NON-STATIONARITY AND STAGE-OF-THE-BUSINESS-CYCLE EFFECTS IN CONSUMPTION-BASED ASSET PRICING RELATIONS *

Wayne E. FERSON
University of Chicago, Chicago, IL 60637, USA

John J. MERRICK, Jr.
New York University, New York, NY 10006, USA

Received December 1984. Final version received May 1986

Empirical tests of Euler equations relating security returns and consumption usually appear to reject the model. Using a common specification of aggregate preferences and instrumental variables, this paper examines some potential reasons for rejections. The evidence indicates that maintained stationarity assumptions of previous tests fail for post-war U.S. quarterly and monthly data. Shifts in model parameters are found across policy regimes (pre-1951 and post-1979) and across stages of the business cycle (recession versus non-recession). Controlling for some of these factors, less evidence is found against a simple consumption-based asset pricing model in non-recession periods.

1. Introduction

A number of recent studies have estimated relations between financial asset returns and measures of aggregate consumption derived from the first-order conditions of a representative agent's optimization problem (hereafter, marginal rate of substitution models). Such models are attractive because they are elegant, permit estimation of fundamental preference parameters, and imply simple testable expressions. Furthermore, these models avoid the task of specifying a complete model of the economy, assuming instead the form of an aggregate utility function and intertemporal optimality of consumer choice. 1

* Financial support from the Center for Research in Security Prices (Ferson) and the L. Glucksman Institute for Research in Securities Markets (Merrick) is gratefully acknowledged. Neal Horrell provided capable research assistance. Earlier versions of this paper were presented at the University of Illinois, the University of Pennsylvania and the Western Finance Association. We thank Stephen Cecchetti, Robert Cumby, Kenneth French, Michael Gibbons, William Greene, David Modest, Stephen Zeldes, Louis K. C. Chan (the referee), and John B. Long (the editor) for constructive comments and suggestions.

1 For theoretical treatments of such models, see, e.g., Lucas (1978), Cox, Ingersoll and Ross (1985), Breeden and Litzenberger (1978), Brock (1980), Richard and Sundareson (1981), and Grossman and Shiller (1982).

However, virtually every empirical study to date has concluded that marginal rate of substitution models are inconsistent with the data. This paper presents new evidence on potential causes of such rejections.

The criteria for judging whether a marginal rate of substitution model ‘explains’ observed fluctuations in macroeconomic and asset return data have been of two types. First, statistical tests of the restrictions implied by such models have been conducted. Second, point estimates of the parameters have been examined to see if the magnitudes are consistent with a reasonable economic interpretation.

The models that are estimated and tested must, of course, involve a joint hypothesis; departures from the assumptions can influence statistical tests and the magnitudes of parameter estimates. The joint hypotheses in most existing tests of marginal rate of substitution models involve three major assumptions. First, economic outcomes are assumed to be the result of dynamic optimization characterized by an interior solution. Second, a specific relation between aggregate marginal utility and observable quantities is assumed. Third, assumptions are maintained about the relevant probability distribution of consumption, asset returns and some instrumental variables.

The earliest papers in this area examined models with very simple utility functions. Subsequent studies have focused on more general specifications of preferences. For example, recent studies have relaxed the assumption that the representative utility function is separable over time and across different goods. This line of inquiry has produced mixed results. In some cases, evidence of complementarities has been uncovered; in many cases, strong evidence against separability is not found. Typically, however, the use of more elaborate preference structures has failed to remedy the apparent inconsistency of marginal rate of substitution models with consumption and return data. That is, the model’s restrictions usually are rejected and the magnitudes of parameter estimates frequently are unreasonable.

---

2 One approach to relaxing the separable utility assumption introduces additional parameters to model preference complementarities, maintaining the assumption that measured consumption expenditures represent the fundamental objects of preference [e.g., Mankiw, Rotemberg and Summers (1984)]. A second approach introduces a technology for transforming measured expenditures into a flow of services from which utility is assumed to be derived [e.g., Kydland and Prescott (1982), Dunn and Singleton (1986), Eichenbaum, Hansen and Singleton (1986)].

3 See, for example, Kydland and Prescott (1982) or Eichenbaum, Hansen and Singleton (1986).

4 For example, see Eichenbaum and Hansen (1985). Also, point estimates of utility function parameters not reliably different from values which imply simpler, separable models are found in the studies of Mankiw, Rotemberg and Summers (1985), Eichenbaum and Hansen (1983), and Dunn and Singleton (1986).

5 Dunn and Singleton (1986) find that a model with complementarity of durables and non-durables consumption and a single treasury bill return is not rejected. Using more than a single asset in the tests, however, results in a rejection.

This paper examines other elements of the joint hypothesis which previous studies have rejected. The evidence indicates that maintained statistical assumptions may be violated in the data. Shifts in model parameters inconsistent with the assumptions of previous tests are found across stages of the business cycle, policy regimes and over calendar time. Controlling for some of these parameter shifts, no strong evidence is found against a simple version of the marginal rate of substitution model during non-recession periods. Thus, the ability to detect departures from the model’s restrictions may be related to consumer behavior during recessions.

The paper is organized as follows. The next section reviews marginal rate of substitution models and further motivates the investigation. Section 3 discusses the data. Section 4 presents the empirical results, and section 5 summarizes the findings and conclusions.

2. Overview of the model

Consider a single-good economy with a representative consumer who chooses at each date investments in financial assets and a planned consumption path to maximize

\[ E\left( \sum_{j \geq 0} e^{-\rho j} U(C_{t+j}) | \Omega_t \right) \],

(1)

where \( U(\cdot) \) is a single-period utility function, \( C_{t+j} \) is aggregate real per capita consumption at date \( t+j \), \( \rho \) is the rate of time discount, and \( E(\cdot | \Omega_t) \) represents mathematical expectation conditional on the information \( \Omega_t \) available at time \( t \).

An expected utility-maximizing consumer at an interior optimum will choose consumption and investment such that the marginal utility of time \( t \) consumption is equal to the discounted expected marginal utility of \( t+j \) consumption attainable through each asset’s payoff. Thus, optimality requires for any asset (setting \( j = 1 \))

\[ E\{e^{-\rho}[U_c(C_{t+1})/U_c(C_t)] R_{t+1} | \Omega_t \} = 1, \]

(2)

where \( U_c(\cdot) \) denotes the partial derivative of \( U(\cdot) \) and \( R_{t+1} \) is the random gross ‘real’ asset return (end of period value divided by time \( t \) price in consumption units).

Empirical tests of marginal rate of substitution models employ a specific representation of aggregate preferences. A common specification and the one examined here is the iso-elastic utility function

\[ U(C) = (1 - \alpha)^{-1} [C^{1-\alpha} - 1], \]

(3)
where \( \alpha > 0 \) is a parameter interpreted either as the coefficient of relative risk aversion for timeless gambles [e.g., Pratt (1964)] or as the inverse of the intertemporal elasticity of substitution [e.g., Hall (1985)]. This specification is chosen because most studies that have rejected the model use preference functions consistent with (3) and because the empirical evidence does not indicate that such rejections are avoided by using more general utility functions.

One approach to estimating and testing marginal rate of substitution models assumes normality and employs maximum likelihood methods on the resulting closed-form relations [e.g., Hansen and Singleton (1983) or Ferson (1983)]. Assume that the conditional distribution of \( \ln(C_{t+1}/C_t) \) and \( \ln(R_{t+1}) \) given \( Z_t \) is normal, where \( Z_t \) is a vector of predetermined, instrumental variables which are elements of \( \Omega_t \) and \( \ln(\cdot) \) denotes natural logarithm. Substituting eq. (3) into (2), applying the law of iterated expectations and then the normal moment generator implies

\[
E(r_{t+1} | Z_t) = \Psi + \alpha E(\Delta \ln C_{t+1} | Z_t),
\]

where

\[
\Psi = -\frac{1}{2} \text{var}(r_{t+1} - \alpha \Delta \ln C_{t+1} | Z_t) + \rho,
\]

\[
\Delta \ln C_{t+1} = \ln(C_{t+1}) - \ln(C_t),
\]

\[
r_{t+1} = \ln(R_{t+1}).
\]

If the joint distribution of \( \{\Delta \ln C_{t+1}, r_{t+1}, Z_t\} \) is an independent stationary normal, then \( \Psi \) in eq. (4) and the parameters of the following bivariate regression system will be constant:

\[
\Delta \ln C_{t+1} = \beta_{c0} + \beta_c Z_t + e_{c,t+1},
\]

\[
r_{t+1} = \beta_{r0} + \beta_r Z_t + e_{r,t+1}.
\]

In eq. (5), \( \beta_c \) and \( \beta_r \) are regression coefficient vectors, \( \beta_{c0} \) and \( \beta_{r0} \) are scalars, and the regression errors \( e_{c,t+1} \) and \( e_{r,t+1} \) are assumed to have contemporaneous covariance matrix \( \Sigma \).

The theoretical model (4) places cross-equation restrictions on the parameters of the regression model (5):

\[
\beta_c = \alpha \beta_r, \quad \beta_{r0} = \Psi + \alpha \beta_{c0}.
\]

The determinant of the residual covariance matrix of the unrestricted system (5) is compared with that of the restricted system, imposing (6), using a
standard likelihood ratio test. If the imposition of the cross-equation restrictions reduces the model’s explanatory power, then the test should produce a ‘large’ value of the likelihood ratio statistic.\(^7\) While the maximum likelihood approach makes strong assumptions about the probability distribution of consumption and real returns, it produces easily interpretable results.\(^8\) For example, the individual regression equations may be examined for diagnostic purposes and for evidence of correlation of expected real returns and consumption growth rates with economic information variables.\(^9\)

In summary, tests of marginal rate of substitution models typically reject a joint hypothesis including (1) a specification of aggregate preferences, (2) interior optimality, and (3) assumptions about the joint distribution of consumption, real returns and the set of instrumental variables. In view of the failure of previous studies to explain rejections by changing the utility function assumptions, the present investigation considers alternative explanations.

3. The data

Both quarterly and monthly data are employed. The quarterly data are aggregate per capita consumption expenditures (non-durables plus services) and real treasury bill returns for the 1949–1983 sample period. The monthly data are for the 1959–1983 period. All of the series correspond closely to data employed in previous studies.

Monthly U.S. treasury bill rates and three-month bill rates sampled quarterly are from Salomon Brothers’ \textit{Analytical Record of Yields and Yield Spreads}. Monthly and quarterly inflation rates are measured using the deflators for personal consumption expenditures (non-durables and services). Real interest rates are the difference between continuously-compounded nominal bill rates and inflation rates. Consumption data are from the \textit{Survey of Current Business} and \textit{Business Statistics Supplements}. Population data are from the Bureau of the Census.

Although previous studies of marginal rate of substitution models have included real rates of return on common stock and long-term bond portfolios, this study focuses on the real return to treasury bills. The low ratio of explained variability to total variability for long-term assets means that tests of constraints on the coefficients of these return equations would be expected to have relatively low power. Recall that much of the early ‘efficient markets’

\(^7\)Typically, studies have employed the asymptotic chi-square distribution to evaluate the sample likelihood ratio statistics. We also follow this convention.

\(^8\)An alternative approach is generalized method of moments as developed by Hansen (1982) and applied by Hansen and Singleton (1982) and others. Evidence relevant to this approach is discussed in section 4.4 below.

\(^9\)See Huizinga and Mishkin (1984) for a recent example of such interpretations of real return equations like the one in system (5).
literature was consistent with the hypothesis that expected stock returns are constant. Studies of marginal rate of substitution models using single stock or long-term bond indexes less often rejected the overidentifying restrictions than when using bills [see, e.g., Hansen and Singleton (1983), Dunn and Singleton (1983, 1986)]. Thus, an explanation for rejections of the model using bills is of particular interest.

There are many well-known deficiencies in the available consumption and inflation data which could, in principle, lead to a rejection of the null hypothesis. For example, in addition to the theoretical deficiencies of price indexes, the deflator for personal consumption expenditures suffers from infrequent data sampling and publication lag. For these data, the publication lag can be seven weeks or longer. The data employed in this and previous studies are seasonally-adjusted, which may be inappropriate. The implicit assumption of the existence of a representative consumer [see Brennan and Kraus (1978) for necessary conditions] is another potential source of misspecification.

The usual specification of the marginal rate of substitution model ignores taxation, but taxable rates of return are measured. Several studies have been sensitive to this potential problem. For example, Grossman and Shiller (1981), Grossman, Melino and Shiller (1985), Mankiw, Rotemberg and Summers (1984), and Rotemberg (1984), have used measures of after-tax returns. These authors report little difference in results from those using before-tax returns. In light of this, and because no clearly appropriate measure of after-tax returns is available, before-tax returns are employed here.

4. Empirical results

4.1. Selection of instrumental variables

A potentially important initial decision in an empirical examination of marginal rate of substitution models is the selection of the instrumental variables $Z_r$. These should be correlated with actual market expectations (given $\Omega_r$) and be public information at time $t$. Most studies to date have included lagged values of the dependent consumption and return variables. As Ferson (1983) and Hall (1985) note, measurement error due to infrequent sampling and data averaging may be a problem. These authors choose not to include the first lag of dependent variables as instruments in an attempt to minimize such potential misspecification and to allow for publication lag. Subsequent tables report results using instruments consisting of the second and third lags of consumption growth and real treasury bill returns, plus a constant intercept. This is representative of the instruments previous studies have employed.\textsuperscript{10}

\textsuperscript{10} Most of the tests reported below were replicated including first lagged values as instruments. The results, where they differ, are noted.
4.2. Regression results

Tests of marginal rate of substitution models have maintained assumptions about the stochastic process of consumption, real returns and the instrumental variables. In particular, a bivariate regression approach assumes that the coefficients in (5) are constant. If this assumption fails, parameter restrictions given by eq. (6) may not hold even if the underlying theory is correct. Previous studies have rejected the joint hypothesis, but few have reported diagnostics verifying their statistical assumptions.

Tables 1 and 2 present some summary statistics and examine the individual consumption growth and real treasury bill return regressions of (5) for changing parameters over calendar time and across stages of the business cycle. This is accomplished by allowing the regression coefficients to shift using dummy variables. For example, to investigate parameter shifts over calendar time, a dummy variable $D_t$ is defined to equal one in the first half of the sample and zero in the second half. Thus, tables 1 and 2 examine regressions of the form

$$\Delta \ln C_{t+1} = \beta_{c0} + \beta_c Z_t + \beta_{dco} D_t + \beta_{dt} D_t Z_t + u_{c,t+1},$$

$$r_{t+1} = \beta_{ro} + \beta_r Z_t + \beta_{dro} D_t + \beta_{dr} D_t Z_t + u_{r,t+1}.$$  \(7\)

Finding reliably non-zero coefficients on the variables $\{D_t, Z_t\}$ in regression (7) indicates that the parameters of regression (5) are different in the first and second halves of the sample.

A second definition of the dummy variable $D_t$ allows the coefficients of (5) to differ in the pre-March 1951 and post-October 1979 period from their values in the intervening period. This period was isolated because of arguments that the underlying behavior of expected real returns on bills may have experienced 'structural shifts' associated with the Treasury–Federal Reserve Accord of 1951 [Fama (1975)] and with the 1979 change in targets employed by the Federal Reserve in the implementation of monetary policy [e.g., Huizinga and Mishkin (1984)].

A third definition of the dummy variable $D_t$ allows the regression coefficients to differ in recessions from their values in non-recession periods. The possibility that asymmetries exist in economic time series over the business cycle has been of recent interest [e.g., Neftci (1984)]. Such patterns in the data may shed further light on the relation of consumption and asset returns. For example, a number of authors have emphasized that market imperfections may destroy the interior optimum assumed in the marginal rate of substitution model.\(^{11}\) Possibly, the effects of imperfections like 'liquidity constraints' are more important (i.e., more easily detected in aggregate data) during recession

---

\(^{11}\)See, for example, Hall (1978), Mankiw, Rotemberg and Summers (1984), Brown (1986), Zeldes (1985), and Rotemberg (1984).
periods than at other times. Flavin (1984) makes a similar argument in the context of testing the permanent income hypothesis.

Recession periods are defined as the inclusive intervals between a business cycle peak and its trough, as designated by the National Bureau of Economic Research (NBER). The dummy variable here indicates a shift in the stage of the business cycle with a one-period delay. For example, $\Delta \ln C_{t+1}$ is anticipated to be a recession or non-recession observation on the basis of the NBER classification prevailing in period $t$. Under this scheme, 36 of the 140 quarterly observations (table 1) are recession quarters. The remaining observations are classified as ‘non-recession’ quarters. For the monthly data (table 2), 59 out of 296 observations are classified as months of recession; the remaining 237 are non-recession months.\footnote{Neftci (1984) emphasizes that ex post sample classification can bias test results if based on sample properties correlated with the effects to be examined. Although the NBER appears to determine business cycle stages ex post, there is evidence that leading indicators and other information can be used to anticipate reliably a change in the business cycle [see, e.g., Moore and Zarnowitz (1982)]. The recession dummy variable employed here embodies a simple ‘no change’ model based on the current stage of the business cycle. This is probably conservative with respect to investors’ forecasting abilities. The approach maintains the assumption that the dummy variable is ex ante information. (See footnote 19.)}

Tables 1 and 2 show (in the left-hand columns) the sample means and standard deviations of consumption growth and real treasury bill returns over selected subsamples. The sample statistics indicate that mean consumption growth is consistently lower in recessions than in non-recession periods; however, average real treasury bill returns do not appear to display a similar pattern. This is characteristic of both monthly and quarterly data.

The finding that mean consumption growth rates are different given a recession is not a surprising result. If the NBER determines business cycle stages by examining the levels of economic time series that rise and fall in a manner similar to consumption levels, then consumption growth rates should be lower between a business cycle peak and its trough than at other times.\footnote{The NBER actually determines business cycle stages by a consensus impression of the levels of several aggregate economic time series representing employment, unemployment, income, sales and output [see, e.g., Moore (1983)].} However, if the model is correct, then changes in anticipated consumption growth rates should be mirrored in anticipated real returns over any subsample. Divergent behavior of expectations of the two variables would suggest a failure of the model’s restrictions associated with the business cycle.

Recall that a maintained assumption of the tests is that the regression coefficients in (5) are constant. The six right-hand columns of tables 1 and 2 test the hypothesis of constant regression parameters in the individual equations of (5) against the alternative that the dummy variables described above can detect shifts in the parameters. The evidence in the upper portion of table 1 indicates that the stage-of-the-business-cycle dummies are reliably non-zero
Table 1

Consumption growth and real treasury bill return regressions: Tests of constant parameters across stages of the business cycle and over calendar time. Quarterly data, 1949:1–1983:4 and 1951:3–1979:3. The regressions of system (5) are nested in OLS regressions of the form:

\[ \Delta \ln \hat{G}_{t+1} = \hat{\beta}_a + \hat{\beta}_r Z_t + \hat{\beta}_{d,a} D_t + \hat{\beta}_{d} D_t Z_t + \hat{u}_{t+1} \]

\[ r_{t+1} = \hat{\beta}_a + \hat{\beta}_r Z_t + \hat{\beta}_{d,a} D_t + \hat{\beta}_{d} D_t Z_t + \hat{u}_{t+1} \]

\(^a\Delta \ln \hat{G}_{t+1}\) is the continuously-compounded growth rate of real non-durables and services consumption expenditures per capita and \(r_{t+1}\) is the real return on a three-month treasury bill. \(Z_t\) is a vector consisting of lagged values of real consumption growth and bill returns at lags 2 and 3. \(\hat{R}^2\) is the adjusted coefficient of determination and \(p\)-value is the right-tail probability value of the likelihood ratio test. NA indicates not applicable to this sample.

\(^bp\)-value less than 0.0001.

<table>
<thead>
<tr>
<th>Dependent variable</th>
<th>Sample mean and standard deviation (in parentheses) of dependent variable (annual percentage rates)</th>
<th>Tests for significance of regression (5) with no dummy variables</th>
<th>Regression (R^2)-squares and likelihood ratio tests of zero coefficients on dummy variables</th>
<th>Calendar time dummies</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Non-recession subsample</td>
<td>Recession subsample</td>
<td>(\hat{R}^2)</td>
<td>(p)-value</td>
</tr>
<tr>
<td>Consumption growth</td>
<td>2.31 (2.08)</td>
<td>0.64 (2.37)</td>
<td>0.052</td>
<td>0.020</td>
</tr>
<tr>
<td>Real treasury bill return</td>
<td>0.43 (2.48)</td>
<td>2.00 (3.12)</td>
<td>0.349</td>
<td>(\ni)</td>
</tr>
</tbody>
</table>

1949:1–1983:4 (104 non-recession and 36 recession observations)

1951:3–1979:3 (91 non-recession and 22 recession observations)
### Table 2


\[
\Delta \ln C_{t+1} = \beta_0 + \beta_1 Z_{t} + \beta_2 D_t + \beta_3 D_tZ_{t} + u_{t, t+1},
\]

\[
r_{t+1} = \beta_0 + \beta_1 Z_{t} + \beta_2 D_t + \beta_3 D_tZ_{t} + u_{t, t+1}.
\]

<table>
<thead>
<tr>
<th>Dependent variable</th>
<th>Sample mean and standard deviation (in parentheses) of dependent variable (annual percentage rates)</th>
<th>Tests for significance of regression (5) with no dummy variables</th>
<th>Regression $R$-squares and likelihood ratio tests of zero coefficients on dummy variables</th>
<th>Calendar time dummies</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Non-recession subsample</td>
<td>Recession subsample</td>
<td>$R^2$</td>
<td>$p$-value</td>
</tr>
<tr>
<td>1959:5–1983:12 (237 non-recession and 59 recession observations)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Consumption growth</td>
<td>2.51</td>
<td>0.51</td>
<td>0.032</td>
<td>0.009</td>
</tr>
<tr>
<td></td>
<td>(5.27)</td>
<td>(4.78)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Real treasury bill return</td>
<td>0.90</td>
<td>1.75</td>
<td>0.209</td>
<td>$a$</td>
</tr>
<tr>
<td></td>
<td>(2.63)</td>
<td>(4.28)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>1959:5–1979:9 (208 non-recession and 37 recession observations)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Consumption growth</td>
<td>2.61</td>
<td>0.66</td>
<td>0.032</td>
<td>0.018</td>
</tr>
<tr>
<td></td>
<td>(5.39)</td>
<td>(4.83)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Real treasury bill return</td>
<td>0.50</td>
<td>$b$</td>
<td>0.090</td>
<td>$a$</td>
</tr>
<tr>
<td></td>
<td>(2.18)</td>
<td>(3.00)</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

$^a$Δ ln $C_{t+1}$ is the continuously-compounded growth rate of real non-durables and services consumption expenditures per capita and $r_{t+1}$ is the real return on a one-month treasury bill. $Z_t$ is a vector consisting of lagged values of real consumption growth and bill returns at lags 2 and 3. $R^2$ is the adjusted coefficient of determination and $p$-value is the right-tail probability value of the likelihood ratio test. NA indicates not applicable to this sample.

$p$-value less than 0.0001.
in the quarterly consumption growth equation. However, the hypothesis of constant parameters over the business cycle cannot be rejected for the quarterly treasury bill equation. The monthly regressions in table 2 show little evidence of parameter shifts associated with recessions.\textsuperscript{14} Tables 1 and 2 indicate parameter shifts over time, more reliably in the real treasury bill return equations. Shifts across Fed ‘policy regimes’ seem to be more strongly indicated than are changes over pure calendar time.\textsuperscript{15}

The lower portions of tables 1 and 2 repeat the tests for constant parameters in shorter subsamples omitting the pre-Accord (March 1951) and post-target shift (October 1979) periods. Stage-of-the-business-cycle effects on the consumption regression coefficients are still evident in the quarterly data over this period (p-value of 0.004), and the monthly treasury bill regressions continue to indicate non-constant parameters over the first and second halves of the sample.\textsuperscript{16} Furthermore, stage-of-the-business-cycle effects on the monthly treasury bill equation are detectable having removed the influence of the pre-1951, post-1979 periods (the dummy is significant at the five percent level).

Some of the previous rejections of marginal rate substitution models reported in the literature are based on data limited to the pre-1979 period. The above evidence indicates that a test including both the pre-1979 and the subsequent period in the sample would have suffered a violation of the maintained assumptions using a typical set of instruments. The evidence also indicates that limiting a sample to the pre-1979 period is not sufficient to avoid these misspecifications. Even in this period, there is evidence of shifts in the parameters of regression (5) associated with the stage of the business cycle and calendar time.

\textsuperscript{14}Three of the eight recessions in the overall sample occur prior to 1959 and are not included in the monthly sample.

\textsuperscript{15}The likelihood ratio test for the significance of the dummy variables assumes a constant error covariance matrix and no residual autocorrelation, which may be inappropriate. However, the individual \(t\)-statistics using White’s (1980) heteroskedasticity-consistent standard errors were very similar in most cases to the OLS \(t\)-statistics. The residual autocorrelations in tables 1 and 2 are often large. [For example, Box–Pierce \(Q\)-statistics for the first three autocorrelations – asymptotically \(\chi^2(3)\) if the true autocorrelations are zero – were 4.75 in the quarterly consumption equation and 26.01 in the real return equations for 1949–1983.] Replicating these tests using instruments which included the first-lagged dependent variables produced smaller and generally insignificant residual autocorrelations (the \(Q\)-statistics were 0.91 and 2.51, respectively, in the above mentioned example). This alternative selection of instruments had little impact on the results, with one exception. The exception is that when the instrumental variables are the first two lags of consumption growth and real returns, the business cycle dummies in the monthly consumption equation of table 2 appear significant at the 0.02 level and the calendar time dummies produce a \(p\)-value of 0.03.

\textsuperscript{16}Since the monthly data begins in 1959, the dummy variable tests only a post-1979 effect in the monthly data. The first-versus-second-half time dummies for the pre-1979 subsample indicate the same dates as those for the full sample (i.e., 1966 : 3 in the quarterly sample and 1971 : 9 in the monthly).
The theory used to derive eq. (4) implies restrictions like eq. (6) for regression coefficients using any set of predetermined instruments for which a regression model like (5) is well-specified. Thus, if the parameters of regression (5) shift, for example, with the stage of the business cycle, a regression like (7) can potentially form the basis of an empirical test.

Regression (7) including the business cycle dummy was examined for evidence of misspecification by extending the dummy variable techniques of the preceding tables. An additional time dummy variable having a value of one in the first half of the sample and zero in the second half was included to allow each of the regression coefficients of (7) to differ in these two subperiods (including the coefficients on \( D_t \) and \( D_t Z_t \), where \( D_t \) is the business cycle dummy). For quarterly data, these tests did not reject the hypothesis that the regression model (7) has constant parameters (the \( p \)-values of this test over 1951:3–1979:3 were 0.447 in the consumption equation and 0.252 in the treasury bill equation). For monthly data (1959:5–1979:9), the consumption equation \( p \)-value was 0.938, but there was still evidence (\( p \)-value less than 0.0001) that the relation of real treasury bill returns to the instruments of regression (7) changes over this sample period. Thus, even a monthly regression like (7) which attempts to control for stage of the business cycle effects may be subject to violations of the maintained stationarity assumptions of the tests.

4.3. Tests of restrictions of a marginal rate of substitution model

Since previous tests of marginal rate of substitution models have rejected a joint hypothesis including stationarity, the results of such tests could be different if a model controlling for some of the observed parameter shifts is employed. This section investigates the sensitivity of the tests to the choice of time period and to the stage of the business cycle. Sensitivity to the time period is examined by conducting tests for the full 1949:1–1983:4 period (1959:5–1983:12 for monthly data) and also for the 1951:3–1979:3 sub-period (1959:5–1979:9 for monthly data). To control for shifts in the coefficients of regression (5) associated with the stage of the business cycle, the bivariate regression (7) where \( D_t \) is the business cycle dummy variable is employed as the alternative hypothesis. The null hypothesis restricts the coefficients of (7) to be proportional like in eq. (6).

In addition to allowing coefficient shifts over the business cycle, regression (7) permits an investigation of the model’s restrictions conditional on a given stage of the business cycle. As previously suggested, the effects of imperfections like ‘liquidity constraints’ or other departures from the model may be more easily detected in aggregate data during recession periods than at other times. Thus, tables 3 and 4 present likelihood ratio tests (LRTs) in which the cross-equation restrictions are selectively imposed. When all of the slope
Preference parameter estimates and likelihood ratio tests of cross-equation restrictions of the marginal rate of substitution model (4) using quarterly data. The unrestricted bivariate regression system is of the form:

\[
\Delta \ln C_{t+1} = \beta_{t0} + \beta_{t} Z_{t} + \beta_{d,t0} D_{t} + \beta_{d,t} D_{t} Z_{t} + u_{t},
\]

\[
r_{t+1} = \beta_{r0} + \beta_{r} Z_{t} + \beta_{r,d0} D_{t} + \beta_{r,d} D_{t} Z_{t} + u_{r,t}.
\]

<table>
<thead>
<tr>
<th>Sample period</th>
<th>Restrictions tested</th>
<th>Point estimate of relative risk aversion, ( \alpha ) (standard error)</th>
<th>Point estimate of intercept, ( \Psi ) (standard error)</th>
<th>P-values for tests of restrictions</th>
</tr>
</thead>
<tbody>
<tr>
<td>1949:1–1983:4</td>
<td>( \beta_{t} = \alpha \beta_{t} ), ( \beta_{r0} = \alpha \beta_{r0} ), ( \beta_{ro} = \Psi + \alpha \beta_{ro} )</td>
<td>-4.50 (3.13)</td>
<td>0.0233 (0.0152)</td>
<td>( \text{---} ) ( b )</td>
</tr>
<tr>
<td>(140 obs.)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1951:3–1979:3</td>
<td>( \beta_{r} = \alpha \beta_{r} ), ( \beta_{r0} = \alpha \beta_{r0} ), ( \beta_{ro} = \Psi + \alpha \beta_{ro} )</td>
<td>3.16 (1.55)</td>
<td>-0.0169 (0.0087)</td>
<td>0.044</td>
</tr>
<tr>
<td>(113 obs.)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

\( Z_{t} \) is the instrumental variables vector consisting of the second and third lagged values of the real three-month treasury bill return, \( r_{t} \), and the continuously-compounded growth rate of real consumption expenditures per capita for non-durable goods plus services, \( \Delta \ln C_{t} \). \( D_{t} \) is the business cycle dummy variable (\( D_{t} = 1 \) if recession, \( D_{t} = 0 \) if non-recession). The alternative hypothesis is the unrestricted bivariate regression shown above and as (7). The null hypothesis generating the restrictions is \( E(r_{t+1} Z_{t}) = \bar{\Psi} + E(\alpha \Delta \ln C_{t+1} Z_{t}) \), where \( \alpha \) is the coefficient of relative risk aversion. The \( p \)-value is the probability, based on the asymptotic chi-square distribution, of obtaining a likelihood ratio larger than the sample value.

\( ab \)P-value less than 0.0001.

coefficients of (7) are restricted, the null hypothesis is that the model holds (possibly with different regression coefficients) in both recession and non-recession periods. This hypothesis is strongly rejected in the full sample. The right-tail \( p \)-values of the LRT are 0.015 in the monthly data and less than 0.0001 in the quarterly data. The 1951:3–1979:3 subperiod using quarterly data produces a \( p \)-value of 0.001; with monthly data (1959:5–1979:9) the \( p \)-value of the LRT is 0.127.\(^{17}\) Note that the estimates of relative risk aversion in tables 3 and 4 appear more reasonable in the shorter subsample than over the full period. The full-period estimates are negative (implying convex or risk-loving aggregate utility). In the shorter subsamples, estimates of risk

\(^{17}\)This \( p \)-value is larger than reported in some previous studies using similar data, perhaps because the model allows the regression coefficients to differ over the business cycle stages. Replicating these tests using eq. (5), which assumes constant regression coefficients over the business cycle, the \( p \)-value of this monthly test for 1959–1979 declines to 0.041.
Preference parameter estimates and likelihood ratio tests of cross-equation restrictions of the marginal rate of substitution model (4) using monthly data. The unrestricted bivariate regression system is of the form:

\[
\Delta \ln C_{t+1} = \beta_{\alpha} + \beta_z Z_t + \beta_{dc} D_t + \beta_{de} D_t Z_t + u_{c,t+1}.
\]

\[
\tau_{t+1} = \beta_{\tau} + \beta_z Z_t + \beta_{d\tau} D_t + \beta_{de} D_t Z_t + u_{\tau,t+1}.
\]

<table>
<thead>
<tr>
<th>Sample period</th>
<th>Restrictions tested</th>
<th>Point estimate of relative risk aversion, ( \alpha ) (standard error)</th>
<th>Point estimate of intercept, ( \Psi ) (standard error)</th>
<th>( p )-values for tests of restrictions</th>
</tr>
</thead>
<tbody>
<tr>
<td>1959:5–1983:12</td>
<td>( \beta_{\alpha} = \alpha \beta_{\alpha} ) ( \beta_{\delta} = \alpha \beta_{\delta} ) ( \beta_{de\alpha} = \alpha \beta_{de\alpha} ) ( \beta_{de} = \Psi + \alpha \beta_{de} )</td>
<td>-57.64 (757.69)</td>
<td>0.1025 (1.335)</td>
<td>0.015</td>
</tr>
<tr>
<td>1959:5–1979:9</td>
<td>( \beta_{\alpha} = \alpha \beta_{\alpha} ) ( \beta_{\delta} = \alpha \beta_{\delta} ) ( \beta_{de\alpha} = \alpha \beta_{de\alpha} ) ( \beta_{de} = \Psi + \alpha \beta_{de} )</td>
<td>4.90 (9.47)</td>
<td>-0.0095 (0.0200)</td>
<td>0.065</td>
</tr>
<tr>
<td>(296 obs.)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>( \beta_{\delta} = \alpha \beta_{\delta} ) ( \beta_{de\alpha} = \alpha \beta_{de\alpha} ) ( \beta_{de} = \Psi + \alpha \beta_{de} )</td>
<td>1.24 (0.74)</td>
<td>-0.0020 (0.0015)</td>
<td>0.127</td>
</tr>
<tr>
<td>1959:5–1979:9</td>
<td>( \beta_{\alpha} = \alpha \beta_{\alpha} ) ( \beta_{\delta} = \alpha \beta_{\delta} ) ( \beta_{de\alpha} = \alpha \beta_{de\alpha} ) ( \beta_{de} = \Psi + \alpha \beta_{de} )</td>
<td>0.31 (0.17)</td>
<td>-0.0003 (0.0004)</td>
<td>0.199</td>
</tr>
<tr>
<td>(245 obs.)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

\( Z_t \) is the instrumental variables vector consisting of the second and third lagged values of the real one-month treasury bill return, \( \tau_{t+1} \), and the continuously-compounded growth rate of real consumption expenditures per capita for non-durable goods plus services, \( \Delta \ln C_{t+1} \). \( D_t \) is the business cycle dummy variable (\( D_t = 1 \) if recession, \( D_t = 0 \) if non-recession). The alternative hypothesis is the unrestricted bivariate regression shown above and as (7). The null hypothesis generating the restrictions is \( E(\tau_{t+1}|Z_t) = \Psi + E(\Delta \ln C_{t+1}|Z_t) \), where \( \alpha \) is the coefficient of relative risk aversion. The \( p \)-value is the probability, based on the asymptotic chi-square distribution, of obtaining a likelihood ratio larger than the sample value.

Aversion equal a more reasonable 3.93 (quarterly data) and 1.24 (monthly data). Thus, the evidence against the joint hypothesis appears slightly weaker when the pre-Accord (1951) and post-target shift (1979) periods are not admitted to the sample.

Tables 3 and 4 also present tests which impose the proportionality restrictions on the parameters of (7) conditional on \( D_t = 0 \) (i.e., \( \beta_{\delta} = \alpha \beta_{\delta} \), \( \beta_{de\alpha} = \Psi + \alpha \beta_{de\alpha} \); the other parameters are unrestricted). In these tests the regression coefficients may differ in recessions from non-recessions and the hypothesis is that the model holds in non-recession periods (but may not hold in recessions). Applying this test to the full samples results in much weaker evidence against the restrictions. The \( p \)-value of the LRT is 0.044 with quarterly data and 0.065 with monthly data.

The lower panels of tables 3 and 4 test the hypothesis that the marginal rate of substitution model holds in non-recessions during the 1951:3–1979:3 period (1959:5–1979:9 for monthly data), allowing the regression coefficients
to be different and unrestricted during recessions. These tests produce virtually no evidence against the null hypothesis. The \( p \)-value of the \( \chi^2 \) statistic is 0.213 in the quarterly data and 0.199 in the monthly data. The estimates of relative risk aversion appear economically reasonable and have relatively small standard errors.\(^{18}\)

So far, the evidence suggests that rejections of the simple marginal rate of substitution model with monthly and quarterly consumption and real treasury bill return data reflect at least two (possibly related) phenomena. The first is a failure of a maintained assumption of constant regression parameters, given a typical set of instrumental variables like the \( Z_t \) examined here. The evidence also points to a second potential reason for previous rejections of the model. It appears that the evidence against the model is much stronger over a full business cycle than during non-recession periods. The tests produce virtually no evidence against the model when the economy is not in recession. These results are particularly striking in view of the simple utility function employed and the evidence that, at least for the monthly data, even regression (7) may violate the statistical assumptions. These sources of potential misspecification could have produced rejections even in non-recession periods.\(^{19}\)

4.4. An alternative approach

A further examination of the consumption and real return data provides additional evidence related to the model’s failure across the stages of the business cycle. Consider the following regression:

\[
X_{t+1} = \gamma_0 + \gamma_x Z_t + \varepsilon_{X_{t+1}},
\]

where

\[
X_{t+1} = \left( \frac{C_{t+1}}{C_t} \right)^{-\alpha} R_{t+1},
\]

\(^{18}\)The point estimates of \( \Psi \), the intercept term in eq. (4), frequently are negative in tables 3 and 4 even when the hypothesis is not rejected by the LRT. Given the sample values of \( \text{var}(r_{t+1} - \alpha \ln C_{t+1} | Z_t) \), these \( \Psi \) estimates imply point estimates of \( \rho \), the rate of time preference, which also are negative. However, these coefficients are not estimated very precisely.

\(^{19}\)Three further checks indicated some robustness of these results. In the first, the tests in tables 3 and 4 were repeated using the first two lags of consumption growth and the real treasury bill returns as the instrumental variables. The results were very similar. The most dramatic difference was that the \( p \)-value of the LRT for the quarterly full sample, non-recession test became 0.144 instead of 0.044. This reinforces the results presented in the text. A second series of experiments did not indicate that differences in the number of observations per se is driving the results. For example, using a dummy variable which selected observations randomly in the same numbers as did the recession dummies \( D_t \) over 1951:3–1977:3 (i.e., \( D_t = 1 \) for 22 randomly selected observations, \( D_t = 0 \) for the other 91 observations), the \( p \)-value of the \( \chi^2 \) test was 0.018 in contrast to the 0.213 reported in table 3 for the non-recession test. A third experiment was to omit sample points where the value of the recession dummy variable changed and to repeat the tests using this subsample. There was little change in results to indicate that these are influential points.
and $\alpha$ is the coefficient of relative risk aversion. The marginal rate of substitution model (2) and (3) implies that the expected value of $X_{t+1}$ given information available to the market at time $t$ is equal to the constant, $e^\rho$. If expectations are rational, then any variation over time in $X_{t+1}$ will be unpredictable using a predetermined $Z_t$. Thus, the model implies $\gamma_x = 0$ and $\gamma_0 = e^\rho$ in regression (8), given the correct value of relative risk aversion. Such tests are reminiscent of early 'random walk' tests of market efficiency, but do not rely on the assumption employed in those studies that expected returns are constant.

A test of the zero restriction on $\gamma_x$ in (8) has the advantage of not assuming that consumption growth rates and real rates of return have a normal distribution, unlike tests based on equation system (5). Neither does such a test require that 'consumption risk' as reflected in the intercept $\Psi$ of (4) or the covariance matrix of the errors in (5) is constant. Thus, eq. (8) can provide a simple check on the stage-of-the-business-cycle effects, using alternative statistical assumptions.

Tables 5 and 6 choose $Z_t$ in (8) to be the recession dummy variable $D_t$. The results indicate that the mean of $X$ in recessions is reliably larger than the mean in non-recessions. Thus, the evidence indicates that consumption growth relative to real treasury bill returns is lower during recessions (higher in non-recession periods) than would be predicted by the marginal rate of substitution model, assuming a representative agent with constant time preference and relative risk aversion.

Note that the mean of $X$ is equal to the product of the means of the marginal utility ratio $(C_{t+1}/C_t)^{-\alpha}$ and the gross real return plus the covariance of the two. Finding that the mean of $X$ is higher in recessions (for any 'reasonable' value of $\alpha$) complements the summary statistics presented in tables 1 and 2. Those tables indicated that low mean consumption growth rates in recessions were not accompanied by low mean real treasury bill returns, in apparent conflict with the model assuming constant second moments. Tables 5 and 6 suggest that the divergent behavior of average consumption growth and real bill returns over the business cycle is not simply explained by offsetting shifts in the covariance.

---

20 If the normality assumed in (4) were correct, then $X_{t+1}$ would have a lognormal distribution in eq. (8).

21 For a given value of $\alpha$ the model (8) is linear, and the generalized method of moments (GMM) estimators [see, e.g., Hansen (1982)] produce OLS coefficient estimates and White's heteroskedasticity-consistent standard errors. (We do not estimate $\alpha$ using non-linear GMM.) The GMM approach assumes that $(X_{t+1}, Z_t)$ is a strictly stationary stochastic process. Exploratory regression evidence (not reported) indicated that the dummy variables can detect shifts in the joint distribution of $(X_{t+1}, Z_t)$ for the instruments $Z_t$ employed in the previous tables.
Table 5
Tests of stage-of-the-business-cycle effects on the mean of \( X_{t+1} \) for quarterly data. OLS regressions are of the form:
\[
X_{t+1} = \gamma_0 + \gamma_D D_t + \varepsilon_{t+1} \quad \text{where} \quad X_{t+1} = (C_{t+1}/C_t)^{-\alpha} R_{t+1}, \quad ^a
\]

<table>
<thead>
<tr>
<th>Value of relative risk aversion</th>
<th>Coefficient estimates (White’s standard errors)</th>
<th>Adjusted R-squared</th>
<th>Durbin–Watson</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>( \gamma_0 )</td>
<td>( \gamma_D )</td>
<td></td>
</tr>
<tr>
<td>( \alpha = 1 )</td>
<td>0.9955</td>
<td>0.0077</td>
<td>0.141</td>
</tr>
<tr>
<td></td>
<td>(0.0008)</td>
<td>(0.0017)</td>
<td></td>
</tr>
<tr>
<td>( \alpha = 2 )</td>
<td>0.9899</td>
<td>0.0117</td>
<td>0.147</td>
</tr>
<tr>
<td></td>
<td>(0.0011)</td>
<td>(0.0025)</td>
<td></td>
</tr>
<tr>
<td>( \alpha = 4 )</td>
<td>0.9788</td>
<td>0.0197</td>
<td>0.129</td>
</tr>
<tr>
<td></td>
<td>(0.0020)</td>
<td>(0.0045)</td>
<td></td>
</tr>
<tr>
<td>( \alpha = 6 )</td>
<td>0.9679</td>
<td>0.0276</td>
<td>0.119</td>
</tr>
<tr>
<td></td>
<td>(0.0030)</td>
<td>(0.0067)</td>
<td></td>
</tr>
<tr>
<td>Sample period: 1949:1–1983:4</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \alpha = 1 )</td>
<td>0.9951</td>
<td>0.0043</td>
<td>0.086</td>
</tr>
<tr>
<td></td>
<td>(0.0005)</td>
<td>(0.0014)</td>
<td></td>
</tr>
<tr>
<td>( \alpha = 2 )</td>
<td>0.9893</td>
<td>0.0081</td>
<td>0.097</td>
</tr>
<tr>
<td></td>
<td>(0.0009)</td>
<td>(0.0026)</td>
<td></td>
</tr>
<tr>
<td>( \alpha = 4 )</td>
<td>0.9778</td>
<td>0.0158</td>
<td>0.093</td>
</tr>
<tr>
<td></td>
<td>(0.0018)</td>
<td>(0.0053)</td>
<td></td>
</tr>
<tr>
<td>( \alpha = 6 )</td>
<td>0.9666</td>
<td>0.0234</td>
<td>0.090</td>
</tr>
<tr>
<td></td>
<td>(0.0027)</td>
<td>(0.0081)</td>
<td></td>
</tr>
<tr>
<td>Sample period: 1951:3–1979:3</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

*The variables \( C_{t+1} \) and \( R_{t+1} \) are quarterly real per capita expenditures for non-durable goods plus services and the gross real return to three-month treasury bills. \( D_t \) is a dummy variable equal to one in recession and zero otherwise. \( \alpha \) is the given value of the coefficient of relative risk aversion.

Taken together, the results of this and the previous section indicate that evidence against the consumption-based pricing model is related to the business cycle; in particular, to factors that differ in recessions from non-recession periods. One potential factor is the existence of ‘liquidity constraints’ that could destroy the interior optimum assumed in the model. Recent studies of consumption at the household level are consistent with the empirical relevance of such imperfections. However, the findings presented here that mean

\(^{22}\)Runkle (1983), for example, examines household panel data from the Denver Income Maintenance Experiment (1972–1976) and is unable to reject Euler equation restrictions for high net worth households; the restrictions can be rejected for low net worth households. Zeldes (1985), using household food consumption data from the Panel Study of Income Dynamics (1968–1982), finds similar results for subsamples stratified by ratios of wealth to income.
Table 6
Tests of stage-of-the-business-cycle effects on the mean of the $X_{t+1}$ for monthly data. OLS regressions are of the form:

$$X_{t+1} = \gamma_0 + \gamma_d D_t + \varepsilon_{x,t+1} \quad \text{where} \quad X_{t+1} = (C_{t+1}/C_t)^{-\alpha} R_{t+1}.$$ 

<table>
<thead>
<tr>
<th>Value of risk aversion</th>
<th>Coefficient estimates (White’s standard errors)</th>
<th>Adjusted $R$-squared</th>
<th>Durbin–Watson</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$\gamma_0$</td>
<td>$\gamma_d$</td>
<td></td>
</tr>
<tr>
<td>Sample period: 1959:5–1983:12</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$\alpha = 1$</td>
<td>0.9987</td>
<td>0.0023</td>
<td>0.037</td>
</tr>
<tr>
<td></td>
<td>(0.0003)</td>
<td>(0.0007)</td>
<td></td>
</tr>
<tr>
<td>$\alpha = 2$</td>
<td>0.9966</td>
<td>0.0040</td>
<td>0.031</td>
</tr>
<tr>
<td></td>
<td>(0.0006)</td>
<td>(0.0012)</td>
<td></td>
</tr>
<tr>
<td>$\alpha = 4$</td>
<td>0.9926</td>
<td>0.0073</td>
<td>0.026</td>
</tr>
<tr>
<td></td>
<td>(0.0011)</td>
<td>(0.0023)</td>
<td></td>
</tr>
<tr>
<td>$\alpha = 6$</td>
<td>0.9886</td>
<td>0.0106</td>
<td>0.024</td>
</tr>
<tr>
<td></td>
<td>(0.0017)</td>
<td>(0.0034)</td>
<td></td>
</tr>
<tr>
<td>Sample period: 1959:5–1979:9</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$\alpha = 1$</td>
<td>0.9983</td>
<td>0.0012</td>
<td>0.005</td>
</tr>
<tr>
<td></td>
<td>(0.0003)</td>
<td>(0.0007)</td>
<td></td>
</tr>
<tr>
<td>$\alpha = 2$</td>
<td>0.9961</td>
<td>0.0028</td>
<td>0.009</td>
</tr>
<tr>
<td></td>
<td>(0.0006)</td>
<td>(0.0014)</td>
<td></td>
</tr>
<tr>
<td>$\alpha = 4$</td>
<td>0.9919</td>
<td>0.0060</td>
<td>0.011</td>
</tr>
<tr>
<td></td>
<td>(0.0012)</td>
<td>(0.0028)</td>
<td></td>
</tr>
<tr>
<td>$\alpha = 6$</td>
<td>0.9878</td>
<td>0.0092</td>
<td>0.012</td>
</tr>
<tr>
<td></td>
<td>(0.0018)</td>
<td>(0.0042)</td>
<td></td>
</tr>
</tbody>
</table>

*The variables $C_{t+1}$ and $R_{t+1}$ are monthly real per capita expenditures for non-durable goods plus services and the gross real return to one-month treasury bills. $D_t$ is a dummy variable equal to one in recession and zero otherwise. $\alpha$ is the given value of the coefficient of relative risk aversion.*

Consumption growth rates appear ‘too low’ relative to bill returns during recessions do not seem to be consistent with a simple liquidity constraint explanation. If current consumption is ‘constrained’ to be below the level that would obtain in a perfect market – given anticipated future consumption and the interest rate – then measured consumption growth rates should be ‘too high’. 23

5. Conclusions

This study finds that the relation of real per capita consumption growth, treasury bill returns and a typical vector of instrumental variables appears to

23Zeldes (1985) estimates that food consumption growth for his low wealth-to-income group of families is higher than it would have been in the absence of constraints.
shift over stages of the business cycle, across policy regimes and over time. Departures from the maintained stationarity assumptions of previous tests of consumption-based models appear in both monthly and quarterly data. The results of tests seem to be sensitive to these violations. The data suggest that consumption growth is lower relative to treasury bill returns in recession periods than in non-recession periods. The model implies that anticipated changes in the two should move together. This inconsistency suggests that failure of the model's restrictions may be associated with the business cycle. Introducing a simple dummy variable model to control for some of the parameter shifts found in the relation of consumption and returns to the instruments, this study finds less evidence against the model in non-recession periods. Although estimates of time preference continue to be negative, estimates of relative risk aversion during these periods appear relatively precise and have magnitudes which are economically reasonable.

References


Brown, D., 1986, Implications of nonmarketable income for models of asset pricing, Unpublished working paper (Indiana University, Bloomington, IN).


Christiano, L., 1984, Time aggregation problems in test of the intertemporal capital asset pricing model, Unpublished working paper (University of Chicago, Chicago, IL).


Flavin, M., 1984, Excess sensitivity of consumption to current income: Liquidity constraints or myopia?, NBER working paper no. 1341.

Hall, R., 1981, Intertemporal substitution in consumption, Unpublished working paper (Stanford University, Stanford, CA).
Hall, R., 1985, Real interest rates and consumption, NBER working paper no. 1694.
Pratt, J., 1964, Risk aversion in the small and in the large, Econometrica 32, 122–137.