Internal vs External Habit Formation: The relative importance for asset pricing

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Abstract

I present a generalized model that structurally nests either "catching up with Joneses" (external habit) or "time non-separable" (internal habit) preference specifications. The model asset pricing implications are confronted with the observed aggregate US consumption and asset returns data to determine the relative importance of "catching up with Joneses" and internal habit formation. Using long-horizon returns, I show that habit persistence with a sufficiently long history of consumption realizations is more consistent with observed aggregate returns properties than "catching up with Joneses" preferences. These results have important implications for researchers attempting to provide microeconomic foundations of habit formation.

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Habit formation models became increasingly successful and important in explaining number of dynamical asset pricing facts, such as equity premium puzzle, see, e.g. Constantinides (1990), Campbell and Cochrane (1999), and Abel (1990), as well as macroeconomics facts, such as output persistence (Boldrin, Christiano, and Fisher (2001)), savings and growth (Carroll, Overland, and Weil (2000)), and response of consumption to monetary shocks (Fuhrer (2000)). Although successful, different studies use alternative types of habit models, external[1] or internal habit formation. Abel (1990), Gali (1994), Campbell and Cochrane (1999) among the first introduce and study models with external habit formation and their implications for asset pricing.[2] In this type of models, past consumption enters into habit process but has no effect on current consumption choice, that is, habit formation is an externality. Ryder and Heal (1973), Dunn and Singleton (1986), Sundaresan (1989), Constantinides (1990), Detemple and Zapatero (1991) introduce and study habit persistence, or internal habit formation, where the past consumption choice enters into habit process and affects current and future consumption choices. As a result, these two types of habits produce different pricing kernels and might lead potentially to different asset pricing implications. Campbell and Cochrane (1999) essentially claim that the difference between the two types is more or less innocuous, and their argument hinges on the fact that if the aggregate endowment process is a random walk and the habit formation process is linear, the marginal rate of substitution of an external habit formation model is proportional to that of an internal habit formation model, which implies that both types of models should have exactly the same asset pricing implications. However, this claim need not to be true on theoretical grounds because the endowment process might differ from random walk (see, e.g. Kandel and Stambaugh (1990), Hansen and Singleton (1983), or Hall (1978)) or the habit process might not be linear. Therefore, an unresolved question in the habit literature is whether the difference between an internal habit formation model and an external habit formation model is empirically relevant. This is the main focus of the current paper.

This debate is important not only for finding correct specification for pricing kernel, but also for theoretical reasons. In the last few years there has appeared several studies attempting to provide microeconomic foundations of habit formation.[3] It is interesting that researchers are concerned more with finding microeconomic foundations for "catching up with Joneses" preferences than for internal habit formation. Based on different model assumptions they (might) come to different conclusions about the nature of habit process. This motivates me to investigate a simpler, but more fundamental question: to what extent the asset pricing data is consistent with either external or internal habit formation preferences?

Empirical studies related to habit formation are rather limited. For example, Ferson and Constantinides (1991) find empirical support for one-lag internal habit formation model using quarterly seasonally non-adjusted data. Heaton (1995) also finds evidence for habit formation in quarterly aggregate consumption data by adopting a multi-lag habit structure. Chen and Ludvigson (2006s) estimate separately external and internal habit formation models using aggregate consumption data and find evidence for internal habit persistence. Likewise, Ravina (2005) provides an evidence for habit persistence using the U.S. credit card accounts data of households located in California. On contrary, Dynan (2000) uses annual household food consumption data and finds no evidence of habit formation in this data set.

To proceed, I begin with the idea first developed by Abel (1990) that the "catching up with the Joneses" (namely external habit formation) and "habit persistence" (namely internal habit formation) behavior can be captured by a more general specification in which a free parameter controls the relative importance of both. For his theoretical development, Abel (1990) adopted a Cobb-Douglas type of specification of the habit based on the lagged value of both individual and aggregate consumption. To facilitate both theoretical interpretation and empirical implementation, I propose a specification that captures the same spirit of Abel's model but extends it along several dimensions. First, I assume that the habit level is a function of individual and aggregate consumption histories. Following Ryder and Heal (1973) and Constantinides (1990) the habit level is an exponentially weighted average function of these two. Second, I extend Abel's one-lag specification to infinite lags. This extension is motivated by Heaton (1995) who shows that habit formation behavior does not kick in for at least several months and is in general highly persistent (meaning that the influence of past consumption on current habit level decays slowly). Third, although an individual's habit level depends on the entire history of the aggregate consumption, I assume that the agent looks back only at finite consumption history in forming her current habit level, meaning that only the finite number of individual consumption lags enter into habit stock. Such a specification is motivated by a reasonable conjecture that individual consumers typically do not keep record or remember their own consumption choices beyond several quarters and almost certainly not beyond several years. The practical benefit of this specification is that the marginal utility of consumption for the individual cuts off naturally to the finite number of lags (as far as the individual looks into her own consumption history), rather than infinite number of lags (or as far as records of past aggregate consumption remain accessible). This makes it a lot easier to implement the econometric estimation and test the model using generalized method of moments (GMM).[4]

Based on the theoretical specification, I derive, in closed form, stochastic Euler equations that restrict aggregate asset pricing behavior. I estimate then the model parameters based on the first moments of asset returns. The central emphasis of my work is the estimation of the degree of relevance of either habit type that is most consistent with the historical asset pricing behavior. The over-identifying moment restrictions on the cross-section of asset returns are used to test the model based on the standard asymptotic distribution theory developed by Hansen (1982).

Main empirical findings emerge. First, I find that US aggregate postwar consumption and stock market data strongly support internal habit preferences. Second, external habit formation model is rejected on the conventional levels of statistical significance. Third, internal habit model is identified and not rejected, but only when habit stock is formed using sufficiently long history of individual consumption. This result is robust to different choices of instrumental variables. It is consistent with Heaton (1995) who found support for long-term effect in habit formation and Ferson and Constantinides (1991) who found empirical support for short-term habit using quarterly and annual data, but not monthly data. Fourth, I find that long horizon returns are necessary to identify the relative importance between external and internal habit preferences. Predictability of long-horizon returns (see Campbell, Lo, and MacKinlay (1997), page 268, e.g.) can potentially serve to distinguish between external and internal habit. I find empirical support for this by examining my "mixture" model using not only quarterly returns, but also annual and 2-year returns. This suggests that internal habit is more consistent with the observed asset return behavior than external habit – other things being equal. These results have important implications for researchers attempting to provide microeconomic foundations of habit formation.

The rest of the paper is organized as follows. I introduce the model specification and derive the stochastic Euler equations in section I. In section II I present empirical setup, discuss methodological issues related to our empirical study and report model estimation results. I conclude in Section III.

I A "Mixture" Habit Formation Model

The economy in the present model is populated by a continuum of identical, competitive agents with total measure 1. At time t, each individual agent consumes c_t. The aggregate (per capita) endowment is denoted as C_t. The individual consumption choice c_t is chosen so as to maximize her expected utility of the form:

$$V_0 = E\left[\sum_{t=0}^{\infty} \rho^t\, u(c_t - x_t) \middle| I_0\right], \tag{1}$$

where $0 < \rho < 1$ is time discount factor, $u(\cdot, \cdot)$ is strictly increasing in the first argument and strictly decreasing in the second argument, and strictly concave in both arguments. I assume the standard CRRA utility function:

$$u(z_t) = \frac{z_t^{1-\gamma} - 1}{1-\gamma}, \qquad \gamma > 0, \tag{2}$$

γ is the utility curvature parameter and is literally a relative risk aversion coefficient in the case of time separable utility. The variable $z_t = c_t - x_t$ is individual's surplus consumption. x_t is interpreted as the individual reference level, in general, and is assumed to be a function of the past history of both individual consumption choices and aggregate endowment. That is, in general, I write

$$x_t = \mathbb{X}(c_s, C_s : s < t). \tag{3}$$

We say that the individual's preference exhibits *internal habit formation behavior* with horizon $j > 0$ if

$$b_{t,j} \equiv \frac{\partial x_{t+j}}{\partial c_t} > 0, \tag{4}$$

and *external habit formation behavior* with horizon $j > 0$ if

$$B_{t,j} \equiv \frac{\partial x_{t+j}}{\partial C_t} > 0. \tag{5}$$

The "habit formation behavior" (either external or internal) is re-interpreted as a durability if $b_{t,j}$ or $B_{t,j}$ are negative.[5] Durability of consumption induces negative autocorrelation in consumption growth. For example, individual who has purchased a car (real estate, etc.) this period, is unlikely to purchase another one in the next period. On the other hand, habit persistence induces positive autocorrelation in the consumption growth because utility-maximizing consumer is smoothing consumption by more than would be optimal with time separable preferences.

In equilibrium, $c_t = C_t$ for all t, and the equilibrium habit process is given by $X_t \equiv \mathbb{X}(C_t, C_t)$.

A *Habit Specification*

The most important aspect of the model is the parametric form of the individual consumers' habit formation processes. I assume that individual consumers form their habit level based on both their own and aggregate (per capita) consumption. I also assume that individual keeps the history of her own consumption up to $J+1$ last periods, but the complete history of aggregate consumption is known to every consumer and she takes it into account when forming habit stock.[6] Formally, I assume that an individual consumer's habit level is determined by

$$x_{t+1} = b \sum_{j=0}^{J} (1-a)^j \{ \omega c_{t-j} + (1-\omega) C_{t-j} \} + b \sum_{j=J+1}^{\infty} (1-a)^j C_{t-j}, \tag{6}$$

where $0 < b < a < 1$, and $J \geq 0$. This habit specification means that x_t is a function of both agent's own consumption and aggregate per capita consumption, in the spirit of Abel (1990).[7] For period t, a "catching up with Joneses" agent compares own current consumption c_t with the past consumption of her/his peers and this is reflected in the fact that s/he maximizes her utility over consumption in excess of the weighted average of the past aggregate per capita consumption C_{t-j}, $j \geq 1$. At the same time, an (internal) habit consumer compares own current consumption c_t with the weighted sum of the own past consumption. Hence, s/he maximizes utility over c_t in the excess of her/his own past weighted average c_{t-j}, $1 \leq j \leq J$. In other words, s/he takes into account the effect of the current consumption choice on future realizations of x_t. This is reflected in the fact that marginal utility of consumption has forward-looking terms, which are the conditional expectations of the future atemporal marginal utilities (The exact form of the marginal utility is presented in the Section B).

The parameter b is a scaling parameter, which indexes the degree of importance of the habit formation level relative to the current consumption level. If $b = 0$, then the standard time separable model applies. The parameter a indexes the degree of persistence, or "memory" in the habit stock. If $a = 1$ then only last period consumption is important. In general, the smaller is a, the further back in history is the habit formation level determined. The parameter ω is referred to as the "mixture" parameter, since, when $0 < \omega < 1$, the model captures a mixture of both internal and external habit formation behavior:

$$b_{t,j} = b_h = \begin{cases} \omega \times b\,(1-a)^{j-1}, & \text{if} \quad j \leq J+1, \\ 0, & \text{if} \quad j > J+1, \end{cases} \tag{7}$$

$$B_{t,j} = B_j = \begin{cases} (1-\omega) \times b\,(1-a)^{j-1}, & \text{if} \quad j \leq J+1, \\ b\,(1-a)^{j-1}, & \text{if} \quad j > J+1. \end{cases} \tag{8}$$

The fact that $b_{t,j} = b_j$ and $B_{t,j} = B_j$ are constant is due to the linearity of the habit specification.

Compared to conventional habit specifications, my model contains two new parameters: $\omega \in \mathbb{R}$ and $J \in \mathbb{Z} \cup \infty$. I refer to ω as "mixture" habit parameter and to J as cutoff parameter. Note that cutoff J means that $J+1$ lags of individual consumption are used in the formation of habit stock. Where no confusion arises, I use number of lags $NLAG = J+1$ and cutoff J interchangeably. I refer to $NLAG$ as the memory of the habit process. ω captures the same "mixture" feature that Abel (1990) introduced, but without the sign restriction. In his model $J = 0$, and the tail sum in the right hand side of equation (6) is absent.[8] However, in my specification, tail sum (6) serves a practical purpose. First, it ensures that the equilibrium habit process is not affected by either ω or J. To see this, note that in equilibrium, $c_t = C_t$. It follows that the equilibrium habit process is given by

$$X_{t+1} = b \sum_{j=0}^{\infty} (1-a)^j \, C_{t-j} = b\, C_t + (1-a)\, X_t, \tag{9}$$

This in turn ensures that the parameters a and b have exactly the same meaning as in Constantinides (1990).[9] Second, it makes the equilibrium habit process a lot smoother than and less correlated (unconditionally) with the equilibrium consumption process, and consequently makes the growth rate of the surplus consumption process a lot more volatile than the consumption growth. This in turn increases the equilibrium market price of risk without increasing the curvature parameter.

Different values of ω and/or J give rise to different asset pricing implications because individual consumers' marginal rates of substitution and hence the equilibrium marginal rate of substitution depend on ω and J. When $\omega \neq 0$, the marginal utility of consumption is a series sum with $J+2$ terms, with the first term representing the standard marginal utility in the absence of temporal dependence (no habit or pure external habit) and the remaining terms representing the contribution to the expected future utility from habit persistence. As we will see explicitly in the next section, the contribution due to internal habit formation is regulated by both ω and J: ω controls the relative magnitude of importance of two habit types and J controls the horizon of temporal dependence. To make such a dependence explicit, it is convenient to write the marginal utility of consumption (MUC) and the marginal rate of substitution (MRS) as $MUC_t^{\omega,J}$ and $MRS_{t,t+1}^{\omega,J}$, respectively, so that the dependence of the stochastic Euler equations on ω and J can be made explicit through (and only through) the MRS:

$$E_t\left[MRS_{t,t+1}^{\omega,J} \times R_{t,t+1}\right] = 1, \tag{10}$$

where $R_{t,t+1}$ is the one-period realized return for any traded security. In principle, the parameter ω and J are identified if the Euler equations hold if and and only if $\omega = \omega^*$ and $J = J^*$, where ω^* and J^* are the true values.[10]

Holding J to a fixed and finite value, the model changes smoothly from one type of habit formation behavior to another as ω varies between 0 and 1. Two ends of $[0,1]$ interval correspond to two special cases characterized by either pure internal or external habit at each horizon (that is, there is no mixture of internal and external habit formation for a given horizon). When $\omega = 0$, no weight is placed on the past individual consumption in the construction of habit process. In this case, aggregate consumption presents a mere externality since agents who increase their consumption do not take into account their effect on the aggregate desire by other agents to "catch up". When $\omega = 1$, agents construct their habit process using individual consumption up to $J+1$ lags and take into account the effect of changing future marginal utility up to $J+1$ lags too because of a higher today's consumption.

For sufficiently large J, these two special cases capture the essence of the Campbell-Cochrane model (literally) and the Constantinides model (as a close approximation). Of course, the Constantinides model obtains when $\omega = 1$ in the limit $J \to \infty$. Thus, these two models wound up as two special cases.

When $\omega < 0$, $b_h < 0$ for any $h \leq J + 1$, which represents *local substitution* at horizon h. When $\omega > 1$, $B_h < 0$ for $h \leq J + 1$, which represents a behavior that may be characterized, with a slight abuse of language, as *external local substitution*.

The following diagram outlines special cases of the model studied earlier:

Parameters	Model	Notes
$b = 0$	Lucas (1978)	time separable model
$J = 0$	Abel (1990)	surplus consumption is a ratio of current consumption to "benchmark" consumption level, which is a weighted average of one-period lag consumer's own and one-period lag average consumption
$\omega = 1,\ J = \infty$	Constantinides (1990)	internal habit formation
$\omega = 0,\ J = \infty$	Campbell and Cochrane (1999)	external habit formation
$\omega = 1,\ a = 1$	Dunn and Singleton (1986)	preferences are non-separable over non-durable and durable goods, habit formation is on the non-durable goods only
$\omega = 1,\ a = 1$	Ferson and Constantinides (1991)	empirical investigation
$\omega = 1,\ J = 19$	Heaton (1995)	weekly frequency, 19 weeks $\approx$ 5 quarters

I estimate the model via Generalized Method of Moments (GMM) developed by **?**. The empirical strategy is to estimate the parameter ω from observed asset returns, holding fixed the parameter J at some representative and finite values. J is not estimated econometrically partly because it takes value only in the set of non-negative integers, and partly because the econometric procedure breaks down when J becomes too large (relative to the sample length). The central focus of my analysis is the parameter ω, which will be estimated under different values of J as a form of robustness check.

As a diagnostic check, I plot Hansen-Jagannathan bounds to examine the differences between external (EH) and internal (IH) habit formation models. Using asset market data

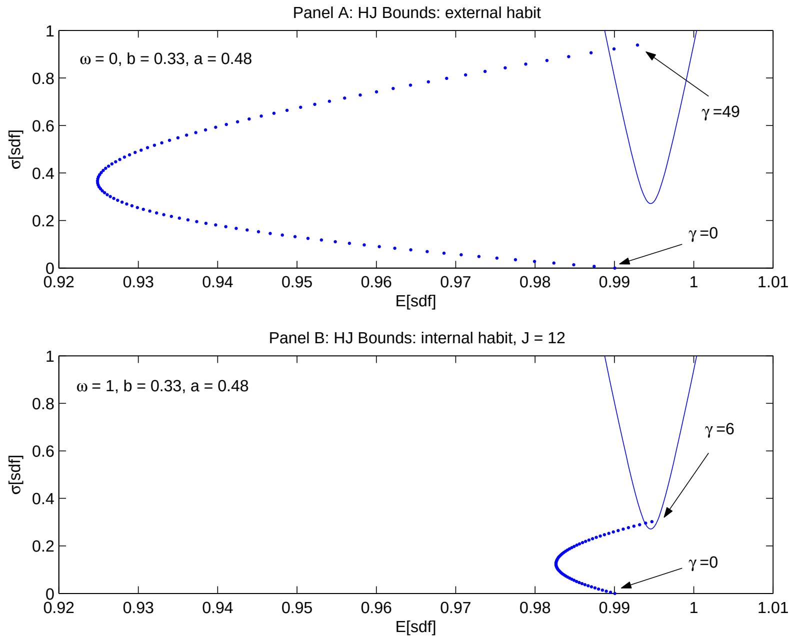

Figure 1: **Hansen-Jagannathan Bounds**.

alone, Hansen and Jagannathan (1991) compute lower bound on the volatility of those stochastic discount factors that correctly price assets under consideration. The volatility bound developed by Hansen and Jagannathan (1991) is constructed by specifying a mean of the marginal rate of substitution and then using asset market data to estimate the lower bound. To calculate the bound, I use real quarterly returns of US Treasury Bills, 5-year maturity Treasury Bond portfolio, and quarterly returns on six size/book-to-market sorted portfolios.[11] The cup-shaped regions on the pictures of Figure 1 give the lower standard deviation bound as a function of the mean of the stochastic discount factor. In order to satisfy a bound, a model's mean-variance pair of the intertemporal marginal rate of substitution must lie in the cup-shaped region. Panel A shows that curvature parameter $\gamma = 49$ is required for external habit to fit HJ-bounds. On the other hand, internal habit formation with 12 consumption lags gets into the bounds for γ as small as 6. This observation motivates the hypothesis that internal habit, which is formed by sufficiently long history of past consumption "fits" better time-varying stock and bond returns.[12] The reason is that when the consumer is "catching up with Joneses", her reference point presents a mere externality and does not affect the optimal consumption choice. From the perspective of an individual consumer, Abel-type preferences are time separable, because a change in the individual consumption level c_t at time t does not affect the marginal utility at time $t+\tau$ with respect to $c_{t+\tau}$. Note that for such preferences the coefficient of relative risk aversion, RRA_t, is equal to:

$$RRA_t = \gamma \frac{c_t}{c_t - x_t}. \tag{11}$$

In this case RRA_t is time-varying and is greater than γ everywhere because $c_t > x_t$.[13]

Alternatively, (internal) habit formation preferences are time non-separable: in a response to a wealth shock at time t agent with habit formation preferences adjusts her state-contingent consumption plan at the future dates in such a way so to adjust optimally future habit stock $x_{t+\tau} : \tau = 1, \ldots, J+1$. Therefore, marginal utility at t is affected not only by the change of consumption c_t, but also by changes in the consumption plan at future dates $t+\tau$, $\tau = 1, \ldots, J+1$. Compared with "catching up with Joneses" preferences, this reduces the impact of a given wealth shock on the objective function and explains why internal habit agent has a lower curvature. Although it is not possible to derive relative risk aversion coefficient in the present setup, Constantinides (1990) and Ferson and Constantinides (1991) find bounds on RRA and prove that RRA coefficient approximately equals γ, but the elasticity of intertemporal substitution might be lower than the inverse of the RRA coefficient.[14] They show that relative risk aversion coefficient is much closer to γ than the one implied by Abel-type preferences. In both cases, relative risk aversion coefficient is time-varying in the present setup when consumer cares about the difference of present consumption and reference level. This is consistent with countercyclical risk premia observed in the historical data.

B *Marginal Utility of Consumption and Stochastic Euler Equations*

In preparation for the model econometric analysis, in particular, the asset pricing implications of ω and J, I derive explicitly stochastic Euler equations (10) or the marginal rate of substitution $MRS_{t,t+1}^{\omega,J}$. To do this, I apply a standard perturbation argument. To this end, consider an arbitrary traded security with price-dividend pair (p_t, d_t). The one-period return from t to $t+1$ is given by $R_{t+1} \equiv \frac{p_{t+1}+d_{t+1}}{p_t}$. Suppose that the economy has achieved equilibrium with the optimal consumption policy of an arbitrary individual consumer given by c_t and the equilibrium consumption process given by C_t. If the consumer tries to trade away from her optimal consumption policy by purchasing α share of the security at t and selling it at $t+1$, the trading strategy must be financed by a reduction in her consumption level at t and will raise her consumption level at $t+1$. In other words, the consumer gives up αp_t consumption at time t, but receives additional consumption $\alpha(p_{t+1} + d_{t+1})$.

The net marginal effect of the increase in consumption on her expected utility at t is given by

$$\frac{\partial}{\partial \alpha} E_t \left[\sum_{j=0}^{\infty} \rho^j \times u(\hat{c}_{t+j} - \hat{x}_{t+j}) \right], \tag{12}$$

where

$$\begin{aligned}
\hat{c}_{t+j} &= c_t - \alpha p_t \text{ for } j = 0, \\
\hat{c}_{t+j} &= c_{t+1} + \alpha(p_{t+1} + d_{t+1}), \text{ for } j = 1, \\
\hat{c}_{t+j} &= 0, \text{ otherwise, } \quad \text{and} \\
\hat{x}_{t+j} &= \mathbb{X}(\hat{c}_s, C_s : s < t+j), j \geq 0
\end{aligned} \tag{13}$$

Since no-trade is optimal at equilibrium, I evaluate (12) at $\alpha = 0$. Then the marginal rate of expected utility loss (with respect to α) is given by $MUC_t^{\omega,J} \times p_t$, where $MUC_t^{\omega,J}$ is the *marginal utility of consumption*, defined by

$$MUC_t^{\omega,J} \equiv E_t \left[\sum_{j=0}^{J+1} \rho^j \times a_{t,j} \times u'(c_{t+j} - x_{t+j}) \right] \tag{14}$$

where

$$a_{t,j} \equiv \left. \frac{\partial}{\partial c_t} (\hat{c}_{t+j} - \hat{x}_{t+j}) \right|_{\alpha=0} = a_j = \begin{cases} 1, & j = 0, \\ -b_j, & 1 \leq j \leq J+1. \end{cases} \tag{15}$$

On the other hand, the net marginal gain in expected utility from selling the security at $t+1$, evaluated at $\alpha = 0$, is given by $E_t\left[\rho \times MUC_{t+1}^{\omega,J} \times (p_{t+1} + d_{t+1})\right]$. At equilibrium, the expected gain must be equal to the expected loss. It follows that

$$MUC_t^{\omega,J} = E_t\left[\rho \times MUC_{t+1}^{\omega,J} \times R_{t+1}\right], \tag{16}$$

which can be re-written as equation (10), by defining $MRS_{t,t+1}^{\omega,J} \equiv \rho \times \frac{MUC_{t+1}^{\omega,J}}{MUC_t^{\omega,J}}$ as the marginal rate of substitution *for the individual consumer*.

While equation (10) is more familiar and is more convenient for economic interpretation, an alternative and equivalent expression is more suitable for the purpose of econometric analysis. Specifically, through repeated use of the law of iterated expectations, equation (16) can be re-written as

$$E_t\left[\Phi_t^{\omega,J} - \rho \times \Phi_{t+1}^{\omega,J} \times R_{t+1}\right] = 0, \tag{17}$$

where $\Phi_t^{\omega,J} \equiv \sum_{j=0}^{J+1} \rho^j \times a_{t,j} \times u'(c_{t+j} - x_{t+j})$. Obviously, $MUC_t^{\omega,J} \equiv E_t\left[\Phi_t^{\omega,J}\right]$. Note that the sample counterpart of equation (17) can be constructed explicitly without evaluating any conditional expectations. In contrast, the sample counterpart of equation (16) can not be easily constructed because $MUC_t^{\omega,J}$ can not be evaluated without specifying the law of motion for the aggregate or equilibrium consumption process.

In equilibrium, equations (10), (16), and (17) must hold with c_t replaced by C_t, x_t replaced by X_t, and $a_{t,j}$ remaining the same as above. Henceforth, I re-define $MUC_t^{\omega,J}$ as the equilibrium marginal utility of consumption, which is given by the same equation (14), except that c_t and x_t are replaced by their aggregate counterparts, C_t and X_t, respectively. Similarly, $MRS_{t,t+1}^{\omega,J}$ and $\Phi_t^{\omega,J}$ are re-defined in terms of equilibrium consumption and habit levels.[15]

II Empirical Analysis

Using observed aggregate consumption and asset return data, I estimate and test the model using GMM based on the Euler equations (17).[16]

A *Econometric Procedure*

To this end, let us collect all model parameters in the vector $\theta = (\rho, \gamma, a, b, \omega; J)$, and denote the vector of n-period returns of K assets by $R_{t,t+n}$ and the vector of M instruments by Z_t. Under the null that the model is correctly specified, the following $K \times M$ orthogonality conditions must hold:

$$E\left[\epsilon_t(\theta)\right] = 0, \tag{18}$$

where

$$\epsilon_t(\theta) \equiv (\Phi_t - \rho^n \times \Phi_{t+n} \times R_{t,t+n}) \otimes Z_t, \tag{19}$$

$$\Phi_t \equiv u'(C_t - X_t) - \omega \times b \times \sum_{j=1}^{J+1} \rho^j \times (1-a)^{j-1} \times u'(C_{t+j} - X_{t+j}). \tag{20}$$

Let $C^T = \{C_t : 1 \leq t \leq T + J + 1 + n\}$ be the observed per capita consumption process of length $\hat{T} = T + J + 1 + n$ quarters, and let $g_T(\theta) \equiv g(C^T; \theta)$ be the sample counterpart of the left hand side of equation (18), that is,

$$g_T(\theta) = \frac{1}{T}\sum_{t=1}^{T} \epsilon_t(\theta), \tag{21}$$

where the effective sample length T takes into account the fact that ϵ_t depends on future consumptions up to $C_{t+J+1+n}$. Then the GMM estimator of θ is given by

$$\theta_T = \arg\min_{\theta} g'_T W_T^{-1} g_T, \tag{22}$$

where W_T is the sample counter-part of the optimal weighting matrix (see **?**), namely,

$$W_T(\theta) = \frac{1}{T}\sum_{t=1}^{T} \epsilon_t(\theta)\epsilon_t(\theta)' + \frac{2}{T}\sum_{t=1}^{T}\sum_{j=1}^{t-1} \epsilon_t(\theta)\epsilon_{t-j}(\theta)', \tag{23}$$

evaluated at any consistent estimator of θ. While estimator (23) is consistent it is not necessarily positive semi-definite in any finite sample when residuals are autocorrelated to some non-zero degree. Such a situation is problematic for two reasons. First, estimated variances and test statistics will be negative for some linear combinations of θ when the estimated covariance matrix is not positive semi-definite. Second, iterative GMM techniques for computing an optimal GMM estimator $\hat{\theta}$ may behave poorly if $\hat{W}_T(\theta)$ is not positive semi-definite. Therefore, I use weighing matrix estimator, which is suggested by Newey and West (1987) and is positive semi-definite.

Under the null, the GMM objective function $TJ_T = T\, g_T' W_T^{-1} g_T$ has an asymptotic χ^2 distribution with degrees of freedom equal to $(K \times M - \dim(\theta))$. This provides an overall goodness-of-fit test and referred to as J_T test.

Note that ϵ_t is in the information set at $t + J + 1 + n$, but not in the information at $t + J + 2 + n$. Eichenbaum, Hansen, and Singleton (1988) show that in the case of time non-separable utility model pricing errors have an autocorrelation structure of a moving average process of order equal to one less the maximum number of leads in the decision variable. In my case it implies that ϵ_t has an $MA(J+n)$ autocorrelation structure. Following many authors, I do not impose this restriction formally in our estimation. As an additional robustness check, I experiment with different autocorrelation structures, which depend on a cut-off horizon J in habit specification.

As I mentioned earlier, J takes discrete values, and therefore cannot be estimated by standard econometric procedures. In addition, I would run into serious small sample problems if J becomes too large. Part of our empirical strategy is therefore to estimate only the continuous parameters ρ, γ, a, b, and ω with J fixed at some representative values. Accordingly, I re-define the parameter vector θ by excluding J: $\theta = (\rho, \gamma, a, b, \omega)$.

B ***Data***

In the empirical part of the study I use standard data set described below. Table I presents descriptive statistics.

[INSERT TABLE I HERE ...]

Consumption: I use quarterly US consumption data because it is known to contain less measurement errors than the monthly consumption data.[17] Quarterly decision interval also allows me to focus on pure habit effect, because Heaton (1995) shows that local substitution is important only for decision intervals much shorter than a quarter. The sample period is from the fourth quarter of 1951 to the fourth quarter of 2002.

Aggregate consumption data is measured as expenditures on non-durable goods and services excluding shoes and clothing.[18] In order to distinguish between long-term habit persistence and short-term seasonality,[19] I use seasonally adjusted data at annual rates, in billions of chain-weighted 2000 dollars. Real per capita consumption is obtained by dividing real aggregates by a measure of U.S. population. The latter is obtained by dividing real total disposable income by real per capita disposable income. Consumption is detrended by the mean consumption growth rate, so that detrended series has a gross consumption growth rate one.[20] Consumption, price deflator and the measure of population are from NIPA (National Income and Product Account) tables.[21]

Asset Returns: For asset returns I use the following US quarterly data:

- 6 size/book-market returns: Six portfolios, monthly returns from January 1947-December 2002. The portfolios, which are constructed at the end of each June, are the inter-sections of 2 portfolios formed on size (market equity, ME) and 3 portfolios formed on the ratio of book equity to market equity (BE/ME). The size breakpoint for year t is the median NYSE market equity at the end of June of year t. BE/ME for June of year t is the book equity for the last fiscal year end in $t-1$ divided by ME for December of $t-1$. The BE/ME breakpoints are the 30th and 70th NYSE percentiles. Data is taken from Kenneth French's website, *http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html*. Real asset returns are deflated by the implicit chain-type price deflator (2000=100).

- 3-month Treasury Bill Portfolio, 5-, and 10-year maturity Treasury Bond Portfolio returns are obtained from CRSP (Center for Research in Security Prices at the University of Chicago) database. Real asset returns are deflated by the implicit chain-type price deflator (2000=100).

Instruments: As instrumental variables I use constant vector of ones and a proxy for consumption-wealth ratio, cay_t, where wealth consists of both human and non-human capital. Lettau and Ludvigson (2001a) find that this variable has a predictive power for asset returns of different horizons.[22] In their (2001b) paper they report that this variable forecasts portfolio returns too. Thus, they primary instrumental variables vector $z_t = (\mathbf{1}, cay_t)$. As an additional robustness check I also consider two other instrumental variables, which are "relative T-bill rate" ($RREL$, which is measured as three month Treasury-Bill rate minus its 4-quarter moving average) and the lagged value of the excess return on Standard&Poor 500 stock market index (S&P500) over the three-month Treasury bill rate. **?**, Hodrick (1992), and Lettau and Ludvigson (2001a) document that $RREL$ has a forecasting power for excess returns at a quarterly frequency. I did not include other popular forecasting variables like dividend-price ratio into the instrumental set because they are found to be driven away by the above variables cay_t, $RREL$, and $SPEX$.[23] Many previous researchers, e.g. **?**, Ferson and Constantinides (1991), to name just a few, include lagged consumption growth rate and lagged returns in the vector of instrumental variables, so I did this as well as an additional robustness check. My conclusions do not change much when $RREL$, and $SPEX$ and lagged variables are added to z_t, so I do not report these results in the paper.[24]

C *Model Estimation*

C.1 *Identification of Curvature* γ

I start empirical analysis with estimating curvature parameter γ. The estimation of this parameter is interesting by itself for the following reasons. Looking at different mixture degrees (ω) and history of consumption in habit stock (J) we learn how curvature is changing with respect to ω and J. In addition, I investigate whether moment conditions associated with long-horizon returns help to identify γ. This is important for future estimation of mixture parameter. I fix time discount factor $\rho = 0.96$, habit long run mean parameter $b = 0.492$, and mean reversion parameter $a = 0.6$. Last two values are taken from Constantinides (1990) as reference point.[25]

[INSERT TABLE II HERE ...]

Table II reports unconditional estimation results using 1-quarter (Panels A and B) and 2-year (Panels C and D) returns of Fama-French size and book-to-market sorted portfolios and short-term Treasury Bill rate, respectively. Estimation results are reported in column 2 (time separable model), column 3 (external habit), columns 4 to 6 (mixture model for varying ω) and column 7 (internal habit). Besides finding plausible values for γ, there are several interesting observations that can be made. First, γ point estimates of time separable model are 41.26 (1-quarter returns) and 26.88 (2-year returns). This is the ubiquitous manifestation of the equity premium puzzle, in which implausibly high level of risk aversion[26] is required to fit the data. Of course, as some form of time non-separability is introduced in the utility function, $\hat{\gamma}$ falls because the volatility of intertemporal marginal rate of substitution is increased through another channel: more volatile surplus consumption growth rate (as opposed to consumption growth rate in the time separable case). Second, γ and ω are negatively correlated in the model: as ω increases, γ falls. The intuition is that the volatility of stochastic discount factor can be increased through higher parameter values of either γ or ω. *Ceteris paribus*, when ω is small, little weight is placed on forward-looking terms in the marginal utility of consumption (see (14)), and the model resembles more a time separable one from the point of view of an individual consumer. This means that higher γ is required to reconcile the volatility of pricing kernel with asset returns. This mechanism is evident both when estimating γ with moment conditions using short- and long-horizon returns. Next, for fixed ω, there is a similar negative relationship between number of consumption lags J and γ. Recall that although equilibrium habit process has infinite-lag structure, marginal utility of consumption is derived on the individual level and, therefore, cuts off to finite number of forward looking terms that correspond to the number of individual consumption lags in the habit process. This means that for fixed habit and mixture levels, the volatility of pricing kernel and asset returns can be reconciled either through increasing J or γ. Although not dramatic effect, as J increases from 1 to 8,[27] γ decreases from 7.32 to 5.67: see Panels A and B, 1-quarter returns, $\omega = 0.5$ (column 5). The effect becomes more dramatic with ω getting closer to 1. Point estimates of γ drop from 4.71(3.93)[28] to 1.41(1.17) in case of 1-quarter returns and from 10.7(2.54) to 7.51(1.90) in case of 2-year returns. Finally, relatively lower standard estimates in the latter case signal that long-horizon returns better identify risk aversion than short horizon returns do.[29] I get to similar conclusions using both conditional and unconditional moment conditions with constant and cay_t taken as instrumental variables.[30]

C.2 ***Pricing Errors***

In this section, I look at the average pricing errors as a function of mixture parameter ω. **?** warrants against using J_T test as a formal comparison between different models because the weighting matrix W_T changes as model does. Indeed, low J_T might be misleading because it can indicate "improved" model fit, but be in fact the result of model estimates blowing up the estimates of W_T. For this reason it is interesting to examine the behavior of average pricing errors, too. Pricing errors are computed in the following way:[31] ϵ_i is an average pricing error for assets i, $i = 1, \ldots, N$, where N is the number of assets in a given portfolio. Portfolio pricing error ϵ_t^P is the square root of the average squared pricing errors of its components, denoted RMSE (root mean square error):

$$\epsilon_t^P = \sqrt{\frac{1}{N}\sum_{i=1}^{N}\epsilon_i^2}. \tag{24}$$

Figure 2 reports RMSE as a function of mixture parameter ω for 1 and 8 consumption lags in the individual habit stock. As in Section 4.3.1 I fix $\rho = 0.96$, $b = 0.492$ and $a = 0.6$

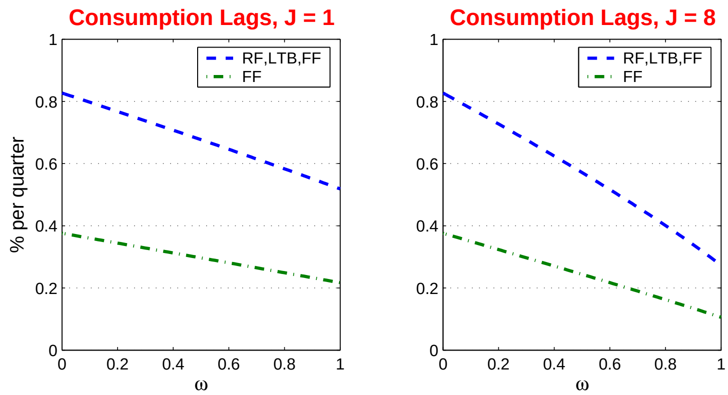

Figure 2: **Model Pricing Errors**.

and estimate γ for different values of ω, varying it from 0 to 1 with the step interval 0.01. Dash-dotted line on each panel of Figure 2 plots pricing errors from moment restrictions on Fama-French portfolio returns along with short-term risk-free rate (Set A), and dashed line corresponds to pricing errors from moment restrictions of Set A and additional moment condition for 5-year Treasury bond returns (Set B). Pricing errors are in % terms per quarter. First, pricing errors monotonically decrease as ω increases for all documented consumption lag structures.[32] Second, I observe lower portfolio pricing errors if risk-free rate and long-term bonds are excluded. It is consistent with earlier evidence that to a large extent CCAPMs with time separable and time non-separable preferences fail to account for time variation in bond returns. The inability of such models to explain the bond returns in the context of consumption based asset pricing models is explored in depth by Singleton (1993) and Heaton (1995). Both the parametric pricing models of Constantinides (1990) and Campbell and Cochrane (1999) side-step the issue of term structure dynamics by imposing restrictions so that the real term structure is constant and flat. Several papers have tried to extend habit formation models in order to accommodate stochastic interest rates.[33] For example, **?** relaxes Campbell and Cochrane (1999)'s assumption of an *iid* consumption growth rate by assuming that the expected consumption growth rate is an autonomous (latent) state variable, and shows that interest rates and risk premium driven by this new state variable have properties broadly consistent with observed bond return predictability.

Dai (2003) relaxes Constantinides (1990)'s assumption of a constant investment opportunity set by allowing the instantaneous short rate to be driven by the level of the habit stock, and shows that the time-varying risk premium implied by the model is capable of explaining the violation of the expectations puzzle. All of these extensions share the common feature that the consumption habit process is no longer locally deterministic. Finally, Dai and Grishchenko (2004) econometrically test Dai (2003)'s model using Treasury bond and broad equity market index returns. They show that stochastic internal habit formation is able to resolve the dichotomy between autocorrelation properties of stochastic discount factor and those of expected returns and provide better explanation of time-variation in expected returns compared to models with either deterministic habit or stochastic external habit. Third, in case of internal habit ($\omega = 1$) the average pricing errors fall from 0.52% to 0.27% for Set A moment conditions and from 0.22% to 0.11% for Set B as cutoff J increases from 1 to 8. Although pricing errors are always lower for Set A than those for Set B moment restrictions the difference between them shrinks with higher J. This suggests (and confirms my earlier results!) that longer-term habit formation is more consistent with time-varying properties of asset returns.

C.3 *Estimation of Preference Parameters: ρ and γ*

Next, I estimate preference parameters.[34] In doing so, I consider two extreme cases of pure external habit ($\omega = 0$) or pure internal habit ($\omega = 1$). Table III reports estimation and test results for time-discount factor ρ and curvature γ.

[INSERT TABLE III HERE ...]

Habit parameters are fixed as before. I use 8 1-quarter portfolio returns, which constitute the benchmark set: 90-day Treasury bill, 5-year maturity Government Bond and 6 Fama-French portfolios. Panel A and B report unconditional and conditional estimation results, respectively. The asymptotic standard errors ϵ_t are assumed to follow $MA(0)$ process in the case of time separable and external, and $MA(J+1)$ in case of $(J+1)$-lag internal habit and 1-period returns.[35] First, consider the case of time separable model (Column 2, Panel A). The point estimate of the time discount factor ρ is 0.95(0.13). The relative risk aversion coefficient estimate $\hat{\gamma}$ is 50.152(117.46). This is consistent with previous findings of **?**, Ferson and Harvey (1992), and Ferson and Constantinides (1991) who also find large but imprecise estimates of γ when they estimate time separable model using only unconditional moments. Consistent with many earlier studies starting from **?**, the model is rejected on the conventional levels of statistical significance: the overall goodness-of-fit value is 21.41 and right tail p-value is 0.002. The rejection of time separable model is not an artifact of seasonally adjusted data since Ferson and Harvey (1992) rejected it too both on seasonally adjusted and unadjusted consumption series. External habit estimates are reported in Column 3: because habit stock is a mere externality from the agent's point of view, the effect of today's consumption choice on future habit is ignored. In this case, intertemporal marginal rate of substitution has the same functional form as in the time separable model. However, the volatility of market price of risk is increased now through two channels: the higher value of γ and the increased volatility of surplus consumption growth because of the presence of externality. This is a reason why lower curvature estimate of a power utility function can "fit" asset returns. Indeed, the point estimate of γ drops to 34.87(12.90). With the model's fit considerably improved, the model is nonetheless rejected. This empirical result is consistent with **?**'s theoretical result that in the multiperiod economy, equilibrium asset prices and returns in an economy with externalities are identical to those of an externality-free economy, with a properly adjusted degree of risk aversion.[36] This means that simply an introduction of externalities cannot possibly account for the observed "excess volatility" in the stock prices, because changes in stochastic discount rates brought up by a modification in the risk aversion parameter fail to account for this volatility. Columns 4 to 7 of Panel A present results for internal habit formation with different degree of persistence, measured by cutoff J. Both point estimates of discount factor and curvature drop, with fit improved significantly as J increases. Internal habit is not rejected at 5% level of statistical significance if habit is formed of more than 8 quarters of lagged consumption. This preliminary estimation shows one important pattern. The importance of habit persistence kicks in only when sufficiently long history of past individual consumption is accounted for in the reference level. I call this effect *long-term* internal habit. Alternatively, I call internal habit with only a few lags (less than 4) *short-term* internal habit. Note that goodness-of-fit statistics are similar for time separable and external habit, on one hand, and short-term internal habit, on the other hand. It is consistent with Ferson and Constantinides (1991), who cannot reject the hypothesis of the time separable model in favor of time non-separable(with only one lag habit stock) one using unconditional moment restrictions. This again points out to the finding documented in Table II and Figure 1. The models with high ω and low γ, on one hand, and low ω and high γ, on the other hand, are observationally equivalent.

Of course, the problem with unconditional moments is that γ is not identified well. I address this problem by looking which instrumental variables improve the identification of γ parameter. Hence, I use $\mathbf{1}$ and cay_t[37] as instrumental variables in the conditional estimation of preference parameters. The model, described in (10), (14), (15), and (16) predicts that the time variation in the real returns, predictable by cay_t is removed when $R_{t,t+1}$ is multiplied by the stochastic discount factor.

Conditional estimation results reported in Panel B, Table III confirm Panel A unconditional results: similar curvature estimates obtained but the asymptotic standard errors are tightened as a result of the conditioning information use. Also, time separable and short-term internal habit models are rejected on the conventional statistical significance levels.[38] Internal habit persistence is not rejected for cutoffs $J \geq 4$.

Using $\mathbf{1}$, one-period lag consumption growth rate and one-period lag benchmark asset returns, I run additional robustness check that proves that long-term internal habit model explains time-varying asset returns better than time separable or external habit model. Therefore, the hypothesis of long-term internal habit persistence is economically reasonable.

C.4 *Estimation of Preference and Habit Parameters:* ρ, γ, b, and a

The next step is to jointly estimate preference parameters ρ, γ and habit parameters b, a and to examine if the same conclusion is warranted. The main, and generic problem in this kind of estimation problems is twofold: (1) habit process is very persistent and (2) it is not observable by an econometrician. Ferson and Constantinides (1991), Hansen and Jagannathan (1991), Eichenbaum and Hansen (1990), Gallant, Hansen, and Tauchen (1990), and Ferson and Harvey (1992) consider only one lag habit models. They define surplus consumption $z_t = c_t - b\, c_{t-1}$, and, accordingly, their $MUC_t = MUC_t(\rho, \gamma, b)$ is a function of three parameters only.[39] My model collapses to theirs with $a = 1$. Ferson and Constantinides find $b = 0.95$ using quarterly consumption data, but their results are mixed given different instrument sets. As an additional robustness check I estimate their model using my data set and unconditional Euler equations. I find that point estimates are quite close to theirs and the model is not rejected: $\hat{\rho} = 0.91(0.11)$, $\hat{\gamma} = 4.97(4.02)$, and $\hat{b} = 0.86(0.14)$. Preference parameters' estimates are not estimated very precisely, but $\hat{b}$ is well identified in this set-up. In my data sample instrumental variables, especially consumption-wealth ratio cay_t, help proper identification of preference parameters.

Unfortunately, the above estimates cannot be used as starting values since it requires the value of a different from 1 for the multiple-lag habit process reestimation. Therefore, I estimate preference and habit parameters in two stages. First, I estimate ρ and γ using grid search over triangular region of pairs of habit parameters: $0 < b < a < 1$. Constantinides (1990) shows that $b < r + a$ is a regularity condition for consumer's optimization problem, where r is the historical mean of the short-term interest rate. This is a necessary condition for a set of admissible policies[40] to be non-empty. My model converges to Constantinides (1990) for sufficiently large J. In this region, the above regularity condition is satisfied.

I search for the global optimum in the following way. To each pair of (b, a) there corresponds an objective function value $f = f(\hat{\rho}, \hat{\gamma}; b, a)$. By comparing these values across different (b, a) pairs I choose $(b^*, a^*) = argmin f(\hat{\rho}, \hat{\gamma}; b, a)$. I use (b^*, a^*) and corresponding $(\hat{\rho}, \hat{\gamma})$ as starting values for the joint estimation of four parameters. This insures the global minimum of the objective function.

Table IV reports results for external and internal habit with 1, 4, 8, 12 and 20 consumption lags. Here I use the same set of instruments and assets as in Table III.

[INSERT TABLE IV HERE ...]

Panel A reports unconditional point estimates, standard errors in parentheses, overall goodness-of-fit statistics, and associated $p-$values. While not significant for most of the consumption lags in this panel,[41] habit parameters are always in accord with Constantinides regularity condition and $a < 1$ for all consumption lags. The best model fit across different J is obtained for $J = 20$: $\chi^2 = 16.72$, but the model is still rejected unconditionally. Conditional estimation results in Panel B are different along several dimensions. First, for high J, the standard errors are reduced significantly, although for most of the specifications, habit parameters cannot be estimated with any adequate precision. Second, model fit improvement is monotonic as J increases. When $J = 20$, conditional habit parameters are estimated quite sharply at the expense of γ identification: $\hat{b} = 0.291(0.157)$ and $\hat{a} = 0.432(0.214)$, but $\hat{\gamma} = 7.446(6.011)$.[42] Thus, habit parameters are significantly different from zero, and second, they imply economically reasonable long-run mean of habit stock, 67%.[43]

These results show that the persistent feature of habit process still remains the achillean heel in such a type of estimations: habit parameters cannot be jointly estimated precisely at least in the current set-up. Since the central question of the study is empirical relevance of distinguishing between external and internal habit formation, I keep habit parameters fixed in the subsequent analysis.

C.5 *Estimation of Mixture Model*

In this section I estimate mixture habit formation model focusing on its most important parameter: ω. Preliminary econometric analysis for external ($\omega = 0$) and internal ($\omega = 1$) habit models yields reasonable estimates for preference and habit parameters. I use them as starting values for joint estimation of (ρ, γ, b, a, ω) or for fixing (some of the) parameters while focusing on ω. In addition, we know from previous estimations that long-horizon returns and/or long-term history of individual consumption pave a way for distinguishing between "catching up with Joneses" and internal habit formation. The next two figures visualize the importance of these factors. I plot GMM objective function (which is minimized in the estimation) as a function of γ and ω and keep other parameters fixed.[44] Formally, a GMM objective function Q is a scalar product of unweighed unconditional moment restrictions:

$$Q(\gamma, \omega; \rho^*, b^*, a^*, J^*) = m' \times m, \tag{25}$$

where $m = m(\gamma, \omega; \rho^*, b^*, a^*, J^*)$. Figure 3 demonstrates the effect of return horizon on ω identification. I plot quadratic form Q constructed out of unconditional moment restrictions using quarterly, annual, 2- and 3-year returns on 90-day T-Bill portfolio and 6 Fama-French portfolios.[45] I assume $J = 8$ in all panels of Figure 3. First, objective function associated with one quarter returns (upper-left panel) is flat meaning that γ and ω cannot be identified jointly using 1-quarter returns. As return horizon increases $Q(\gamma, \omega; \rho^*, b^*, a^*, J^*)$ becomes more convex along ω dimension, suggesting better identification properties for mixture parameter. Although their goal is different,[46] Daniel and Marshall (1999) also document that longer horizon returns capture better time non-separability in the spirit of Constantinides (1990). The reason is that habit persistence generates forward-looking intertemporal marginal rate of substitution, which is more consistent with predictable returns.[47] Predictability can be induced not only through persistence in surplus consumption ratio (like in external habit model) but also through persistence in the marginal rate of substitution. Note that objective function remains flat along γ dimension though. This means that γ and ω cannot be identified jointly using the above moment conditions. *Ceteris paribus*, γ and ω address the same properties of the intertemporal marginal rate of substitution. Holding the level of equilibrium habit process constant (which is controlled by fixed b and a), the volatility of MRS can be increased either by increasing γ or ω. The latter makes the MRS more persistent (and consequently, more volatile) by placing addi-

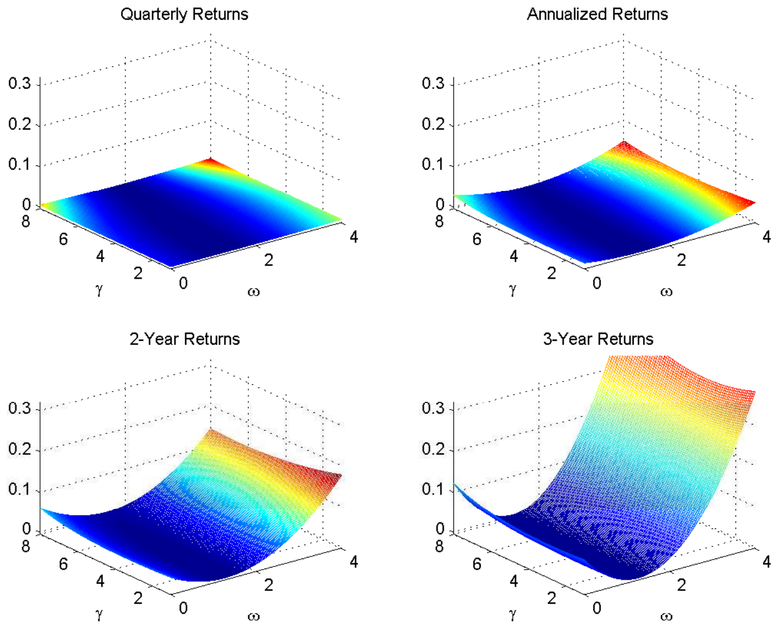

Figure 3: **Effect of Return Horizon on ω Identification.**

tional weight on the forward-looking terms in (14). The same point is made implicitly by Hansen-Jagannathan bounds' plots in Figure 1. Recall that I plot HJ bounds for external ($\omega = 0$) and internal ($\omega = 1$) cases. Higher values of the curvature parameter γ are needed for external habit stochastic discount factor to fit HJ bounds than for internal habit SDF. In other words, external habit with high γ and internal habit with low γ are observationally equivalent. Campbell and Cochrane (1999) also find that a slow moving and persistent habit stock allows to match time series properties of long horizon returns. Next, I fix 3-year return horizon and plot objective function Q given by (25), as a function of γ and ω for different cutoff values, $J = 1$ and $J = 8$, corresponding to 1-quarter and 2-year habit persistence. Two corresponding panels on Figure 4 demonstrate the effect of habit cutoff J on the identification of ω. Clearly, more persistent habit formation preferences (given by higher J) induce more convex shape of the objective function along ω direction and yield better identification properties of ω. It is interesting that optimal ω is closer to 1 for higher J, but off (roughly magnitude of 2) for low J. This can be a manifestation of the model misspecification for $J = 1$. Because only one consumption lag is used in forming habit

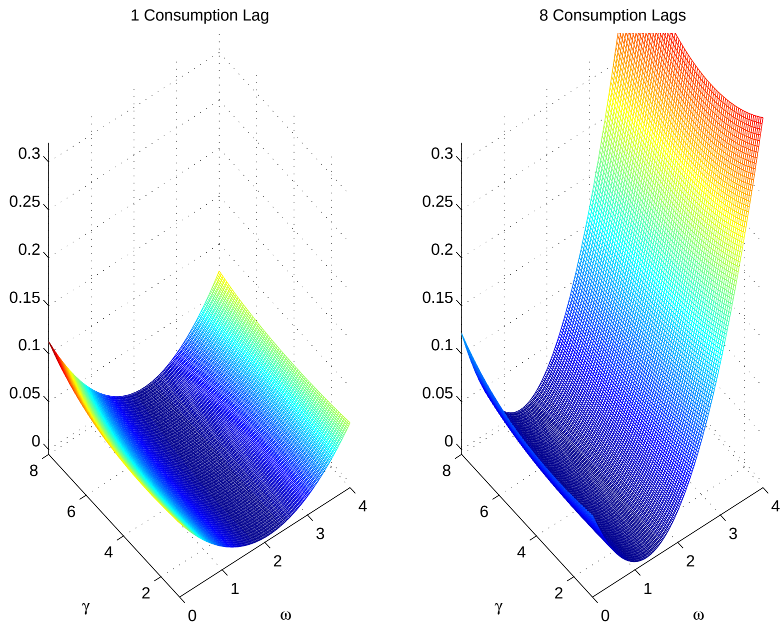

Figure 4: **Effect of Cutoff J on ω Identification.**

stock, higher ω is required make the forward-looking nature of marginal rate of substitution more important, and consequently, to "fit" the model.

Table V reports joint estimation results for γ and ω for different return horizons and cutoff values J.[48] Panels A and B report estimation results using quarterly and annual returns, respectively. Each panel presents estimation results for 1, 4 and 8 consumption lags corresponding to one quarter, one and two years in habit.

[INSERT TABLE V HERE ...]

In all cases ω is greater than 1, with fairly small standard errors, supporting intuition illustrated by Figures 3 and 4. Also, huge standard errors of γ show that curvature cannot be well identified along with the mixture parameter. The model's fit is the best for moment restrictions on annual returns and $J = 4$: $\chi^2(12) = 5.812$ with $\hat{\omega} = 1.191(0.293)$. In general, the standard errors of point estimates $\hat{\omega}$ in the case of 1-quarter returns (Panel A) are significantly larger (on average, by a factor of 2) than in the case of 1-year returns (Panel B, column (5)). This indicates that both long-horizon returns and longer-term habit formation are crucial for mixture parameter identification. Although $\hat{\omega}$ point estimates are higher than 1, they are not significantly different from 1 for long-horizon returns. This might indicate a slight model misspecification because habit and other parameters are chosen plausibly but not optimally.

Nevertheless, it is interesting to interpret the magnitude of the mixture point estimate. When $\omega > 1$, investors assign positive weight to own past consumption and negative weight to past aggregate consumption. This means that individual consumption is complimentary over time (in the long-run) and aggregate consumption is substitutable over time, which is not unreasonable. Indeed, while consumers get addicted to certain goods in the long-run, they treat aggregate consumption as a substitute good intertemporally. For example, if consumers in the economy have bought on average a lot of durable goods, like cars, in the current period, it is less likely that they will buy as many of the same kind of durable goods in the next period. It induces durability effect, which is reflected in their preferences through the magnitude of the mixture estimate.

In general, this estimation indicates that long-run habit persistence is more consistent with observed patterns in asset pricing data than either "catching up with Joneses" or short-run habit persistence.[49] However, this result is obtained by having some of the model parameters fixed.

I present full-fledged conditional estimation of the model in Table VI using quarterly (Panel A) and annual returns (Panel B) each with 1, 4, 8, 12, and 20 consumption lags. I include "long" cutoffs (J = 12, 20) to see when mean-reversion parameter a is better identified: different values of J can indicate best corresponding estimates for a. Three empirical findings emerge.

[INSERT TABLE VI HERE ...]

First, model's fit is the best when long horizon returns in the moment conditions: one year returns for cutoff $J = 20$ yield the lowest $\chi^2 = 88.50$.[50] Second, for longer cutoffs $J = 12$ and 20 habit parameters a and b are sharply estimated with a right magnitude. Thus, point estimates associated with one year returns and $J = 20$ (column 6, Panel B) are $\hat{a} = 0.76(0.12)$ and $\hat{b} = 0.66(0.13)$ and they imply long-run mean of habit stock roughly $0.86 \approx 0.66/0.76$[51] of the long-run mean aggregate consumption level. Third, mixture point estimate $\hat{\omega} = 1.03(0.16)$ is in the ball-park and also sharply estimated. There is no difference $J = 12$ and $J = 20$ in the habit stock estimates because the decay's parameter in the habit stock $(1-a)^J \approx 0$ for these values of J. However, the standard errors are smaller for higher J. This implies that autocorrelation structure of Newey-West residuals, which takes into account more lags, captures better the autocorrelation properties of long-horizon returns and internal habit persistence. Fourth, consistent with earlier evidence presented in Figures 3, 4 and Table V, I cannot identify curvature γ jointly with ω. Overall, these results are expected and they show the benefits of using long horizon returns and long-term habit persistence the estimation of habit models.

In addition, I find that 2-year holding period returns and 20 lags work the best for identifying jointly two key model parameters:[52] $\hat{\omega} = 0.782$ with standard error 0.431, and $\hat{\gamma} = 6.402$ is a reasonable estimate with standard error 2.866. However, 20 consumption lags seem to be at odds with $\hat{a} = 0.46$ because the effect of past consumption dies off much earlier with such a persistence parameter. This might be an evidence of small sample bias, too.

III Conclusion

In this paper, I have proposed a generalized asset pricing model that nests both "catching up with Joneses" (external habit formation) and "time non-separable" (internal habit formation) preference specifications. An agent forms her subsistence level based on both average and individual past consumption levels. I have empirically estimated a mixture habit parameter, which controls the degree of relevance of the individual past consumption levels as opposed to the importance of past aggregate consumption levels in the agent's preferences. In the econometric analysis, I have used seasonally adjusted detrended quarterly consumption expenditures on non-durable goods and services, short-term rate, Fama-French portfolios and Treasury long-term bond portfolio returns of different horizons. Using long horizon aggregate stock market returns I have found strong support for internal habit formation preferences, which decays slowly over time. My empirical findings suggest that such a habit persistence becomes empirically relevant only after sufficiently long history of individual consumption (8 and more consumption lags) is accounted for in the formation of habit stock and help explain why Ferson and Constantinides (1991) were unable to estimate internal habit model with small number of consumption lags. My empirical results have important implications for researchers attempting to provide microeconomic foundations of habit formation.

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Notes

[1]although the term "external habit" is widely accepted after it was first used by Campbell and Cochrane (1999), the intellectual lineage of this type of preference specification can be traced to the "catching up with the Joneses" specification of Abel (1990) and "keeping up with the Joneses" specification of Jordi Gali

[2]See also Bekaert, Engstrom, and Grenadier (2005), Wachter (2006), Menzly, Santos, and Veronezi (2004)

[3]E.g., Chetty and Szeidl (2004), Xu (2004) and Yogo (2004).

[4]With the exception of Heaton (1995), almost all existing econometrics estimations and tests of habit formation models (which are almost always based GMM) are based on small number of lags in the dependence of habit level on past consumption. The limited number of lags is motivated solely by computational feasibility, and represents a strong theoretical restriction. Using simulated method of moments, Heaton was able to side step the difficulty of multiple lags, and presented evidences that such a restriction can be soundly rejected. Later, I will elaborate further why a habit formation model with small number of lags has fundamentally different properties from a model with a large number of lags.

[5]The general form of my habit specification (3) implies that the partial derivatives of the current habit level with respect to the current consumption levels are identically zero. This restriction is imposed to rule out the possibility of "consumption bunching" behavior. **?** show that, otherwise, households would themselves choose to periodically destroy endowments. Therefore, optimally chosen consumption would not be the same as exogenously given endowment process.

[6]I considered and estimated alternative finite lag habit specification, in which I assume the agent forms own habit stock based on the same number of lags in either individual or aggregate consumption series. The empirical results are essentially identical to the present formulation because equilibrium infinite lag habit stock is well approximated by finite lag habit stock for large enough J.

[7]Here I mean aggregate consumption scaled by population. Sometimes it is referred to as average consumption. On the partial equilibrium level it is different from individual consumption.

[8]Abel has a Cobb-Douglas specification of the reference level.

[9]A key parameter that helps habit formation models resolve the equity premium puzzle is the long-run value of the surplus consumption ratio. When the equilibrium habit process is given by (9), the long-run value of the surplus consumption ratio is given by $\frac{C_t - X_t}{C_t} = 1 - \frac{b}{a} \in (0,1)$, and is independent of either J or ω. In contrast, if the habit specification (6) is not padded by the tail sum, the steady-state surplus consumption ratio is given by $\frac{C_t - X_t}{C_t} = 1 - \frac{b}{a} + \frac{b}{a} \times \left[1 - (1-a)^{J+1}\right]$. In order for models with different cut-offs J to have the same ability to explain the equity premium puzzle, the parameters a and b must be adjusted accordingly based on the values of J across different models. This is cumbersome both for interpretation and for econometric implementation.

[10]Chen and Ludvigson (2006s) consider external and internal habit models as two different models, which

are not structurally nested. Their findings suggest that internal habit formation is more consistent with aggregate stock market behavior.

[11]For detailed data description please go to section B.

[12]I have also looked how marginal rate of substitution fits into the HJ-bounds in case of internal habit formation with smaller number of lags in the habit stock and/or $\omega \in (0, 1)$. The curvature has to be quite large (on the magnitude of 20) to fit the HJ-bounds.

[13]Braun, Constantinides, and Ferson (1993) note that for the representative utility to be well-identified marginal utility should be positive with probability one. Chapman (1998) also notes that additional distributional assumptions (beyond moment restrictions) should be made in the endowment economy with habit formation or one should consistently check that marginal utility is non-negative everywhere that guarantees positive state-prices. In the empirical analysis I check consistently that at the optimum $z_t^* = c_t - x_t^*$ is always positive, giving positive marginal utility of surplus consumption.

[14]Recall that both setups are the special cases of the "mixture" habit formation model.

[15]Since all agents are identical in our model, we can speak, whenever convenient, of a "representative agent", who has the same utility specification as individual agents and who consumes, *in equilibrium*, the aggregate endowment process C_t. It is important to note that such a representative agent need not be the same as a central planner. According to some authors (e.g., Lars Ljungqvist), a central planner can not have an external habit: by definition, the central planner will take into account any effect of the aggregate consumption on future habit levels and therefore any habit is necessarily internal. The "representative agent" I have in mind here does exhibit external habit formation behavior.

[16]For expositional purposes, I often write equation (17) as (10) so long as no confusion arises.

[17]See Ferson and Constantinides (1991) and Heaton (1995) for more details.

[18]I exclude shoes and clothing from expenditures on non-durable goods because I would like to abstract from any durability effect, which is contained in these series. The exclusion of shoes and clothing follows the paper of Blinder, Grossman, and Wang (1985), p.473.

[19]Using seasonally-unadjusted data, Ferson and Harvey (1992) find that quarterly seasonality may induce "quarterly" habit persistence, in the sense that the habit level is determined by consumption lagged four quarters. I wish to abstract away from this effect.

[20]The estimation with detrended series is more stable than the one with undetrended series. In addition, model parameters in the detrended and undetrended economies can be easily transformed from one to another because of the linear habit stock.

[21]Source: Bureau of Economic Analysis (http://www.bea.gov).

[22]This variable is measured as the cointegrating residual between log consumption, log asset wealth and log labor income and called cay_t. See Lettau and Ludvigson (2001a) for more details.

[23]See Lettau and Ludvigson (2001a) for further details.

[24]However, measurement errors and other data problems can result in the spurious correlation between the consumption growth rate and asset returns and lead to the spurious rejection of the Euler equations and biased point estimates.

[25]Recall that when habit is defined as in (9), the long-run mean of the habit is given by $X_{t+1} = \frac{b}{a}C_t = 0.82C_t$. Alternatively, Cochrane and Hansen (1992) set this value at 0.5 and 0.6.

[26]Recall that in case of time separable model the curvature of the power utility function is equal to relative risk aversion coefficient.

[27]I do not consider J beyond 8 lags because the effect of past consumption dies off by that time in the case of fixed habit parameters. Recall that mean reversion parameter $a = 0.6$ implies decay rate 0.4^J, which is $0.4^8 = 0.0006$.

[28]From now on, where it is the case, point estimates are followed by standard errors in the parentheses.

[29]Intermediate case of 1-year returns falls in between for all preceding comparative analysis.

[30]Results are available upon request.

[31]The computation is in the spirit of Lettau and Ludvigson (2001b).

[32]Although I present cases with 1 and 8 consumption lags, the same evidence applies for different cutoff specification.

[33]Campbell and Cochrane (1999) show how to extend their model by relaxing a parametric restriction on the specification of the surplus consumption ratio. However, interest rates and bond risk premium in their model are perfectly correlated with consumption growth shock, which is counter-factual.

[34]Formally, all parameters are preference parameters in our model. However, where no confusion arises, I refer to parameters ρ and γ as "preference parameters", and to b and a - as "habit parameters". ω is referred to as "mixture" parameter.

[35]Eichenbaum, Hansen, and Singleton (1988) show that autocorrelation structure of pricing errors has moving average structure of the order one less the maximum number of leads in the decision variable.

[36]His definition of externality is slightly different from ours: by externality Gali means that current period consumption is valued by the consumer along with his own consumption in the utility function, thus, "keeping up with Joneses" concept.

[37]Lettau and Ludvigson (2001a) have shown that the proxy for log consumption-wealth ratio cay_t forecasts quarterly real asset and portfolio returns and drives away other popular forecasting variables like dividend-price ratio etc.

[38] External habit model is not rejected on 5% significance level.

[39]One of the reasons why Ferson and Constantinides could not estimate more than one lag habit model with any precision is that consumption expenditures are highly correlated, and therefore, it is empirically difficult to resolve the issue what is the most optimal lag structure.

[40] Admissible policies are consumption and investment policy.

[41]Note that parameters a and J are not independent of each other: higher a implies faster decay and is consistent with lower J and vice versa.

[42]I also estimated ρ, γ, b, a using constant, lagged consumption growth rate and lagged asset returns. In this case the results are mixed and unstable. However, the relative risk aversion coefficient is much lower in the internal than in external habit model.

[43] Long-run mean of habit stock is given by $\bar{x} = b\,\frac{1-(1-a)^{J+1}}{a}\,\bar{c}$. This estimate is lower than Constantinides values of long-run mean of habit stock. This might be due to higher estimate of γ, which results in a lower average of habit stock.

[44]$\rho = 0.96$, $b = 0.492$, $a = 0.6$, $J = 8$.

[45]All of the following conclusions are warranted if long-term bond moment restrictions are included. Because habit stock is linear and deterministic, these moments do not have any nontrivial implications for identification of either parameters, they just increase overall pricing errors, as illustrated on Figure 2.

[46]In particular, they examine whether the rejection of consumption-based models is due to market frictions that are more important for the short-horizon returns. They estimate separately Abel (1990) and Constantinides (1990) models.

[47]I am grateful to Martin Lettau for pointing this to me.

[48]I do not include long-term bond returns in the set of moment conditions because they do not affect ω identification.

[49]As obvious from habit process (6) construction, the ability of conditional moments to identify internal habit refers to deterministic habit formation models only.

[50]There are 13 degrees of freedom for this estimation, and so the model is rejected nevertheless.

[51]Broadly consistent with Constantinides (1990) calibration of infinite-horizon habit process.

[52]Not reported, but available upon request.

Figure Captions

Figure1. Hansen-Jagannathan Bounds. The figure presents Hansen-Jagannathan bounds and mean-volatility pairs of marginal rates of substitution in the external and internal habit cases for varying curvature parameter γ.

Figure2. Model Pricing Errors. The figure presents the model pricing errors with 1 and 8 consumption lags as a function of "mixture" parameter ω. Dashed lines correspond to moment conditions associated with risk-free rate, 5-year Government bond an Fama-French portfolios. Dot-dashed lines correspond to those associated with Fama-French portfolios only.

Figure3. Effect of Return Horizon on ω Identification. The figure plots the objective function as a function of curvature γ and "mixture" ω constructed of unconditional moments associated with 1-quarter, 1-year, 2-year and 3-year returns on 90-day T-bill portfolio and Fama-French portfolios. $J = 8$ is assumed here.

Figure 4. Effect of Cutoff J on ω Identification. The figure plots the objective function as a function of curvature γ and "mixture" ω constructed of unconditional moments associated with 1 and 8 consumption lags in the habit stock. Return horizon is 3 years. Assets used are 90-day T-bill portfolio and Fama-French portfolios.

Table I: Descriptive Statistics

	Mean	St.Dev.	Corr with Δc_t	ρ_1	ρ_2	ρ_3
Consumption Growth	0.021	0.009	1.000	0.331	0.182	0.196
Fama-French Portfolio Returns						
Big-Low	0.081	0.179	0.134	0.072	-0.060	-0.006
Big-Medium	0.087	0.148	0.126	0.081	-0.070	-0.028
Big-High	0.107	0.173	0.166	0.137	-0.051	-0.027
Small-Low	0.091	0.265	0.140	-0.024	-0.049	-0.046
Small-Medium	0.121	0.210	0.148	-0.016	-0.058	-0.058
Small-High	0.140	0.221	0.151	-0.036	-0.098	-0.061
Treasury Bond Returns						
90-day Treasury Bill	0.022	0.012	0.045	0.690	0.657	0.594
5-year Treasury Bond	0.032	0.060	-0.118	0.000	0.069	0.126
10-year Treasury Bond	0.030	0.083	-0.118	0.055	0.032	0.108

This table reports annualized means, standard deviations, correlations with real per capita consumption growth rate (Δc_t) and autocorrelations (ρ_i, $i = 1, \ldots, 5$) of real quarterly per capita consumption, real quarterly returns on Fama-French portfolios, 90-day Treasury Bill portfolio, and long-term Government Bond portfolios. Consumption is measured as expenditures of non-durable goods and services minus consumption of clothing and shoes. Classification of Fama-French portfolios is standard. For example, Big-Low portfolio stands for portfolio Big in size and Low in book-to-market value, etc. Nominal returns are converted by the growth rate of the seasonally unadjusted CPI. There are 204 observations in the sample. The data is from 1952:Q1 to 2002:Q4.

Table II: Estimation of curvature γ

	Time Separable $b \equiv 0$	External Habit	Internal Habit ω			
		0	0.25	0.5	0.75	1
(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: 1 Consumption lag, 1 quarter returns						
γ	41.257	7.555	7.506	7.317	6.689	4.711
s.e.	(18.760)	(2.627)	(2.858)	(3.177)	(3.565)	(3.931)
$\chi^2(6)$	41.23	27.25	27.46	27.79	28.29	28.85
p-value	0.000	0.011	0.011	0.010	0.008	0.007
Panel B: 8 Consumption lags, 1 quarter returns						
γ		7.555	6.329	5.673	3.958	1.413
s.e.		(2.627)	(1.602)	(1.742)	(1.728)	(1.166)
$\chi^2(6)$		27.25	15.41	15.64	15.95	15.23
p-value		0.011	0.282	0.269	0.252	0.293
Panel C: 1 Consumption lag, 2-year returns						
γ	26.883	12.270	11.998	11.664	11.244	10.700
s.e.	(10.099)	(3.042)	(2.933)	(2.811)	(2.679)	(2.541)
$\chi^2(6)$	12.37	13.20	13.24	13.27	13.29	13.29
p-value	0.054	0.040	0.039	0.039	0.039	0.039
Panel D: 8 Consumption lags, 2-year returns						
γ		12.270	11.621	10.743	9.487	7.505
s.e.		(3.042)	(2.720)	(2.442)	(2.152)	(1.902)
$\chi^2(6)$		13.20	9.20	9.17	9.04	8.64
p-value		0.040	0.163	0.164	0.171	0.195

Euler equations:

$$
\begin{aligned}
E\left[\mathrm{MRS}_{t,t+k}(1+R^f_{t,t+k})\right] &= 1, \\
E\left[\mathrm{MRS}_{t,t+k}(1+R^i_{t,t+k})\right] &= 1, \quad i = 1,\ldots,6
\end{aligned}
$$

where $R^f_{t,t+k}$ is k-period compounded three-month T-bill rate (known at t), $R^i_{t,t+k}$ is k-period quarterly holding-period returns on 6 Fama-French portfolios. Moment conditions in Panels A and B assume $k = 1$, and those in Panels C and D assume $k = 8$. Sample is from 1952:Q1 to 2002:Q4. In the time-separable model specification long-run mean habit $b \equiv 0$. In the external and internal habit parameter specification time-discount factor $\rho = 0.96$, habit parameters $b = 0.492$, $a = 0.6$, γ is the curvature parameter of power utility function. The error terms are assumed to follow $MA(k-1)$ and $MA(J+k)$ processes when external and J-lag internal habit model are estimated, respectively. Asymptotic standard errors are reported in parentheses below point estimates. $p-$value is the probability value that a χ^2 exceeds the minimized sample value of GMM criterion function. The real consumption expenditures are per capita non-durable goods and services excluding shoes and clothing and deflated using chain-type price deflator (2000=100). Identity weighting matrix. One-stage GMM estimation.

Table III: Joint estimation of preference parameters

	Time Separable $b \equiv 0$	External Habit	Internal Habit Consumption Lags			
			1	4	8	12
(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Instrumental variables: unit vector						
ρ	0.950	0.640	0.892	0.877	0.885	0.885
	(0.130)	(0.199)	(0.115)	(0.144)	(0.149)	(0.156)
γ	50.152	34.870	4.999	4.453	4.172	4.104
	(117.456)	(12.899)	(3.925)	(4.681)	(4.906)	(5.141)
$\chi^2(6)$	21.41	14.00	21.40	15.86	12.59	10.68
p-value	0.002	0.030	0.002	0.015	0.050	0.099
Panel B: Instrumental variables: $\{\mathbf{1}, cay_t\}$						
ρ	0.966	0.875	0.945	0.904	0.871	0.870
	(0.022)	(0.082)	(0.034)	(0.051)	(0.030)	(0.029)
γ	50.150	21.692	2.883	2.348	4.549	4.523
	(16.939)	(7.681)	(1.863)	(1.963)	(1.105)	(1.089)
$\chi^2(14)$	39.40	23.47	30.22	8.60	15.30	12.38
p-value	0.000	0.053	0.007	0.856	0.358	0.575

Euler equations:

$$
\begin{aligned}
E_t\left[\mathrm{MRS}_{t,t+1}(1+R^f_{t,t+1})\right] &= 1, \\
E_t\left[\mathrm{MRS}_{t,t+1}(1+R^{b,5}_{t,t+1})\right] &= 1, \\
E_t\left[\mathrm{MRS}_{t,t+1}(1+R^i_{t,t+1})\right] &= 1, \quad i = 1, \ldots, 6
\end{aligned}
$$

where $R^f_{t,t+1}$ is three-month T-bill rate (known at t), $R^{b,5}_{t,t+1}$ is quarterly holding period return on 5-year long-term bond, $R^i_{t,t+1}$ is quarterly holding-period returns on 6 Fama-French portfolios. Sample is from 1952:Q1 to 2002:Q4. In the time-separable model specification long-run habit parameter $b \equiv 0$. In the external and internal habit parameter specification $b = 0.492$, $a = 0.6$. ρ is the time-discount factor, γ is the curvature parameter of power utility function. The error terms are assumed to follow $MA(0)$ process when time-separable ($b \equiv 0$) or external habit model is estimated, and a $MA(J+1)$ process when J-lag internal habit is estimated. Asymptotic standard errors are reported in parentheses below point estimates. $p-$value is the probability value that a χ^2 exceeds the minimized sample value of GMM criterion function. The real consumption expenditures are per capita non-durable goods and services excluding shoes and clothing and deflated using chain-type price deflator (2000=100). Initial Weighting Matrix is the inverse of $Z'Z$, where Z is the vector of instrumental variables. Optimal weighting matrix is the inverse of spectral density matrix Newey-West corrected for autocorrelated residuals.

Table IV: Joint estimation of preference and habit parameters

	External Habit	Internal Habit Consumption Lags				
		1	4	8	12	20
(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Instrumental variables: unit vector						
ρ	0.948	0.976	0.981	0.939	0.939	0.972
	(0.103)	(0.035)	(0.011)	(0.087)	(0.083)	(0.046)
γ	11.027	3.955	6.099	8.464	8.460	8.459
	(263.803)	(113.358)	(53.694)	(7.670)	(8.455)	(17.370)
b	0.138	0.334	0.222	0.492	0.492	0.239
	(0.874)	(0.193)	(1.906)	(0.501)	(0.508)	(0.626)
a	0.171	0.397	0.431	0.600	0.600	0.600
	(0.657)	(1.961)	(5.911)	(0.593)	(0.603)	(0.603)
$\chi^2(4)$	26.18	72.08	43.70	26.92	38.20	16.72
p-value	0.000	0.000	0.000	0.000	0.000	0.002
Panel B: Instrumental variables: $\{\mathbf{1}, cay_t\}$						
ρ	0.978	0.981	0.979	0.977	0.979	0.977
	(0.021)	(0.007)	(0.013)	(0.013)	(0.008)	(0.008)
γ	11.027	4.252	5.332	7.648	7.457	7.446
	(30.804)	(8.371)	(16.814)	(6.465)	(6.550)	(6.011)
b	0.371	0.220	0.287	0.291	0.277	0.291
	(1.952)	(0.165)	(0.165)	(0.433)	(0.237)	(0.157)
a	0.649	0.435	0.399	0.432	0.447	0.432
	(2.873)	(2.558)	(1.013)	(0.381)	(0.483)	(0.214)
$\chi^2(12)$	232.64	315.78	271.69	213.25	116.04	76.89
p-value	0.000	0.000	0.000	0.000	0.000	0.000

Euler equations:

$$
\begin{aligned}
E_t\left[\text{MRS}_{t,t+1}(1+R^f_{t,t+1})\right] &= 1, \\
E_t\left[\text{MRS}_{t,t+1}(1+R^{b,5}_{t,t+1})\right] &= 1, \\
E_t\left[\text{MRS}_{t,t+1}(1+R^i_{t,t+1})\right] &= 1, \quad i = 1, \dots, 6
\end{aligned}
$$

where $R^f_{t,t+1}$ is three-month T-bill rate (known at t), $R^{b,5}_{t,t+1}$ is quarterly holding period return on 5-year long-term bond, $R^i_{t,t+1}$ is quarterly holding-period returns on 6 Fama-French portfolios. Sample is from 1952:Q1 to 2002:Q4. ρ is the time-discount factor, γ is curvature parameter of power utility function, b is long-run mean of habit stock, a is habit mean-reversion parameter. The error terms are assumed to be $MA(0)$ for external habit and $MA(J+1)$ for J-lag internal habit. Asymptotic standard errors are reported in parentheses below point estimates. $p-$value is the probability value that a χ^2 exceeds the minimized sample value of GMM criterion function. The real consumption expenditures are per capita non-durable goods and services excluding shoes and clothing and deflated using chain-type price deflator (2000=100). Initial Weighting Matrix is the inverse of $Z'Z$, where Z is the vector of instrumental variables. Optimal weighting matrix is the inverse of spectral density matrix Newey-West corrected for autocorrelated residuals.

Table V: Joint Estimation of curvature γ and mixture ω

Lag	$\hat{\gamma}$	s.e.	$\hat{\omega}$	s.e.	$\chi^2(12)$	p-val	RMSE
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
			Panel A: 1-quarter return				
1	4.593	(6.155)	1.306	(0.664)	6.806	0.870	0.001
4	4.555	(5.112)	1.128	(0.549)	7.491	0.824	0.001
8	4.550	(3.859)	1.096	(0.386)	13.828	0.312	0.001
			Panel B: annual returns				
1	4.584	(4.975)	1.388	(0.307)	22.471	0.033	0.002
4	4.556	(4.646)	1.191	(0.293)	5.812	0.925	0.003
8	4.552	(4.479)	1.161	(0.268)	8.414	0.752	0.003

Euler equations:

$$\begin{aligned} E_t\left[\mathrm{MRS}_{t,t+k}(1+R^f_{t,t+k})\right] &= 1, \\ E_t\left[\mathrm{MRS}_{t,t+k}(1+R^i_{t,t+k})\right] &= 1, \quad i=1,\ldots,6 \end{aligned}$$

where $R^f_{t,t+k}$ is k-period compounded three-month T-bill rate (known at t), $R^i_{t,t+k}$ is k-period quarterly holding-period returns on 6 Fama-French portfolios. Sample is from 1952:Q1 to 2002:Q4. Time-discount factor $\rho = 0.96$, habit parameters $b = 0.492$, $a = 0.6$, γ is the curvature parameter of power utility function. The error terms are assumed to follow a $MA(J+k)$ process when J-lag habit is estimated. Asymptotic standard errors are reported in parentheses next to point estimates. $p-$value is the probability value that a χ^2 exceeds the minimized sample value of GMM criterion function. The real consumption expenditures are per capita non-durable goods and services excluding shoes and clothing and deflated using chain-type price deflator (2000=100). Initial Weighting Matrix is the inverse of $Z'Z$, where $Z = (\mathbf{1},\, cay_t)$ is the vector of instrumental variables. Optimal weighting matrix is the inverse of spectral density matrix corrected for autocorrelated residuals. RMSE stands for the square root of the average squared pricing errors of moment conditions.

Table VI: Estimation of mixture model

	Consumption Lags				
	1	4	8	12	20
(1)	(2)	(3)	(4)	(5)	(6)
			Panel A: 1-quarter returns		
ρ	0.956	0.923	0.967	0.925	0.920
	(0.036)	(0.062)	(0.020)	(0.041)	(0.037)
γ	6.343	6.343	6.342	6.343	6.343
	(34.092)	(20.015)	(16.182)	(13.668)	(11.986)
b	0.651	0.650	0.616	0.651	0.650
	(0.710)	(0.302)	(0.451)	(0.250)	(0.182)
a	0.791	0.770	0.826	0.770	0.770
	(0.720)	(0.344)	(0.383)	(0.230)	(0.157)
ω	1.022	1.023	1.038	1.023	1.022
	(1.129)	(0.567)	(0.848)	(0.380)	(0.302)
$\chi^2(13)$	513.4941	832.1697	1624.5995	459.2633	287.9532
p-value	0.000	0.000	0.000	0.000	0.000
			Panel B: 4-quarter returns		
ρ	0.970	0.671	0.785	0.937	0.935
	(0.016)	(0.119)	(0.133)	(0.025)	(0.021)
γ	6.343	6.257	6.262	6.343	6.343
	(10.037)	(5.031)	(3.720)	(8.611)	(8.887)
b	0.656	0.956	0.852	0.661	0.661
	(1.113)	(0.560)	(0.407)	(0.145)	(0.134)
a	0.793	1.823	1.843	0.760	0.760
	(1.357)	(0.105)	(0.103)	(0.135)	(0.121)
ω	1.029	2.650	2.562	1.031	1.030
	(1.698)	(1.450)	(1.081)	(0.175)	(0.162)
$\chi^2(13)$	724.3351	175.5248	389.6149	410.4004	88.5030
p-value	0.000	0.000	0.000	0.000	0.000

Euler equations:

$$\begin{aligned} E_t\left[\mathrm{MRS}_{t,t+k}(1+R^f_{t,t+k})\right] &= 1, \\ E_t\left[\mathrm{MRS}_{t,t+k}(1+R^{b,n}_{t,t+k})\right] &= 1, \quad n = 5,\ 10 \\ E_t\left[\mathrm{MRS}_{t,t+k}(1+R^{i}_{t,t+k})\right] &= 1, \quad i = 1, \ldots, 6 \end{aligned}$$

where $R^f_{t,t+k}$ is three-month T-bill rate (known at t) compounded for k quarters, $R^{b,n}_{t,t+k}$ is k-quarter holding period return on n-year long-term bond, $R^i_{t,t+k}$ is k-quarter holding-period returns on 6 Fama-French portfolios. Sample is from 1952:Q1 to 2002:Q4. ρ is time-discount factor, γ - curvature parameter of power utility function, b is long-run mean of habit stock, a is mean-reversion parameter of habit stock, ω is mixture parameter. The error terms are assumed to follow $MA(J+n)$ process when J-lag model is estimated using $n-$period returns. Number of Lags is equal to $J-1$. Asymptotic standard errors are reported in parentheses below point estimates. $p-$value is the probability values that a χ^2 exceeds the minimized sample value of GMM criterion function. The real consumption expenditures are per capita non-durable goods and services excluding shoes and clothing and deflated using chain-type price deflator (2000=100). One-stage GMM estimation. Instrumental variables are unit vector and cay_t.