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THE REAL TERM STRUCTURE AND CONSUMPTION GROWTH*

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One version of the consumption-based asset pricing model implies a linear relation between expected returns and expected consumption growth. This paper provides evidence that the expected real term structure contains information that can be used to forecast consumption growth. The evidence is strongest for the 1970s and 1980s. The real term structure contains more information than two alternative measures: lagged consumption growth and lagged stock returns. Further, the real term structure appears to have slightly more forecasting power than the leading commercial econometric models.

1. Introduction

The foundations of the consumption-based asset pricing model of Rubinstein (1976), Breeden and Litzenberger (1978), Lucas (1978), and Breeden (1979) reach back to Fisher (1907). Fisher suggests that, in equilibrium, the one-year interest rate will reflect the marginal value of income today in relation to its marginal value next year. The intuition is straightforward. If a recession is expected next year, there is an incentive to sacrifice today to buy a one-year bond that pays off in the bad times. The demand for the bond will bid up the price and lower the yield. The theory implies that current real interest rates contain information about expected economic growth.

Most studies of the consumption-based asset pricing model have tested restrictions on the time-series behavior of real consumption and real asset returns. For example, Hansen and Singleton (1983) derive a time-series representation of asset returns and consumption that is consistent with the consumption-based model, time-separable isoelastic utility, and lognormally

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distributed data. They estimate the parameters of the utility function and test the model's implied restrictions. Some of the distributional assumptions are relaxed in the tests of Hansen and Singleton (1982, 1984). Recently, Dunn and Singleton (1986) have allowed for nonseparability in the utility function and for durable as well as nondurable goods.

My approach is quite different. I pursue the following insight. One version of the consumption-based asset pricing model implies that expected returns and expected consumption growth are linearly related. If expected Treasury bill returns can be estimated, those estimates should contain information about expected consumption growth. Accordingly, regression models in this paper attempt to document the comovement between the real term structure and consumption growth.

There are many reasons why this might be interesting, Kessel (1965) observes that the term structure moves with the business cycle. He shows that the difference between annualized yields on long-term and short-term bonds tends to be small immediately before a recession. This difference becomes larger before and during recovery. Recently, Fama (1986, p. 176) has noticed the changes from 'upward sloping term structures during good times to humped and inverted term structures of expected returns during recessions'. He submits that this phenomenon 'produces challenging evidence for eventual explanation by term structure models'. The consumption-based asset pricing model implies that cyclical movements in personal consumption should be reflected in cyclical movements in expected returns.

The evidence that the real term structure does contain information that can be used to forecast consumption growth is strongest in the 1970s and 1980s. The real term structure contains more information about future consumption than two alternative measures, lagged consumption growth and lagged stock returns, and appears to have slightly more explanatory power than the leading commercial econometric models.

The paper is organized as follows. Section 2 presents the framework that links expectations of consumption growth to the expected real term structure. Section 3 documents the data sources. The empirical tests are presented in section 4. Section 5 offers some concluding remarks.

2. The model

2.1. The consumer's planning problem

Consider a representative consumer with additively separable utility receiving a stochastic endowment in an exchange economy. This consumer can choose to consume this endowment or invest $\$P_{ij}$ in $=1,\ldots,N$ assets with $j=1,\ldots,k$ maturities. Expectations at time t are conditioned on the information set $\mathbf{I}_t$, which contains all the information about the environment available at t. Consumption C_t is required to be measurable at t with respect to $\mathbf{I}_t$. The consumer maximizes the following objective:

$$\max_{\{C_t,\{P_{ij,t}\}\}_{t=0}^{\infty}} \sum_{t=0}^{\infty} \delta^t \mathrm{E}[U(C_t)|\mathbf{I}_0], \qquad 0<\delta<1, \tag{1}$$

where δ is the consumer's constant time discount factor. The consumer is constrained to spend at time t only his endowment and proceeds from the sale of assets already owned to finance current consumption and new asset acquisitions.

The first-order conditions that characterize the solution to this problem can be written:

$$\mathrm{E}\left[\delta^j \frac{U'(C_{t+j})}{U'(C_t)}(1+R_{ij,t})-1\middle|\mathbf{I}_t\right]=0, \tag{2}$$

$$i=1,\dots,N, \quad j=1,\dots,k,$$

where $R_{ij,t}$ is the real j-period return on asset i from time t to time $t+j$. There are $N \times k$ of these conditions, corresponding to the N assets available and k holding periods. To concentrate the analysis, I examine only Treasury bills. $R_{j,t}$ represents the real yield (or return) on a j-period bill. The Euler eqs. (2) provide the necessary conditions for the intertemporal consumption-based asset pricing model. Necessary and sufficient conditions to derive this relation are provided in a number of papers, such as Lucas (1978) and Breeden (1986).

Eq. (2) depicts a nonlinear relation between marginal utility ratios and interest rates. The real interest rate, $R_{j,t}$, represents the return over the period t to $t+j$. If this value is known at time t and the parameters are also known, it is possible to solve for the expected marginal utility ratio. With some utility functions, the marginal utility ratio can be linked to the growth rate in consumption. With this specification, the real interest rate should forecast future economic growth.

In practice, the real interest rate is not known at time t. Below, I estimate expected real rates and test whether they contain information about future economic growth.

2.2. A linear specification

Let utility be represented by the constant relative risk aversion class:

$$\begin{aligned} U(C,\alpha) &= \frac{C^{1-\alpha}-1}{1-\alpha} \quad \text{if} \quad \alpha>0, \quad \alpha\neq 1, \\ &= \ln(C) \qquad \text{if} \quad \alpha=1. \end{aligned} \tag{3}$$

With this convenient form, we can rewrite the initial first-order conditions as

$$\mathrm{E}_t\left[\delta^j\left\{\frac{C_t}{C_{t+j}}\right\}^\alpha(1+R_{j,t})\right]=1, \qquad j=1,\dots k, \tag{4}$$

where E_t is the conditional expectation operator. Following Hansen and Singleton (1983), suppose that consumption and returns are stationary jointly lognormally distributed. Then (4) implies

$$\begin{aligned}\ln \mathrm{E}_t\left[\delta^j\left\{\frac{C_t}{C_{t+j}}\right\}^\alpha(1+R_{j,t})\right]&=\mathrm{E}_t\left[\ln\left(\delta^j\left\{\frac{C_t}{C_{t+j}}\right\}^\alpha(1+R_{j,t})\right)\right]\\&\quad+\tfrac{1}{2}\mathrm{var}_t\left[\ln\left(\delta^j\left\{\frac{C_t}{C_{t+j}}\right\}^\alpha(1+R_{j,t})\right)\right]\\&=0.\end{aligned} \tag{5}$$

The right-hand side of (5) can be rearranged to bring expected consumption growth to the left-hand side.

$$\mathrm{E}_t\left[\ln\frac{C_{t+j}}{C_t}\right]=\frac{j}{\alpha}\ln\delta+\frac{v_j}{2\alpha}+\frac{1}{\alpha}\mathrm{E}_t\left[\ln(1+R_{j,t})\right], \tag{6}$$

where v_j is the conditional variance term in (5), which is assumed constant.

The coefficient $1/\alpha$ can be interpreted as an elasticity as well as one over the coefficient of relative risk aversion. In the life cycle–permanent income hypothesis literature, this coefficient is sometimes referred to as the *elasticity of intertemporal substitution*. It can be interpreted as the sensitivity of consumption growth to changes in expected real rates. Hall (1988, p. 350) has argued that this elasticity is very small and perhaps even zero, which may imply very high levels of risk aversion. A small or zero elasticity also means that there is little or no information in the expected real rate that is relevant for forecasting real consumption growth. This paper looks beyond the single short-term rate to consider the term structure of rates.

In the spirit of Kessel's (1965) analysis, consider writing (6) in terms of a yield curve measure. The measure Kessel uses is the difference between annualized yields of different maturities and an annualized short-term yield. This type of specification can easily be accommodated by taking the difference between (6) for $j=j$ and (6) for $j=1$. It is possible to estimate

$$\Delta c_{t+1:t+j}=\beta_0^j+\beta_1\mathrm{E}_t[ys_{j,t}]+\beta_2\mathrm{E}_t[r_{1,t}]+u_{j,t+j}, \tag{7}$$

where

$$\Delta c_{t+1:t+j} = \ln \frac{C_{t+j}}{C_{t+1}},$$

$$ys_{j,t} = \ln \frac{1+R_{j,t}}{(1+R_{1,t})^{j}},$$

$$r_{1,t} = \ln(1+R_{1,t})^{j-1},$$

and $u_{j,t+j}$ is the forecast error. The coefficients should equal

$$\beta_0^j = \frac{1}{\alpha}\ln\delta^{j-1} + \frac{v_j - v_1}{2\alpha}, \qquad \beta_1, \beta_2 = \frac{1}{\alpha}.$$

2.3. Estimation issues

Most of the empirical results are focused on eq. (7). Note that the conditional expectation of the real rate appears on the right-hand side. A two-step estimator is used. Out-of-sample forecasts of the inflation rate are made at each point in the time series and parameters are reestimated at each date. The out-of-sample inflation forecasts are then subtracted from the nominal interest rates to form forecasts of the real rate, $\hat{r}_{j,t}$, and the real yield spreads, $\widehat{ys}_{j,t}$. Both $\hat{r}_{j,t}$ and $\widehat{ys}_{j,t}$ are substituted for the expectations in (7), and then the parameters are estimated by least squares.[1]

The error process $\{u_{j,t+j}: t \geq 1\}$ will not be independently distributed because of the temporal aggregation of the consumption data [see, for example, Hall (1988)] and an overlapping dependent variable. The standard errors on the regression coefficients need to be corrected for an induced moving average process of order j in the residuals. Following Hansen and Hodrick (1980), all standard errors are corrected for the moving average process.

Consistent estimates of the parameters and standard errors can also be obtained in one step, using an instrumental variables technique. The distur-

[1]A formal proof of the asymptotic normality of the estimator that uses a rolling procedure in the first stage is not available. But the results in Marcet and Sargent (1986, 1987a, 1987b) suggest that the expectations derived from the rolling procedure should be close to those obtained with the entire sample. So it is expected that the asymptotic behavior of estimators that use the rolling procedure should be the same as using the whole sample. Pagan (1984, theorem 3) has shown that a two-step estimator that uses the whole sample in the first stage will deliver consistent estimators of the coefficients and standard errors (when the standard errors are calculated using the residual variance obtained by using the actual values of the regressors). I have rerun all of the results in table 3 using the entire sample in the first stage (so Pagan's theorem 3 applies), and the results are extremely similar.

bance in eq. (6) is

$$e_{j,t+j}(x_{t+j},\theta_0)=r_{j,t}-\alpha(\Delta c_{t:t+j})-\psi_j, \qquad j=1,\ldots,4, \tag{8}$$

where x_{t+j} represents the consumption and interest rate data, θ_0 is the true value of the parameter vector, and ψ_j is the sum of the rate of time preference and the (constant) conditional variance. The parameters can be estimated with Hansen's (1982) generalized method of moments (GMM). The implication of the model is that $\mathrm{E}[e_{t+j}|\mathbf{I}_t]=0$ and it follows that

$$\mathrm{E}\left[e(x_{t+j},\theta_0)\mathbf{Z}_t\right]=\mathrm{E}\left\{\mathrm{E}\left[e(x_{t+j},\theta_0)|\mathbf{I}_t\right]\mathbf{Z}_t\right\}=0, \tag{9}$$

for any vector $\mathbf{Z}_t$ in consumers' information set $\mathbf{I}_t$. Letting

$$\mathbf{G}_T(\theta)=\frac{1}{T}\sum_{t=1}^{T}e(x_{t+j},\theta)\mathbf{Z}_t, \tag{10}$$

the GMM estimator proposed by Hansen (1982) is obtained by minimizing

$$J_T(\theta)=\mathbf{G}_T(\theta)'\mathbf{W}_T\mathbf{G}_T(\theta), \tag{11}$$

by choice of θ, where $\mathbf{W}_T$ is a positive semi-definite matrix with dimension equal to the number of orthogonality conditions in (10). An estimate of the parameter vector θ_T is necessary to solve for the weighting matrix. The standard estimation technique proceeds in two stages. First, a suboptimal choice of the weighting matrix, such as the identity matrix, is chosen and parameters are estimated by minimizing the objective function (11). The initial parameter estimates are used to solve for the optimal weighting matrix[2] $\mathbf{W}_T^*$ given by

$$\mathbf{W}_T^*=\left[\mathbf{R}_T(0)+\sum_{q=1}^{j}\left[\mathbf{R}_T(q)+\mathbf{R}_T(q)'\right]\right]^{-1},$$

where

$$\mathbf{R}_T(q)=\frac{1}{T}\sum_{t=1+q}^{T}\left[e(x_{t+j},\theta_T)\mathbf{Z}_t\right]\left[e(x_{t+j-q},\theta_T)\mathbf{Z}_{t-q}\right]'. \tag{12}$$

[2] Note that the summation in (12) runs from $q=1$ to j rather than $j-1$ in Hansen and Singleton (1982, p. 1277). This reflects the extra moving average term induced by time aggregation.

This matrix is used[3] in the objective function and the final parameter vector θ_T^* is solved for. The results reported in the next section continue this process by using θ_T^* to solve for a new weighting matrix. Parameters are then reestimated. This multistage procedure is repeated until the objective function changes by less than 0.001.

The number of observations times the minimized value of the objective function in (11) is distributed χ^2 with degrees of freedom equal to the number of orthogonality conditions less than the number of parameters [Hansen (1982)]. This statistic provides a goodness-of-fit test of the model.[4]

3. Data sources

My empirical analysis uses the National Income and Product Accounts (NIPA) quarterly personal consumption data, which incorporate the 1985 revision in the National Accounts. The consumption data are in 1982 dollars and are seasonally adjusted by the Department of Commerce.

Many problems are associated with the NIPA consumption data. They omit some components of consumption and the seasonal adjustment method creates potential difficulty. The X-11 program extracts seasonal factors by comparing the original series with data smoothed with a sequence of centered moving averages. It is not clear that the representative consumer filters the data in this way when making consumption–investment decisions. Further, since information about the future is used in calculating the seasonal factors, lagged values of consumption may not be admissible instruments for the GMM estimation. For these reasons, the empirical work considers both the seasonally adjusted and the unadjusted data.

The NIPA divide the consumption data into three categories: durables, nondurables, and services. The empirical work in the next sections uses the combined measure of nondurables and services. All variables are transformed into per capita terms with the Department of Commerce's population estimate.

[3] In a finite sample, there is no guarantee that $\mathbf{W}_T^*$ is positive definite when $j \geq 1$. Where $\mathbf{W}$ is not positive definite, the weighting matrix suggested by Eichenbaum, Hansen, and Singleton (1987) is used. For this procedure, it is necessary to estimate the coefficients of a Wold decomposition of the process $e_{t+j}\mathbf{Z}_t$. As suggested by the authors, Durbin's procedure (with the number of autoregressive parameters truncated to the sample size to the power of one-third) is used to construct the one-step-ahead forecast errors. In the cases with multiple equations and higher-order moving average processes, the method suggested by Newey and West (1987) is used. The reason for the switch of technique is the small size of the data set. In the multiequation estimation, the disturbance associated with a particular instrument has about 100 observations. However, there are (number of orthogonality conditions $\times$ order of moving average process) parameters to estimate for this disturbance. With a MA(4) and 36 orthogonality conditions, this is not feasible.

[4] The interpretation of this test is provided in Hansen and Singleton (1982, p. 1278).

The consumption of nondurables and services reported by the Commerce Department is approximately an average consumption over the quarter,[5] so in the empirical analysis I match these data with average interest rates.[6]

I obtain the interest rate data from the Selected Interest Rates and Bond Prices table of the *Federal Reserve Bulletin*. These widely quoted data are used as input variables in many commercial econometric models.[7] The data used are the yields on three-month, six-month, and nine-month Treasury bills and one-year Treasury bonds. The monthly data published by the Federal Reserve represent the average of daily closing bid yields for at least five dealers. All bills are quoted on a bank discount basis; throughout the analysis I have adjusted them to true yield. The quarterly yield data are the average of the logarithms of the monthly data.

4. Empirical results

4.1. Calculating real interest rates

According to Fisher, the expected real interest rate is the difference between the nominal interest rate and the expected inflation rate. The expected real rate can be forecasted directly, or it can be calculated by forecasting the inflation rate and subtracting it from the nominal interest rate. This analysis follows the later approach.

The inflation measure I use is the logarithmic difference in the personal consumption deflator for nondurables and services. These data are available from the second quarter of 1947 (hereafter, 1947:2) through 1987:1 (160 observations). The time series appears to follow an IMA(1,1) process.[8] The full-sample estimates (standard errors) are

$$INF_t = INF_{t-1} \underset{(0.00013)}{- 0.00003} \underset{(0.0604)}{- 0.6469}\varepsilon_{t-1} + \varepsilon_t, \qquad R^2 = 54\%.$$

[5] The average is approximate in the following sense. Although almost all of the personal consumption expenditures on nondurables are sampled monthly, the sampling for the services is different. Roughly 35% of the personal consumption expenditures on services are sampled annually and trended, 5–10% of the services are sampled quarterly, and 55–60% are sampled monthly. The quarterly consumption reported by Commerce is a sum of the levels of the monthly data.

[6] The averaging problem has been considered by Christiano (1984), Breeden, Gibbons, and Litzenberger (1986), Grossman, Melino, and Shiller (1987), Hall (1988), and Litzenberger and Ronn (1986).

[7] For example, the Data Resources, Inc. model uses these data. This model serves as a benchmark in section 5.

[8] I examined many forecasting models of inflation, including a single-input (bond returns) transfer function model and a vector autoregression that included inflation, consumption, and bond returns (five lags). I chose the univariate time-series model because it had the lowest one-step-ahead root mean square error.

Table 1

Inflation forecast evaluation: 1953:1–1987:1.

Evaluation of forecasts from a differenced first-order moving-average process[a] [IMA(1,1)] model of inflation using quarterly data in the period 1953–1987. The inflation variable is the rate of change in the personal consumption expenditures deflator for nondurables and services.

Time span	No. obs.	Mean	Std. dev.	Root mean squared error	Mean absolute error
		One-quarter forecasts			
53:1–87:1	137	0.011537	0.006716		
53:1–87:1	137	0.011357	0.006251	0.003606	0.002777
		Two-quarter forecasts			
53:1–86:4	136	0.023044	0.012696		
53:1–86:4	136	0.022772	0.012529	0.006595	0.004761
		Three-quarter forecasts			
53:1–86:3	135	0.034588	0.019201		
53:1–86:3	135	0.034277	0.018813	0.010025	0.007429
		Four-quarter forecasts			
53:1–86:2	134	0.046153	0.025365		
53:1–86:2	134	0.045870	0.025103	0.013788	0.010157

[a] Parameters of the IMA(1,1) model are estimated on data from 1947:2–1954:2 and reestimated at every point in the time series of 1987:1. Forecasts from 1953:1–1954:2 are obtained from a random-walk model of inflation.

The residuals appear to be random. The Ljung and Box (1978) statistic that tests the null hypothesis that the first twelve residual autocorrelations are zero has a right-tail p-value of 0.42, which does not provide evidence against the null.

Summary statistics for the out-of-sample forecasting performance of this inflation model are provided in table 1. The model is initially estimated over the 1947:2–1954:2 period and j-step-ahead forecasts are made. As new data are added to the sample, parameters are reestimated and forecasts are obtained. The last one-period forecast is made for 1987:1 on the basis of estimates from the 1947:2–1986:4 period.[9] Table 1 provides two forecast evaluation statistics: root mean square error (RMSE) and mean absolute error (MAE). The mean absolute errors are approximately 21–24% of the mean inflation rate. These errors are considerably lower than the errors obtained from a random-walk model.

[9] The forecast evaluation statistics cover the 1953:1–1987:1 period. The inflation forecasts are based on the IMA(1,1) model for the 1954:3–1987:1 period (131 forecasts) and a random-walk model for 1953:1–1954:2 (6 forecasts).

Table 2

Summary statistics: 1953:1–1987:1.

Summary statistics for real consumption growth, real yield spreads, and the real short-term interest rate based on quarterly data.[a]

Variable	Time span	No. obs.	Mean	Std. dev.	Autocorrelation ρ_1	ρ_2	ρ_3	ρ_4	ρ_8	ρ_{12}
				One-quarter measures						
Consumption growth[b]	59:1–87;1	111	0.00488	0.00522	0.25	0.16	0.25	0.11	−0.17	−0.01
Real spread[c]	59:1–87:1	111	0.00260	0.00202	0.51	0.15	0.07	0.03	0.03	−0.08
Real three-month bill[d]	59:1–87:1	111	0.01451	0.02081	0.88	0.74	0.71	0.65	0.46	0.32
				Two-quarter measures						
Consumption growth	59:4–87:1	107	0.00989	0.00834	0.66	0.31	0.29	0.12	−0.10	−0.03
Real spread	59:4–87:1	107	0.00341	0.00410	0.66	0.32	0.23	0.15	−0.07	−0.19
Real three-month bill	59:4–87;1	107	0.01458	0.02108	0.88	0.74	0.71	0.65	0.44	0.31
				Three-quarter measures						
Consumption growth	53:2–87;1	132	0.01447	0.01063	0.80	0.56	0.28	0.15	−0.11	−0.02
Real spread	53:2–87:1	132	0.00260	0.00421	0.70	0.41	0.33	0.25	−0.04	−0.19
Real three-month bill	53:2–87:1	132	0.01037	0.2118	0.87	0.74	0.71	0.65	0.42	0.31

[a] The starting date for each measure corresponds to the earliest availability of the data.

[b] Consumption growth is real per capita growth of personal consumption of nondurables and services.

[c] Real spread is the difference between expected real rates of interest (annualized) between a j-quarter maturity and a one-quarter maturity.

[d] Real three-month bill is the yield on three-month bill less expected inflation.

The expected real rates are calculated by subtracting the out-of-sample forecasts of inflation from the nominal interest rates. The expected yield spreads are differences in the expected real rates for various maturities.

4.2. Summary statistics

Some summary statistics for real consumption growth, expected yield spreads, and expected real short-term rates are provided in table 2. Each panel selects the sample that is used in the regression analysis.

In the version of the model in which the conditional variance is assumed to be constant, the expected interest rate variables should be linearly related to expected consumption. Autocorrelation patterns should be similar across the different series. The comparison is not exact because actual consumption growth is compared with the expected interest rate variables. Nevertheless, the autocorrelation patterns of the consumption growth measures are remarkably similar to that of the expected yield spreads. The greatest similarity occurs

(a)

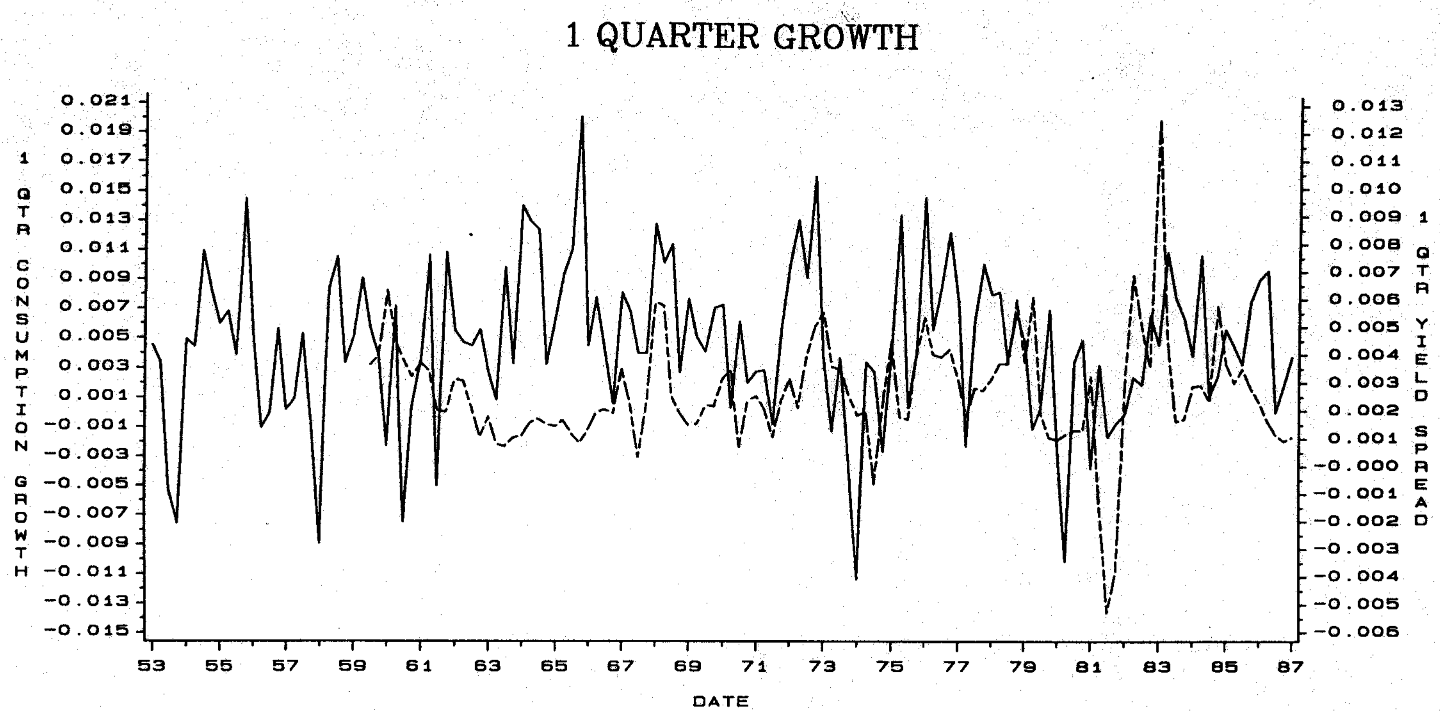

Fig. 1. Time series plots: 1953:1–1987:1.

Expected real yield spreads (dash) and consumption growth (line). The series are aligned so that consumption growth from $t+1$ to $t+j$ is matched with the expected yield spread at time t.

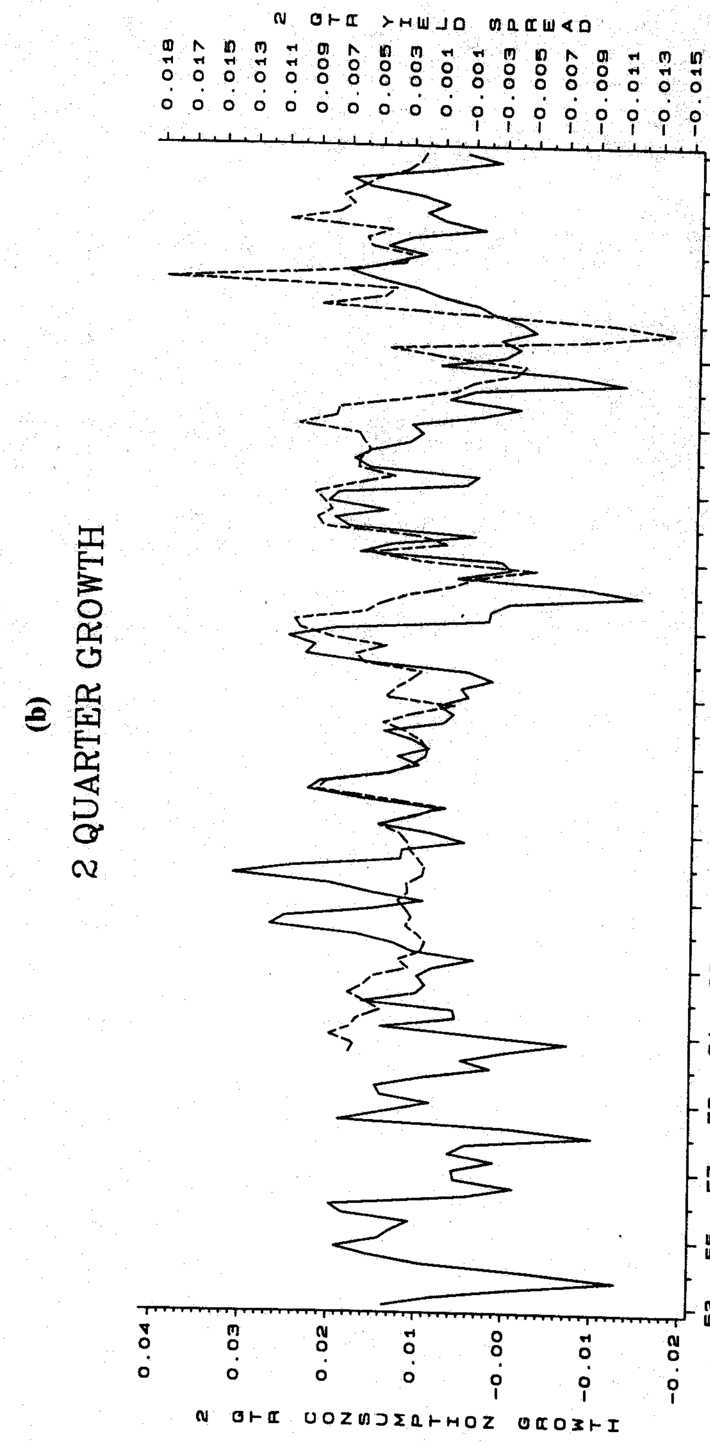

Fig. 1 (continued)

(c)

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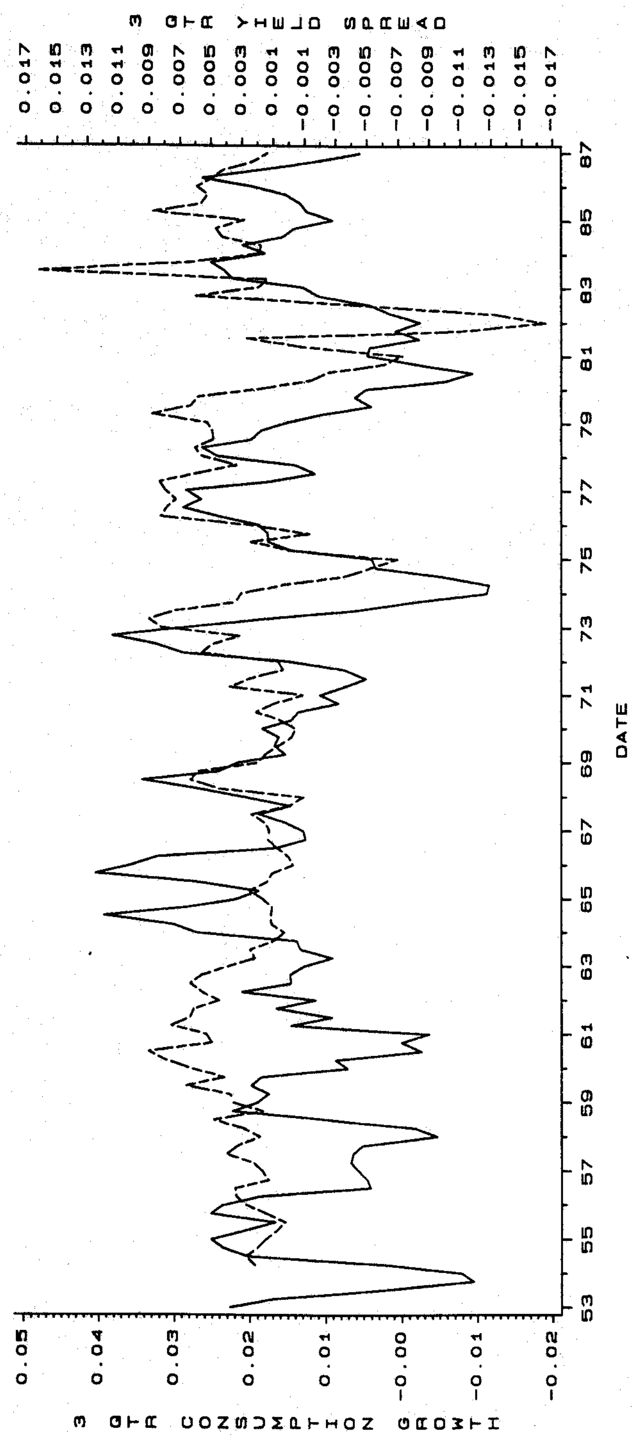

Fig. 1 (continued)

with the two-quarter growth. The difference in the autocorrelations between the consumption growth and the yield spread is never more than 0.05. The expected real rate series takes longer to decay to zero in all of the samples.

The consumption growth variable and the expected yield spread are plotted in fig. 1 so as to match the implied linearity in eq. (7). That is, the yield spread variable is lagged. If the yield spread variable exactly coincides with consumption growth, it provides a perfect forecast. If there is an unexpected consumption growth, the consumption line should look noisier than the expected spread. Fig. 1 reveals some similarity in the two series, particularly in the two- and three-quarter growth horizons. These series clearly move together from 1966 to 1987. The information in the real term structure is formally measured in the next section.

4.3. Regression tests

The regression tests of eq. (7) are presented in table 3. There is a panel for each growth horizon used in the regression. Further, each panel is divided into two subperiods with a break point of 1971:4 and an overall period.

The results from the overall period indicate that the coefficient on the yield spread variable is not different from zero at conventional levels of significance for the one-quarter regressions. The coefficient is 2.2 and 1.5 standard errors from zero in the two- and three-quarter horizons. The point estimates on this coefficient range from 0.52 to 0.93, implying a coefficient of relative risk aversion from 1.42 to 1.91. The coefficient on the short-term expected real rate is always indistinguishable from zero. Not surprisingly, a test of the null hypothesis (not reported) that these coefficients are equal does not provide any evidence against the null. The explanatory power of the one-quarter regressions is low. The R^2s for the two- and three-quarter regressions are 12% and 9%, respectively.

The subperiod results provide a check on the stability of the relation over different samples. The results of the first subperiod, which runs through 1971:4, suggest that there is little or no significant relation between the interest rate variables and consumption growth. The coefficient on the yield spread is never more than 1.4 standard errors from zero.

The results from the second subperiod (1972:1–1987:1) show a stronger relation between the yield spread variable and consumption growth. The coefficient on the yield spread is close to three standard errors from zero in both the two- and three-quarter regressions. The point estimates range from 0.96 to 1.10, which imply a coefficient of relative risk aversion ranging from 0.91 to 1.04. These estimates are very close to the value of 1.00 implied by logarithmic preferences. The R^2 of these regressions ranges from 9% for the one-quarter regressions to 28% and 31% in the two- and three-quarter regres-

Table 3

Regression results: 1953:1–1987:1.

Regressions of real consumption growth on the expected real yield spread and the expected short-term interest rate based on quarterly data.

$$\Delta c_{t+1:t+j} = \beta_0^j + \beta_1 \widehat{ys}_{j,t} + \beta_2 \hat{r}_{1,t} + u_{j,t+j}, \qquad j = 1,2,3. \qquad (7)$$

$\Delta c_{t+1:t+j}$ denotes per capita growth in real consumption of nondurables and services, $\widehat{ys}_{j,t}$ is the spread between expected real yields of j-quarter maturity and one-quarter maturity, and $\hat{r}_{1,t}$ is the expected real (yield) interest rate. Standard errors (in parentheses) are corrected for the moving-average process induced by temporal aggregation and the overlapping dependent variable [Hansen and Hodrick (1980)].

Time span	No. obs.	β_0	β_1	β_2	$\bar{R}^2$
		One-quarter growth			
59:1–87:1	111	0.0047	0.5242	−0.1332	0.01
		(0.0011)	(0.5466)	(0.1141)	
59:1–71:4	50	0.0046	−0.7539	0.5068	0.04
		(0.0019)	(1.0969)	0.2805)	
72:1–87:1	61	0.0036	0.9603	−0.1922	0.09
		(0.0013)	(0.6293)	(0.1209)	
		Two-quarter growth			
59:4–87:1	107	0.0079	0.9345	−0.0597	0.12
		(0.0022)	(0.4270)	(0.1360)	
59:4–71:4	46	0.0092	−0.2009	0.3656	0.01
		(0.0048)	(1.1555)	(0.3939)	
72:1–87:1	61	0.0063	1.1011	−0.1015	0.28
		(0.0022)	(0.3895)	(0.1197)	
		Three-quarter growth			
53:2–87:1	132	0.0131	0.7035	−0.0571	0.09
		(0.0029)	(0.4830)	(0.1371)	
53:2–71:4	71	0.0188	−1.1750	0.0144	0.05
		(0.0032)	(0.8033)	(0.1699)	
72:1–87:1	61	0.0109	1.0513	−0.0686	0.31
		(0.0031)	(0.4232)	(0.1271)	

sions. Without a benchmark, however, it is difficult to assess the forecasting power of these regressions.

4.4. Generalized method-of-moments estimation

The results of the single-equation generalized method-of-moments estimation of (8) are presented in table 4. Note that the interest-rate variables are different from those in the least-squares regressions. In the regressions, the

Table 4

GMM results with seasonally adjusted data: 1953:1–1987:1.

Generalized method of moments estimation[a] of the restrictions implied by the model between expected real consumption growth and expected real returns; seasonally-adjusted quarterly data.

$$e_{j,t+j} = r_{j,t} - \alpha \Delta c_{t:t+j} - \psi_j, \qquad j = 1,\dots,4. \qquad (8)$$

$\Delta c_{t:t+j}$ is per capita growth in real consumption of nondurables and services and $r_{j,t}$ is the real j-quarter return. The standard errors (in parentheses) for each equation are corrected for the moving average process in the disturbance terms.

Time span	No. obs.	α	ψ	$\chi^{2\,b}$	P-value[c]
		One-quarter growth			
54:2–87:1	131	−0.5744 (0.2998)	0.0027 (0.0018)	19.01	0.008
		Two-quarter growth			
61:1–87;1	104	0.9692 (0.3250)	−0.0078 (0.0029)	10.92	0.142
		Three-quarter growth			
60:1–87:1	97	0.4342 (0.4964)	−0.0022 (0.0081)	7.26	0.400
		Four-quarter growth			
57:2–87:1	119	0.3366 (0.6060)	−0.0013 (0.0116)	6.02	0.538

[a]Estimation based on Hansen's (1982) generalized method of moments. The instrumentation consists of a constant, four lagged values of the real return, and four lagged values of real consumption growth (seasonally adjusted).
[b]χ^2 is the minimized value of the GMM criterion function.
[c]P-value is the probability that a χ^2 variate exceeds the sample value of the statistic. There are two parameters and nine orthogonality conditions, which implies that there are seven overidentifying conditions to be tested.

explanatory variables are out-of-sample forecasts of the interest-rate variables. In the GMM estimation, orthogonality conditions are formed with the ex post values of the interest rates.

The instrumentation consists of a constant, second to fifth lagged values of the ex post real return, and the second to fifth lagged values of seasonally adjusted consumption growth. The first lag is an inadmissible instrument because of the temporal averaging [see Hall (1988)].[10]

The point estimates of the α coefficient range from −0.57 to 0.96. The estimates are imprecise, with only the α coefficient in the two-quarter growth specification being more than two standard errors from zero. The point estimates of the coefficient of relative risk aversion are not out of line with the

[10]Ferson (1983) and Ferson and Merrick (1987) argue that the consumption instruments should be rolled back because of a lag in the publication of the personal consumption expenditures.

results of other studies.[11] For example, using monthly data, Hansen and Singleton (1984) report estimates from -1.3 to 1.6.

The GMM also provides a test of specification. Since there are nine instruments and two parameters, there are seven overidentifying restrictions to be tested. The χ^2 statistics reported in table 4 indicate that this test provides evidence against the specification only with the one-period rate.[12]

As noted earlier, the method of seasonal adjustment imposed on the data is potentially problematic. For example, the GMM requires all instruments to be predetermined. The forward-looking part of the seasonal filter causes some of the lagged values of consumption not to be predetermined. As is suggested by Miron (1986) and Singleton (1988), this could lead to spurious rejection of the orthogonality conditions. For this reason, an alternative formulation of the model using unadjusted consumption data is also presented.

Following Miron, the utility function is restated[13] to include seasonal *taste shifters*. The version of (8) with the seasonal taste shifters is

$$\varepsilon_{j,t+j} = r_{j,t} - \alpha \Delta c_{t:t+j} - \psi_j + (1-\alpha)\boldsymbol{d}_t \boldsymbol{\phi}_j, \qquad j = 1,\ldots,4, \tag{13}$$

where $\boldsymbol{d}_t$ is a row vector of dummy variables representing the quarters and $\boldsymbol{\phi}_j$ is

$$\boldsymbol{\phi}_1 = \begin{pmatrix} -\gamma_4 \\ \gamma_2 \\ \gamma_3 - \gamma_2 \\ \gamma_4 - \gamma_3 \end{pmatrix}, \qquad \boldsymbol{\phi}_2 = \begin{pmatrix} -\gamma_3 \\ \gamma_2 - \gamma_4 \\ \gamma_3 \\ \gamma_4 - \gamma_2 \end{pmatrix},$$

$$\boldsymbol{\phi}_3 = \begin{pmatrix} -\gamma_2 \\ \gamma_2 - \gamma_3 \\ \gamma_3 - \gamma_4 \\ \gamma_4 \end{pmatrix}, \qquad \boldsymbol{\phi}_4 = \begin{pmatrix} 0 \\ 0 \\ 0 \\ 0 \end{pmatrix}.$$

[11] Using monthly data, Brown and Gibbons (1985) estimate a range for the coefficient of relative risk aversion of 0.09 to 7.00. Dunn and Singleton (1986) report values between 1.22 and 1.91 for single-equation and 2.50 to 3.45 for multiple-equation estimates. With quarterly consumption data, Mankiw, Rotemberg, and Summers (1985) estimate values from 0.09 to 0.51 in the case where utility is assumed to be separable between consumption and leisure. Mankiw (1985) reports a range between 2.43 and 5.26 using fourth-quarter data for each year. Using seasonally unadjusted data, Miron (1986) estimates a range of 0.02 to 1.71 for the risk aversion parameter. All of these studies use the GMM technique to obtain parameter estimates.

[12] Tests of a multiple-equation version of (8) are also estimated. That is, all four error terms are stacked into a system. The instrumentation contains the second to fifth large on the one-period rate and the second to fifth lag of the one-period seasonally adjusted consumption growth. The estimate of coefficient of relative risk aversion is 2.41, which is larger than the estimates obtained with the single equations. The test of the overidentifying restrictions does not provide evidence against the specification.

[13] With isoelastic utility, the utility function is $U_{t,s} = [C_t e^{\gamma_s}]^{1-\alpha}/(1-\alpha)$, where γ_s is the taste shift parameter for quarter s.

Table 5

GMM results with unadjusted data: 1953:1–1986:4.

Generalized method-of-moments estimation[a] of the restrictions implied by the linear consumption-based model between expected real consumption growth and expected real returns using unadjusted quarterly data.

$$\varepsilon_{j,t+j} = r_{j,t} - \alpha \Delta c_{t,t+j} - \psi_j + (1-\alpha) \boldsymbol{d}_t \boldsymbol{\phi}_j, \qquad j = 1,\dots,4. \qquad (13)$$

$\Delta c_{t,t+j}$ is per capita growth in real consumption of nondurables and services and $r_{j,t}$ is the real j-quarter return. $\boldsymbol{d}_t$ are dummy variables representing the four quarters. The parameter vector $\boldsymbol{\phi}_j$ is

$$\boldsymbol{\phi}_1 = \begin{pmatrix} -\gamma_4 \\ \gamma_2 \\ \gamma_3 - \gamma_2 \\ \gamma_4 - \gamma_3 \end{pmatrix}, \quad \boldsymbol{\phi}_2 = \begin{pmatrix} -\gamma_3 \\ \gamma_2 - \gamma_4 \\ \gamma_3 \\ \gamma_4 - \gamma_2 \end{pmatrix}, \quad \boldsymbol{\phi}_3 = \begin{pmatrix} -\gamma_2 \\ \gamma_2 - \gamma_3 \\ \gamma_3 - \gamma_4 \\ \gamma_4 \end{pmatrix}, \quad \boldsymbol{\phi}_4 = \begin{pmatrix} 0 \\ 0 \\ 0 \\ 0 \end{pmatrix}.$$

The standard errors (in parentheses) for each equation are corrected for the moving-average process in the disturbance terms.

Time span	No. obs.	α	γ_2	γ_3	γ_4	ψ	χ^2 [b]	P-value[c]
			One-quarter growth					
54:2–86:4	130	−0.2405	−0.0061	−0.0052	−0.0155	0.0015	13.25	0.010
		(0.2105)	(0.0044)	(0.0045)	(0.0114)	(0.0014)		
			Two-quarter growth					
61:1–86:4	103	3.7329	0.0634[d]	−0.0433		−0.0320	4.37	0.497
		(1.1925)	(0.0079)	(0.0063)		(0.0132)		
			Three-quarter growth					
60:1–86:4	96	3.0323	−0.0448	−0.0479	−0.1122	−0.0368	5.52	0.243
		(1.2709).	(0.0102)	(0.0109)	(0.0238)	(0.0225)		
			Four-quarter growth					
57:2–86:4	118	2.9172				−0.0450	4.80	0.685
		(0.9389)				(0.0210)		

[a] Estimation based on Hansen's (1982) generalized method of moments. The instrumentation consists of a constant, four lagged values of the real return, and four lagged values of real consumption growth (not seasonally adjusted).

[b] χ^2 is the minimized value of the GMM criterion function.

[c] P-value is the probability that a χ^2 variate exceeds the sample value of the statistic. In the one-quarter growth model there are five parameters and nine orthogonality conditions, leaving four overidentifying restrictions. In the two-quarter growth model, there are four parameters, which implies five overidentifying restrictions. In the three-quarter growth model, there are five parameters and four overidentifying restrictions. Finally, in the four-quarter model, there are two parameters and seven overidentifying conditions.

[d] Only the difference in the parameters γ_2–γ_4 is identified in the two-quarter growth model.

The first quarter is the reference quarter and γ_1 is set to zero. Further, in the two-quarter formulation, only the difference in the parameters, $\gamma_2 - \gamma_4$, is identified in a single-equation estimation. The instrumentation consists of a constant, second to fifth lagged values of the ex post real return, and the second to fifth lagged values of unadjusted consumption growth.

The results of estimating (12) are presented in table 5. As for the seasonally-adjusted counterpart in table 4, there is evidence against the model specification with the one-quarter growth specification[14] and little evidence against the specification when longer horizons are examined. Consistent with the intuition that using seasonally-adjusted data might induce correlation between the disturbances and the instruments, the probability values of the chi-square statistics are higher (providing less evidence against the model) in three of four cases using unadjusted data in table 5 than when the model is estimated with seasonally-adjusted data in table 4. The model parameters are also more precisely estimated with the unadjusted data. A multiple-asset version provided results consistent with the last three panels in table 5.

4.5. *Alternative consumption-prediction models*

The quality of the interest-rate variables' prediction of consumption growth is difficult to evaluate without some benchmark models. Two alternative formulations are suggested. The first includes lagged consumption growth, which Hall (1978) and Hansen and Singleton (1983) find important in explaining future consumption. The second allows stock returns to predict consumption growth. Fama (1981) suggests that stock returns lead changes in real activity. The usual measure of real activity, gross national product, contains personal consumption expenditures as well as investment and government spending. Fama shows that stock returns can predict real GNP growth, and it is possible that they can also predict real consumption growth.

To be comparable to the interest-rate regressions, information should be available at time t that is used to forecast growth from time $t+1$ to $t+j$. As a result, one-quarter consumption growth and stock returns[15] from time $t-1$ to t are used to predict consumption growth.

The results of these regressions are presented in table 6. The samples are chosen to match the least-squares results presented in table 3. The performance of the alternative models is inferior to that of the interest-rate-based models. In all the samples, the highest R^2 is 6% and the average R^2 is less than 2%. Further, in only one regression is the coefficient on the predictor

[14] These results are also consistent with the results presented in Ferson and Harvey (1988), who use different asset data and allow for time variation in the ψ parameter as well as seasonal taste shifters.

[15] The stock returns variable is the real return on the New York Stock Exchange value-weighted index.

Table 6

Regression results for alternative predictors of consumption growth: 1953:1–1987:1.

Regressions of real consumption growth on lagged consumption growth and the lagged return on the New York Stock Exchange value-weighted stock portfolio; quarterly data.

$$\Delta c_{t+1:t+j} = \beta_0 + \beta_1 x_{t-1,t} + \varepsilon_{t+j}, \qquad j = 1,2,3.$$

$\Delta c_{t+1:t+j}$ denotes per capita growth in real consumption of nondurables and services, $x_{t+1,t}$ is either lagged real one-quarter consumption growth of the lagged one-quarter real return on the value-weighted stock index. Standard errors (in parentheses) are corrected for the moving average process induced by the overlapping dependent variable [Hansen and Hodrick (1980)].

Predictor	Time span	No. obs.	β_0	β_1	$\bar{R}^2$
		One-quarter growth			
Lagged consumption	59:1–87:1	111	0.0041 (0.0007)	0.1559 (0.0977)	0.02
Lagged stock returns	59:1–87:1	111	0.0049 (0.0006)	0.0026 (0.0061)	−0.01
Lagged consumption	59:1–71:4	50	0.0041 (0.0011)	0.2585 (0.1395)	0.05
Lagged stock returns	59:1–71:4	50	0.0056 (0.0008)	−0.0017 (0.0090)	−0.02
Lagged consumption	72:1–87;1	61	0.0040 (0.0010	0.0647 (0.1376)	−0.01
Lagged stock returns	72:1–87:1	61	0.0043 (0.0009)	0.0051 (0.0081)	−0.01
		Two-quarter growth			
Lagged consumption	59:4–87:1	107	0.0079 (0.0017)	0.4112 (0.1850)	0.06
Lagged stock returns	59:4–87:1	107	0.0098 (0.0015)	0.0051 (0.0109)	−0.01
Lagged consumption	59:4–71:4	46	0.0097 (0.0023)	0.3235 (0.2295)	0.03
Lagged stock returns	59:4–71:4	46	0.0115 (0.0020)	0.0029 (0.0146)	−0.02
Lagged consumption	72:1–87:1	61	0.0068 (0.0023)	0.4169 (0.2658)	0.04
Lagged stock returns	72:1–87:1	61	0.0086 (0.0021)	0.0059 (0.0151)	−0.01
		Three-quarter growth			
Lagged consumption	53:1–87:1	132	0.0125 (0.0024)	0.4105 (0.2481)	0.04
Lagged stock returns	53:1–87:1	132	0.0141 (0.0022)	0.0223 (0.0145)	0.02
Lagged consumption	53:1–71:4	71	0.0144 (0.0029)	0.2718 (0.2814)	0.01
Lagged stock returns	53:1–71:4	71	0.0153 (0.0025)	0.0265 (0.0187)	0.03
Lagged consumption	72:1–87:1	61	0.0107 (0.0038)	0.5250 (0.4186)	0.05
Lagged stock returns	72:1–87:1	61	0.0127 (0.0035)	0.0173 (0.0215)	0.02

Table 7

Forecasting performance of the three models: 1976:1–1987:1.

An evaluation of the out-of-sample forecasting performance[a] of consumption-growth forecasting models that use the yield spread, lagged real consumption growth, and lagged real stock returns. Yield spread denotes the forecasts from the model based on eq. (7), which has the expected yield spread and expected real rate as explanatory variables. Lagged consumption represents the model with lagged real consumption growth as the explanatory variable. Lagged stock returns is the model that uses the lagged real return on a portfolio of New York Stock Exchange value-weighted stocks as an explanatory variable.

Predictor	Forecast horizon	No. forecasts	Mean absolute error	Root mean squared error
Yield spread	1 quarter	45	0.003747	0.004642
Lagged consumption	1 quarter	45	0.003670	0.004781
Lagged stock returns	1 quarter	45	0.003635	0.004759
Yield spread	2 quarter	45	0.005188	0.006487
Lagged consumption	2 quarter	45	0.005984	0.007426
Lagged stock returns	2 quarter	45	0.006139	0.007712
Yield spread	3 quarter	45	0.006262	0.007964
Lagged consumption	3 quarter	45	0.007740	0.009451
Lagged stock returns	3 quarter	45	0.007995	0.009968

[a]Parameters of each model are reestimated at each point in the time series during 1975:4–1986:4. These parameters are used to forecast the 1976:1–1987:1 period.

variable greater than two standard errors from zero. Lagged consumption tends to have slightly more explanatory power than lagged stock returns.

The table 6 regressions suggest consumption growth is difficult to predict. Comparing table 3 with table 6, it seems that the interest-rate variables can explain more of the variation in consumption growth than the other two candidate predictors.

Another benchmark can be obtained by comparing the out-of-sample forecasting performance of these three models. These results are contained in table 7. All models are initially estimated up to 1975:4, and one- to three-quarter-ahead forecasts are made. The models are reestimated with data up to 1976:1 and forecasts are calculated. This procedure is repeated to the end of the sample. Two forecast-evaluation statistics are provided in table 7: the root mean square error and the mean absolute error. Although the models differ little with the one-quarter horizon, the interest rate model outperforms the others in the two- and three-quarter horizons in MAE and RMSE.

Plots of the out-of-sample forecasts are provided in fig. 2. The lagged consumption model and the lagged stock returns model roughly track the mean of the consumption growth series. The interest rate model picks up more of the peaks and troughs and appears to provide better out-of-sample forecasts.

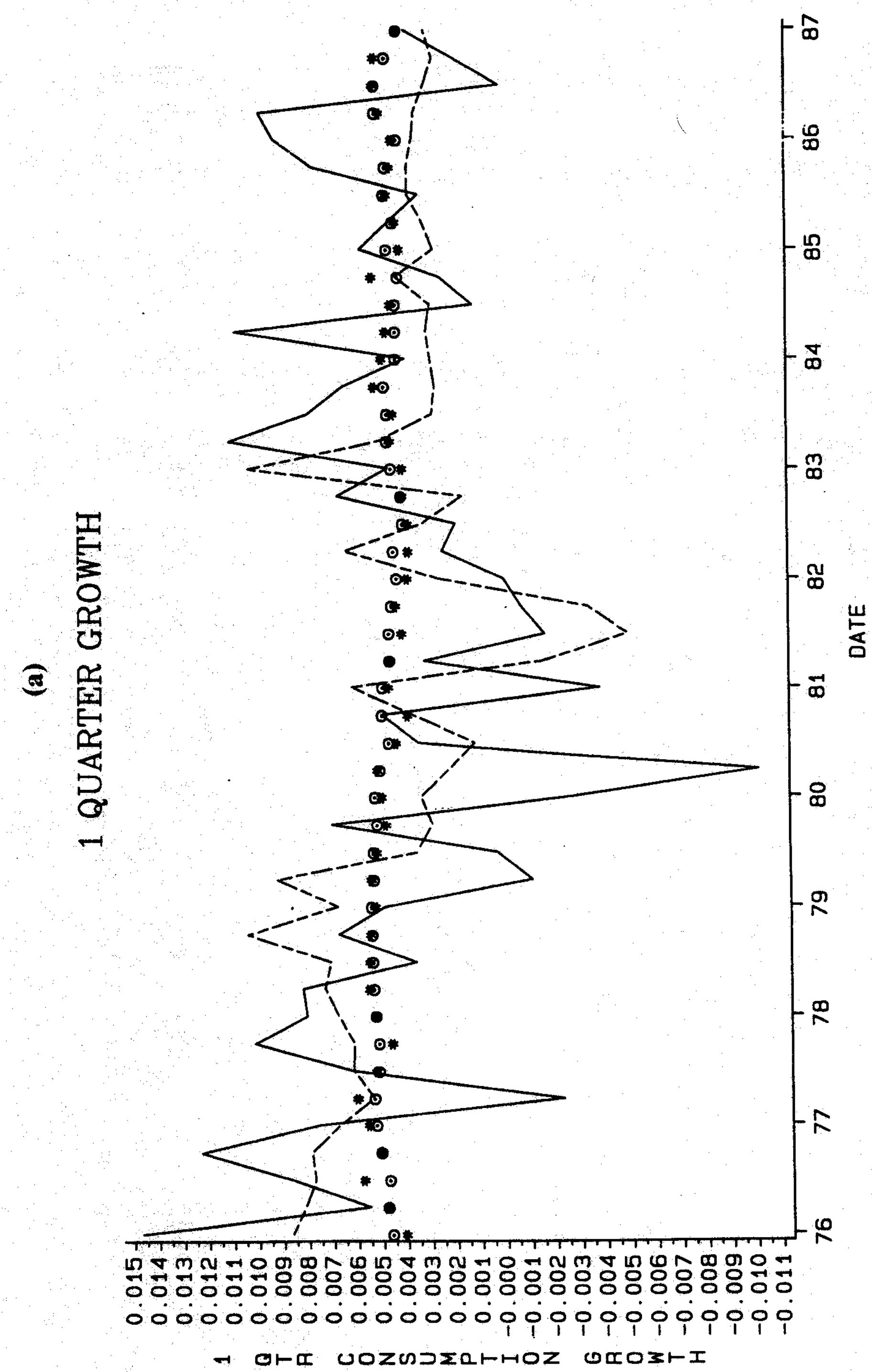

Fig. 2. Out-of-sample forecasting performance: 1976:1–1987:1.

The symbols are: actual consumption growth (line), yield spread forecasts (dash), consumption autoregression forecasts (star), and stock returns forecasts (circle).

(b)

2 QUARTER GROWTH

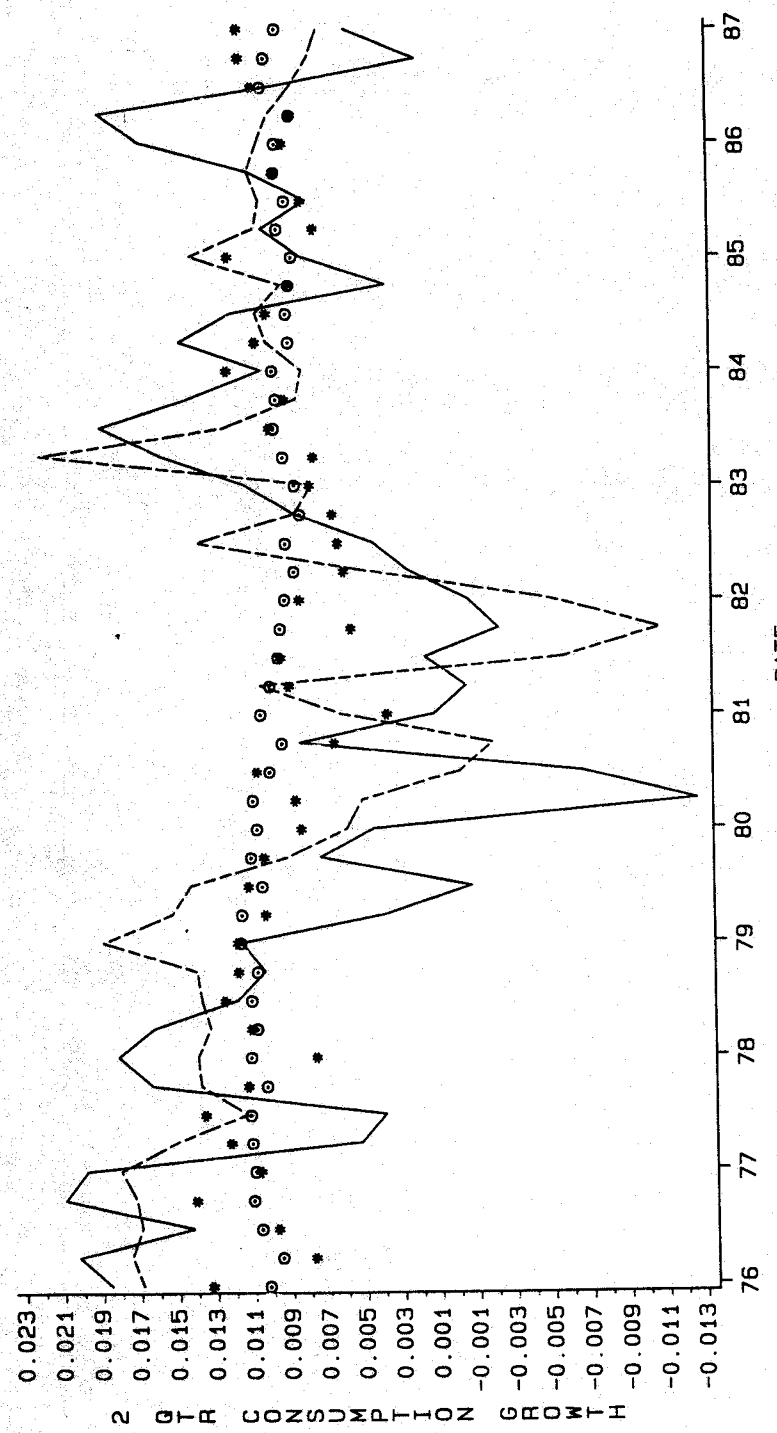

Fig. 2 (continued)

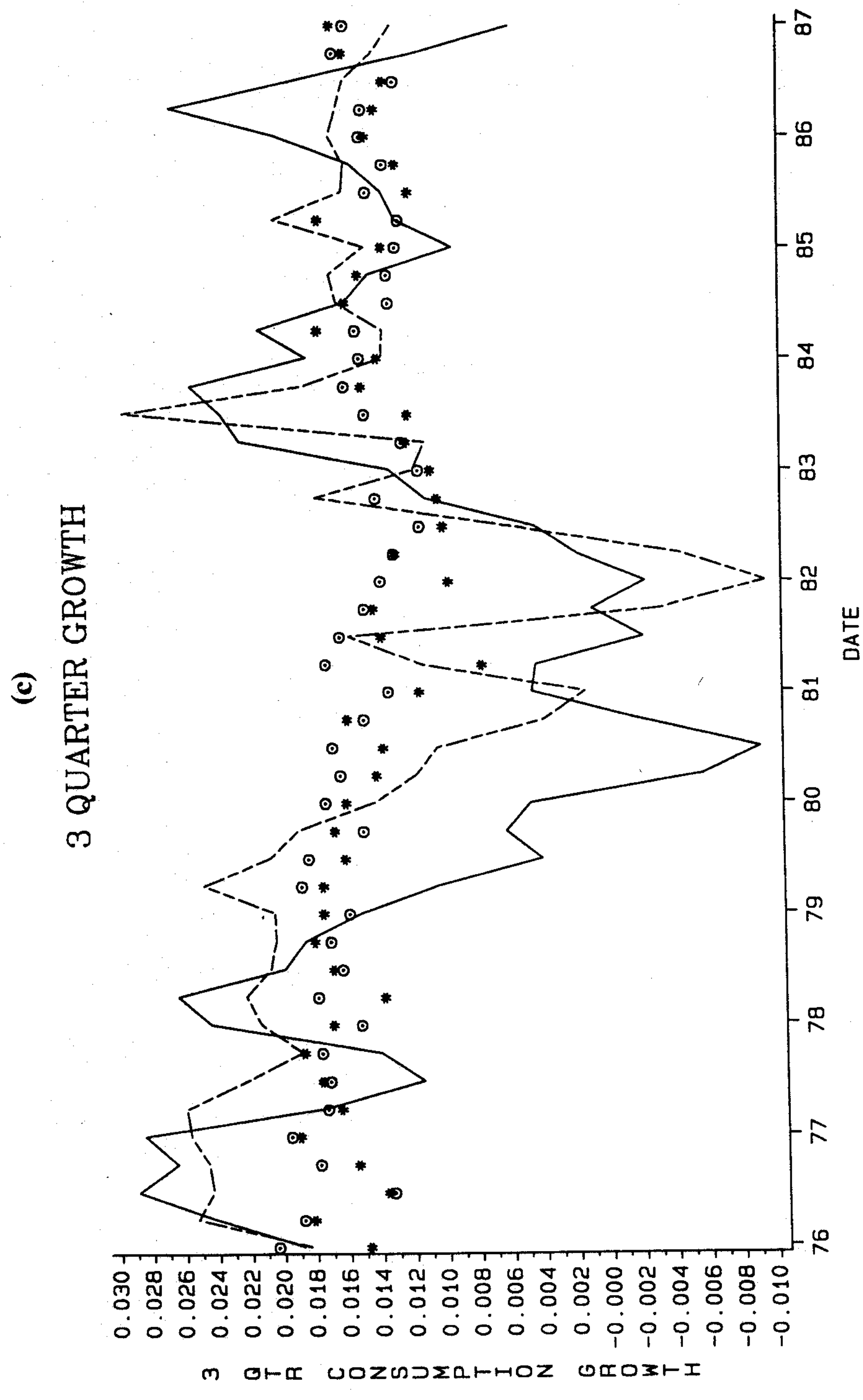

Fig. 2 (continued)

The higher quality of the two- and three-quarter forecasts in comparison with the one-quarter forecasts could be consistent with some measurement error in the data. If the reported growth in consumption is the true growth plus some stochastic measurement error, the two- and three-quarter growth rates will have a higher 'signal to noise' ratio, because some of the measurement error will cancel out when the one-quarter growth is aggregated to obtain the longer-term rates.

The final benchmarks examined are forecasts from seven leading econometric models. The data for this comparison are from unpublished research by McNees (1985) on the performance of the forecasting services. The forecasts from six of these models – Chase Econometric Asssociates, Data Resources, Inc., Economic Forecasting Project (Georgia State University), Research Seminar on Quantitative Economics (University of Michigan), Townsend-Greenspan and Co., and Wharton Econometric Forecasting Associates, Inc. – are sold for a fee. The Bureau of Economic Analysis forecasts are for internal use in the Department of Commerce.

Many problems arise in comparing forecasts from these models with the interest-rate model. The first is the timing issue. The yield-spread model uses information at the end of time t to forecast growth from $t+1$ to $t+j$. Although the *early* quarter forecasts are used, the econometric models have more recent information. A second factor is revisions in the data. The yield-spread model is fit with revised data, whereas the commercial models do not have these data available.

Table 8 presents the summary statistics for the predictions.[16] All of the growth measures are annualized and in percentage terms. The yield-spread model's forecast evaluation statistics are generally lower than or equal to those of the econometric models, except in the one-quarter horizon. In the two- and three-quarter horizons, none of the seven econometric forecasting services has lower mean absolute errors or root mean squared errors. For example, in the three-quarter horizon the average consumption growth is 3.17% from 1976:1 to 1984:2. The mean absolute error for the yield-spread model is 0.75%. The MAEs for the other services range from 0.87% for Chase to 1.11% for the Bureau of Economic Analysis. Although the difference between the yield-spread model's predictions and the best of the econometric models predictions is probably not economically significant, one must pay for the commercial models' forecasts.

In summary, the model suggested by the consumption-based theory shows some ability to forecast growth in the economy. It generally outperforms other candidate predictor variables such as lagged consumption growth and real

[16]For this comparison, the yield-spread model was estimated using an older set of consumption data available in 1985, so that all of the forecasts could be compared with the same realized consumption data.

Table 8

Forecasting performance of other econometric models: 1976:1–1984:2.

The out-of-sample forecasting performance of the yield-spread model and the commercial econometric forecasting sevices.[a] All figures are in annualized percentage growth. Yield spread denotes forecasts based on eq. (7), BEA is Bureau of Economic Analysis, Chase is Chase Econometric Associates, Inc., DRI represents Data Resources, Inc., EFP is the Econometric Forecasting Project at Georgia State University, RSQE denotes the Research Seminar on Quantitative Economics at the University of Michigan, TG is Townsend-Greenspan and Co., Inc., and WEFA represents Wharton Econometric Forecasting Associates, Inc. The forecasts evaluation statistics for the seven models are from McNees (1985). The forecast evaluation statistics for the model are based on consumption data available in mid-1985.

Model	Forecast horizon	No. forecasts	Mean absolute error	Root mean squared error
Yield spread	1 quarter	34	1.57	2.04
BEA	1 quarter	34	1.59	2.05
Chase	1 quarter	34	1.47	1.91
DRI	1 quarter	34	1.60	2.10
EFP	1 quarter	34	1.44	1.89
RSQE	1 quarter	34	1.51	1.89
TG	1 quarter	34	1.70	2.04
WEFA	1 quarter	34	1.52	2.01
Yield spread	2 quarter	34	1.01	1.31
BEA	2 quarter	34	1.27	1.66
Chase	2 quarter	34	1.09	1.37
DRI	2 quarter	34	1.24	1.62
EFP	2 quarter	34	1.23	1.50
RSQE	2 quarter	34	1.09	1.31
TG	2 quarter	34	1.29	1.51
WEFA	2 quarter	34	1.13	1.52
Yield spread	3 quarter	34	0.75	1.00
BEA	3 quarter	34	1.11	1.37
Chase	3 quarter	34	0.87	1.06
DRI	3 quarter	34	0.98	1.27
EFP	3 quarter	34	0.97	1.14
RSQE	3 quarter	34	0.87	1.07
TG	3 quarter	34	1.09	1.38
WEFA	3 quarter	34	0.88	1.16

[a]The parameters of each model are reestimated at each point in the time series during 1975:4–1984:1. These parameters are used to forecast annualized percentage growth in the 1976:1–1984:2 period.

stock returns in the within-sample and out-of-sample evaluations. The explanatory power seems nontrivial considering how well it fares against the commercial econometric models.

5. Conclusions

A common strategy in financial market research is to find variables that explain movements in prices. This paper takes a different approach: expected returns are estimated and the information in them about expected consumption growth is extracted.

Kessel (1965) and Fama (1986) have documented that the term structure moves with the business cycle. The cyclicality of the term structure is a direct implication of the linear consumption-based model. Indeed, it may be an implication of other asset pricing models. For example, in Merton's (1973) multi-beta asset pricing model, expected returns may move with state variables that are correlated with movements in the business cycle. The advantage of the consumption-based model is that all of these state variables are summarized by one measure – per capita personal consumption.

The tests carried out in this paper show that there is information about future consumption growth in the real term structure. Expected real interest-rate variables forecast consumption growth better than lagged consumption or real stock returns, both within-sample and in out-of-sample tests. The interest rate model also shows promise in out-of-sample forecasting (1976–1984) comparisons with the leading commercial econometric models.

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