Consumption-Based Asset Pricing: Durable Goods, Adjustment Costs, and Aggregation

Abstract

In this paper, we investigate the implications of non-separable preferences over durable and nondurable consumption for asset pricing tests when adjusting durable consumption is costly. In an economy without adjustment costs, in which a frictionless rental market exists for the durable good, the standard Euler equation with respect to nondurable consumption will hold for each individual agent as well as for aggregate data. If the adjustment of the durable good is costly, however, aggregation generally fails. We use aggregate data to find substantial deviations from the frictionless model, consistent with the presence of non-convex adjustment costs for the durable good. We also show how empirical asset pricing tests that use aggregate data can be affected by these deviations. We then propose and implement asset pricing tests that are robust to the presence of adjustment costs by relying on microeconomic data. Using household-level observations of nondurable food and durable housing consumption, our estimation results suggest that preferences are indeed non-separable in the two consumption goods and that reasonable structural parameters characterize agents' intertemporal utility optimizations.

1 Introduction

Consumption-based asset pricing models, despite their many empirical failures, have been at the heart of asset pricing research for over two decades. Naturally, the starting point of these models is a particular assumption on what consumption is and here researchers have, for the most part, restricted themselves to measuring consumption as consumption of nondurable goods and services. Recently though, there has been renewed interest in an old topic, namely the role of durable consumption in asset pricing tests.[1] Introducing durable consumption into the agent's utility function brings new modelling issues to the asset pricing literature. One such issue is whether changing durable consumption is subject to adjustment costs and how these adjustment costs potentially affect asset prices. This question is particularly important if one uses aggregate data to test asset pricing implications. On the one hand, standard aggregation arguments will generally not hold if individual agents are subject to adjustment costs. Asset pricing tests using aggregate data therefore risk being misspecified if they do not account for consumption heterogeneity across individual agents. On the other hand, however, aggregation introduces smoothness that could make infrequent adjustments at the micro level irrelevant for understanding aggregate phenomena. In other words, macroeconomic data might behave as if generated by a representative agent in a frictionless economy. In this paper we therefore investigate whether asset pricing tests that use aggregate data can safely ignore adjustment costs, either because they are irrelevant at the individual level or, more likely, because frictions at the micro level do not affect the relationship between aggregate data and asset returns.

Our tests rely on the fact that in a frictionless economy with a rental market for the durable good, the ratio of marginal utilities with respect to the two consumption goods equals the relative user cost at any point in time. By analyzing the implied cointegration relationship between durable and nondurable consumption and the relative user cost for aggregate data, we are able to determine whether the data are consistent with the assumption

[1]See Dunn and Singleton (1986) and Eichenbaum and Hansen (1990) for early work on this topic. See Piazessi, Schneider, and Tuzel (2006), Pakos (2003), Yogo (2006) and Lustig and Van Nieuwerburgh (2005) for recent work.

of a representative agent that can adjust his durable consumption costlessly. Our empirical results suggest that this is not the case. Moreover, we find that deviations from the cointegration relationship contain information about asset returns, indicating that the standard Euler equation with respect to nondurable consumption does not hold in aggregate data. At the same time, though, the Euler equation still has to hold at the individual level, even if non-convex adjustment costs with respect to the durable good are present. In the second part of the paper, we therefore use micro-level data to evaluate a model with non-separable preferences and adjustment costs. Unlike models estimated from aggregate data, we obtain reasonable structural estimates. The variability of the implied pricing kernel in our model is, however, still too low to to be reconciled with the observed US equity premium.

Early empirical work on aggregate data by Dunn and Singleton (1986) and Eichenbaum and Hansen (1990) finds that non-separable preferences over durable and nondurable consumption do not significantly improve asset pricing models based on nondurable consumption only. Piazessi, Schneider, and Tuzel (2006), however, show that a representative agent with CES preferences over nondurable and housing service consumption implies an Euler equation that after linearization is consistent with observed equity returns. Focusing on services from durable goods such as vehicles and furniture, Yogo (2006) finds that such a model is able to price the cross-section of equity excess returns. Finally, Lustig and Van Nieuwerburgh (2005) study a model with bankruptcy similar to the one in Piazzesi et. al. (2006), where the housing good serves as collateral. They find that the aggregate housing stock to income ratio helps to explain time series as well as cross-sectional variation of equity returns. All these papers assume that there are no frictions with respect to the durable good and that standard first order conditions can be tested on aggregate data.

There is substantial empirical evidence in the economics literature, however, that durable goods are subject to non-convex adjustment costs. Eberly (1994) finds that expenditures on private vehicles are consistent with models of non-convex adjustments costs. Similarly, Caballero (1993), Carroll and Dunn (1997), and Attanasio (2000) provide evidence that the time series dynamics of aggregate durable purchases can better be explained by models in which individuals face non-convex adjustment costs than by frictionless models. Kullmann and Siegel (2005) study homeowners' portfolio choices and find them to be consistent with state-dependent risk aversion, as implied by a model with non-convex adjustment costs with respect to residential housing. Grossman and Laroque (1990), Beaulieu (1992), and Damgaard, Fuglsbjerg, and Munk (2000) study partial equilibrium models in which an individual consumer faces non-convex costs for adjusting the size of the durable good. Continuous adjustment of the durable good is infinitely expensive, resulting in infrequent adjustments of the durable good stock. In this case, the ratio of marginal utilities of the two goods is no longer equated to the relative user cost at every moment. This deviation from the frictionless model provides the motivation for the empirical tests in the first part of this paper.

Marshall and Parekh (1999) explore the consequences of adjustment costs for asset prices when preferences are separable across different consumption goods. They find that adjustment costs with respect to nondurable consumption can partially reconcile nondurable consumption data with the observed equity premium. However, several empirical papers suggest that the separability between durable and nondurable consumption is difficult to maintain. Eichenbaum and Hansen's (1990) analysis with macroeconomic data supports the assumption that preferences over nondurable and durable good consumption are not separable. Using household-level data to explicitly account for non-convex adjustment costs, Martin (2001) also finds support for non-separability. If preferences are non-separable, adjustment costs for the durable good have a direct impact on the marginal utility with respect to both durable and nondurable consumption. When different agents adjust their durable stock at different points in time, we can no longer aggregate across individuals to use a representative agent framework. That is, even though the Euler equation with respect to nondurable consumption will continue to hold for every individual agent, it will generally not hold for aggregate consumption data. This motivates the use of micro-level data to perform asset pricing tests in the second part of the paper.

Household-level data have been used in recent studies to explore market incompleteness and idiosyncratic risks when preferences are separable in the two consumption goods. Among those, Jacobs (1999) is most related to our estimation approach. In the permanent income hypothesis literature, a few panel data studies have included durable goods in their tests.

Hayashi (1985) and Padula (1999) find that the inclusion of durable goods improves the performance of their models. Following the tradition of the saving-consumption literature, however, they focus exclusively on the Euler equation with respect to the risk-free asset. Since they do not assume a specific form for their utility function, they furthermore do not provide estimates of structural preference parameters. In a current research paper, Flavin and Nakagawa (2004) use a micro-level data set similar to ours to study non-separable preferences in a somewhat different setting. We discuss their findings together with our own results at the end of the second part of the paper.

The remainder of this paper is organized as follows. The next section presents the underlying model for an economy with and without adjustment costs and discusses the consequences for aggregation and empirical asset pricing tests. In Section 3, we test for the presence of adjustment costs using macroeconomic data. To account for the existence of adjustment costs, Section 4 presents asset pricing tests performed on household level data. Section 5 concludes.

2 Model

We first describe a simple frictionless economy with many agents that differ only in their relative endowments of the durable good. We show that if the durable good can be adjusted costlessly, this initial heterogeneity is without consequence: Aggregation across agents is straightforward and empirical asset pricing tests can be carried out on macroeconomic consumption data. We then introduce non-convex adjustment costs with respect to the durable good to show that aggregation across agents is no longer trivial.

2.1 The frictionless economy

To be specific, suppose N agents, subscripted $i = 1, ..., N$, have identical preferences that are non-separable in a nondurable consumption good $C_{i,t}$ and the service flow $S_{i,t}$ from a durable good $D_{i,t}$. For simplicity, we assume the service flow to be proportional to the size of the durable good, i.e. $S_{i,t} = \phi D_{i,t}$:

$$U(C_{i,t}, D_{i,t}) = \frac{u(C_{i,t}, D_{i,t})^{1-\gamma}}{1-\gamma} \tag{1}$$

For the intraperiod utility $u(C_{i,t}, D_{i,t})$, we assume CES preferences:

$$u(C_{i,t}, D_{i,t}) = \left(C_{i,t}^{\omega} + \alpha(\phi D_{i,t})^{\omega}\right)^{\frac{1}{\omega}} \tag{2}$$

Note that this includes the special case of Cobb-Douglas preferences, for which the elasticity of substitution between the two goods, $\frac{1}{1-\omega}$, is equal to one:

$$u(C_{i,t}, D_{i,t}) = C_{i,t}^{\frac{1}{1+\alpha}}(\phi D_{i,t})^{\frac{\alpha}{1+\alpha}} \tag{3}$$

Assume that at time $t = 0$, each agent in the economy is endowed with initial positive holdings of L risky financial assets, $A_{i,0}$, with gross return, R_t and with a positive amount of the durable good, $D_{i,0}$. $A_{i,t}$ and R_t are expressed in units of the nondurable consumption good, $C_{i,t}$. P_t is the relative price of the durable consumption good. While agents are identical with respect to their preferences and beliefs, they differ in their initial endowment of the durable good relative to their initial wealth $W_{i,0}$, where $W_{i,0} = \imath' A_{i,0} + P_0 D_{i,0}$ and $\imath$ denotes a vector of ones.

Assume that the durable good depreciates each period by $(1-\delta)D_{i,t-1}$ and that an agent can adjust the size of the durable consumption good only by selling the existing durable good and buying a new one. For now though, assume that this transaction is costless.

Each agent chooses the sequence of nondurable consumption, the size of the durable good, as well as the holdings of the available financial assets that solve the following problem:

$$\max_{\{C_{i,t}, D_{i,t}, A_{i,t}\}_{t=1}^{\infty}} E_0 \left[\sum_{t=1}^{\infty} \beta^t U(C_{i,t}, D_{i,t})\right] \tag{4}$$

subject to the constraint:

$$C_{i,t} + P_t D_{i,t} + \imath' A_{i,t} \leq R_t' A_{i,t-1} + P_t \delta D_{i,t-1} \tag{5}$$

From the Lagrangian

$$L = E_0 \left[\sum_{t=1}^{\infty} (\beta^t U(C_{i,t}, D_{i,t}) - \mu_{i,t}\left(C_{i,t} + P_t D_{i,t} + \imath' A_{i,t} - R_t' A_{i,t-1} - P_t \delta D_{i,t-1}\right))\right] \tag{6}$$

we obtain the following first order conditions:

$$\beta^t U_{Ci,t} = \mu_{i,t} \tag{7}$$

$$\beta^t U_{Di,t} = P_t \mu_{i,t} - E_t \left[P_{t+1} \delta \mu_{i,t+1} \right] \tag{8}$$

$$\mu_{i,t} \imath = E_t \left[\mu_{i,t+1} R_{t+1} \right] \tag{9}$$

In equilibrium, prices must be such that the standard stochastic Euler equation with respect to nondurable consumption holds for each agent i:

$$E_t \left[\beta \frac{U_{C_{i,t+1}}}{U_{C_{i,t}}} R_{t+1} \right] = \imath \tag{10}$$

Similarly, for each agent the ratio of marginal utilities, $\frac{U_{D_{i,t}}}{U_{C_{i,t}}}$, will equal the period's user cost for the durable good in terms of the non-durable consumption good (the numeraire). Note that we assume that the user cost Q_t is the same across agents or equivalently that the frictionless economy is identical to one with a perfect rental market for the durable good:

$$\frac{U_{D_{i,t}}}{U_{C_{i,t}}} = P_t - \delta \beta E_t \left[\frac{U_{C_{i,t+1}}}{U_{C_{i,t}}} P_{t+1} \right] \equiv Q_t \tag{11}$$

Under the assumptions of CES or Cobb-Douglas utility and homogenous beliefs, standard aggregation theorems for effectively complete markets apply (see, for example, Rubinstein (1974), or Brennan and Kraus (1978)). Intuitively, when markets open, agents will undo the heterogeneity in the initial relative durable holdings by adjusting the size of the durable good so that (11) holds. Across all agents, nondurable and durable consumption will then evolve in a perfectly synchronous way. Hence equations (10) and (11) will hold for aggregate consumption data $C_t = \sum_i C_{i,t}$ and $D_t = \sum_i D_{i,t}$ as well.

2.2 Introducing adjustment costs

Now assume that adjusting the size of the durable good requires the payment of an adjustment cost $\Lambda_{i,t}$ that is proportional to the size of the durable good at the beginning of the period, i. e. $\Lambda_{i,t} = \lambda \delta P_t D_{i,t-1}$.

When solving (4), the agent now faces a case dependent constraint set. If the durable stock is not adjusted at t, the constraints are:

$$\begin{aligned} C_{i,t} + \imath' A_{i,t} &\leq R_t' A_{i,t-1} && (12) \\ D_{i,t} &= \delta D_{i,t-1} && (13) \end{aligned}$$

If the durable stock is adjusted at t, then the constraint is:

$$C_{i,t} + P_t D_{i,t} + \imath' A_{i,t} \leq R_t' A_{i,t-1} + (1-\lambda)\delta P_t D_{i,t-1} \tag{14}$$

For this problem, we can write the Lagrangian as:

$$\begin{aligned} L = E_0 \Bigg[\sum_{t=1}^{\infty} (\beta^t U(C_{i,t}, D_{i,t}) - \mu_{i,t}^{NA} \left(C_{i,t} + \imath' A_{i,t} - R_t' A_{i,t-1} \right) - \xi_{i,t}^{NA} \left(D_{i,t} - \delta D_{i,t-1} \right) \\ - \mu_{i,t}^{A} \left(C_{i,t} + P_t D_{i,t} + \imath' A_{i,t} - R_t' A_{i,t-1} - (1-\lambda)\delta P_t D_{i,t-1} \right)) \Bigg] \end{aligned} \tag{15}$$

where $\mu_{i,t}^{NA}$ and $\xi_{i,t}^{NA}$ are zero if the durable good is adjusted and $\mu_{i,t}^{A}$ is zero if the durable good is not adjusted. The first order conditions are:

$$\begin{aligned} \beta^t U_{Ci,t} &= \mu_{i,t}^{NA} + \mu_{i,t}^{A} && (16) \\ \beta^t U_{Di,t} &= \xi_{i,t}^{NA} - E_t \left[\xi_{i,t+1}^{NA} \delta \right] + \mu_{i,t}^{A} P_t - E_t \left[\mu_{i,t+1}^{A} P_{t+1} \delta (1-\lambda) \right] && (17) \\ \mu_{i,t}^{NA} \imath + \mu_{i,t}^{A} \imath &= E_t \left[R_{t+1} (\mu_{t+1}^{NA} + \mu_{t+1}^{A}) \right] && (18) \end{aligned}$$

(16) and (18) imply that the standard Euler equation with respect to nondurable consumption (10) still holds for each agent i whether the durable good is adjusted or not. Notice, though, that (11) only holds if the durable good were to be adjusted each period.

Due to the nature of the adjustment costs, frequent adjustment of the size of the durable good is not be optimal. Grossman and Laroque (1990) were the first to show, in a simpler framework with the durable good as the only consumption good, that the agent will adopt an (s, S) policy, in which adjustments of the durable good only occur when its actual size deviates from the optimal target size far enough that paying the adjustment cost is offset by a sufficient increase in utility due to the better alignment between total wealth and the durable good. This result has subsequently been extended to the case of non-separable preferences considered here by Beaulieu (1992) and Damgaard, Fuglsbjerg, and Munk (2003). Within the no-adjustment boundaries, the size of the durable good changes only due to depreciation. Due to the non-separability of preferences, both the agent's relative risk aversion and his nondurable consumption growth are a function of his position between the (s, S) adjustment points and therefore vary endogenously over time.[2]

Given the initial distribution of relative durable holdings across agents, only agents whose initial holdings are outside of the (s, S) boundaries that describe the non-adjustment region will adjust their durable stock to the same target level and from thereon behave in a perfectly synchronous way. For the remaining agents, adjustments will occur at different times and the ratios $\frac{C_{i,t+1}}{C_{i,t}}$, $\frac{D_{i,t+1}}{D_{i,t}}$, and $\frac{D_{i,t}}{C_{i,t}}$ will remain different across agents.[3] Consequently, the Euler equation, while valid for each individual agent, does not generally hold in aggregate data.

3 Macro Evidence

3.1 Testing implications of the frictionless economy

In the absence of adjustment costs, all agents equate the ratio of their marginal utilities with respect to nondurable and durable consumption to the relative user cost Q_t. Hence in equilibrium, equation (11) will have to hold exactly for aggregate data:

$$\frac{U_{D_t}}{U_{C_t}} = P_t - \delta\beta E_t\left[\frac{U_{C_{t+1}}}{U_{C_t}}P_{t+1}\right] = P_t\left[1 - \delta\beta E_t\left[\frac{U_{C_{t+1}}}{U_{C_t}}\frac{P_{t+1}}{P_t}\right]\right] \equiv Q_t \tag{19}$$

Note that the absence of adjustment costs is a sufficient, but not a necessary condition for (19) to hold. Even in the presence of adjustment costs at the micro level, aggregate data *could* behave as if generated by a representative agent in a frictionless economy.

[2]Note that the existing solutions solve partial equilibrium problems, in which all conditional moments of the return and price processes are assumed to be constant. This assumption is unlikely to be consistent with the general equilibrium outcome. The general equilibrium of economies with adjustment costs has been studied by a few authors (e.g. Detemple and Giannikos (1996) and Mamaysky (2001)), but not with heterogeneous agents.

[3]For the cross-sectional distribution not to collapse as $t \to \infty$, a source of idiosyncratic risk in addition to aggregate risk is needed (see Caballero (1993) for a detailed discussion).

In the case of CES preferences with $\frac{U_{D_t}}{U_{C_t}} = \alpha\phi^{\omega}\left(\frac{D_t}{C_t}\right)^{\omega-1}$, (19) implies:

$$
\begin{aligned}
lnD_t &= \frac{-ln\alpha - \omega ln\phi}{\omega - 1} + lnC_t + \frac{1}{\omega - 1} lnQ_t && (20) \\
&= [\iota, lnC_t, lnQ_t]\,\kappa && (21)
\end{aligned}
$$

where $\kappa = [\frac{-ln\alpha-\omega ln\phi}{\omega-1}, 1, \frac{1}{\omega-1}]'$. Equation (20) implies that the log of the durable good, the log of nondurable consumption and the log of the relative user cost of the durable good are cointegrated, stochastically *as well as* deterministically.[4] Furthermore, since (20) is an equality, any deviation from this relationship should have no economic content. Finding a cointegration relationship is necessary, but not sufficient evidence for the frictionless model. Even in an economy with adjustment costs, durable and nondurable consumption will have to reflect the evolution of the relative user cost *in the long run.* Therefore, important test is to examine whether deviations from (20) display an economically significant pattern or not.

Our empirical tests first examine whether the cointegration relationship indeed holds. We use Park 's (1992) Canonical Cointegration Regression (CCR) tests to examine whether both a stochastic and a deterministic cointegration relationship exist. The test is performed under the null hypothesis of cointegration and tests whether spurious linear as well as nonlinear deterministic trends have jointly insignificant coefficients. Let $H(l, m)$ denote a Wald test of the null hypothesis that $\eta_{l+1} = \eta_{l+2} = ... = \eta_m = 0$, where the cointegration regression is:[5]

$$lnD_t = [\iota, lnC_t, lnQ_t]\,\kappa + \sum_{j=1}^{m} \eta_j t^j + res_t \qquad (24)$$

$H(0, 1)$ is then a χ^2_1 distributed statistic for a test of deterministic cointegration and $H(1, m)$ are statistics for tests of stochastic integration that follow a χ^2_{m-1} distribution.

[4]Assume X_t and Y_t are I(1) processes with nonzero drift:

$$
\begin{aligned}
X_t &= X_{t-1} + \pi_X + \nu_{X,t} && (22) \\
Y_t &= Y_{t-1} + \pi_Y + \nu_{Y,t} && (23)
\end{aligned}
$$

where $\nu_{X,t}$ and $\nu_{Y,t}$ are stationary. X_t and Y_t are *stochastically* cointegrated with cointegrating vector β if $X_t - \beta Y_t - \pi t$ is stationary. If $\pi = \pi_X - \pi_Y\beta = 0$, then X_t and Y_t are also *deterministically* cointegrated.

[5]Both the dependent and explanatory variables are transformed first to control for endogeneity. The transformed regression can then be estimated by OLS. See Park (1992) for details.

Next, we use Stock and Watson's (1993) Dynamic Least Square (DLS) method to estimate the cointegrating vector κ'. DLS estimation includes leads and lags of the right hand side variables and has been shown to possess better small sample properties than OLS (see for example Caballero (1994)). Using this cointegrating vector, we derive the cointegration residual res_t as:

$$res_t = lnD_t - [\iota, lnC_t, lnQ_t]\,\kappa \tag{25}$$

Finally, we test whether the cointegration residual is white noise or whether it contains, contrary to the assumptions of the frictionless model, *but* consistent with the presence of adjustment costs, relevant economic information. To do this, we proceed in two steps. First, we investigate whether the cointegration residual has statistical power in predicting the evolution of the economy. We will regress y_{t+k}, for example the cumulative growth rate of durable consumption over k periods, on the cointegration residual:

$$y_{t+k} = c_0 + c_{res} res_t + \varepsilon_{t+k} \tag{26}$$

According to the frictionless model, c_{res} should be insignificantly different from zero for any y and any horizon k.

In a second step, we include changes in the cointegration residual in standard cross-sectional asset pricing tests of the form:

$$E_t\left[\beta\frac{U_{C_{t+1}}}{U_{C_t}}R_{t+1}\right] = \imath \tag{27}$$

where we use a linear approximation of the pricing kernel $M_{t+1} = \beta\frac{U_{C_{t+1}}}{U_{C_t}}$:[6]

$$\beta\frac{U_{C_{t+1}}}{U_{C_t}} \approx 1 + ln\beta + b2\Delta lnC_{t+1} + b3\Delta lnD_{t+1} \tag{30}$$

[6]For CES preferences, the approximation is around $\omega = 0$. $b2$ and $b3$ are defined as:

$$b2 = \frac{1-\gamma-\omega}{1+\alpha\phi^{\omega}} - 1 + \omega \tag{28}$$

$$b3 = \frac{\alpha\phi^{\omega}(1-\gamma-\omega)}{1+\alpha\phi^{\omega}} \tag{29}$$

Yogo (2006) employs (30) to test the frictionless model on aggregate data. Using (20), we can substitute out ΔlnD_t to obtain the following specification:

$$\beta\frac{U_{C_{t+1}}}{U_{C_t}} \approx 1 + ln\beta + (b2 + b3)\Delta lnC_{t+1} - b3\Delta lnQ_{t+1} + b4\Delta res_{t+1} \tag{31}$$

Under the assumptions of the frictionless model, any economically relevant variation of the durable growth factor, ΔlnD_{t+1}, should be entirely explained by the variation of ΔlnC_{t+1} and ΔlnQ_{t+1}. Δres_{t+1} should be an additional factor that does not help to price the cross-section of returns. Therefore, evidence that changes in the cointegration residual help explain the cross-section of asset returns would provide evidence against the frictionless model.

3.2 Data

3.2.1 Consumption data

We will perform the outlined tests of the frictionless model on two separate data sets.

In our primary data set, we measure nondurable consumption C as the sum of real nondurables and service expenditures as reported by the Bureau of Economic Analysis (BEA) in the National Income and Product Accounts (NIPA). We construct the evolution of the durable stock D by iterating forward the initial real value of the durable stock reported by the BEA, by adding the real durable expenditure and subtracting a proportional depreciation amount. Like Ogaki and Reinhart (1998), we assume a quarterly depreciation rate of 6%. The flow of nondurable consumption and the stock of durables are expressed as *per capita* values, where the size of the US population is also from the BEA. As we do not observe the relative user cost Q directly, we will have to use a transformation of the relative price P (see details below). P is constructed by dividing the implicit deflator for durable expenditures by the implicit deflator for nondurable and service expenditures. The data set covers 1952:I and 2002:IV at a quarterly frequency.

Durable goods considered here include mainly vehicles (on average 44% of the nominal quarterly expenditure) and furniture (on average 40% of the nominal quarterly expenditure). Both goods are potentially subject to non-convex adjustment costs, especially when we also consider search times and associated opportunity costs. Using microeconomic data, Eberly (1994) and Attanasio (2000) provide evidence of the relevance of adjustment costs for expenditures on private automobiles. Residential real estate transactions (for owners as well as for renters) are even more likely subject to economically relevant adjustment costs.

In our second data set, we measure the service flow S as the real housing service consumption, which is imputed by the Bureau of Economic Analysis and reflects the service flow from the underlying residential real estate asset representing the durable good in this data set.[7] We define nondurable consumption C as the remaining real service expenditure together with nondurable expenditures. Corresponding to the category of housing services, the BEA constructs a price deflator that, after normalization by the implicit nondurables deflator, serves as our measure of the relative user cost Q. Following Lustig and Van Nieuwerburgh (2004), we divide nondurable consumption and the service flow from housing by the number of *households* as reported by the US Bureau of the Census.[8] The data set covers 1929 to 2002 at an annual frequency.

Table 1 Panel A displays the time-series mean, standard deviation and extrema for the growth rate, expressed as the log difference, of our measures of nondurable consumption, (the service flow from) durables and the relative price. Most importantly, note that the variation of the growth rate of (the service flow from) durables is roughly of the same magnitude as the variation of nondurable consumption growth in both data sets. The Augmented Dickey-Fuller (ADF) tests presented in Panel B cannot reject the hypothesis that all series are individually integrated processes of order one, a necessary prerequisite for any cointegration test.

[7]The Bureau of Economic Analysis uses observed market rents for tenant-occupied housing and an imputed “space rent” for owner-occupied housing. See BEA (1990) for details.

[8]Annual data on the number of US households become available only in 1947. Before we have data on the number of households for 1920, 1930, and 1940. Using annual data for the size of the US population, we infer the number of households for the years with missing observations by linear interpolation of the average household size.

3.2.2 Return data

The sources and construction of our return data are identical for both data sets. The data differ only in their frequency, quarterly in the first data set and annual in the second data set. All returns are expressed in units of the nondurable consumption good.

Table 2 Panel A reports summary statistics for the risk-free rate R_f, the CRSP value weighted market return R_m as well as the excess market return R_m^e. The risk-free rate and the market return are from CRSP.[9] For the predictive regressions we calculate long horizon returns for k quarters or years as cumulative log returns, i.e. $r_{m,t+k} = \sum_{j=1}^{k} lnR_{m,t+j}$. Panel B contains summary statistics for the 25 Fama-French portfolios that we use as test assets in the cross-sectional asset pricing tests. The data are from Kenneth French's web site.

3.3 Results

3.3.1 Tests using quarterly durable goods data

The Relationship between the relative price P and the relative user cost Q

Our quarterly data set contains observations on the relative price P, but not on the relative user cost Q. As we have shown above, P and Q are related in the following way:

$$P_t \left[1 - \delta \beta E_t \left[\frac{U_{C_{t+1}}}{U_{C_t}} \frac{P_{t+1}}{P_t} \right] \right] = Q_t \tag{32}$$

As a first approximation, we assume that $E_t \left[\frac{U_{C_{t+1}}}{U_{C_t}} \frac{P_{t+1}}{P_t} \right]$ is time-invariant, in which case we can write the relationship between P and Q as:[10]

$$\frac{U_{D_t}}{U_{C_t}} = qP_t = Q_t \tag{33}$$

With this additional assumption of constant proportionality, we can rewrite the cointegration relationship as:

$$lnD_t = \frac{lnq - ln\alpha - \omega ln\phi}{\omega - 1} + lnC_t + \frac{1}{\omega - 1} lnP_t \tag{34}$$

[9]To obtain the risk-free rate at annual frequency for our second data set, we calculate the return of four consecutive investments in 90-day T-Bills.

[10]For robustness, we later relax this assumption and allow q to vary as a function of the risk-free rate.

The cointegration residual is now defined as:

$$res_t = lnD_t - \frac{-ln\alpha - \omega ln\phi}{\omega - 1} - lnC_t - \frac{1}{\omega - 1}lnP_t \tag{35}$$

The approximate linear pricing kernel in (31), can we be written as:

$$\beta\frac{U_{C_{t+1}}}{U_{C_t}} \approx 1 + ln\beta + (b2 + b3)\Delta lnC_{t+1} - b3\Delta lnP_{t+1} + b4\Delta res_{t+1} \tag{36}$$

Cointegration

Table 3 Panel A presents results from the Canonical Cointegration Regressions (CCR). Neither deterministic ($H(0,1)$) nor stochastic ($H(1,m)$) cointegration can be rejected for durable consumption, nondurable consumption, and the relative price. As a robustness check, we repeat the tests while imposing the unit coefficient on nondurable consumption. Again, p-values are high.

Panel B reports the cointegrating vector estimated by DLS. For comparison, we also report simple OLS estimates as well as the CCR estimates. All estimates appear relatively similar. Below we will focus on those obtained by DLS. The coefficient estimates on the relative price imply that the elasticity of substitution between durable and nondurable consumption, $\frac{1}{1-\omega}$, is about 0.47, a value lower than the 1.17 estimated by Ogaki and Reinhart (1998), who use quarterly nondurable and durable expenditures together with the relative price to infer ω. Notice, that the coefficient estimate on nondurable consumption is significantly above 1. When we impose a unit coefficient for nondurable consumption, as implied by the model, the implied elasticity of substitution increases to 0.80. For the following analysis, our focus will be on the cointegration residual obtained from the unconstrained DLS estimation.

The economic content of the level of the residual

While the durable stock, nondurable consumption and the relative price seem to be cointegrated as predicted by (34), the cointegration residual displays a distinct pattern that seems inconsistent with the frictionless model. Figure 1 graphs the cointegration residual (scaled to a mean of zero) over time and relative to the US business cycle as measured by the NBER. The residual peaks during contractions and bottoms out at some point during expansions. Note that, unlike the frictionless model, an economy with adjustment costs for the durable good is consistent with such a pattern: At the beginning of an expansions, wealth increases and agents increase their nondurable consumption immediately,[11] while only few agents increase the size of the durable good at any point in time, leading initially to a negative residual in the aggregate data. When times continue to be good, more and more agents adjust the size of their durable good, pushing the residual back towards zero. Similarly, at the beginning of a recession, wealth decreases and the opposite effect occurs: Nondurable consumption is cut back immediately, while there is a temporary "overhang" of the durable good.

We explore the economic content of the cointegration residual further by looking at the contemporaneous correlation with typical state variables. Table 4 shows that the residual is negatively correlated with the market price-earnings ratio (-0.37) and positively with the market dividend yield (0.41) and the yield spread between AAA- and BAA-rated bonds (0.24), confirming that the residual tends to be high in "bad" times and low in "good" times. As Table 4 shows, the residual is also negatively correlated with consumer expectations (-0.41), as measured by the University of Michigan, and positively so with Stock and Watson's Experimental Recession Index (XRI) (0.39), which measures the probability that the economy enters a recession within six months. For robustness, column II of Table 4 also reports the same correlations for the case where the unit coefficient is imposed when we estimate the cointegration residual. If anything, the correlations are even stronger.

Results from several predictive regressions are presented in Panel A of Table 5. Deviations from the cointegration relationship have strong predictive power of future growth in durable consumption at all horizons. Specifically, an increase of the cointegration residual of one standard deviation (0.0267) will lower the cumulative growth rate of durable consumption by -1.28 percentage points over the following year and by as much as -4.22 percentage points over the following three years. Increases in the cointegration residual are similarly

[11]This follows from the concavity of the utility and value functions and the fact that $U_C = V_W$.

associated with decreased future growth of nondurable consumption, GDP, and disposable income. With respect to future returns, we find that a high level of the residual seems to predict a lower risk-free rate at horizons of up to four quarters, consistent with the finding that the residual increases before and during economic contractions. Overall, though, the cointegration residual seems to have little predictive power for future returns. Given the clear association with known predictors such as the market PE ratio or dividend yield, this appears puzzling. But even for purely financial state variables, like the PE ratio or the dividend-yield, no definite conclusion on the ability to predict returns has been reached (see for example Ang and Bekaert (2005)). Panel B reports the corresponding results when the unit coefficient on nondurable consumption is imposed in the estimation of the cointegration residual. While the point estimates are generally somewhat lower, the pattern of results is again very similar. Overall, these findings contradict the implications of the frictionless model but are consistent with the existence of relevant adjustment costs.

The economic content of changes of the residual

The linear approximation of the pricing kernel described above allows us to directly test whether changes in the cointegration residual contain information that is useful in explaining the cross-section of asset returns. It also links our investigation closely to recent work by Yogo (2006), who uses the same macro-level consumption data to test the two factor model for the frictionless economy. Table 6 columns I and II show the results when the pricing kernel is defined as in (30):

$$\beta \frac{U_{C_{t+1}}}{U_{C_t}} \approx 1 + ln\beta + b2\Delta lnC_{t+1} + b3\Delta lnD_{t+1} \tag{37}$$

The test assets are the risk-free rate and the 25 Fama-French portfolios, in the first column, and excess returns on the 25 Fama-French portfolios in the second column. Similarly to Yogo (2006), we find that, judging by the J-test, the model appears to perform relatively well when estimated on excess returns. The coefficient estimates on both aggregate nondurable and durable consumption growth are statistically significant. Together, they imply a very high point estimate of the structural parameter γ of about 107.8, where $\gamma = -(b2 + b3)$. Note that both factors command a significantly positive risk premium.

Columns III and IV present the corresponding results after substituting out $ln\Delta D_{t+1}$, so that the pricing kernel is now as defined in (36):

$$\beta\frac{U_{C_{t+1}}}{U_{C_t}} \approx 1 + ln\beta + (b2+b3)\Delta lnC_{t+1} - b3\Delta lnP_{t+1} + b4\Delta res_{t+1} \tag{38}$$

Parker and Julliard (2005) conjecture that Yogo's (2006) results might be driven by the fact that durable growth contains information on relative prices. Our results support this conjecture. In addition, though, our results show an important role for changes in the cointegration residual. The coefficient on $ln\Delta res_{t+1}$ is highly significant for both sets of assets. We find a positive risk premium for the nondurable consumption growth factor, while changes in the relative price and cointegration residual command negative risk premia. While the risk premium on the cointegration residual is not significant, its sign is consistent with the pattern and correlations of the residual reported above: Assets that perform well when the residual is increasing can be used to hedge against bad times and hence are sought after, leading to a lower required rate of return. When we repeat the estimations for the case where the cointegration residual is estimated with the unit coefficient imposed on nondurable consumption (columns V and VI), the coefficient estimates are very similar.

Robustness: Allowing time-variation in the relationship between the relative price P and the relative user cost Q

So far, we have made the assumption that the ratio between the relative price P and the relative user cost Q is constant over time. While any empirical evidence with respect to the durable goods considered here (mainly vehicles and furniture) is scarce, we want to verify that our results are robust to possible time variation in this ratio.

We therefore assume that the relationship varies as a function of the risk-free rate. Specifically, we assume

$$Q_t = (R_{f,t} - \delta)P_t \tag{39}$$

where we use the constant depreciation rate of 6% and the time-varying risk-free rate $R_{f,t}$.[12]

[12] For a more detailed discussion of different approximations of user costs, see e.g. Murray and Sarantis

We repeat all the above estimations using (39). The empirical findings (not shown here) are qualitatively similar to the results discussed above: The cointegration residual displays the same cyclical pattern and is again correlated with the same typical state variables. The predictive regressions as well as the cross-sectional asset pricing tests also yield similar results, suggesting that our results are robust to time-variation in the relationship between P and Q.

3.3.2 Using annual housing consumption data

Our second data set contains data on annual housing services as well as the relative user cost of these services. It therefore allows us to relax the assumption that the service flow from the durable good is proportional to the size of the durable good. More importantly, it provides us with a direct observation of the relative user cost. We can therefore write (20) as:

$$lnS_t = \frac{-ln\alpha}{\omega - 1} + lnC_t + \frac{1}{\omega - 1} lnQ_t \tag{40}$$

After confirming deterministic as well as stochastic cointegration and estimating the cointegrating vector,[13] we focus on the predictive regressions and the cross-sectional asset pricing tests.

Table 7 demonstrates that a one standard deviation increase of the cointegration residual (0.1114) lowers the growth rate of housing services by 2.64 percentage points over the course of the following four years and by 6.22 percentage points over the following ten years. With respect to growth of nondurable consumption, GDP, and disposable income, the cointegration residual seems to have significant predictive power only for horizons of five years and beyond. At those long horizons, an increased residual today predicts higher future growth rates. When using quarterly data, a high cointegration residual predicts lower growth rates of those variables over the next five years (see Table 5). The predictability found using annual data, however, applies to horizons of five to twelve years that are beyond the normal business cycle frequency. With respect to financial returns, we find that for horizons of eight years

(1999).

[13]Results are not reported, but available upon request.

and below, the residual has a marginally significant positive correlation with future risk-free rates. Finally, at longer horizons, i.e. seven years and beyond, the residual is associated with lower cumulative future excess returns.

In Table 8, we present the results from the cross-sectional asset pricing tests, again using the 25 Fama-French portfolios and the risk-free asset as test assets. Columns I and II show the results for the base case before substituting out the service flow factor, $ln\Delta S_{t+1}$. When all test assets are excess returns (column II), both coefficient estimates have the same negative sign and the corresponding risk premia are significantly positive, as before. When the risk-free rate is included as a test asset (column I), though, the sign of the coefficient on the service flow factor becomes positive instead of negative and the corresponding risk premium becomes negative. Note also that the implied parameter value of γ is about 46.2 (column II), less than half the size as before. Turning to the case where the pricing kernel contains the growth rate of the relative price and the change in the cointegration residual as factors, our findings (columns III and IV) are basically the same as before. Changes in the cointegration residual have highly significant coefficient estimates and demand a statistically significant negative risk premium in both cases, confirming that increases in residual are associated with deteriorating economic conditions.

These results also offer a different view of the composition risk suggested by Piazessi et. al. (2006). Using a very similar data set and the same set-up of a fritionless economy with perfect rental markets, they introduce the ratio ϱ_t, defined as nondurable consumption over total consumption expenditure, as a state variable:

$$\varrho_t \equiv \frac{C_t}{C_t + Q_t S_t} \tag{41}$$

Assuming that there are no adjustment costs, and that equation (19) holds every period, they interpret changes in ϱ_t as composition risk that demands a positive risk premium. In the presence of frictions, aggregate composition risk might reflect deviations from (19) caused by adjustment costs. Our results that test the validity of (19) suggest that this might indeed be the case.

3.4 Discussion

Our findings using aggregate data provide evidence against the null hypothesis of a frictionless model in which agents equate the ratio of their marginal utilities with the relative user cost in each period: Deviations from (19) display a distinct pattern in relation to the business cycle. Furthermore, these deviations have significant power in predicting the evolution of the economy and explaining the cross-sectional variation in asset returns.

Our results are consistent with the frictions that could be caused by non-convex adjustment costs with respect to the durable good. In such an economy, deviations from (19) contain information on the state of the economy. Intuitively, a positive residual indicates that desired durable consumption is lower than actual durable consumption, while a negative residual signals the opposite. If adjustment costs are present, tests of the Euler equation with respect to nondurable consumption as in equation (10) that use aggregate data are likely to be misspecified. Any approach that constructs a representative agent for use with aggregate data has to reflect, in some way, the heterogeneity across agents with respect to their nondurable and durable consumption. Under restrictive assumptions, it is possible to capture the cross-sectional heterogeneity in a relatively simple way (for one example, see Bar-Ilan and Blinder (1988)), but in general, as pointed out by Caballero (1993) this will not be possible.

One approach to dealing with this problem would be to include moments of the relevant cross-sectional distribution in an aggregate pricing kernel. Such an approach has been taken in recent work by Cogley (2002), Balduzzi and Yao (2007), and Jacobs and Wang (2004), who consider the case of CRRA preferences over nondurable consumption when agents face uninsurable idiosyncratic risks. The empirical results of these tests, however, are ambiguous and depend on the exact approximation used.

An alternative approach is to use micro-level data, thereby avoiding aggregation of $C_{i,t}$ and $D_{i,t}$ as well as approximation of the pricing kernel. As discussed above, the Euler equation for nondurable consumption will continue to hold for individual agents even if adjustment costs are present. We pursue this approach further in the next section.

4 Micro Evidence

4.1 Testing Euler equations with household-level data

The main problem with aggregation in an economy with adjustment costs is that the Euler equation will not hold for changes in *aggregate consumption* data, while it continues to hold for *aggregate marginal utility* growth. Since the Euler equation holds for each individual agent we can average them across agents to obtain a valid aggregate Euler equation of the following form:

$$\frac{1}{N}\sum_{i=1}^{N} E_t\left[\beta\frac{U_{Ci,t+1}}{U_{Ci,t}}R_{t+1}\right] = \imath \tag{42}$$

Jacobs (1999) uses (42) to test the Euler equation in an economy where agents have CRRA preferences over nondurable consumption only and markets are possibly incomplete, invalidating the Euler equation for aggregate consumption data. Brav, Constandinides, and Geczy (2002), using a similar set-up as Jacobs, calibrate (42) on household data to match the equity premium. We extend Jacobs' approach to our case of non-separable preferences over nondurable and durable consumption. Under rational expectations, (42) implies that:

$$\lim_{T\to\infty}\frac{1}{T}\sum_{s=1}^{T}\frac{1}{N}\sum_{i=1}^{N}\left(\beta\frac{U_{Ci,s+1}}{U_{Ci,s}}R_{s+1}-\imath\right)\otimes Z_{i,s} = \lim_{T\to\infty}\frac{1}{T}\sum_{s=1}^{T}\frac{1}{N}\sum_{i=1}^{N} v_{i,s+1}(\theta)\otimes Z_{i,s} = 0 \tag{43}$$

where R_{t+1} is an $Lx1$ vector of asset returns, $\imath$ is a vector of ones, $Z_{i,t}$ is a K x 1 vector of instruments that are known to the agent at time t but independent of the expectational error $v_{i,t+1}$ and θ is the vector of the structural parameters that we want to estimate. It is important to note that in a world with aggregate shocks, the error $v_{i,t+1}$ will be correlated across different agents, in which case we cannot rely on cross-sectional asymptotics, i.e. a large number of agents in the economy, but need to rely on a sufficiently long time dimension (see Chamberlain (1984)).

Let $g(\theta)$ denote the LK moment conditions implied by equation (43):

$$g(\theta) = \frac{1}{T}\sum_{s=1}^{T}\frac{1}{N}\sum_{i=1}^{N} v_{i,s+1}\otimes Z_{i,s} \tag{44}$$

Provided that the number of unknown parameters is less or equal to the number of moments, GMM estimation will identify θ as:

$$\hat{\theta} = \arg \min_{\theta \epsilon \Theta} g(\theta)' W g(\theta) \tag{45}$$

where W is a positive definite weighting matrix.[14] If the model is overidentified, Hansen's (1982) J-test allows us to test the overall performance of the model.

In the general case of CES preferences, the individual pricing kernel $M_{i,t+1}$ is given by:

$$M_{i,t+1} = \beta \frac{U_{Ci,t+1}}{U_{Ci,t}} = \beta \left(\frac{C_{i,t+1}^{\omega} + \alpha(\phi D_{i,t+1})^{\omega}}{C_{i,t}^{\omega} + \alpha(\phi D_{i,t})^{\omega}} \right)^{\frac{1-\gamma}{\omega}-1} \left(\frac{C_{i,t+1}}{C_{i,t}} \right)^{\omega-1} \tag{46}$$

Note that the parameters α and ϕ are not identified individually. We will only be able to estimate $\alpha\phi^{\omega}$. In the Cobb-Douglas case, the pricing kernel reduces to:

$$M_{i,t+1} = \beta \frac{U_{Ci,t+1}}{U_{Ci,t}} = \beta \left(\frac{C_{i,t+1}}{C_{i,t}} \right)^{\frac{1}{1+\alpha}(1-\gamma)-1} \left(\frac{D_{i,t+1}}{D_{i,t}} \right)^{\frac{\alpha}{1+\alpha}(1-\gamma)} \tag{47}$$

For both preference specifications, the parameters will enter the corresponding first order conditions of (45) in a nonlinear way, so that closed form solutions will not be available. We therefore have to solve (45) numerically. While this is computationally straightforward,[15] the small sample properties of nonlinear GMM estimators are not well known. For robustness, we therefore also consider the linear approximation of the pricing kernel $M_{i,t+1}$ (presented above for aggregate data) so that:

$$M_{i,t+1} \approx 1 + ln\beta + b2\Delta lnC_{i,t+1} + b3\Delta lnD_{i,t+1} \tag{48}$$

In the case of CES preferences, γ is identified by $-(b2 + b3)$, where:

$$b2 = \frac{1-\gamma-\omega}{1+\alpha\phi^{\omega}} - 1 + \omega \tag{49}$$

$$b3 = \frac{\alpha\phi^{\omega}(1-\gamma-\omega)}{1+\alpha\phi^{\omega}} \tag{50}$$

[14] The weighing matrix W that we use is the inverse of $I_L \otimes (\sum_{t=1}^{T} Z_t Z_t)$, where $Z_t = \frac{1}{N}(Z_{1,t}, ..., Z_{N,t})$ and I_L is the LxL identity matrix.

[15] We use the Broyden-Fletcher-Goldfarb-Shanno algorithm and interact it with MATLAB's simplex search method *fminserach*. To identify global minima, we repeat all estimations for a set of different starting values.

So far, we have assumed that adjustment costs with respect to the durable good are the only possible source of frictions in the economy so that the Euler equation will continue to hold for individual agents. However, there is strong evidence that in reality borrowing constraints as well as entry costs to stock-ownership exist (e.g. Zeldes (1989), Vissing-Jørgensen (2002)). These additional frictions would invalidate the Euler equation for those households with corner solutions. We control for these additional frictions by only selecting households that are unlikely to be affected by them. We now turn to the sources of our micro data and the selection criteria we apply.

4.2 Data

4.2.1 Micro-level consumption data

We use household-level data from the Panel Study of Income Dynamics (PSID) for 17 years between 1978 to 1997.[16] The PSID is the longest and broadest panel data set for US households and provides detailed information on household demographics, income sources, food consumption expenditures, as well as housing choices. Most importantly, the Wealth Supplements collected in 1984, 1989, and 1994 give information on a household's wealth composition, allowing us to identify households that hold stocks. While the PSID data have severe limitations, in particular the short time dimension as well as the fact that food consumption is the only measure of nondurable consumption, it has been used before to study intertemporal consumption choices of individual households. Mankiw and Zeldes (1990), for example, use PSID data to compare the variability of food consumption between stockholders and non-stockholders over a period of 13 years. Jacobs (1999) performs nonlinear Euler equation tests of the standard CRRA model using 12 years of PSID food consumption growth.

In this paper, we use total household food expenditure as a measure of nondurable consumption $C_{i,t}$. Furthermore, we proxy for the durable good $D_{i,t}$ with the self-reported

[16]Nondurable consumption data were not collected in 1988 and 1989, so that we cannot calculate nondurable consumption growth for 1988 to 1990. The instrument set will also include data from 1975 to 1977.

value of the owner-occupied house. We deflate the food expenditure data with the Bureau of Labor Statistics (BLS) Food Price Index and the reported house value with the House Price Index (HPI) for the relevant US state as provided by the Office of Federal Housing Enterprise Oversight (OFHEO). For each household, we keep the first reported deflated house value constant until the household moves to a new house. That means that we set $D_{i,0} = D_{i,1} = D_{i,2} = ... = D_{i,t-1}$ until the household moves in period t. Said differently, we assume that the service flow from the house remains unchanged over time. To change the utility from housing a household has to move to a new house.

While our model is based on decisions taken by an individual, our data describe the food and housing decisions of an entire household, regularly consisting of several members. To control for this in our estimation, we include an exponential function of household demographics in the preference specification. These preference shifters are the number of household members ($num_{i,t}$) as well as the number of children under 18 years of age living in the household ($kids_{i,t}$). Therefore, we use:

$$U(C_{i,t}, D_{i,t}) = \frac{u(C_{i,t}, D_{i,t})^{1-\gamma}}{1-\gamma} \exp(d1num_{i,t} + d2kids_{i,t}) \tag{51}$$

as our preference specification, where i now stands for the entire household.

Finally, we apply the following sample selection criteria: First, since we observe house values only for homeowners, we drop all renters from the data set.[17] Next, we correct for outliers by dropping the bottom and top 1% of households with respect to the real per capita food consumption and the real house value.[18] We then select those households that have positive net wealth as wealth as positive holdings of stocks in 1984, 1989, and 1994. The set of stockholders contains 352 households with 3,536 year-household observations.

Table 9 provides summary statistics for the set of households with positive stock holdings. We observe that the mean growth rates of nondurable food and durable housing consumption

[17] While the preference specifications should also hold for renters, we lack comparable information on the value of the house renters occupy.

[18] We also drop households that report a house value below 4,000 dollars or a family income growth rate above 100.

are similar. The total volatility of both growth rates is 0.35 and 0.29 respectively. The corresponding standard deviations of the cross-sectional averages are only 0.026 and 0.023, reasonably close to the macro counterparts shown in Panel A of Table 1. Most importantly, the average cross-sectional standard deviation of $D_{i,t}/C_{i,t}$ is 0.98, while the overall mean of $D_{i,t}/C_{i,t}$ is around 1.7. This shows that there is considerable variation across households in a given year.

4.2.2 Return data

The Euler equation in (42) should hold for *any* gross return. While the saving consumption literature focuses on the predetermined risk-free rate, interest in finance lies in the joint determination of the risk-free rate and the equity market return as well as the returns of equity portfolios formed on certain firm characteristics such as the book-to-market ratio or size.

We will therefore consider the risk-free rate measured as the return on the 1-Year Treasury Bill, the annual value-weighted CRSP return, and the annual return of six size and book-to-market portfolios from Kenneth French's web site. Since the PSID data are mostly collected during the first quarter of the year, all returns reflect price changes between the beginning of the second quarter and the end of the first quarter of the following year. Finally, all returns are deflated with the BLS Food Price Index. Table 10 provides times-series statistics for the real return data.

4.3 Results

An aggregate benchmark

In a first step, we attempt to verify that our data resembles aggregate data used in recent research. To this end, we average food consumption and housing values across all home-owners at t and $t-1$.[19] We then calculate growth rates of aggregate food and housing consumption, thereby obtaining a time-series with 17 annual observations. Panel A of Table

[19]Food consumption is expressed per capita and housing values per household. Housing values are also allowed to vary year by year as reported by the household.

11 reports statistics for this aggregate time series. Notice that the standard deviations of the two annual consumption growth rates, 0.018 for food and 0.015 for housing consumption, are slightly below the standard deviations of the annual aggregate data set (see Table 1 for the summary statistics). Panel B of Table 11 reports GMM estimates for the case of CES and Cobb-Douglas preferences and Panel C reports GMM estimates for the linear approximation of those preferences. The test assets are the risk-free rate and the six Fama-French portfolios. While noisy, the estimates appear in line with our own findings above and those of Yogo (2006), who reports estimates of γ around 200.

We now turn to the main results of this section of the paper, the estimation of the Euler equation on household-level data as described in Section 4.1. Unlike the GMM estimation results reported so far, we will report only 1-step results, as 2-step GMM is often infeasible when more than one test asset is used.[20] We report two versions of Hansen's overidentification test: J-test (1) is based on raw model errors, while J-test (2) is based on demeaned model errors. Finally, all standard errors are robust to heteroscedasticity and correlation across households in a given year as well as to autocorrelation (for each unit and across units) of up to two years.[21]

CES pricing kernel

Table 12 contains the estimation results when preferences are of the CES form. The instrument set includes household specific, state-specific, as well as macroeconomic variables. In particular we use the following nine instruments: a constant, food consumption growth (lagged twice), the ratio of housing over food consumption (lagged twice), income growth (lagged twice), a dummy variable that is one if the household expects to move in the near future (lagged twice), the change in the number of household members and in the number of children under 18 years of age living in the household, the state unemployment rate (lagged once), and the risk-free rate (lagged once).

[20]When both 1-step and 2-step GMM are feasible, results are very similar.

[21]We use Newey-West (1987) weights in the construction of the covariance matrix.

The four columns show the results for different combinations of the test assets: the risk-free rate, the market return, and the six Fama-French portfolios. The coefficient estimates are very similar for these different test assets. The curvature parameter γ is estimated at around 3, slightly lower when the only test asset is the risk-free asset (2.75), and slightly higher when the market return by itself or the six Fama-French portfolios together with the risk-free rate are used (3.4 in both of those cases). The standard error is about 0.7 in all cases. For ω, the parameter that governs the elasticity of substitution between the two goods, we obtain point estimates between 0.35 (for the risk-free rate alone) and 0.18 (for the risk-free rate and the six Fama-French portfolios), implying rates of substitution, $\frac{1}{1-\omega}$, between 1.54 and 1.23. In neither case can we reject that preferences are Cobb-Douglas, i.e. that ω is significantly different from zero. The estimate of the time preference parameter β appears quite low, ranging between 0.73 and 0.88. This is the case even after accounting for the effect of the two demographic controls, $exp(d1\Delta num_{t+1} + d2\Delta kids_{t+1})$. As noted above, α and ϕ are not identified individually. The estimate of $\alpha\phi^{\omega}$ is between 1.66 and 1.81 and in all cases significantly different from zero. A Wald test, reported at the end of each column, in all cases strongly rejects that preferences are separable in the two goods, i.e. we reject that $\omega = 1-\gamma$. Finally, the two tests of overidentification, J-test (1) and J-test (2), do not reject the model if only one test asset is used. When two or more test assets are used, and the number of moment conditions increases accordingly, the first J-test continues to not reject the model, while the more stringent second J-test, based on demeaned errors, strongly rejects the model.

CD and linearized pricing kernel

We impose the restriction that the elasticity of substitution between the two goods is equal to one, i.e. $\omega = 0$, and reestimate the model under this assumption of Cobb-Douglas preferences. Table 13 reports the results, confirming an estimate for γ of around 3 as well as a low estimate for β. The weights on food and housing consumption, $\frac{1}{1+\alpha}$ and $\frac{\alpha}{1+\alpha}$ respectively, are now identified, implying a weight on food consumption between 0.18 (using the risk-free return alone) and 0.30 (using the risk-free return together with the six Fama-French assets). The two overidentification tests have again a very similar pattern, contradicting each other in the case of more than one test asset.

Finally, we report results for the linearized pricing kernel. In this case we can solve for the parameter vector in closed form. This offers a robustness check on our numerical solutions in the above cases of nonlinear pricing kernels. As Table 14 shows, the implied parameter of $\gamma = -(b2 + b3)$ is slightly higher than before and also covers a wider range, with the lowest estimate of 2.64 when the risk-free asset is the only test asset and the highest estimate of 4.58 in case of the market portfolio by itself. The precision is also lower, with standard errors between 1.51 and 2.9. The very precisely estimated time preference parameter $\beta \approx 1 + ln\beta$ is higher than in the nonlinear cases, ranging from 0.90 to 0.98, possibly indicating a downward bias in the nonlinear estimations. If preferences are assumed to be Cobb-Douglas, the weights on food and housing consumption can also be identified. As the last line in Table 14 shows, the estimated weight on food consumption ranges between 0.56 and 0.69, significantly above the values we have estimated above. Finally, note that both J-tests display the same pattern as before.

4.4 Discussion

In this section, we discuss the implications of our estimation results with respect to key parameters such as the intratemporal elasticity of substitution between durable and non-durable consumption as well as the intertemporal elasticity of substitution of nondurable consumption between time t and time $t + 1$. Finally, we assess the overall performance of the estimated models.

Intratemporal substitution between nondurable and durable consumption

The elasticity of substitution between nondurable food and durable housing consumption $\frac{1}{1-\omega}$ is a potentially important parameter. Piazessi et. al. (2006) show that a linearized model with CES preferences can be calibrated on aggregate data to match the equity premium with a value for γ of 5 as long as ω is assumed to be slightly larger than zero (0.05 in their case). The particular linear approximation considered by them results in the following pricing kernel:

$$M_{t+1} \approx \beta \left(1 - \gamma \Delta ln C_{t+1} - \frac{\gamma + \omega - 1}{\omega} \Delta ln \varrho_{t+1}\right) \tag{52}$$

where ϱ_{t+1} is the nondurable expenditure share as defined in (41). Our point estimates of ω (see Table 12) fall between 0.18 and 0.35, far above the 0.05 assumed by Piazzesi et. al. (2006). Given the size of the standard errors, however, we cannot reject that ω is equal to 0.05. Also, our micro data estimates are close to the estimates from a cointegration analysis on aggregate nondurable and durable expenditure performed by Ogaki and Reinhart (1998), but above the results from our own cointegration analysis in Section 3. Ogaki and Reinhart (1998) estimate ω to be 0.19, while our own estimates from quarterly aggregate data (see Table 3 Panel B) are -1.13 and -0.25.[22]

These results are very different, though, from the value for ω reported by Flavin and Nakagawa (2004). They estimate the Euler equation implied by CES preferences over food and housing consumption on PSID data, but estimate ω to be -6.5 with a standard error of 1.8. Comparing their results with our results is difficult, however, as their sample period is shorter and they do not control for liquidity constraints and stock-ownership. Also different from our approach, they measure the amount of housing consumption by using the estimated size (in sq. ft.) of the household's residence.

Our results confirm that even if adjustment costs are present, the elasticity of substitution between nondurable and durable consumption can be estimated consistently from aggregate data as long as cointegration analysis is used, because, as argued above, the two consumption goods and relative user costs cannot evolve independently from each other in the long run.

Intertemporal substitution

The EIS with respect to nondurable consumption is given by:

$$-\frac{U_{Ci,t}}{C_{i,t} U_{CCi,t}} = \frac{1 + \alpha \phi^{\omega} \left(\frac{D_{i,t}}{C_{i,t}}\right)^{\omega}}{\gamma - \alpha \phi^{\omega} \left(\frac{D_{i,t}}{C_{i,t}}\right)^{\omega} (\omega - 1)} \tag{53}$$

[22]Remember that the coefficient estimate for the log of the relative price is $\frac{1}{\omega - 1}$.

In the absence of adjustment costs, the ratio $\frac{D_{i,t}}{C_{i,t}}$ will be the same for all households, implying that at each point in time the EIS will be the same across households. The economy-wide EIS will change over time to the degree that the relative user cost for the service flow from the durable good changes. In an economy with non-convex adjustment costs, the EIS will be different across households as they differ in the ratio $\frac{D_{i,t}}{C_{i,t}}$. For each household, the EIS will change as this ratio changes unless $\omega = 0$ (the Cobb-Douglas case) or $\omega = 1 - \gamma$ (the case of separable preferences). If $\gamma > 1$, the EIS will decrease in $\frac{D_{i,t}}{C_{i,t}}$ as long as ω is in $(1 - \gamma, 0)$, otherwise the EIS will increase as $\frac{D_{i,t}}{C_{i,t}}$ increases. The economy-wide EIS will therefore depend on the cross-sectional distribution of the ratio $\frac{D_{i,t}}{C_{i,t}}$ at any time.

Taking the median $\frac{D_{i,t}}{C_{i,t}}$ of 1.52 and our estimates of ω (0.285), $\alpha\phi^{\omega}$ (1.81), and γ (3.01) from Table 12, the implied EIS is 0.68. Notice that this is about twice the value of the EIS implied by the same value for γ if preferences were separable and of the CRRA form.

The equity premium and HJ-bounds

Our results suggest that estimation of the Euler equation using micro-level data leads to reasonable parameter values, especially a low value for γ. Different from existing research on asset pricing models using micro-level data, we incorporate durable consumption in the form of consumption of housing services and find strong evidence that preferences are not separable. In this sense, we present a more comprehensive preference specification, consistent with the renewed interest in two factor models as well as with the presence of adjustment costs. At the same time, the overidentification tests cast some doubt on the overall performance of the model.

In order to understand if the pricing kernel suggested in (42) can provide an answer to the equity premium puzzle, we need to gain a better understanding of its time-series properties. With only 17 years of annual observations, this is difficult in our setting. Nevertheless, we evaluate the "aggregate" pricing kernel implied by our model (see (42)):

$$M_{t+1} = \frac{1}{N} \sum_{i=1}^{N} E_t \beta \frac{U_{Ci,t+1}}{U_{Ci,t}} \tag{54}$$

Hansen and Jagannathan (1991) show that for any valid pricing kernel M, the following inequality has to hold for any excess return R^e:

$$\frac{\sigma(M)}{E(M)} \geq \frac{\mid E(R^e) \mid}{\sigma(R^e)} \tag{55}$$

That is, the Hansen-Jagannathan (HJ) bound, the left-hand side of the inequality, has to be at least as large as the Sharpe ratio of any asset. For our sample of all stockholders, the time-series standard deviation of M is 0.11 and the mean is 0.90, implying a Hansen-Jagannathan (HJ) bound of 0.12. While this value is above the HJ-bound of 0.09 implied by parameter estimates in Jacobs (1999)[23], it is still below the Sharpe ratio of the market excess return, which is about 0.5. As we decrease ω, the HJ-bound decreases as well. When $\omega = 0.1$, for example, the implied HJ-bound is 0.08. On the other hand, when we increase ω, the HJ-bound increases (for $\omega = 0.6$, for example, the point estimate increases to 0.277). These results show that additional tests of the model using a longer time series dimension will be valuable.

5 Conclusions

A number of recent papers have tested the asset pricing implications of non-separable preferences over nondurable and durable consumption. These papers assume that a frictionless rental market for the durable good exists and that a representative agent can be constructed so that tests can be carried out on aggregate data. We have tested one important implication of these assumptions, namely that the ratio of aggregate marginal utilities equals the relative user cost period by period. Our evidence strongly rejects this implication of the frictionless model, while it is consistent with the presence of non-convex adjustment costs. Asset pricing results based on aggregate data are affected by this finding and have to be re-evaluated in the light of the evidence presented here.

Since aggregation in the traditional sense breaks down, we use micro-level data to test the Euler equation for nondurable consumption. Thereby, we effectively average marginal

[23]We use the coefficient estimates from Panel A of Table III in Jacobs (1999) and apply them to our data set.

utility growth across households. Our estimation results imply an elasticity of substitution between food and housing consumption of around 1.40, with a standard error of 0.44 so that we cannot rule out Cobb-Douglas preferences. Our results also imply a value for the intertemporal elasticity of substitution (0.68) that is significantly higher than values derived from aggregate data. At the same time, however, we find that the variability of our implied aggregate pricing kernel is still too low to be reconciled with the observed US equity premium.

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Table 1

Summary Statistics: Aggregate Data

Panel A

Variable	**Mean**	**St. Dev.**	**Min**	**Max**
Quarterly Data (203 observations)				
1952: II until 2002: IV				
Changes in Log Durable Stock	0.0095	0.0057	-0.0036	0.0213
Changes in Log Nondurable Consumption	0.0054	0.0049	-0.0163	0.0206
Changes in Log Relative Price (Durables / Nondurables)	-0.0054	0.0067	-0.0321	0.0140
Annual Data (73 annual observations)				
1930 until 2002				
Changes in Log Housing Service Flow	0.0177	0.0190	-0.0367	0.0669
Changes in Log Nondurable Consumption (excl. Housing Services)	0.0126	0.0257	-0.0993	0.0746
Changes in Log Relative User Cost (Housing Services / Nondurables)	-0.0045	0.0366	-0.1692	0.0668

Quarterly durable stock and nondurable consumption data are per capita. Annual housing service flow and nondurable consumption data are per household.

Panel B

Augmented Dickey-Fuller Tests		**Z(t)**
Quarterly Data (204 observations)		
1952: I until 2002: IV		
	10% Critical Value	*-3.140*
Log Durable Stock		-2.501
Log Nondurable Consumption		-1.045
Log Relative Price (Durables / Nondurables)		0.436
Annual Data (74 annual observations)		
1929 until 2002		
	10% Critical Value	*-3.169*
Changes in Log of Service Flow		-0.663
Changes in Log of Nondurables		-2.515
Changes in Log of Relative Price		-1.931

Quarterly durable stock and nondurable consumption data are per capita. Annual housing service flow and nondurable consumption data are per household. The Augmented Dickey Fuller (ADF) tests use four lags of the differenced variable.

Table 2

Summary Statistics: Return Data

Panel A

Variable	Quarterly Data (203 observations) *1952: II until 2002: IV* Mean	St. Dev.	Min	Max	Annual Data (73 observations) *1930 until 2002* Mean	St. Dev.	Min	Max
Risk-free Rate	1.004	0.005	0.988	1.024	1.008	0.044	0.887	1.158
Market Return	1.020	0.083	0.737	1.222	1.080	0.200	0.629	1.607
Market Excess Return	0.016	0.083	-0.258	0.222	0.072	0.204	-0.518	0.583

All returns are measured in units of the nondurable consumption good.

Panel B

	Quarterly Data (203 observations) *1952: II until 2002: IV* Mean	St. Dev.	Min	Max	Annual Data (73 observations) *1930 until 2002* Mean	St. Dev.	Min	Max
11P	1.013	0.155	0.592	1.461	1.046	0.369	0.313	2.313
12P	1.028	0.133	0.658	1.487	1.110	0.367	0.332	2.334
13P	1.030	0.117	0.713	1.417	1.148	0.343	0.349	2.135
14P	1.036	0.112	0.706	1.406	1.189	0.441	0.437	3.824
15P	1.039	0.121	0.697	1.490	1.188	0.360	0.436	2.418
21P	1.017	0.140	0.601	1.385	1.082	0.310	0.448	1.792
22P	1.025	0.116	0.698	1.415	1.124	0.295	0.490	2.242
23P	1.031	0.103	0.701	1.355	1.147	0.300	0.459	2.288
24P	1.033	0.102	0.744	1.364	1.156	0.320	0.494	2.499
25P	1.035	0.110	0.720	1.449	1.160	0.320	0.380	2.170
31P	1.020	0.125	0.644	1.331	1.095	0.297	0.496	2.422
32P	1.026	0.102	0.724	1.348	1.118	0.269	0.480	2.191
33P	1.027	0.095	0.694	1.287	1.125	0.263	0.509	2.062
34P	1.031	0.094	0.778	1.364	1.140	0.271	0.550	2.030
35P	1.033	0.104	0.734	1.383	1.150	0.323	0.440	2.201
41P	1.022	0.114	0.704	1.442	1.086	0.230	0.529	1.713
42P	1.021	0.095	0.715	1.285	1.098	0.247	0.550	2.251
43P	1.028	0.089	0.757	1.298	1.120	0.257	0.551	2.132
44P	1.029	0.090	0.760	1.326	1.128	0.270	0.462	1.921
45P	1.031	0.104	0.729	1.382	1.139	0.341	0.508	2.669
51P	1.020	0.091	0.667	1.259	1.078	0.207	0.580	1.471
52P	1.020	0.080	0.755	1.243	1.074	0.191	0.625	1.503
53P	1.023	0.074	0.765	1.170	1.092	0.215	0.446	1.837
54P	1.023	0.078	0.801	1.269	1.100	0.252	0.478	2.076
55P	1.023	0.088	0.748	1.341	1.121	0.306	0.478	2.346

All returns are measured in units of the nondurable consumption good. 11P stands for the value-weighted return on a portfolio with small size and low B/M stocks, 15P stands for the value-weighted return on a portfolio with small size and high B/M stocks and so on.

Table 3

Cointegration

Time Period: 1952: I until 2002: IV (204 observations)

Panel A

Cointegrated Variables	**H(0,1)** χ^2	**H(0,1)** p-value	**H(1,2)** χ^2	**H(1,2)** p-value	**H(1,3)** χ^2	**H(1,3)** p-value
Log Durable Stock Log Nondurable Consumption Log Relative Price	0.207	0.649	0.025	0.875	0.427	0.808
Log Durable Stock / Nondurable Consumption Log Relative Price	0.089	0.765	0.618	0.432	1.496	0.473

Panel A reports results from Canonical Cointegration Regressions (CCR). The left hand side variable is log durable stock in the first line and log (durable stock / nondurable consumption) in the second line of the results. H(0,1) is a test of deterministic cointegration. H(1,2) and H(1,3) are tests of stochastic cointegration. The null hypothesis is cointegration. Durable and nondurable consumption are expressed as per capita.

Panel B

Log Durable Stock	**I**	**II**	**III**	**IV**	**V**	**VI**
Estimation Methods	**DLS**	OLS	CCR	**DLS**	OLS	CCR
Constant	**-1.536**	-1.483	-1.769	**-0.703**	-0.739	-0.678
	0.128	*0.123*	*0.243*	***0.016***	*0.008*	*0.014*
Log Nondurable Consumption	**1.297**	1.259	1.358	**1**	1	1
	0.044	*0.043*	*0.084*			
Log Relative Price	**-0.465**	-0.491	-0.378	**-0.797**	-0.766	-0.928
	0.048	*0.047*	*0.098*	***0.022***	*0.016*	*0.005*

Panel B reports the cointegrating vectors obtained by different estimation techniques. In columns IV through VI, the unit coefficient on nondurable consumption is imposed. The DLS estimation includes two leads and eight lags of the differenced right hand side variables. OLS and DLS standard errors are Newey-West standard errors with lag length 4.

Table 4

The Cointegration Residual and other State Variables

Time Period: 1952: I until 2002: IV

Variable	I	II
Log Price Earnings Ratio	-0.37	-0.67
Dividend Yield	0.41	0.55
Yield Spread (AAA - BAA)	0.24	0.44
Consumer Expectations (U. Michigan)	-0.41	-0.62
Experimental Recession Index (XRI)	0.39	0.39

The table reports correlations between the cointegration residual from DLS regressions and different "state" variables. Correlations in column I are for the unrestricted cointegration residual. Correlations in column II are for the cointegration residual when the unit coefficient on nondurable consumption is imposed. The price earnings ratio and the dividend yield are for S&P's composite common index. The yield spread is the difference between yields on AAA and BAA rated bonds, as reported by the Federal Reserve Bank. Consumer expectations are from the University of Michigan. The experimental recession index (XRI) is from Stock and Watson. The index measures the probability that the economy enters a recession within six months.

Table 5

Predictive Regressions

Time Period: 1952: II until 2002: IV

Panel A

Durable Stock Growth	*k=1*	*k=4*	*k=8*	*k=12*	*k=16*	*k=20*
Residual	-0.10	-0.48	-1.04	-1.58	-2.02	-2.36
	0.02	*0.06*	*0.13*	*0.18*	*0.20*	*0.18*
Adj. R^2	21%	36%	51%	64%	72%	76%

Nondurable Growth	*k=1*	*k=4*	*k=8*	*k=12*	*k=16*	*k=20*
Residual	-0.04	-0.15	-0.31	-0.46	-0.57	-0.67
	0.02	*0.05*	*0.11*	*0.15*	*0.18*	*0.17*
Adj. R^2	4%	10%	18%	26%	29%	31%

GDP Growth	*k=1*	*k=4*	*k=8*	*k=12*	*k=16*	*k=20*
Residual	-0.12	-0.32	-0.51	-0.65	-0.74	-0.75
	0.03	*0.10*	*0.16*	*0.21*	*0.24*	*0.22*
Adj. R^2	10%	11%	15%	18%	20%	16%

Disposable Income Growth	*k=1*	*k=4*	*k=8*	*k=12*	*k=16*	*k=20*
Residual	-0.16	-0.60	-0.99	-1.31	-1.57	-1.86
	0.03	*0.11*	*0.22*	*0.34*	*0.42*	*0.45*
Adj. R^2	21%	30%	35%	43%	44%	44%

Risk-free Rate	*k=1*	*k=4*	*k=8*	*k=12*	*k=16*	*k=20*
Residual	-0.07	-0.22	-0.31	-0.34	-0.28	-0.16
	0.01	*0.08*	*0.20*	*0.29*	*0.38*	*0.46*
Adj. R^2	11%	9%	5%	3%	1%	0%

Market Excess Return	*k=1*	*k=4*	*k=8*	*k=12*	*k=16*	*k=20*
Residual	0.03	0.21	0.70	1.06	1.70	2.52
	0.22	*0.63*	*1.00*	*1.21*	*1.40*	*1.52*
Adj. R^2	0%	0%	0%	1%	2%	4%

Panel B

Durable Stock Growth	*k=1*	*k=4*	*k=8*	*k=12*	*k=16*	*k=20*
Residual	-0.06	-0.28	-0.62	-1.00	-1.30	-1.51
(when Nondur = 1)	*0.01*	*0.07*	*0.15*	*0.20*	*0.23*	*0.25*
Adj. R^2	12%	21%	30%	40%	44%	43%

Nondurable Growth	*k=1*	*k=4*	*k=8*	*k=12*	*k=16*	*k=20*
Residual	-0.02	-0.07	-0.18	-0.31	-0.44	-0.54
	0.01	*0.05*	*0.10*	*0.15*	*0.18*	*0.19*
Adj. R^2	1%	4%	9%	19%	25%	28%

GDP Growth	*k=1*	*k=4*	*k=8*	*k=12*	*k=16*	*k=20*
Residual	-0.05	-0.15	-0.26	-0.40	-0.51	-0.55
	0.02	*0.09*	*0.16*	*0.22*	*0.25*	*0.27*
Adj. R^2	3%	4%	6%	11%	13%	12%

Disposable Income Growth	*k=1*	*k=4*	*k=8*	*k=12*	*k=16*	*k=20*
Residual	-0.05	-0.22	-0.43	-0.70	-0.98	-1.28
	0.02	*0.11*	*0.23*	*0.34*	*0.44*	*0.51*
Adj. R^2	3%	7%	10%	19%	25%	29%

Risk-free Rate	*k=1*	*k=4*	*k=8*	*k=12*	*k=16*	*k=20*
Residual	-0.03	-0.10	-0.13	-0.10	0.00	0.17
(when Nondur = 1)	*0.01*	*0.07*	*0.18*	*0.31*	*0.44*	*0.55*
Adj. R^2	3%	3%	1%	0%	-1%	0%

Market Excess Return	*k=1*	*k=4*	*k=8*	*k=12*	*k=16*	*k=20*
Residual	-0.04	-0.24	-0.50	-1.01	-1.27	-1.42
(when Nondur = 1)	*0.18*	*0.58*	*1.09*	*1.29*	*1.35*	*1.34*
Adj. R^2	0%	0%	0%	1%	2%	1%

Panel A and B report results from predictive regressions at different horizons of k quarters. The left hand side variables are the cumulative per capita growth rate of log durable stock, log nondurable consumption, log GDP, log disposable income, the cumulative risk-free return, and the cumulative log market excess return. In Panel A, the right hand side variables are a constant (not reported) and the cointegration residual when the cointegration estimation is unrestricted. Panel B shows results for the cointegration residual when the unit coefficient on nondurable consumption is imposed in the cointegration estimation. Standard errors are Newey-West standard errors with lag length k. The sample size decreases from 203 (when $k=1$) to 184 observations (when $k=20$)

Table 6

Cross-Sectional Asset Pricing Tests

Time Period: 1952: II until 2002: IV (203 observations)

	I	II	III	IV	V	VI
Test assets	RF, FF25	FF25	RF, FF25	FF25	RF, FF25	FF25
Constant	1.985		3.470		3.470	
	0.492		*0.764*		*0.764*	
Δ log Nondurable Consumption	-102.868	-51.545	-335.465	-120.544	-289.759	-105.407
	37.545	*13.931*	*107.690*	*12.511*	*91.758*	*11.782*
Δ log Durable Stock	-45.214	-56.274				
	42.040	*9.197*				
Δ log Relative Price			134.977	33.608	185.566	50.361
			44.522	*8.990*	*60.199*	*9.660*
Δ Residual			-154.059	-51.020	-154.059	-51.020
			67.237	*11.208*	*67.237*	*11.208*
J - test	111.28	19.08	87.57	16.95	87.57	16.95
	0.00	*0.70*	*0.00*	*0.77*	*0.00*	*0.77*
HJ Distance	0.67	0.32	0.63	0.31	0.63	0.31
	0.00	*0.00*	*0.00*	*0.00*	*0.00*	*0.00*
Risk Premia:						
Δ log Nondurable Consumption	0.0029	0.0095	0.0051	0.0105	0.0051	0.0105
	0.0010	*0.0014*	*0.0017*	*0.0018*	*0.0017*	*0.0018*
Δ log Durable Stock	0.0025	0.0123				
	0.0014	*0.0010*				
Δ log Relative Price			-0.0042	-0.0047	-0.0042	-0.0047
			0.0017	*0.0025*	*0.0017*	*0.0025*
Δ Residual			-0.0012	-0.0026	-0.0011	-0.0010
			0.0018	*0.0028*	*0.0018*	*0.0029*

The table reports results of 2-step GMM estimation of E(t)[β Uc(t+1)/Uc(t) R(t+1)] - 1 = 0, where the pricing kernel is linearized. The test assets are the risk-free rate and the 25 Fama-French portfolio (columns I, III, and V) and excess returns on the 25 Fama-French portfolios only (columns II, IV, and VI). Standard errors are Newey-West standard errors with lag length 3. Risk premia are defined as 1/E[M] Cov(f) b, where f=[Δ lnC Δ ln D] or [Δ lnC Δ lnP Δ res] and b=[b2 b3]' or [(b2+b3) -b3 b4]'.

Table 7

Predictive Regressions

Time Period: 1930 until 2002

Housing Service Growth	*k=1*	*k=2*	*k=3*	*k=4*	*k=5*	*k=6*	*k=7*	*k=8*	*k=9*	*k=10*	*k=11*	*k=12*
Residual	-0.06	-0.12	-0.18	-0.24	-0.30	-0.37	-0.43	-0.47	-0.52	-0.56	-0.60	-0.62
	0.03	*0.04*	*0.05*	*0.06*	*0.07*	*0.08*	*0.09*	*0.09*	*0.09*	*0.09*	*0.09*	*0.09*
Adj. R^2	13%	14%	18%	22%	27%	30%	32%	32%	32%	33%	34%	35%

Nondurable Growth	*k=1*	*k=2*	*k=3*	*k=4*	*k=5*	*k=6*	*k=7*	*k=8*	*k=9*	*k=10*	*k=11*	*k=12*
Residual	-0.01	0.02	0.06	0.12	0.16	0.19	0.21	0.23	0.24	0.25	0.27	0.29
	0.05	*0.08*	*0.10*	*0.09*	*0.08*	*0.07*	*0.07*	*0.06*	*0.05*	*0.05*	*0.04*	*0.05*
Adj. R^2	-1%	-1%	0%	3%	8%	12%	15%	16%	16%	18%	20%	23%

GDP Growth	*k=1*	*k=2*	*k=3*	*k=4*	*k=5*	*k=6*	*k=7*	*k=8*	*k=9*	*k=10*	*k=11*	*k=12*
Residual	0.01	0.09	0.18	0.26	0.29	0.31	0.35	0.44	0.56	0.70	0.79	0.83
	0.09	*0.15*	*0.17*	*0.16*	*0.14*	*0.13*	*0.11*	*0.12*	*0.20*	*0.29*	*0.36*	*0.38*
Adj. R^2	-1%	0%	1%	3%	4%	5%	6%	9%	14%	20%	26%	30%

Disposable Income Growth	*k=1*	*k=2*	*k=3*	*k=4*	*k=5*	*k=6*	*k=7*	*k=8*	*k=9*	*k=10*	*k=11*	*k=12*
Residual	0.02	0.09	0.18	0.25	0.28	0.30	0.34	0.40	0.47	0.53	0.56	0.56
	0.07	*0.11*	*0.12*	*0.11*	*0.11*	*0.11*	*0.10*	*0.11*	*0.15*	*0.19*	*0.21*	*0.20*
Adj. R^2	-1%	1%	3%	7%	9%	11%	13%	16%	21%	28%	34%	39%

Risk-free Rate	*k=1*	*k=2*	*k=3*	*k=4*	*k=5*	*k=6*	*k=7*	*k=8*	*k=9*	*k=10*	*k=11*	*k=12*
Residual	0.18	0.30	0.38	0.45	0.53	0.60	0.64	0.64	0.60	0.54	0.46	0.37
	0.07	*0.14*	*0.19*	*0.24*	*0.27*	*0.29*	*0.30*	*0.33*	*0.37*	*0.42*	*0.47*	*0.52*
Adj. R^2	20%	16%	14%	13%	13%	13%	11%	9%	6%	4%	2%	0%

Market Excess Return	*k=1*	*k=2*	*k=3*	*k=4*	*k=5*	*k=6*	*k=7*	*k=8*	*k=9*	*k=10*	*k=11*	*k=12*
Residual	-0.07	-0.20	-0.32	-0.38	-0.60	-0.88	-1.22	-1.59	-1.88	-2.10	-2.25	-2.38
	0.24	*0.41*	*0.51*	*0.48*	*0.47*	*0.51*	*0.53*	*0.52*	*0.53*	*0.60*	*0.71*	*0.82*
Adj. R^2	-1%	-1%	-1%	0%	2%	6%	12%	17%	22%	24%	24%	24%

The table reports results from predictive regressions at different horizons of k years. The left hand side variables are the cumulative per household growth rate of log housing services, log nondurable consumption, log GDP, log disposable income, the cumulative risk-free return, and the cumulative log market excess return. Standard errors are Newey-West standard errors with lag length k. The sample size decreases from 73 (when $k=1$) to 62 observations (when $k=12$)

Table 8

Cross-Sectional Asset Pricing Tests

Time Period: 1930 until 2002 (73 observations)

	I	II	III	IV
Test assets	RF, FF25	FF25	RF, FF25	FF25
Constant	0.889		0.876	
	0.282		*0.291*	
Δ log Nondurable Consumption	-35.780	-24.205	39.380	-38.109
	12.733	*2.465*	*21.569*	*9.113*
Δ log Service Flow	55.491	-22.031		
	17.858	*2.998*		
Δ log Relative User Cost			-34.252	18.826
			14.364	*6.831*
Δ Residual			57.439	-17.632
			18.164	*7.885*
J - test	118.35	21.03	120.44	26.48
	0.00	*0.58*	*0.00*	*0.23*
HJ Distance	1.07	1.37	1.04	0.74
	0.00	*0.00*	*0.01*	*0.04*
Risk Premia:				
Δ log Nondurable Consumption	0.0072	0.0703	0.0080	0.0324
	0.0055	*0.0047*	*0.0055*	*0.0080*
Δ log Service Flow	-0.0079	0.0454		
	0.0042	*0.0031*		
Δ log Relative User Cost			0.0024	-0.0355
			0.0065	*0.0137*
Δ Residual			-0.0172	-0.0411
			0.0068	*0.0124*

The table reports results of 2-step GMM estimation of E(t)[b Uc(t+1)/Uc(t) R(t+1)] - 1 = 0, where the pricing kernel is linearized. The test assets are the risk-free rate and the 25 Fama-French portfolio (columns I and III) and excess returns on the 25 Fama-French portfolios only (columns II and IV). Standard errors are Newey-West standard errors with lag length 3. Risk premia are defined as 1/E[M] Cov(f) b, where f=[Δ lnC Δ ln S] or [Δ lnC Δ lnQ Δ res] and b=[b2 b3]' or [(b2+b3) -b3 b4]'.

Table 9

Summary Statistics: Household-Level Data

	All Homeowners with positive holdings of stocks in 1984, 1989 & 1994 (352 units, 3,536 obs., 17 years)				
Variable	**Median**	**Mean**	**Std Dev**	**Min**	**Max**
Across all households & years:					
Nondurable Food Consumption	4,802	5,203	2,200	560	27,505
Durable Housing Consumption (in 10 USD)	7,500	7,957	3,725	453	23,428
Durable / Nondurable Ratio	1.52	1.72	1.02	0.10	13.85
Nondurable Food Consumption Growth	1.000	1.047	0.350	0.101	5.792
Durable Housing Consumption Growth	1.000	1.034	0.293	0.101	9.571
Family Income Growth	1.027	1.085	1.253	0.004	70.573
Probability of Moving (1 if likely, 0 otherwise)	0%	15%	36%	0%	100%
Age of Household Head	44	44	9	25	61
Number of Family Members	3.00	3.38	1.22	1.00	8.00
Number of Children in Household	1.00	1.15	1.12	0.00	6.00
Change in Number of Family Members	0.00	-0.02	0.46	-3.00	4.00
Change in Number of Children in Household	0.00	-0.03	0.38	-3.00	3.00
Household Head is married	100%	94%	25%	0%	100%
Wife of Head has College Degree	100%	64%	48%	0%	100%
Household owns Business	0%	26%	44%	0%	100%
State Unemployment Rate	7%	7%	2%	2%	16%
Across all households in given year:					
Total Net Wealth in 1984 (in 1,000 USD)	115	172	206	7	1,889
Total Net Wealth in 1989 (in 1,000 USD)	133	219	344	1	4,377
Total Net Wealth in 1994 (in 1,000 USD)	162	239	234	8	1,531
Dollars invested in Stocks in 1984 (in 1,000 USD)	10	31	61	0	498
Dollars invested in Stocks in 1989 (in 1,000 USD)	13	35	60	0	417
Dollars invested in Stocks in 1994 (in 1,000 USD)	28	67	95	0	529
Time series mean of cross-sectional moments:					
Mean of Nondurable Food Consumption Growth		1.045	0.026	0.996	1.097
Mean of Durable Housing Consumption Growth		1.034	0.023	1.012	1.093
Std. Dev. of Durable / Nondurable Ratio		0.980	0.156	0.700	1.413

Table 10

Summary Statistics: Returns

Annual Returns	Mean	Std Dev	Min	Max
Risk-free Rate	1.038	0.040	0.951	1.110
Market Return	1.123	0.152	0.829	1.424
Small / Growth Stocks (FF 1)	1.117	0.232	0.719	1.644
Small / Neutral Stocks (FF 2)	1.178	0.188	0.921	1.666
Small / Value Stocks (FF 3)	1.196	0.172	0.964	1.638
Big / Growth Stocks (FF 4)	1.119	0.181	0.790	1.447
Big / Neutral Stocks (FF 5)	1.128	0.148	0.831	1.370
Big / Value Stocks (FF 6)	1.132	0.132	0.937	1.377

The returns are for 17 years: 1978 - 1987, 1991- 1997. All returns are measured from the beginning of the second quarter until the end of the first quarter of the following year. All returns are deflated with BLS Food Price Index. The risk-free rate is the return on a 1-Year Treasury Bill. The market return is the value-weighted CRSP return. The six size and book-to-market portfolios are from Kenneth French's web site.

Table 11

Aggregate Benchmark

Panel A

Variable	**Mean**	**Std Dev**	**Min**	**Max**
Nondurable Food Consumption	1,324	42	1,265	1,409
Durable Housing Consumption (in 10 USD)	5,811	187	5,440	6,174
Nondurable Food Consumption Growth	1.004	0.018	0.979	1.036
Durable Housing Consumption Growth	1.015	0.015	0.994	1.048

In each year we sum the nondurable and durable values (in levels) across all households. Nondurable consumption is expressed per capita, durable consumption is expressed per household.

Panel B

Parameter / Test	**CES**	**CD**
β	0.792	0.807
	0.187	*0.173*
ω	-0.907	
	3.160	
$\alpha\phi^{\omega}$ if CES; α if CD	1.440	0.723
	6.570	*0.038*
γ	174.000	179.000
	21.000	*18.400*
J-test	4.70	5.10
	0.91	*0.93*
<u>Structural Parameters:</u>		
$1/(1-\omega)$	0.524	
	0.869	
$1/(1+\alpha)$		0.580
		0.013
Wald Test: $\omega = 1-\gamma$	58.200	
	0.000	

Panel C

Parameter / Test	**Linear**
$1+\ln\ \beta$	1.190
	0.189
$b2$ (log nondurable growth)	-34.800
	3.630
$b3$ (log durable growth)	-13.800
	8.680
J-test	5.10
	0.93
<u>Structural Parameters:</u>	
γ	48.600
	11.700
$1/(1+\alpha)$ (for $\omega = 0$)	0.710
	0.113

Panel B and C report results from the 2-step GMM estimation of E(t)[β Uc(t+1)/Uc(t) R(t+1)] - 1 = 0. The underlying utility function is of the CES form or the Cobb-Douglas (CD) form. Results in Panel C are for the linearized pricing kernel. The test assets are the risk-free rate and the six Fama-French portfolios in all cases. The instrument set is a vector of ones and the risk-free rate (lagged once). All standard errors are robust to heteroscedasticity and autocorrelation of up to two lags. Model errors are not demeaned.

Table 12

CES Pricing Kernel

	I	II	III	IV
Test Assets	Rf	Rm	Rf, Rm	Rf, FF6
β	0.881	0.753	0.820	0.738
	0.059	*0.095*	*0.076*	*0.083*
ω	0.349	0.257	0.285	0.184
	0.225	*0.231*	*0.230*	*0.207*
$\alpha\phi^{\omega}$	1.780	1.680	1.810	1.660
	0.974	*0.815*	*1.040*	*0.773*
γ	2.750	3.400	3.010	3.400
	0.688	*0.745*	*0.729*	*0.577*
d1 (d_num)	0.117	0.075	0.102	0.086
	0.040	*0.061*	*0.046*	*0.054*
d2 (d_kids)	-0.054	-0.097	-0.071	-0.107
	0.051	*0.095*	*0.070*	*0.087*
J -test (1)	0.67	0.75	6.00	6.30
	0.88	*0.86*	*0.91*	*1.00*
J -test (2)	0.79	0.82	240.00	2900.00
	0.85	*0.85*	*0.00*	*0.00*
Structural Parameters:				
$1/(1-\omega)$	1.536	1.346	1.399	1.225
	0.531	*0.418*	*0.450*	*0.311*
Wald Test:	8.730	12.900	9.650	18.600
$\omega = 1-\gamma$	*0.003*	*0.000*	*0.002*	*0.000*

The table reports results from the 1-step GMM estimation of E(t)[b Uc(t+1)/Uc(t) R(t+1)] - 1 = 0 for the sample of homeowners that have positive holdings of stocks. The underlying utility function is of the CES form in all cases. The test assets are the risk-free rate (Rf), the market return (Rm) and the six Fama-French portflolios (FF6) as indicated in the first line of the table. The instrument set is a constant, food consumption growth (lagged twice), the ratio of housing over food consumption (lagged twice), income growth (lagged twice), a dummy variable that is one if the household expects to move in the near future (lagged twice), the change in the number of houshold members and childrern under 18 years of age living in the household, the state unemployment rate (lagged once) and the risk-free rate (lagged once). All standard errors are robust to heteroscedasticity and to correlation across households and time (up to two lags). J-test (1) reports the test of overidentification when model errors are not demeaned, J-test (2) when model errors are demeaned.

Table 13

Cobb-Douglas Pricing Kernel

	I	II	III	IV
Test Assets	Rf	Rm	Rf, Rm	Rf, FF6
β	0.893	0.769	0.831	0.750
	0.050	*0.087*	*0.068*	*0.083*
$1/(1+\alpha)$	0.179	0.261	0.227	0.302
	0.232	*0.195*	*0.218*	*0.168*
γ	2.410	3.120	2.740	3.200
	0.606	*0.680*	*0.657*	*0.567*
d1 (d_num)	0.119	0.070	0.103	0.085
	0.046	*0.073*	*0.055*	*0.068*
d2 (d_kids)	-0.056	-0.100	-0.074	-0.110
	0.048	*0.089*	*0.067*	*0.083*
J -test (1)	0.25	0.58	5.80	6.30
	0.88	*0.75*	*0.76*	*1.00*
J -test (2)	0.26	0.60	120.00	4600.00
	0.88	*0.74*	*0.00*	*0.00*

The table reports results from the 1-step GMM estimation of E(t)[β Uc(t+1)/Uc(t) R(t+1)] - 1 = 0 for the sample of homeowners that have positive holdings of stocks. The underlying utility function is of the Cobb-Douglas form. The test assets are the risk-free rate (Rf), the market return (Rm) and the six Fama-French portflolios (FF6) as indicated in the first line of each table. The instrument set is a constant, income growth (lagged twice), a dummy variable that is one if the household expects to move in the near future (lagged twice), the change in the number of houshold members and childrern under 18 years of age living in the household, the state unemployment rate (lagged once) and the risk-free rate (lagged once). All standard errors are robust to heteroscedasticity and to correlation across households and time (up to two lags). J-test (1) reports the test of overidentification when model errors are not demeaned, J-test (2) when model errors are demeaned.

Table 14

Linearized Pricing Kernel

	I	II	III	IV
Test Assets	Rf	Rm	Rf, Rm	Rf, FF6
1+ln β	0.977	0.908	0.939	0.902
	0.020	*0.038*	*0.028*	*0.034*
b2 (log nondurable growth)	-1.920	-3.470	-2.630	-3.370
	1.120	*2.220*	*1.650*	*1.810*
b3 (log durable growth)	-0.716	-1.110	-0.896	-1.140
	0.463	*0.816*	*0.626*	*0.736*
d1 (d_num)	0.272	0.487	0.371	0.472
	0.174	*0.340*	*0.254*	*0.282*
d2 (d_kids)	-0.054	-0.101	-0.076	-0.095
	0.046	*0.102*	*0.073*	*0.085*
J-test (1)	0.15	0.00	4.90	6.30
	0.70	*0.96*	*0.68*	*1.00*
J-test (2)	0.16	0.00	24.00	1900.00
	0.69	*0.96*	*0.00*	*0.00*
Structural Parameters:				
γ	2.640	4.580	3.530	4.510
	1.510	*2.900*	*2.180*	*2.380*
$1/(1+\alpha)$ (for $\omega = 0$)	0.563	0.690	0.646	0.675
	0.199	*0.121*	*0.145*	*0.121*

The table reports results from the 1-step GMM estimation of E(t)[β Uc(t+1)/Uc(t) R(t+1)] - 1 = 0 for the sample of homeowners that have positive holdings of stocks. The underlying utility function is of the linearized form. The test assets are the risk-free rate (Rf), the market return (Rm) and the six Fama-French portflolios (FF6) as indicated in the first line of each table. The instrument set is a constant, a dummy variable that is one if the household expects to move in the near future (lagged twice), the change in the number of houshold members and childrern under 18 years of age living in the household, the state unemployment rate (lagged once) and the risk-free rate (lagged once). All standard errors are robust to heteroscedasticity and to correlation across households and time (up to two lags). J-test (1) reports the test of overidentification when model errors are not demeaned, J-test (2) when model errors are demeaned.

Figure 1
Cointegration Residual (*1952:I to 2002:IV*)