplus consumption ratio and stock returns are negatively related at any horizon. What is the economic explanation for this result? Intuitively, recessions periods in (C)CAPM model with habit formation are characterized by both (i) low consumption and (ii) low consumption relative to the habit stock. Therefore, risky stocks are defined as assets that do not insure the consumer against either a decease in

[2] It is worth noting that variance terms are missing in Equation 5. Actually, the aim of this approximation is to provide a theoretical support to the empirical study of the *linear* (inverse) relation between the surplus consumption ratio and expected stock returns at long horizons using actual data. This inverse–relation is well–established and tested by CAMPBELL and COCHRANE [1999] using simulated data rather than actual data.

his consumption or a decrease in his consumption compared to his reference level. In the CAMPBELL and COCHRANE [1999] model, the quantity of risk is measured by the covariance of stock returns with consumption growth and surplus consumption ratio. Moreover, the price of a unit of risk is measured by counter–cyclical risk aversion. In recession periods, both the quantity and the price of risk increase, implying an increase in risk premium.

Moreover, the expression of the inter–temporal rate of substitution (4) suggests that the surplus consumption ratio should forecast changes in asset prices and therefore explains the cross sectional average returns. Indeed, plugging the expression of the surplus consumption ratio (2) into the stochastic discount factor (4), we obtain:

$$m_{t,t+1} = \tau_0(s_t) + \tau_1(s_t)\Delta c_{t+1} \tag{7}$$

where:

$$\begin{aligned}\frac{\tau_0(s_t)}{\beta} &= 1 - \theta(1-\phi)(\overline{s} - s_t) + \theta g \lambda(s_t) \\ \frac{\tau_1(s_t)}{\beta} &= -\left(1 - \lambda(s_t)\right)\theta \Delta c_{t+1}\end{aligned}$$

Hence, the stochastic discount factor associated with the CAMPBELL and COCHRANE [1999] model can be written as a linear beta pricing model [3] with time varying coefficients τ_0 and τ_1. The source of the variation of these parameters is the surplus consumption ratio s_t.

As suggested by LETTAU and LUDVIGSON [2001b], a linearization of τ_0 and τ_1 allows us to rewrite the linear beta model with time–varying coefficients (7) as a linear beta model with constant coefficients. Assuming $\tau_0 \simeq \eta_0 + \iota_0 s_t$ and $\tau_1 \simeq \eta_1 + \iota_1 s_t$, the stochastic discount factor $M_{t,t+1}$ can be written as follows:

$$M_{t,t+1} \simeq b_0 + b_1 s_t + b_2 \Delta c_{t+1} + b_3 s_t \Delta c_{t+1} \tag{8}$$

Plugging expression (8) into the Euler equation (3), we obtain:

$$1 = \mathbb{E}\left[\left(b_0 + b_1 s_t + b_2 \Delta c_{t+1} + b_3 s_t \Delta c_{t+1}\right) R_{j,t+1}\right] \tag{9}$$

[3] Following COCHRANE [2005], we refer to pricing models of the form:

$$1 = \mathbb{E}\left[mR\right]$$

where

$$m = a + b'f$$

is a linear beta pricing model or a beta representation model. The variables f denote the risk factors. These models imply the following cross-sectional representation:

$$\mathbb{E}\left[R_{i,t}\right] = \mathbb{E}\left[R_{f,t}\right] + \beta_{i,f}\lambda_f$$

where λ_f denotes the prices of risk corresponding to the risk factors f. See COCHRANE [2005], chapter 6 for more details.

for each asset j. It is straightforward to show that equation (9) implies the following unconditional beta representation:

$$\mathbb{E}\left[R_{i,t}\right] = \mathbb{E}\left[R_{f,t}\right] + \beta_{i,\Delta c}\lambda_{\Delta c} + \beta_{i,s_{-1}}\lambda_{s_{-1}} + \beta_{i,s_{-1}\Delta c}\lambda + \beta_{i,}\lambda_{s_{-1}\Delta c} \tag{10}$$

where $\mathbb{E}$ denotes the unconditional mean. Hence, the CAMPBELL and COCHRANE [1999] model can be written as an unconditional mutli–factor model. The risk factors are consumption growth, lagged surplus consumption ratio and their product.

To summarize, this section presents two theoretical implications of the CAMPBELL and COCHRANE [1999] model. First, the surplus consumption ratio is a good candidate to forecast future excess returns at any horizon as mentioned by CAMPBELL and COCHRANE [1999] and LI [2005]. Moreover, the CAMPBELL and COCHRANE [1999] model implies a linear three–factor model that rivals the conditional (C)CAPM model proposed by LETTAU and LUDVIGSON [2001b] and the FAMA and FRENCH [1993] three–factor model in explaining the cross–section of expected returns. Both implications will be evaluated empirically in the next section.

III. Empirical Investigation

This section explores empirically the time–series and the cross–sectional relations between the surplus consumption ratio and excess stock returns. As a benchmark, we consider the Consumption–Based Asset Pricing Model (C)CAPM with external habit formation proposed by CAMPBELL and COCHRANE [1999]. Despite the fact that the surplus consumption ratio s_t is not observable in the CAMPBELL and COCHRANE [1999] model, equation (2) can be used to generate a time–series for s_t. This requires to set ϕ, g, σ and θ. The model is calibrated at annual frequency. The utility curvature parameter, θ is set to 2, a commonly used value in the literature. The parameters g and σ are estimated using annual real consumption data, implying $g = 2.01\%$ and $\sigma = 1.14\%$. The parameter ϕ is set to match the first order serial correlation of the price–dividend ratio, implying $\phi = 0.89$. All these values are close to those used by CAMPBELL and COCHRANE [1999]. The initial value of the times–series for the surplus consumption ratio is set to its steady-state value, $\overline{s}$. To check the robustness of our empirical results, we evaluate the sensitivity of the predictive power of the surplus consumption ratio to alternative values of (i) the degree of curvature of the utility function θ and (ii) the initial value of the time–series for the surplus consumption ratio.

Additionally, the forecasting power of the surplus consumption ratio will be compared to the well–documented predictive power of the (log) price–dividend ratio $p_t - d_t$ and the (log) consumption to aggregate wealth ratio $c_t - w_t$. Following LETTAU and LUDVIGSON [2001a and 2005], we use the deviation from the estimated shared trend among consumption, asset holdings and labor income –denoted by cay_t– as a proxy for the unobservable consumption–wealth ratio.

III.1. Long horizon Regressions

This section studies empirically the role of fluctuations in the surplus consumption ratio for predicting excess stock returns. The macroeconomic and financial data used in this study are borrowed from LETTAU and LUDVIGSON [2005][4] and GARCIA, MEDDAHI and TEDONGAP [2008][5]. The data used are annual US data from 1948 to 2001. The financial data include (*i*) the real U.S three–month treasury bill as proxy for the risk–free rate, (*ii*) the real value weighted returns on CRSP index (which includes the NYSE, AMEX and NASDAQ) as proxy for the market return and (*iii*) the corresponding price–dividend ratio. The macroeconomic data are (*i*) the real per capita consumption for nondurables and services, excluding shoes and clothing and (*ii*) the *cay* as a proxy for the unobservable consumption–wealth ratio. *cay* is measured as follows. First, LETTAU and LUDVIGSON [2001a] define the aggregate total wealth as the sum of human and non–human wealth. Therefore, (log) aggregate wealth may be approximated as a weighted average of asset holdings a_t and labor income y_t. Aggregate U.S. asset holdings a_t are defined as the household net worth series provided by the Board of Governors of the Federal Reserve, and U.S. labor income y_t is defined as wages and salaries plus transfer payments plus other labor income minus personal contributions for social insurance, minus taxes[6]. Then LETTAU and LUDVIGSON [2001a] show that aggregate consumption, asset holdings and labor income share a common long–term trend, but may deviate substantially from one another in the short run. This “trend” deviation, so–called *cay*, is a good proxy for the unobservable consumption–wealth ratio.

We explore the predictive power of the (log) surplus consumption ratio s_t, the (log) price–dividend ratio $p_t - d_t$ and the (log) consumption–wealth ratio $c_t - w_t$ at annual frequency.

Table 1 presents summary statistics for $p_t - d_t$, cay_t and s_t. Two main results emerge. First, the price–dividend ratio and the surplus consumption ratio are highly persistent. Their first–order autocorrelations are 0.89 and their second–order autocorrelations are 0.75 and 0.71 respectively. As documented by LETTAU and LUDVIGSON [2005], cay_t is less persistent and its autocorrelations die out more quickly. Its first–order correlation is about 0.57 and its second–order correlation is 0.14. Second, s_t is weakly correlated to other indicators. The correlations between s_t and cay_t or $p_t - d_t$ are -0.14 and -0.25, respectively.

A common way to investigate the predictive power of the surplus consumption ratio at long horizons is to run regressions for the compounded (log) excess returns $er_{t,t+k}$ on s_t evaluated at several lags:

$$er_{t,t+k} = \alpha_k + \beta_k s_t + u_{t+k,t} \tag{11}$$

where $u_{t+k,t}$ is drawn from a Gaussian distribution with mean zero and constant standard deviation. By construction, the surplus consumption ratio is very persistent. Therefore, several econometric issues arise when assessing the forecasting power of the surplus con-

[4] More details on the data can be found in the appendix to LETTAU and LUDVIGSON(2005), downloadable from http://www.econ.nyu.edu/user/ludvigsons/dappendixe.pdf.

[5] More details on data can be found in GARCIA, MEDDAHI and TEDONGAP [2008].

[6] See the appendix in LETTAU and LUDVIGSON [2001a] for a detailed data description.

	Autocorrelations				Correlation Matrix		
	ρ_1	ρ_2	ρ_3	ρ_4	$p-d$	cay	s
$p-d$	0.89	0.75	0.63	0.53	1.0	-0.53	-0.25
cay	0.57	0.14	0.05	0.01		1.0	-0.14
s	0.89	0.71	0.52	0.34			1.0

Table 1: Summary Statistics using Annual Data

Note: This table reports summary statistics for the (log) price–dividend ratio $p-d$, the (log) surplus consumption ratio s, and the proxy for the (log) consumption–wealth ratio cay. The sample is annual and spans 1948 to 2001.

sumption ratio s_t. As documented by STAMBAUGH [1999] and VALKANOV [2003] among others, highly persistent explanatory variables, and the existence of a strong correlation between unexpected returns and innovations of the explanatory variables ought to distort Ordinary Least Squares (OLS) estimators in a finite sample. In order to investigate this issue, we follow VALKANOV [2003] and run a Monte Carlo experiment under the null of no predictability, assuming that the explanatory variable s_t follows a Gaussian $AR(1)$ process. More precisely, we generate data for the excess returns under the null of no predictability ($\beta_k = 0$ in Equation (11)):

$$er_{t,t+k} = \alpha_k + e_{t+k,t} \tag{12}$$

where α_k is the mean of the compound excess return and $e_{t+k,t}$ is drawn from a Gaussian distribution with mean zero and standard deviation σ_k^e. We generate data for the surplus consumption ratio, assuming that s_t is represented by a Gaussian $AR(1)$:

$$s_{t+1} = \overline{s} + \rho s_t + \upsilon_{t+1} \tag{13}$$

where υ_{t+1} is drawn from a Gaussian distribution with mean zero and standard deviation σ^υ. Let $\sigma^{e,\upsilon}$ denote the correlation between unexpected returns $e_{t+1,t}$ and the innovations of the explanatory variable υ_{t+1}. The parameters α_k, σ_k^e, $\overline{s}$, ρ, σ^υ and $\sigma^{e,\upsilon}$ are estimated from the annual data. We generate $100,000$ samples of the same size as the actual data[7] and each sample is used to estimate Equation (11). Such a procedure enables us to recover (i) the distribution of the estimates of the regressors β_k and the coefficients of determination R^2 under the null of no predictability and (ii) the distributions of the NEWEY–WEST t–statistics and the rescaled $t/\sqrt{T}$–statistics proposed by VALKANOV [2003]. For comparison, we also run the same Monte Carlo experiment when the explanatory variable is the (log) price–dividend ratio $p_t - d_t$ or the (log) consumption–wealth ratio $c_t - w_t$. Note that by construction, the estimated autoregressive coefficient $\hat{\rho} = 0.89$ is the same for both the price–dividend ratio and the surplus consumption ratio. However, the estimated contemporaneous correlation $\hat{\sigma}^{e,\upsilon}$ between the unexpected returns and

[7] We actually generate $T+200$ where T is the size of the actual data, the 200 first observations being discarded from the sample.

	Panel A		Panel B		Panel C		Panel D	
	$x_t = s_t$				$x_t = p_t - d_t$		$x_t = cay_t$	
	$\rho = 0.89$		$\rho = 0.999$		$\rho = 0.89$		$\rho = 0.57$	
	$\sigma^{e,v} = -0.06$		$\sigma^{e,v} = -0.06$		$\sigma^{e,v} = 0.61$		$\sigma^{e,v} = -0.52$	
k (year)	β_k	R^2	β_k	R^2	β_k	R^2	β_k	R^2
1	0.0016	0.01	0.02	0.01	-0.04	0.01	0.21	0.01
	(0.05)		(0.16)		(0.057)		(1.31)	
2	0.003	0.01	0.04	0.01	-0.08	0.02	0.38	0.01
	(0.09)		(0.332)		(0.16)		(2.28)	
3	0.04	0.02	0.06	0.02	-0.11	0.03	0.54	0.02
	(0.13)		(0.47)		(0.21)		(3.10)	
4	0.005	0.03	0.08	0.03	-0.15	0.05	0.67	0.02
	(0.17)		(0.61)		(0.25)		(3.80)	
5	0.006	0.03	0.10	0.04	-0.18	0.06	0.81	0.02
	(0.21)		(0.75)		(0.30)		(4.42)	
6	0.007	0.04	0.12	0.05	-0.22	0.07	0.94	0.02
	(0.25)		(0.89)		(0.34)		(4.97)	

Table 2: Predictability Bias – Annual Data

Note: This table reports the simulation results of long–horizon regressions for simulated compounded (log) excess returns $er_{t,t+k}$ on the simulated (log) surplus consumption ratio ($x_t = s_t$), (log) consumption–wealth ratio ($x_t = cay_t$) or (log) price–dividend ratio ($x_t = p_t - d_t$): $er_{t,t+k} = \sum_{i=1}^{k}(r_{t+i} - r_{f,t+i}) = \alpha_k + \beta_k x_t + u_{t+k,t}$. Simulated compounded (log) excess returns are generated under the null of no predictability: $er_{t,t+k} = \alpha_k + e_{t+k,t}$. Simulated dependent variables x_t are generated under the assumption of Gaussian $AR(1)$: $x_{t+1} = \overline{x} + \rho x_t + v_{t+1}$. The table reports the average values of the OLS estimates of the regressors β_k and coefficient of determination R^2 obtained from 100.000 simulations. Standard errors in parentheses.

innovations of the surplus consumption ratio on the one hand and the price–dividend ratio on the other are respectively -0.06 and 0.61. When the explanatory variable is the annual (log) consumption–wealth ratio, the estimated retrogressive coefficient $\hat{\rho}$ and the estimated contemporaneous correlation $\hat{\sigma}^{e,v}$ are 0.57 and -0.52 respectively. Moreover, as documented by STAMBAUGH [1999], the estimate of the autocorrelation of the price–dividend ratio is most likely biased downward. Therefore, we also run the same Monte Carlo experiment when the simulated surplus consumption ratio is generated by highly persistent Gaussian AR(1) (Equation (13)) by setting $\rho = 0.999$).

The simulations results are reported in Table 2. As we can see in Panels A, B et D of Table 2, the surplus consumption ratio and the consumption–wealth ratio present similar results. First, the average values of the estimated β_k coefficients are upward biased. However, the bias remains small and not statistically significant at any horizon. Moreover, the average value of R^2 is close to 0 at a 1–year horizon and remains low at long horizons. For instance, the average value of R^2 does not exceed 0.04 and 0.02 at a 6–year horizon, when s and *cay* are the explanatory variables respectively. However, the R^2 is larger at any horizon in the case of $p_t - d_t$. This indicates that s_t and cay_t appear to be more immune to bias that the conventional $p_t - d_t$.

Table 3 reports the results of univariate long–horizon regressions of excess returns using actual annual data s_t, $p_t - d_t$ and cay_t. For each regressor, Table 3 reports (i) the OLS estimates of the regressors, (ii) the NEWEY–WEST t–statistics associated to the null of the absence of predictability and the associated empirical size, (iii) the modified $t/\sqrt{T}$–statistics proposed by VALKANOV [2003] and the associated empirical size and (iv) the coefficient of determination R^2. The empirical sizes are obtained from the Monte Carlo simulations. When s_t is used as the regressor, the estimated coefficients $\hat{\beta}_k$ have the right negative sign. – i.e a higher surplus consumption ratio predicts lower excess returns. This is in line with the theoretical implications of the CAMPBELL and COCHRANE [1999] model. Moreover, the R^2 increases with horizon and exceeds the average values obtained from the Monte Carlo experiment. For instance, at a 5–year horizon, the coefficient of determination R^2 and the average value obtained under the null of no predictability (see Table 2, Panel A) are respectively 0.31 and 0.03. Moreover, the empirical sizes corresponding to the NEWEY–WEST t–statistics and the rescaled $t/\sqrt{T}$–statistics proposed by VALKANOV [2003] (obtained from the Monte Carlo experiments when $\rho = 0.89$ and 0.999) have similar conclusions: the surplus consumption ratio is statistically significant (at usual levels) at any horizon. The second part of Table 3 presents the results of univariate long–horizon regressions of excess returns on the (log) consumption–wealth ratio evaluated at several lags. When cay_t is used as the regressor, the estimated coefficients $\hat{\beta}_k$ have a positive sign as in LETTAU and LUDVIGSON [2001a]. Furthermore, the coefficients of determination R^2 increase with horizon. For instance, the (log) consumption–wealth ratio explains about 37% of the variations of excess stock returns at a 5–year horizon. Based on NEWEY–WEST and VALKANOV [2003] t–statistics, the estimated coefficients slopes are statistically significant. In contrast to the surplus consumption ratio and the consumption–wealth ratio, the price–dividend ratio is never statistically significant. This finding is in line with the those of MANKIW and SHAPIRO [1986] and STAMBAUGH [1999]. Indeed, when both contemporaneous correlation $\sigma^{e,v}$ and autoregressive parameter ρ are high, the results based on the standard distributions of the test statistics may lead us to reject the absence of predictability too often.

$er_{t,t+k} = \sum_{i=1}^{k}(r_{t+i} - r_{f,t+i}) = \alpha_k + \beta_k x_t + \varepsilon_{t,t+k}$						
k (year)	1	2	3	4	5	6
(log) surplus consumption ratio $x_t = s_t$						
β_k	-0.13	-0.22	-0.31	-0.43	-0.58	-0.67
t_{NW}	-3.49	-3.62	-3.86	-3.84	-4.03	-4.00
	(0.03)	(0.03)	(0.03)	(0.03)	(0.025)	(0.03)
	{0.03}	{0.03}	{0.03}	{0.03}	{0.025}	{0.03}
$t/\sqrt{T}$	-0.47	-0.49	-0.52	-0.52	-0.54	-0.54
	(0.009)	(0.018)	(0.006)	(0.003)	(0.002)	(0.01)
	{0.002}	{0.004}	{0.006}	{0.01}	{0.018}	{0.03}
R^2	0.07	0.10	0.19	0.31	0.35	0.32
(log) consumption–wealth ratio $x_t = cay_t$						
β_k	5.87	10.50	11.93	12.54	16.30	21.65
t_{NW}	3.74	4.61	7.64	7.03	6.46	7.91
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
$t/\sqrt{T}$	0.50	0.62	1.03	0.95	0.87	1.07
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
R^2	0.26	0.44	0.40	0.33	0.37	0.51
(log) price–dividend ratio $x_t = p_t - d_t$						
β_k	-0.14	-0.24	-0.27	-0.34	-0.53	-0.75
t_{NW}	-2.13	-1.72	-1.35	-1.17	-1.32	-1.58
	(0.4)	(0.5)	(0.7)	(0.8)	(0.8)	(0.8)
$t/\sqrt{T}$	-0.29	-0.23	-0.18	-0.16	-0.18	-0.21
	(0.15)	(0.35)	(0.7)	(0.8)	(0.8)	(0.8)
R^2	0.09	0.12	0.11	0.12	0.18	0.25

Table 3: UNIVARIATE LONG-HORIZON REGRESSIONS - EXCESS STOCK RETURNS

Note: This table reports the results of long–horizon regressions for the compounded (log) excess returns $er_{t,t+k}$ on (i) the (log) surplus consumption ratio ($x_t = s_t$), (ii) the (log) consumption–wealth ratio ($x_t = cay_t$) and (iii) the (log) dividend–price ratio ($x_t = p_t - d_t$): $er_{t,t+k} = \sum_{i=1}^{k}(r_{t+i} - r_{f,t+i}) = \alpha_k + \beta_k x_t + \varepsilon_{t,t+k}$. For each regression, the table reports the OLS estimates of the regressors, the NEWEY–WEST t–statistics associated with the null of the absence of predictability t_{NW}, the modified $t/\sqrt{T}$–statistics proposed by VALKANOV [2003] and the coefficient of determination R^2. Empirial sizes were obtained from the 100.000 Monte Carlo simulations. Empirical size obtained from the Monte Carlo experiment when $\rho = 0.91$ in parentheses and empirical size from Monte Carlo experiment when $\rho = 0.999$ in curly brackets. The standard size of the NEWEY-WEST t-statistics in brackets. The sample is annual and spans the period 1948 to 2001.

	Asset Holding Growth $a_{t+k} - a_t = \alpha_k + \gamma_k x_t + \varepsilon_{t,t+k}$						Consumption Growth $c_{t+k} - c_t = \alpha_k + \gamma_k x_t + \varepsilon_{t,t+k}$					
k (year)	1	2	3	4	5	6	1	2	3	4	5	6
(log) surplus consumption ratio: $x_t = s_t$												
γ_k	-0.02	-0.05	-0.07	-0.11	-0.15	-0.17	0.00	-0.00	-0.00	-0.01	-0.02	-0.03
t_{NW}	**-2.24****	**-2.61****	**-3.24*****	**-3.37*****	**-4.81*****	**-4.44*****	0.13	-0.60	-1.01	-1.33	**-1.68***	**-2.08****
$t/\sqrt{T}$	-0.31	-0.35	-0.44	**-0.59****	**-0.65****	**-0.60****	0.01	-0.08	-0.13	-0.18	-0.22	-0.28
R^2	0.05	0.09	0.19	0.35	0.47	0.52	0.00	0.00	0.02	0.05	0.07	0.11
(log) consumption–wealth ratio $x_t = cay_t$												
γ_k	0.95	2.16	2.38	2.35	2.76	3.30	-0.20	-0.17	-0.25	-0.40	-0.53	-0.51
t_{NW}	**3.21*****	**3.65*****	**3.39*****	**2.58****	**2.32****	**3.16*****	**-2.94*****	-1.16	-1.15	-1.67	-2.01	-2.33
$t/\sqrt{T}$	0.43	0.49	0.46	0.35	0.31	0.43	-0.40	-0.15	-0.15	-0.22	-0.27	-0.31
R^2	0.13	0.29	0.24	0.17	0.20	0.28	0.07	0.02	0.02	0.03	0.04	0.03
(log) price–dividend ratio $x_t = p_t - d_t$												
γ_k	0.01	0.00	-0.01	-0.04	-0.03	-0.02	-0.00	-0.00	-0.01	-0.02	-0.02	-0.02
t_{NW}	1.01	0.36	-0.59	-0.86	-0.56	-0.27	-1.54	-1.63	-1.43	-1.23	-1.06	-0.96
$t/\sqrt{T}$	0.09	0.04	-0.09	-0.16	-0.13	-0.07	-0.21	-0.22	-0.19	-0.16	-0.14	-0.13
R^2	0.01	0.00	0.00	0.02	0.01	0.00	0.02	0.01	0.02	0.03	0.03	0.03

Table 4: UNIVARIATE LONG-HORIZON REGRESSIONS: ASSET HOLDING GROWTH AND CONSUMPTION GROWTH

Note: This table reports the results of long–horizon regressions for consumption growth $(c_{t+k} - c_t)$ and asset holding growth $(a_{t+k} - a_t)$. Dependent variables are the (log) surplus consumption ratio, (log) price–dividend ratio and (log) consumption–wealth ratio. For each regression, the table reports the OLS estimates of the regressors, the NEWEY–WEST t–statistics t_{NW} associated to the null of the absence of predictability, the modified $t/\sqrt{T}$–statitsics proposed by VALKANOV [2003] and the coefficients of determination R^2. Significance at the level 10%, 5% and 1% for (i) theNEWEY–WEST t test using the standard t–test and for (ii) the modified $t/\sqrt{T}$ test using VALKANOV's [2003] critical values (Table 4, pp. 215, case 1, $c = -1$ and $\delta = 0$) is indicated by *, ** and ***. The sample is annual and spans the period 1948 to 2001.

To further investigate the predictive power of s, cay and $p-d$, we run regressions of the asset holdings growth and the consumption growth on each of the macroeconomic indicators at long horizons. The economic intuition for this additional test can be described as follows. Investors who want to maintain a flat consumption path over time will be more willing to adjust their asset holdings as a response to time–variation in expected returns. When excess returns are expected to be lower (higher) in the future, these investors will react by decreasing (increasing) current consumption and saving less (more), implying a decrease (an increase) in future asset holdings growth. Accordingly, if the surplus consumption ratio (or alternative indicators) can predict excess returns, it should forecast asset holdings growths. As expected consumption growth is not so volatile, the surplus consumption ratio should fail to predict future consumption growth. We use standard size to evaluate the NEWEY–WEST t–statistics and VALKANOV's [2003] critical values to evaluate the modified $t/\sqrt{T}$–statistics [8]. The first part of Table 4 reports OLS results on regressions for the asset holding growth evaluated at several lags. When the explanatory variable is the surplus consumption ratio, the estimated coefficients slopes β_k are statistically significant and have the right sign according to the economic intuition described above. An increase in the surplus consumption ratio implies a decrease in both expected future excess returns (see Table 3) and expected future asset holdings growths (see Table 4). Moreover, the estimated coefficient slopes β_k increase (in absolute value) with horizon. The coefficient of determination R^2 increases with horizon to reach 51% at the 5–year horizon. When the explanatory variable is the consumption–wealth ratio, we reach the same conclusions. The estimated coefficient slopes are statistically significant at any horizon and have the right sign, i.e. higher consumption–wealth ratio predicts higher future asset holdings growth. Moreover, the statistic R^2 increases with horizon. For instance, the consumption–wealth ratio explains about 25% of the variation of excess stock returns at a 3–year horizon. In contrast to s_t and cay_t, the price–dividend ratio $p_t - d_t$ is never statistically significant and the corresponding R^2 is almost close to 0. Note that all s_t, cay_t and $d_t - p_t$ fail to predict consumption growth at any horizon.

The main conclusion to be retained from Tables 3 and 4 is that the surplus consumption ratio and the consumption–wealth ratio are strong predictors of excess stock returns at annual frequency. In contrast with LI [2005], the price–dividend ratio fails to forecast excess returns at long horizons.

The predictive power of s_t is now compared to cay_t. Table 5 reports the results of multivariate regressions of long–horizon excess returns using s_t and cay_t. Consistent with previous results, s_t remains statistically significant at long horizons when we add cay_t as a dependent variable, and the sign of the regression coefficients corresponding to s_t is unchanged. Moreover, the introduction of s_t increases R^2 especially at long horizons. For instance, R^2 increases from 33% when we consider only cay_t as a predictive variable (see Table 3) to 44% when we add s_t at a 4–year horizon. The results reported in Table 5 suggest that there is a component of *long–horizon* expected returns — captured by the surplus consumption ratio — that moves independently of cay_t.

Note that for comparison purposes, we studied the predictive power of the surplus con-

[8] VALKANOV [2003] provides the critical values of the modified $t/\sqrt{T}$ test in Table 4, pp. 215.

k–period Regression: Excess Returns						
$er_{t,t+k} = \sum_{i=1}^{k}(r_{t+i} - r_{f,t+i}) = \alpha_k + \beta_k x_t + \varepsilon_{t,t+k}$						
	k (year)					
x_t	1	2	3	4	5	6
cay_t	5.54	9.93	10.55	9.51	11.72	17.43
	(4.23)	(5.36)	(9.17)	(4.55)	(5.73)	(9.12)
s_t	-0.09	-0.15	-0.21	-0.32	-0.40	-0.37
	(-2.85)	(-3.26)	(-6.38)	(-3.31)	(-2.98)	(-3.18)
$\overline{R}^2$	[0.27]	[0.48]	[0.47]	[0.45]	[0.49]	[0.57]

Table 5: Multivariate Long-horizon Regressions - Excess Stock Returns

Note: This table reports the results of long–horizon regressions for the compounded (log) excess returns $er_{t,t+k}$ on the variables listed in the first column. cay_t is the proxy for the consumption–wealth ratio proposed by LETTAU and LUDVIGSON [2001 a, b and 2005]. s_t is the (log) surplus consumption ratio. For each regression, the table reports the OLS estimates of the regressors, theNEWEY-WEST t–statistics associated to the null of the absence of predictability t_{NW} (in parentheses) and the adjusted $\overline{R}^2$ statistics (in brackets). The standard size of the t–test is used to evaluate the NEWEY–WEST t–statistics. The annual sample spans 1948 to 2001.

Table 6: Out-of-Sample Regressions: Excess Returns

First forecast period	1958		1968		1978	
horizon k (year)	β_k	R^2	β_k	R^2	β_k	R^2
1	-0.09	0.07	-0.13	0.12	-0.06	0.03
2	-0.16	0.13	-0.20	0.16	-0.12	0.08
3	-0.23	0.24	-0.25	0.26	-0.18	0.18
4	-0.32	0.39	-0.34	0.39	-0.20	0.18
5	-0.43	0.44	-0.45	0.45	-0.23	0.21
6	-0.49	0.44	-0.50	0.44	-0.25	0.27

Note: This table reports the results of long–horizon regressions for the compounded (log) excess returns $er_{t,t+k}$ on the (log) surplus consumption ratio (s_t) for several lags k (year): $er_{t,t+k} = \sum_{i=1}^{k}(r_{t+i} - r_{f,t+i}) = \alpha_k + \beta_k s_t + \varepsilon_{t,t+k}$. For each regression, the table reports the OLS estimates β_k and the coefficient of determination R^2. The first forecast period presents the first period of the out–of sample regressions.

sumption ratio, the consumption–wealth ratio and the price–dividend ratio at quarterly frequency[9]. The main conclusion to be retained from our empirical results is that, when the appropriate testing procedures are used, the evidence of the predictive power of the consumption–wealth ratio and the surplus consumption ratio at quarterly frequency is not as strong as the predictive power of those indicators at annual frequency. This is due to the fact that when data are sampled at quarterly frequency, they are more prone to (i) high persistence of the explanatory variables[10] and (ii) measurement errors that arise from seasonality and other measurement problems.

The rest of this section presents some additional empirical results to evaluate the robustness of the predictive power of the surplus consumption ratio to various issues.

Robustness

As documented by LETTAU and LUDVIGSON [2001a], a look–ahead bias may arise from the fact that the coefficients ϕ, g and σ used to generate the (log) surplus consumption ratio are estimated from the whole sample. To address this issue, Table 6 reports results for out–of–sample predictions. The results are consistent with previous experiments, regardless of the starting date of the out–of–sample regressions. The estimated coefficients $\hat{\beta}_k$ are negative and increase with horizon. The coefficient of determination R^2 starts low then increases substantially at 5 and 6–year horizons. This result confirms that the surplus consumption ratio is a good predictor of long–horizon excess returns.

Moreover, to check the robustness of the empirical results presented above, we evaluate the sensitivity of the predictive power of the surplus consumption ratio to the degree of

[9] Empirical results obtained at quarterly frequency are provided on request.

[10] At quarterly frequency, the first–order autocorrelations of s and cay are about 0.93 and 0.87 and their second–order autocorrelations are about 0.85 and 0.79, respectively.

Horizon	$\theta = 0.5$		$\theta = 1.5$		$\theta = 5$	
year	β	R^2	β	R^2	β	R^2
1	-0.06	0.05	-0.12	0.07	-0.19	0.07
	(-3.14)		(-3.32)		(-3.74)	
2	-0.11	0.08	-0.19	0.10	-0.33	0.12
	(-2.94)		(-3.38)		(-4.04)	
3	-0.14	0.13	-0.27	0.18	-0.48	0.23
	(-3.08)		(-4.57)		(-3.55)	
4	-0.19	0.19	-0.37	0.29	-0.67	0.36
	(-2.86)		(-3.51)		(-4.76)	
5	-0.27	0.22	-0.51	0.33	-0.88	40
	(-2.67)		(-3.62)		(-5.06)	
6	-0.32	0.20	-0.59	0.30	-1.01	0.37
	(-2.41)		(-3.58)		(-4.98)	

Table 7: Sensitivity Test

Note: This table reports the results of long–horizon regressions for the compounded (log) excess returns $er_{t,t+k}$ on the (log) surplus consumption ratio ($x_t = s_t$) for several lags k (year): $er_{t,t+k} = \sum_{i=1}^{k}(r_{t+i} - r_{f,t+i}) = \alpha_k + \beta_k x_t + \varepsilon_{t,t+k}$.
For each regression, the table reports the OLS estimates β, theNEWEY–WEST t–statistics associated to the null of the absence of predictability t_{NW} (in parentheses) and the coefficient of determination R^2. Long–horizon regressions are run for the different values of the curvature of the consumer's utility $\theta = 0.5$, 1.5 and 5.

curvature of the utility function θ. Indeed, the time–series for the surplus consumption ratio is generated using Equation (2) and therefore depends on the value of the curvature of the utility function θ. Therefore, we gauge the ability of the model to replicate the long–horizon predictability of the surplus consumption ratio on excess returns for different values of θ. This experiment is reported in Table 7 for values of $\theta = 0.5$, 1.5 and 5. As shown in Table 7, we recover the same pattern whatever the value of θ. Indeed, the negative relationship between excess returns and the surplus consumption ratio remains unchanged. Moreover, s_t is statistically significant at any horizon. In addition, predictability is an increasing function of horizon. The longer the prediction horizon, the higher the measure of fit R^2.

Finally, we study the robustness of our empirical results by using alternative initial values of the surplus consumption ratio. Indeed, the time–series for s_t used in the previous empirical studies is generated using the specification (2) by imposing $\overline{s} = \log \overline{S} = \log\left(\sigma\sqrt{\frac{\theta}{1-\phi}}\right)$ as an initial value. The parameters σ, θ and ϕ are set to 1.14%, 2 and 0.89 respectively, implying $\overline{s} = -3.02$. Note that the maximum and the minimum of the benchmark time–series for s_t are respectively 0.08 and -3.84. Therefore, we test different initial values ranking between -5 and 5. As conclusions remain unchanged whatever the chosen initial value, Table 8 only reports results relative to the initial values $\overline{s}$ and the extreme values

-5 and 5. Moreover, different starting dates for forecast periods (1948, 1958 and 1968) are tested. When the starting point is 1948, the results of the univariate regressions depend on the choice of the initial value. For instance, at a 4–year horizon, the coefficient of determination R^2 shifts from 31% to only 12% when the initial value is set to 5 rather than $\overline{s}$. Focusing on 1958 and 1968 as starting dates and comparing the results obtained with various initial values of the time–series for the surplus consumption ratio imposed at date 1948, it can be noticed that the R^2 statistics remain high. Additionally, there is a very small change in both estimated slope coefficients and their standard errors (or equivalently the corresponding t–statistics) at all horizons, suggesting that our empirical results are robust to the initial value of the surplus consumption ratio.

III.2. Cross–section of Expected Stock Returns

This section provides the main results of our paper. We explore the ability of the surplus consumption ratio to explain the cross–sectional variations in expected returns. More precisely, we estimate the linear three–factor model when the risk factors are consumption growth, the lagged surplus consumption ratio, and their product. We compare the performance of CAMPBELL and COCHRANE [1999]'s (C)CAPM model with habit formation to alternative models: (*i*) the well–documented FAMA–FRENCH three–factor model, (*ii*) the unconditional version of the Capital Asset Pricing Model CAPM, (*iii*) the unconditional version of the Consumption–based Asset Pricing Model (C)CAPM and (*iv*) the conditional (C)CAPM proposed by LETTAU and LUDVIGSON [2001b]. As a benchmark, the surplus consumption ratio is generated using specification (2) proposed by CAMPBELL and COCHRANE [1999]. Then we evaluate the sensitivity of our empirical results to (*i*) the degree of curvature of the utility function[11] θ, (*ii*) the initial value of the time–series for the surplus consumption ratio generated using Equation (2) and (*iii*) alternative specifications of the level of habit stock.

The financial data used in this cross-section study are borrowed from the web site of Kenneth FRENCH [12]. We use data on (*i*) the value weighted returns of 25 Portfolios on the NYSE, AMEX and NASDAQ sorted by size and book-to-market value, (*ii*) the value weighted returns R_{vw} on the NYSE, AMEX and NASDAQ, (*iii*) the three–month treasury bill as proxy for the risk–free rate and (*iv*) the two excess returns capturing the value and the size premia, denoted respectively SMB and HML. We convert the nominal returns to real returns using the consumer price index (CPI) borrowed from $NIPA$. Then we convert the monthly real returns to quarterly real data spanning the first quarter of 1952 to the first quarter of 2005, that is, 212 observations for each of the 25 portfolios. Table 9 reports the well-established empirical fact that expected returns vary across stocks. More precisely, it summarizes the size and book–to–market effects. Stocks with low prices relative to their book values (*Book–to–market value*) or stocks with high market values (*size*) provide higher average returns. The challenge of the asset pricing models is to develop credible models that can account for the cross–sectional variations

[11] As mentioned in the beginning of section 2, parameter θ is set to 2.

[12] We refer the reader to the FAMA and FRENCH articles [1992, 1993 and 1996] for more details.

Table 8: Sensitivity Test: Alternative initial values of the time–series s_t

h	Starting date 1948 β_k	R^2	1958 β_k	R^2	1968 β_k	R^2
$s(t=1948)=\overline{s}=-3.02$						
1	-0.13	0.07	-0.10	0.06	-0.13	0.09
	(-3.49)		(-3.43)		(-2.97)	
2	-0.22	0.10	-0.22	0.18	-0.26	0.21
	(-3.62)		(-4.08)		(-3.33)	
4	-0.43	0.31	-0.42	0.18	-0.45	0.47
	(-3.84)		(-5.34)		(-4.33)	
6	-0.67	0.32	-0.65	0.49	-0.70	0.49
	(-4.00)		(-5.00)		(-3.74)	
$s(t=1948)=5$						
1	0.01	0.04	-0.02	0.01	-0.11	0.12
	(2.13)		(-1.18)		(-4.83)	
2	0.03	0.07	-0.05	0.05	-0.21	0.22
	(1.90)		(-1.43)		(-5.58)	
4	0.05	0.12	-0.11	0.17	-0.41	0.59
	(1.72)		(-1.60)		(-8.11)	
6	0.08	0.19	-0.15	0.16	-0.63	0.66
	(1.87)		(-1.50)		(-6.48)	
$s(t=1948)=-5$						
1	-0.07	0.12	-0.06	0.04	-0.13	0.09
	(-7.66)		(-2.19		(-2.97	
2	-0.14	0.25	-0.13	0.11	-0.26	0.21
	(-7.55)		(-2.26)		(-3.33)	
4	-0.24	0.49	-0.23	0.26	-0.45	0.47
	(-7.14)		(-2.44)		(-4.34)	
6	-0.34	0.52	-0.31	0.26	-0.70	0.49
	(-7.23)		(-2.58)		(-3.74)	

Note: This table reports the results of long–horizon regressions for the compounded (log) excess returns $er_{t,t+k}$ on the (log) surplus consumption ratio (s_t) for several lags k (year): $er_{t,t+k} = \sum_{i=1}^{k}(r_{t+i} - r_{f,t+i}) = \alpha_k + \beta_k x_t + \varepsilon_{t,t+k}$.
For each regression, the table reports the OLS estimates β, the NEWEY–WEST t–statistics associated to the null of the absence of predictability t_{NW} (in parentheses) and the coefficient of determination R^2. Long–horizon regressions are run for different initial values of the surplus consumption ratio at date $t = 1948$.

	Book–to–Market				
Size	Low	2	3	4	High
Small	1.13	2.66	2.77	3.44	3.80
2	1.50	2.34	2.93	3.11	3.45
3	1.79	2.47	2.46	2.90	3.18
4	1.92	1.94	2.63	2.72	2.86
Big	1.64	1.75	2.00	1.95	2.16

Table 9: Average excess returns (in % on 25 FAMA–FRENCH Portfolios

Note: This table reports the quarterly mean excess returns (in %) on 25 FAMA-FRENCH portfolios sorted by size and book–to–market characteristics. "Small size" refers to the portfolios with the smallest firm, while "Big size" includes the largest firms. Similarly, "low book–to–market" includes firms with the loWEST book-to-market ratio and "high book–to–market" the highest. Data are quarterly and spans the first quarter of 1952 to the first quarter of 2005.

in average returns on portfolios sorted by size and book–to–market value.

The macroeconomic data are borrowed from the web site of Martin LETTAU[13]. We use quarterly data on (i) the real per capita consumption data for nondurables and services, excluding shoes and clothing [14] and (ii) the *cay* as a proxy for the unobservable consumption to aggregate wealth ratio. Data span the first quarter of 1952 to the first quarter of 2005.

We use the beta representation of each model as the basis of the empirical work:

$$\mathbb{E}\left[R^e_{i,t}\right] = \mathbb{E}\left[R_{i,t} - R_{f,t}\right] = \lambda_0 + \beta_i' \lambda \tag{14}$$

$$R^e_{i,t} = \beta_{i,0} + F_t \beta_i + u_{i,t} \tag{15}$$

where $R^e_{i,t}$ denote excess returns on the 25 Fama–FRENCH portfolios over the risk–free rate $R_{f,t}$, λ is the $K \times 1$ vector of the market price of risk corresponding to the vector of K risk factors F_t.

The linear beta representation is estimated by the 2 pass FAMA–MACBECH regressions. As mentioned by LETTAU and LUDVIGSON [2001b] and JAGANNATHAN et al. [2005], among others, the FAMA–MACBECH procedure is well–adapted to a moderate number of quarterly time–series observations and a reasonably large number of asset returns. As the model is evaluated using excess stock returns $(R^e_{i,t})$, a well–specified asset pricing model produces intercept λ_0 that is indistinguishable from zero. For each portfolio i, the pricing error is given by:

$$\hat{\mathbb{E}}\left[R^e_{i,t}\right] - \mathbb{E}_T\left[R^e_{i,t}\right]$$

[13]We refer the reader to the LETTAU and LUDVIGSON articles [2001a, 2001b and 2005] for more details.

[14]The same results are obtained when we use quarterly real per capita consumption data for nondurables and services borrowed from NIPA.

where $\hat{\mathbb{E}}\left[R^e_{i,t}\right]$ is the average excess return on portfolio i fitted by the estimated model, and $\mathbb{E}_T\left[R^e_{i,t}\right]$ is the empirical average[15] excess return.

Table 10 reports the estimated coefficients, their uncorrected and SHANKEN–corrected t-statistics, the R^2 and the adjusted $\overline{R}^2$ for the cross–sectional regressions. Table 11 presents the square root of the average squared pricing errors across 10 aggregated portfolios formed on the basis of the size and book–to–market quintiles.

Moreover, the empirical performance of each model is evaluated using the asymptotic χ^2(Wald) test of the null hypothesis that all the pricing errors are jointly zero. The χ^2–statistic is reported in Table 11.

We first examine the unconditional Capital Asset Pricing Model CAPM. The single factor F_t in the unconditional CAPM is the market portfolio $R_{m,t}$ as a proxy for the total wealth return. It is commonly well–assumed that the value weighted return R_{vw} on the NYSE, AMEX and NASDAQ is a good proxy for the market portfolio return R_m. The cross-sectional implication of the model is given by:

$$\mathbb{E}\left[R^e_{i,t}\right] = \lambda_0 + \beta_{i,R_m}\lambda_{R_m} \tag{16}$$

As shown in the first row of Table 10, the unconditional CAPM fails to explain the cross–sectional variation in expected excess returns. The coefficient of determination R^2 of the regression is only 6% and the adjusted $\overline{R}^2$ is about 2%. Moreover, the estimated coefficient $\widehat{\lambda}_{R_w}$ is negative and therefore has the wrong sign according to SHARPE [1964]. Additionally, the estimated market risk price is not significantly different to zero. Table 11 shows that the model is statistically rejected according to the test of the null hypothesis that all of the pricing errors are jointly zero.

[15]The empirical average is defined as $\mathbb{E}_T = \frac{1}{T}\sum_{t=1}^{T}$.

					factors					R^2
Row	cst	R^e_m	SMB	HML	s_{-1}	cay_{-1}	Δc	$s_{-1}\Delta c$	$cay_{-1}\Delta c$	$(\overline{R}^2)$
CAPM	**3.46**	-0.91								0.06
	(3.79***)	(-0.85)								0.02
	(3.77***)	(-0.85)								
FFF	**3.63**	-1.83	0.47	**1.27**						0.76
	(3.23***)	(-1.45)	(1.26)	(3.29***)						0.74
	(3.05***)	(-1.38)	(1.26)	(3.28***)						
(C)CAPM	**1.53**						0.30			0.14
	(2.73***)						(1.56)			0.10
	(2.28***)						(1.31)			
LL(2001)	**3.86**					**-1.12**	0.21		-0.12	0.62
	(5.22***)					(-2.78***)	(1.00)		(-0.40)	0.56
	(3.45***)					(-1.86*)	(0.66)		(-0.26)	
CC(1999)	1.44				**-1.17**		-0.09	-0.15		0.49
	(2.13**)				(-4.18***)		(-0.70)	(-0.81)		0.42
	(1.00)				(-1.99**)		(-0.33)	(-0.83)		

Table 10: CROSS–SECTIONAL REGRESSIONS: FAMA-MACBETH REGRESSIONS USING 25 FAMA–FRENCH PORTFOLIOS

Note: This table reports the resutls of the 2–pass FAMA–MACBETH regressions for the avreage excess returns across the FAMA–FRENCH (25) portfolios sorted by size and market to book characteristics on the factors F_t:
Model CAPM: $F_t =[\ cst\ \Delta R^e_{m,t}]$
Model FFF: $F_t =[\ cst\ R^e_{m,t}\ SMB\ HML]$
Model (C)CAPM: $F_t =[\ cst\ \Delta c_t]$
Model LL(2001): $F_t =[\ cst\ cay_{t-1}\ \Delta c_t\ cay_{t-1}\ \Delta c_t]$
Model CC(1999): $F_t =[\ cst\ s_{t-1}\ \Delta c_t\ s_{t-1}\Delta c_t]$
where R^e_m is the excess return on the market portfolio, s_{-1} is the lagged (log) surplus consumption ratio, Δc the consumption growth and cay_{-1} the lagged proxy for the consumption–wealth ratio proposed by LETTAU and LUDVIGSON [2001b]. For each regression, the table reports the cross sectional coefficients. For each coefficient, two t-statistics are reported in parentheses. The top statistic uses the uncorrected FAMA–MACBETH standard errors. The bottom statistic uses the JAGANNATHAN and WANG [1998] correction. Significance at the 10%, 5% and 1% levels using the standard t–test are indicated by *, ** and ***. Data are quarterly and span the first quarter of 1952 to the first quarter of 2005. Δc and cay are expressed in 100 basis points.

The second row of Table 10 presents results relative to FAMA–FRENCH Model [1992 and 1993]. The three factors of the FAMA–FRENCH model are the market portfolio R_m and the two excess returns capturing the value and the size premia SMB and HML, implying the following cross-sectional model:

$$\mathbb{E}\left[R^e_{i,t}\right] = \lambda_0 + \beta_{i,R_m}\lambda_{R_m} + \beta_{i,HML}\lambda_{HML} + \beta_{i,SMB}\lambda_{SMB} \tag{17}$$

Table 10, row 2, presents the well–established results for the three factors of FAMA and FRENCH (1992, 1993]. The model explains 76% of the cross-sectional variability of expected returns. In addition, the t-statistic on the HML factor is highly statistically significant even after correction for sampling errors. However, contrary to economic theory, the intercept comes out statistically significant with corrected t–statistic equal to 3.05.

We now turn to the Consumption–based Asset Pricing Models. We first examine the unconditional (C)CAPM model given by:

$$\mathbb{E}\left[R^e_{i,t}\right] = \lambda_0 + \beta_{i,\Delta c}\lambda_{\Delta c}$$

The single risk factor is consumption growth. As can be seen in row 3 of Table 10, the unconditional (C)CAPM model has little power to explain the cross–section of expected returns. Indeed, the market price of consumption growth risk $\widehat{\lambda}_{\Delta c}$ is not statistically significant, indicating that the beta $\beta_{\Delta c}$ is unable to account for the cross–sectional variation in average excess returns. The failure of the standard (C)CAPM is also summarized by (i) a very low cross–sectional $\overline{R}^2$ which does not exceed 10% and (ii) a large χ^2–statistic rejecting the model at usual significance levels.

Table 10, row 4, reports cross–sectional regression results for the conditional (C)CAPM model proposed by LETTAU and LUDVIGSON [2001b]. The factors are the consumption growth, the lagged (log) consumption–wealth ratio and their product, implying the following linear model:

$$\mathbb{E}\left[R^e_{i,t}\right] = \lambda_0 + \beta_{i,cay_{-1}}\lambda_{cay_{-1}} + \beta_{\Delta c}\lambda_{\Delta c} + \beta_{cay_{-1}\Delta c}\lambda_{cay_{-1}\Delta c}$$

As can be seen in Table 10, the conditional (C)CAPM model performs far better than the unconditional version. As documented by LETTAU and LUDVIGSON [2001b], the scaling variable cay_{t-1} is statistically significant, indicating that the $\beta_{cay_{-1}}$ accounts for a part of the cross–sectional variation in average returns. Moreover, the conditional (C)CAPM model explains about 56% of the cross–sectional variations in returns. Furthermore, Table 11 shows that the model is not statistically rejected, with a low χ^2–statistic. However, the estimated intercept $\hat{\lambda}_0$ remains statistically significant as in the unconditional (C)CAPM model, which is contrary to the economic intuition.

The last panel of Table 10 reports the empirical results for the implied CAMPBELL and COCHRANE [1999] model given by:

$$\mathbb{E}\left[R^e_{i,t}\right] = \lambda_0 + \beta_{i,s_{-1}}\lambda_{s_{-1}} + \beta_{i,\Delta c}\lambda_{\Delta c} + \beta_{i,s_{-1}\Delta c}\lambda_{s_{-1}\Delta c}$$

	CAPM	FFF	(C)CAPM	LL(2001)	CC(1999)
$S1$	0.90	0.48	0.98	0.60	0.65
$S2$	0.62	0.20	0.64	0.29	0.43
$S3$	0.37	0.18	0.47	0.33	0.37
$S4$	0.34	0.30	0.34	0.27	0.26
$S5$	0.76	0.35	0.37	0.43	0.53
$BTM1$	0.90	0.48	0.98	0.60	0.65
$BTM2$	0.75	0.23	0.71	0.35	0.50
$BTM3$	0.55	0.23	0.54	0.36	0.35
$BTM4$	0.44	0.31	0.41	0.35	0.25
$BTM5$	0.49	0.22	0.42	0.11	0.34
$Avr.$(%)	0.63	0.32	0.61	0.40	0.47
χ^2	47.71	**8.30*****	80.0	**15.20*****	**23.05*****

Table 11: Pricing Errors (%) of Aggregated Portfolios

Note: This Table reports the error pricings from the FAMA–MacBeth regressions presented in Table 10. The models are described in table 10. For each model, we report the square root of the average squared pricing errors for aggregated portfolios. $S1$ refers to the portfolios with the smallest firm, while $S5$ includes the largest firms. Similarly, $BTM1$ includes firms with the lowest book-to-market ratio and $BTM5$ the highest. We also report the average squared pricing errors across all portfolios ($Avr.$) and a χ^2–statistic for a test of the null hypothesis that all of the pricing errors are jointly zero. This statistic is computed using the SHANKEN [1992] correction. *, ** and *** indicate that the all of the pricing errors are statistically different from zero at the 10%, 5% and 1% levels .

The risk factors are consumption growth, the lagged (log) surplus consumption ratio and their product. The model performs far better than the unconditional standard (C)CAPM model, as it explains about 45% of the cross–sectional variations of expected returns. Moreover, the estimated risk price $\widehat{\lambda}_{s_{-1}}$ associated to the risk factor s_{t-1} has the right sign and is statistically significant. Additionally, the χ^2–statistic reported in Table 11 confirms the ability of the lagged surplus consumption ratio to account for the cross–sectional variation in average excess returns. Indeed, the corresponding χ^2–statistic is low and therefore the test of the null hypothesis that all the pricing errors are jointly zero does not reject the model at conventional significance levels. Finally, consistent with economic theory, the intercept comes out statistically insignificant. To summarize, Tables 10 and 11 show that both CAMPBELL and COCHRANE [1999]'s unconditional (C)CAPM model with external habit and LETTAU and LUDVIGSON [2001b]'s conditional (C)CAPM model perform far better than the unconditional standard (C)CAPM model in accounting for the cross-sectional variation in average stock returns. The (lagged) surplus consumption ratio s_{-1} and the (lagged) consumption–wealth ratio cay_{-1} are respectively the relevant risk factors that improve the empirical implications of the standard Consumption–based Asset Pricing Model. However, LETTAU and LUDVIGSON [2001b]'s conditional (C)CAPM model performs better than the CAMPBELL and COCHRANE [1999]'s model in explaining the cross–sectional variations in average excess returns on the 25 FAMA–FRENCH portfolios sorted by size and book–to–market value. The two first rows of table 12 illustrate these conclusions.

Table 12, row 1, reports results for the linear beta asset pricing model where the single risk factor is the (lagged) surplus consumption ratio:

$$\mathbb{E}\left[R^e_{i,t}\right] = \lambda_0 + \beta_{i,s_{-1}}\lambda_{s_{-1}}$$

This one–factor model explains about 28% of the cross–sectional variation in average excess returns. According to the low χ^2–statistic, the model is not statistically rejected at conventional significance level (10%). Consistent with economic theory, the estimated risk price has the right sign and is statistically significant. However, the intercept is not statistically significant.

The conclusions are similar when we consider the following linear beta model:

$$\mathbb{E}\left[R^e_{i,t}\right] = \lambda_0 + \beta_{i,cay_{-1}}\lambda_{cay_{-1}}$$

where the single risk factor is the lagged (log) consumption–wealth ratio. As documented by LETTAU and LUDVIGSON [2001b], the ability of the lagged cay to account for the cross–sectional variation in average excess returns is summarized by (i) a high adjusted coefficient of determination ($\overline{R}^2 = 52\%$), (ii) a low χ^2 statistic, implying the non–rejection of the model and (iii) a statistically significant estimated price of risk $\hat{\lambda}_{cay_{-1}}$.

Table 12, row 3, presents empirical results for the linear beta model when both the lagged surplus consumption ratio and the lagged consumption–wealth ratio are considered as risk factors, simultaneously. The cross–sectional implication of this model is given by:

$$\mathbb{E}\left[R^e_{i,t}\right] = \lambda_0 + \beta_{i,s_{-1}}\lambda_{s_{-1}} + \beta_{i,cay_{-1}}\lambda_{cay_{-1}}$$

Factors F_t						R^2	$Avr.(\%)$
cst	s_{-1}	cay_{-1}	Δc	$s_{-1}\Delta c$	$cay_{-1}\Delta c$	$(\overline{R}^2)$	χ^2
0.90	**-0.96**					0.31	0.54
(1.53)	(-2.37***)					0.28	**39.17***
(0.98)	(-1.92**)						
5.24		**-1.93**				0.54	0.44
(5.61***)		(-3.68***)				0.52	**13.57*****
(2.95***)		(-1.95**)					
3.85	-0.41	-1.47				0.64	0.39
(3.50***)	(-0.84)	(-2.28***)				0.60	**24.65*****
(2.01***)	(-0.48)	(-1.31)					
3.89	-0.50	-1.44	0.00			0.64	0.39
(3.33)	(-1.52)	(-2.39***)	(0.06)			0.59	**20.23*****
(1.88)	(-0.86)	(-1.36)	(0.03)				
2.42	-0.40	-0.34	0.15	0.15	0.03	0.74	0.30
(2.71)	(-1.44)	(-1.04)	(1.14)	(0.83)	(0.19)	0.67	41.94
(1.72)	(-0.93)	(-0.68)	(0.74)	(0.54)	(0.12)		

Table 12: Cross–sectional Analysis: FAMA–MACBETH Regressions Using 25 FAMA-FRENCH Portfolios

Note: This table reports the results of the 2–pass FAMA–MACBETH regressions for the average returns across the FAMA–FRENCH (25) portfolios sorted by size and market to book characteristics on different combinations of risk factors. s_{-1} is the lagged (log) surplus consumption ratio, Δc the consumption growth and cay_{-1} the lagged proxy for the consumption–wealth ratio proposed by LETTAU and LUDVIGSON [2001a, 2001b and 2005]. For each regression, the table reports the cross–sectional coefficients. For each coefficient, two t-statistics are reported in parentheses. The top statistic uses the uncorrected FAMA–MACBETH standard errors. The bottom statistic uses the JAGANNATHAN and WANG [1998] correction. Significance at 10%, 5% and 1% levels using the standard t–test are indicated by *, ** and ***. Data are quarterly and span the first quarter of 1952 to the first quarter of 2005. Δc and cay are expressed in 100 base points.

Despite the fact that the estimated risk prices $\hat{\lambda}_{s_{-1}}$ and $\hat{\lambda}_{cay_{-1}}$ are not statistically significant, the adjusted $\overline{R}^2$ reaches 60% and the model is not statistically rejected. The conclusions remain unchanged when consumption growth is added as a risk factor. Indeed, as shown in row 4 of Table 12, the model explains about 59% of the cross–sectional variation in the average excess model. Moreover, the χ^2–statistic remains low and the model cannot be statistically rejected at the 1% level of significance.

As the scope of this paper is to evaluate the cross–section empirical implications of the CAMPBELL and COCHRANE [1999] model, we propose additional empirical regressions testing for the ability of the surplus consumption ratio to explain variations in average returns across stocks. First, we investigate the model misspecification. As suggested by JAGANNATHAN and Wang [1998], the model missspecification can be tested for by including firm characteristics as additional variables:

$$\mathbb{E}\left[R^e_{i,t}\right] = \beta_{i,F}\lambda_F + Z_i\lambda_Z$$

for i= 1 ... 25. The variables F denote the risk factors and Z the firm characteristics. A large t–statistic for the additional characteristics suggests that the model may be misspecified. As the 25 FAMA–FRENCH portfolios are sorted by size and book–to–market value, we re–estimate the Campbell and Cochrane model by including either portfolio size (*SIZE*) or portfolio book–to–market value (BTM). The portfolio size is measured as the time–series average of the (log) market equity for each portfolio. The book–to–market characteristic is measured as the time–series average of the book–to–market ratio for each portfolio. The results are reported in the first panel of Table 13. Two main results emerge. Firstly, the CAMPBELL and COCHRANE [1999]'s model has no difficulty in eliminating the residual size effect as shown in row 1. The t–statistic and the corrected t–statistic corresponding to the *SIZE* characteristic are not statistically significant. Moreover, the adjusted $\overline{R}^2$ does not increase substantially once the size effect is included. However, the model is not able to account for the book–to–market effect. As we can see in Table 13, row 2, the additional book–to–market variable is statistically significant, and the adjusted $\overline{R}^2$ increases dramatically from 42% (Table 10, row 5) to 73%. As a comparison, we report the results of testing for the misspecification of the conditional (C)CAPM model when the scaled variable is the lagged consumption–wealth ratio. As can be seen in the second panel of Table 13, similar conclusions are obtained. Indeed, the BTM characteristic is highly significant when included in the conditional(C)CAPM model and the adjusted $\overline{R}^2$ increases by more than 10%. However, when the *SIZE* characteristic is added, the corresponding $\hat{\lambda}_{SIZE}$ is not statistically significant, implying that the model is unable to capture the size effect.

In conclusion, Table 13 shows that the linear (C)CAPM model with external habit performs far better than the unconditional standard (C)CAPM in accounting for the cross-sectional variation in average stock returns. The key risk factor, i.e. the lagged surplus consumption ratio, is the relevant factor that mimics the risk factors related to the size effect.

	CC(1999)			LL(2001)	
	Size	*BTM*		*Size*	*BTM*
cst	4.31	1.60	cst	**5.07**	**3.15**
	(3.17***)	(2.33***)		(4.75***)	(4.61***)
	(1.53)	(1.58)		(3.42***)	(3.64***)
s_{-1}	-67.38	**-85.71**	cay_{-1}	**-0.06**	-0.08
	(-2.76***)	(-3.10***)		(-1.77)	(-1.89)
	(-1.36)	(-2.12***)		(-2.29***)	(-1.50)
Δc	-0.2	-0.15	Δc	-0.04	0.20
	(-1.38)	(-1.02)		(-0.30)	(0.93)
	(-0.68)	(-0.70)		(-0.22)	(0.74)
$s_{-1}\Delta c$	0.19	-0.35	$cay_{-1}\Delta c$	0.00	-0.00
	(1.28)	(-2.04)		(0.74)	(-1.68)
	(0.63)	(-1.40)		(0.54)	(-1.34)
CHARAC	-2.21	**0.70**	*CHARAC*	-0.14	**0.62**
	(-1.93*)	(2.69***)		(-1.39)	(2.49***)
	(-0.93)	(1.82*)		(-1.00)	(1.96**)
R^2	0.53	0.77	R^2	0.67	0.69
$\overline{R}^2$	0.44	0.73	$\overline{R}^2$	0.61	0.63
Avr.(%)	0.44	0.31	*Avr.*(%)	0.37	0.36
χ^2	**14.00*****	**24.83*****	χ^2	**31.91****	**34.47****

Table 13: CROSS–SECTIONAL REGRESSIONS: FAMA-MACBETH REGRESSIONS INCLUDING CHARACTERISTICS

Note: This table reports the results of the 2 pass FAMA–MacBeth regressions for the average excess returns across the FAMA–FRENCH (25) portfolios including size or book–to–market ratio as characteristics. See Table 10 for the description of models CC(1999) and LL(2001).
For each regression, the table reports the cross sectional coefficients. For each coefficient, two t–statistics are reported in parentheses. The top statistic uses the uncorrected FAMA–MACBETH standard errors. The bottom statistic uses the JAGANNATHAN and WANG [1998] correction. Significance at 10%, 5% and 1% levels using the standard t–test are indicated by *, ** and ***. We also report the average squared pricing errors across all portfolios (*Avr.*) in % and a χ^2–statistic for the test of the null hypothesis that all of the pricing errors are jointly zero. This statistic is computed using the SHANKEN [1992] correction. *, ** and *** indicate that the average pricing error is statistically different from zero at the level 10%, 5% and 1%. Data are quarterly and span the first quarter of 1952 to the first quarter of 2005. Δc and *cay* are expressed in 100 base points.

	cst	s_{-1}	Δc	$s_{-1}\Delta c$	R^2 $\overline{R}^2$
$\theta = 0.5$	1.66	-3.44	-0.07	-0.61	0.59
	(2.35)	(-4.46)	(-0.53)	(-1.26)	0.53
	(1.08)	(-2.07)	(-0.25)	(-0.58)	
$\theta = 2$	1.53	-1.17	-0.09	-0.15	0.49
	(2.20)	(-4.18)	(-0.70)	(-0.81)	0.42
	(1.03)	(-1.99)	(-0.33)	(-0.38)	
$\theta = 5$	1.35	-0.51	-0.04	-0.02	0.35
	(1.89)	(-3.23)	(-0.32)	(-0.23)	0.26
	(1.01)	(-1.75)	(-0.17)	(-0.12)	

Table 14: Sensitivity Analysis: Alternative values for θ

Note: This table reports the results of the 2 pass FAMA–MacBeth regressions for average returns across the FAMA–FRENCH (25) portfolios sorted by size and market to book value on the factors $F_t = [\ cst\ s_{t-1}\ \Delta c_t\ s_{t-1}\Delta c_t\ ]$ where s_{-1} is the lagged (log) surplus consumption ratio and Δc consumption growth. The table reports the FAMA-MacBeth cross–sectional coefficients. For each coefficient, two t–statistics are reported in parentheses. The top statistic uses the uncorrected FAMA–MACBETH standard errors. The bottom statistic uses the JAGANNATHAN and WANG [1998] correction. The term R^2 denotes the cross-sectional R^2–statistic and the $\overline{R}^2$ adjusted for the degree of freedom. Data are quarterly and span the first quarter of 1952 to the first quarter of 2005. Δc is expressed in 100 base points.

Robustness

As in the time–series analysis presented in Section 2.1, we study the robustness of our cross–sectional empirical results by using (i) alternative values of the utility curvature parameter θ and (ii) alternative initial values of the time–series surplus consumption ratio.

Firstly, we test the empirical implications of alternative values of the coefficient of relative risk aversion $\theta = 0.5$, 2 and 5. Table 14 shows that the empirical implications of the CAMPBELL and COCHRANE [1999] model remain unchanged whatever the value of $\theta = 0.5$, 2 and 5. Indeed, the estimated risk price associated to the lagged surplus consumption ratio is always negative and statistically significant, even after correction for errors. Moreover, the adjusted $\overline{R}^2$ remains substantially high compared to the adjusted $\overline{R}^2$ implied by the standard Consumption–based Asset Pricing model.

Secondly, we check the robustness of our empirical results to alternative initial values of the time–series surplus consumption ratio. Several initial values of the surplus are tested. As the conclusions remain unchanged, only a few of them are reported in Table 15 as in the benchmark case[16]. As already noted in Section 2.1, the predictive power of the (lagged) surplus consumption ratio is remarkably stable over the restricted sub–sample 1962 : $Q1$

[16]Results corresponding to intermediate initial values are provided upon request.

	Whole sample 1952:Q1–2002:Q4			Sub Sample 1962:Q1–2002:Q4		
initial value	$\lambda_{s_{-1}}$	t–stat	R^2 $\overline{R}^2$	$\lambda_{s_{-1}}$	t–stat	$\overline{R}^2$ $\overline{R}^2$
-10	-1.32	-3.23	0.31	-1.24	-4.06	0.42
		-1.70	0.21		-1.76	0.34
-5	-1.18	-4.36	0.43	-1.22	-4.05	0.41
		-2.20	035		-1.77	0.33
0	-1.31	-3.24	0.69	-1.22	-4.05	0.41
		-1.16	0.65		-1.77	0.33
5	-1.93	-2.98	0.78	-1.18	-4.03	0.41
		-0.94	0.74		-1.78	0.33
10	-2.44	-2.61	0.78	-1.18	-4.03	0.41
		-0.76	0.75		-1.78	0.32

Table 15: SENSITIVITY ANALYSIS: ALTERNATIVE VALUES FOR THE INITIAL VALUE

Note: This table reports the results of the 2 pass FAMA–MACBETH regressions for the average excess returns across the FAMA–FRENCH (25) portfolios sorted by size and market to book value on the factors $F_t = [\ cst\ s_{t-1}\ \Delta c_t\ s_{t-1}\Delta c_t\ ]$ where s_{-1} is the lagged (log) surplus consumption ratio and Δc consumption growth. For each estimated coefficient slope, two t-statistics are reported. The top statistic uses the uncorrected FAMA–MACBETH standard errors and The bottom statistic uses the JAGANNATHAN and WANG [1998] correction. The surplus consumption ratio is generated using Equation (2). Initial values imposed at the first quarter of 1952 are reported in the first row. Data are quarterly and span the first quarter of 1952 to the first quarter of 2005. Δc is expressed in 100 base points.

– $2002 : Q4$. Indeed, in this case, the estimated risk price associated to the surplus does not fluctuate substantially in response to changes in the initial values of the time–series for the surplus consumption ratio, imposed in the first quarter of 1952. For instance, the estimated coefficients shift from -1.28 to -1.18 when initial values vary between -10 and 10. Additionally, both the coefficient of determination R^2 and the adjusted R^2 remain almost unchanged. This suggests that our cross–sectional results are not sensitive to the initial values of the surplus consumption ratio.

IV. Concluding Remarks

This paper investigates the role of surplus consumption ratio in predicting excess returns. We show empirically that the surplus consumption ratio is a good predictor of excess returns at long horizons. Additionally, as CAMPBELL and COCHRANE [1999]'s Consumption–Based Asset Pricing Model with habit formation implies a three–macroeconomic–factors model, we test the empirical performance of the model in accounting for the cross–sectional variations in average excess returns. The risk factors are consumption growth, the lagged surplus consumption ratio and their product. We show empirically that the surplus consumption ratio is the key risk factor that helps to explain the cross–section of average returns on the 25 FAMA–FRENCH portfolios sorted by size and book–to–market value.

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