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THE PERMANENT INCOME HYPOTHESIS, BUSINESS CYCLES, AND REGIME SHIFTS: EVIDENCE FROM EIGHT COUNTRIES

BY

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1 INTRODUCTION

Marginal rate of substitution models represent a theoretically rigorous and analytically simple framework for the analysis of the consumption behavior of economic agents. Utility-maximizing agents will substitute between current and future consumption up to the point where the marginal rate of substitution equals the marginal rate of return on saving and investment. This basic model is shared by finance and macroeconomics and has been used to study the joint determination of consumption and asset prices. Combining the first-order conditions of the optimizing problem with the hypothesis of rational expectations creates interesting and testable implications. For instance, the model implies that future consumption, or, more precise, expected marginal utility of future consumption cannot be predicted using current or lagged information once the current consumption decision has been accounted for.[1] Unfortunately, the accumulated literature seems to indicate a formal statistical rejection, at least of the rational expectations restrictions of this model.[2] It is unclear, however, what factors cause these restrictions to be rejected. Among the explanations for rejection that are often suggested are (temporal) aggregation, habit persistence, liquidity constraints, capital market imperfections, and regime shifts. The relative importance of these factors is not known. Moreover, international evidence on the life cycle-permanent income hypothesis (LCPIH) which could provide insights in the shortcomings of the model across countries is lacking. In this paper we try to fill

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1 Sargent (1978), Hall (1978) and Flavin (1981) pioneered the consumption-behavior aspects of the rational expectations restrictions (assuming a constant real interest rate). Hansen and Singleton (1982, 1983) were the first to test the rational expectations model of stochastic consumption for changes in asset returns.

2 See Deaton (1987), Hayashi (1987), Hall (1989) and Speight (1990) for a review of evidence and issues in life cycle-permanent income consumption models.

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this gap and provide international evidence on the LCPIH in a unified framework. We use the same model and estimation period and a consistent data set for non-durables consumption in eight countries in order to ease comparison. This way we are able to sort out whether rejections of the LCPIH are general or more specific and, in fact, limited to only a few countries.

Our tests of the LCPIH-RE restrictions use the Euler equation approach. This is to say that the consumption time-series implications are derived from the first-order conditions of the consumer optimization problem and not the closed-form solution or consumption function. The Euler equation approach allows a wider range of utility functions and the simultaneous presence of effects of interest rate and income uncertainty. Also, alternative tests measuring the excess sensitivity of consumption to current income depend on estimates of the expected present value of income innovations. This introduces complicated, and as yet unresolved, issues concerning the time-series specification of the income process and the consumer planning horizon. The consumption-interest rate equation is estimated for eight countries: Canada, France, Germany, Japan, The Netherlands, Sweden, the United Kingdom, and the United States. We use quarterly data from the sample period 1970 : 1–1989 : 4. Because of the frequent previous rejections of the basic model we reexamine the evidence when in an extented model the consumption-interest rate relationship is allowed to change over time. This is motivated by Ferson and Merrick (1987, p. 144) who conclude 'that the relation of real per capita consumption growth, treasury bill returns and a typical vector of instrumental variables appears to shift over stages of the business cycle, across policy regimes and over time. Departures from the maintained stationarity assumptions of previous test of consumption-based models appear in [the] data. The results of tests seem to be sensitive to these violations.'

The set-up of this paper is as follows. In section 2 we derive the rational expectations restrictions for the consumption-interest relationship. We use the basic representative consumer life cycle-permanent income (LCPIH-RE) model with constant relative risk aversion and a time-varying real interest rate. In section 3 we examine some properties of the time series used in this paper, in particular the non-U.S. time-series. Section 4 contains the empirical results for the basic model. Next, we address possible effects of the time-varying, non-stationary character of economic time series. We employ simple dummy variables to account for effects of the stage of the business cycle and a regime shift on the joint behavior of consumption and interest rates. A summary and concluding remarks are in section 5.

2 THE REPRESENTATIVE CONSUMER MODEL

A representative consumer is faced with the trade-off between current and future consumption. This trade-off is formalized in a maximization problem, where the consumer maximizes the expected discounted utility value of current and future consumption. The optimal consumption strategy will conform to the Euler equation

$$E_t[\beta(1 + r_{t+1})U'_{t+1}/U'_t)] = 1\,. \tag{1}$$

E_t is the expectations operator conditional on information available at time t, U'_t denotes the marginal utility function for consumption evaluated at time t, β is the utility factor. The real return on accumulated savings, r_{t+1}, is defined as $(1 + r_{t+1}) = (1 + i_t)P_t/P_{t+1}$, where i_t denotes the nominal rate of interest and P_t the price level. The constant relative risk aversion (CRRA) utility function $U_t = c_t^{1-\gamma}/(1-\gamma)$, with $\gamma > 0$, implies

$$E_t[\beta(1 + r_{t+1})(c_{t+1}/c_t)^{-\gamma}] = 1\,. \tag{2}$$

We assume (marginal) utility to be a function of consumption only. Alternative utility functions may include variables such as leisure or labor (Mankiw, Rotemberg, and Summers (1985)), public or government consumption (Bean (1986)), or wealth (Deaton (1972)). We also assume intertemporally separable preferences. A recent strand in the literature has shifted attention to intertemporally dependent preferences or habit persistence (*i.e.* Eichenbaum, Hansen and Singleton (1988), Muellbauer (1988)).[3]

To estimate this relationship between consumption and the real rate of interest we follow Mankiw (1981, 1985). With rational expectations we know that the conditional expectation of any variable x_{t+1} deviates from its realization by a random expectational error only. Thus, the expectation $E_t[x_{t+1}] = 1$ is identical to $x_{t+1} = 1 + \varepsilon_{t+1}$ for $E_t[\varepsilon_{t+1}] = 0$. Using the rational expectations hypothesis, a logarithmic transformation, and the Taylor approximation $log(1 + \varepsilon_t) \simeq \varepsilon_t - (\varepsilon_t)^2/2$, equation (2) yields

$$\Delta log\, c_{t+1} = \alpha_0 + \alpha_1\, log(1 + r_{t+1}) + \eta_{t+1}\,. \tag{3}$$

Equation (3) uses only realized values of consumption growth and the real rate of interest, where

$$E(\varepsilon_t^2) = \sigma^2\,,$$

$$\alpha_0 = [\sigma^2/2 + log\beta]/\gamma\,,$$

$$\alpha_1 = 1/\gamma\,,$$

$$\eta_{t+1} = [(\varepsilon_{t+1})^2/2 - \sigma^2/2 - \varepsilon_{t+1}]/\gamma\,, \text{ and } E(\eta_t) = 0\,.$$

3 This is an attempt to account for the autocorrelation found in consumption growth.

We find that the optimal time path of consumption implies that the (expected) growth rate of consumption is higher for increased levels of (expected) real interest rates r_{t+1}, a larger elasticity of intertemporal substitution α_1 (or lower index of risk aversion γ), and increased uncertainty represented by σ^2. Note that without explicity deriving the consumption function, we can say nothing about the level of the consumption time series. While the (expected) real interest rate determines the (expected) slope of the consumption time path, the general level of the consumption path will also depend on expected future income and wealth. Our equation represents an estimate of planned current and future consumption relatives. Complete solutions to the optimization problem for utility functions other than the simple quadratic type (exhibiting certainty equivalence) are unavailable or cumbersome when the consumer faces an uncertain income stream.

The rational expectations restrictions in this model inform us that the only legitimate role of lagged information variables in explaining future consumption is in predicting real interest rates. This requires that when we estimate the regression equation

$$\Delta log\, c_{t+1} = a_0 + a_1\, log(1 + r_{t+1}) + \mathbf{b}\mathbf{Z}_t + \eta_{t+1}\,,$$

where $\mathbf{Z}_t$ represents a vector of variables included in the time t information set of the consumer, we find $\mathbf{b} = 0$. This exclusion test applies to all variables in the time t information set. However, the power of the test is 'evidently' greater for those variables expected to be more closely related to the consumption decision. Natural candidates then are lags of consumption, income, wealth and interest rates. The data used in this paper are described in section 3.

3 THE DATA[4]

We compiled quarterly data series for eight countries – Canada, France, Germany, Japan, The Netherlands, Sweden, the United Kingdom, and the United States – from quarterly national accounts data covering the period 1970 : 1 (first quarter 1970) through 1989 : 4. We follow the usual practice of limiting our consumption data to a measure of non-durables consumption, and define c_t as the per capita non-durables consumption of goods and services (NDS). Hence, we assume that utility is separable in non-durable and durable consumption components. In the exclusion tests we include lagged income as one of the instruments. The only income variable that is readily available and consistent for all selected countries is gross domestic/national product. Although in the LCPIH framework this is not the most appropriate concept of income, as we mentioned above, the exclusion test includes all variables dated time t or

4 A separate data appendix is available from the authors.

earlier and this includes an income measure based on domestic/national product.[5]

We use seasonally adjusted per capita consumption and income series. Where only unadjusted data were available we used the Census X-11 correction for seasonality as this seems to be the most commonly used method of seasonal adjustment.[6]

The real interest rate we selected is a short-term nominal interest rate adjusted for changes in the consumer price index. (Our interest rate (and income) variables are not adjusted for taxes.) Note that for the consumption-based capital asset pricing model (CCAPM) interpretation of equation (2), the choice for returns on any specific asset (short-term or long-term, stocks or bonds) is of limited consequence, because the model applies to all assets and interest rates. In the LCPIH interpretation the interest rate represents the marginal return on the representative consumer's accumulated savings and, therefore, implies a particular assumption about the representative investment portfolio. If we assume that the use of quarterly data corresponds to the frequency with which the average consumer re-evaluates its consumption-saving/investment decisions then quarterly (short-term) returns are appropriate.[7]

In Table 1 we present summary statistics for consumption growth, income growth, and the (*ex post*) real interest rates used in this paper.

In the 1970-89 period average growth rates of consumption range from a low of 1.2 percent (annual rate) in Sweden to a high of 3.1 percent in Japan. GNP/GDP income growth rates range from 1.7 percent in Sweden and The Netherlands to 3.5 percent in Japan. Average real interest rates were low in the United Kingdom (0.5 percent) and high in France (8.5 percent). In France and Sweden the sample variance of consumption growth exceeds the sample variance of income growth. This appears to contradict the PIH consumption-smoothing proposition.[8] However, the difference is not significant at the 10 percent level.

The Bera-Jarque test statistic indicates that several of the sample distributions of consumption growth, income growth and real interest rates exhibit significant deviations from normality. Both skewness and kurtosis can be explained by the occurrence of a change in the mean of the data-generating process and/or the

5 Of course, the power of the test is conditional on the measure used. Quarterly data on disposable (personal) income series are published for four countries in our sample. For these countries we examined alternative results using these series.

6 Miron (1986) suggested that using seasonally adjusted data could bias our results. English, *et al.* (1989) show that this effect is not in the data.

7 Estimates of intertemporal substitution generally appear to be more significant for short-term than for long-term interest rates (Koedijk and Kool (1990)). On the other hand, the rational expectations restrictions seem to be rejected more often using the short-term interest rate.

8 Deaton (1987) suggests that consumption innovations should exceed income innovations when the income process exhibits permanent shocks that are positively autocorrelated.

TABLE 1 – SUMMARY OF DESCRIPTIVE STATISTICS OF THE DATA, 1970:2–1989:4

	Mean	St.Dev.	Autocorrelations 1	2	3	4	5	6	7	8	$Q(8)$	Bera-Jarque (1980)
Canada												
$\Delta log\ c$	2.569	2.906	0.16	0.25	0.14	0.11	0.16	0.12	0.11	−0.04	14.60[c]	0.424
$\Delta log\ y$	2.661	3.997	0.30	0.09	0.09	−0.04	−0.12	0.02	0.03	−0.15	12.29	0.455
$log(1+r)$	2.387	3.971	0.77	0.62	0.59	0.60	0.55	0.51	0.47	0.47	228.01[a]	2.046
France												
$\Delta log\ c$	2.484	2.822	−0.28	0.08	0.26	−0.11	0.18	−0.10	0.15	0.03	19.50[b]	0.978
$\Delta log\ y$	2.377	2.490	0.13	0.32	0.10	0.08	0.13	0.02	0.12	−0.08	14.22[c]	19.698[a]
$log(1+r)$	8.547	3.790	0.82	0.74	0.73	0.70	0.67	0.64	0.64	0.62	333.30[a]	3.073
Germany												
$\Delta log\ c$	1.945	1.590	0.53	0.24	0.21	0.23	0.08	−0.11	0.02	0.10	38.71[a]	0.555
$\Delta log\ y$	2.294	4.418	−0.09	0.01	0.13	0.10	−0.23	−0.03	0.02	−0.11	8.80	3.081
$log(1+r)$	3.014	2.790	0.51	0.07	0.19	0.35	0.08	−0.21	−0.07	0.11	40.92[a]	1.167
Japan												
$\Delta log\ c$	3.087	2.307	0.67	0.49	0.40	0.23	0.03	−0.08	−0.16	−0.08	78.36[a]	1.965
$\Delta log\ y$	3.471	3.531	0.08	0.06	0.16	0.04	−0.18	−0.04	−0.10	−0.22	11.42	45.415[a]
$log(1+r)$	1.738	5.649	0.46	0.53	0.41	0.44	0.15	0.23	0.06	0.24	82.90[a]	262.450[a]
Netherlands												
$\Delta log\ c$	1.896	2.108	0.69	0.65	0.55	0.56	0.42	0.32	0.22	0.19	158.07[a]	0.071
$\Delta log\ y$	1.758	6.960	−0.24	−0.10	−0.03	0.01	0.08	0.05	0.00	−0.17	9.10	71.863[a]
$log(1+r)$	2.019	4.012	0.59	0.67	0.43	0.64	0.35	0.47	0.28	0.49	174.67[a]	4.870[c]

Sweden												
$\Delta log\ c$	1.220	4.719	−0.27	0.06	0.01	0.07	0.06	−0.02	0.04	0.02	7.33	91.666[a]
$\Delta log\ y$	1.713	3.750	−0.03	−0.10	0.07	0.18	−0.11	−0.17	0.09	0.06	9.07	2.648
$log(1+r)$	1.112	5.131	0.51	0.41	0.53	0.54	0.40	0.31	0.32	0.48	135.61[a]	1.372
United Kingdom												
$\Delta log\ c$	2.529	3.744	0.14	0.22	0.17	−0.14	0.33	0.06	0.05	0.03	19.98[b]	0.379
$\Delta log\ y$	2.193	5.743	−0.16	0.01	−0.01	−0.13	0.04	0.18	0.00	−0.10	7.32	22.676[a]
$log(1+r)$	0.518	6.329	0.53	0.45	0.36	0.55	0.35	0.38	0.38	0.48	133.50[a]	39.752[a]
United States												
$\Delta log\ c$	1.677	2.014	0.29	0.08	0.32	0.09	−0.26	0.01	−0.03	−0.25	28.53[a]	2.696
$\Delta log\ y$	1.710	4.220	0.25	0.22	0.00	0.04	−0.08	−0.09	−0.10	−0.26	17.62[b]	3.344
$log(1+r)$	1.520	3.462	0.79	0.67	0.71	0.63	0.56	0.49	0.45	0.36	242.56[a]	2.389

Notes:
Superscripts a, b, c, denote significance at 1, 5 and 10 percent levels. We define c, y and r as real per capita consumption of non-durables and services, real per capita GNP/GDP income, and the short-term real interest rate (see data appendix). Variables are percentages at annual rates. Estimated standard error of the autocorrelations is 0.113. $Q(8)$ is the Ljung-Box Q-statistic testing for autocorrelations up to order 8. Skewness and excess kurtosis are combined in the $\chi^2(2)$ distributed Bera-Jarque (1980) test for normality.

presence of outliers. In our analysis below we suggest that these effects might be related to the business cycle or a regime shift.

The autocorrelation function for consumption growth presents us with some early indications of the possible success or failure of the rational expectations tests. A standard approach is to ignore the (significant) correlations at lag 1 by referring to problems related to time-averaging (Hall (1988), Haug (1991)), (semi-durability of consumer expenditures (Mankiw (1982)), measurement error (see Wilcox (1992) for a review of the construction of U.S. consumption data) and publication lags. However, for all countries except Sweden we observe significant autocorrelations at other lags as well. More important, however, is the lack of correspondence between the autocorrelations in consumption growth and real interest rates. Autocorrelations in consumption growth are entirely consistent with the model if they are directly related to the autocorrelations in the real interest rate. On the other hand, if autocorrelations in consumption growth were found to be absent we might conclude that revisions in expectations, or 'news,' dominated the consumption time-series process. Our problem concerns the observation that consumption growth autocorrelations appear at specific lags other than lag one.

4 EMPIRICAL RESULTS

4.1 *Estimation Issues*

The transformations of the CRRA utility function allow us to use logarithmic growth rates of consumption and income in our empirical work. At this stage, we assume that this suffices to obtain stationary time series.[9] The contemporaneous correlation between r_{t+1} and η_{t+1} means we cannot simply estimate (3) or (4) using ordinary least squares. But, we can use appropriate variables dated time t or earlier as instruments in an instrumental-variables procedure. For our tests of the rational expectations restrictions $\mathbf{b} = 0$ we use a Lagrange Multiplier test. The standard procedure to avoid the problems associated with time aggregation, measurement error, transitory consumption, semi-durability, and publication lags is to use instruments lagged at least two periods (see Hall (1988), Campbell and Mankiw (1989, 1991)). To check on the effects of this choice we will also briefly refer to alternative results with the one-period lagged instru-

9 In our extended model below, we add dummy variables to correct for specific examples of nonstationarity in the growth rate of consumption. In the main text we do not explicitly address the ongoing debate on unit roots. Mankiw and Shapiro (1985) have shown that detrending consumption and income time series, while, in fact, they follow a random walk, may bias the tests towards rejecting the random walk consumption hypothesis. Other studies show that discriminating between trend stationary and difference stationary processes is not straightforward. Tests developed so far, have low power and depend on the exact specification of the null regarding the time series creating process. On theoretical grounds we assume the real interest rate to be a stationary series, that is, at least in the long run, related to economic growth. Whether nominal interest rates are stationary is a different matter.

ments. A better solution would have been to model or correct for these (auto)correlations explicitly. Recent research on habit persistence, durability, and non-time-separable utility functions is taking this direction. The range of problems to be modeled is, however, extensive and we prefer at this stage to compare our results with the existing literature.

TABLE 2 - INSTRUMENT SETS USED IN IV/2SLS ESTIMATION AND TESTS OF RESTRICTIONS

Set	Instruments	Total number
	constant[a] and	
I	$log(1 + r_{t-2})$, $\Delta log\, c_{t-i}$	5
II	$log(1 + r_{t-2})$, $\Delta log\, y_{t-i}$	5
III	$log(1 + r_{t-i})$, $\Delta log\, c_{t-i}$	7
IV	$log(1 + r_{t-i})$, $\Delta log\, c_{t-i}$, Δi_{t-i}	10
V	$log(1 + r_{t-i})$, $\Delta log\, c_{t-i}$, $\Delta log\, y_{t-i}$, $\Delta log\, s_{t-i}$	13
VI	$log(1 + r_{t-i})$, $\Delta log\, c_{t-i}$, $\Delta log\, y_{t-i}$, $log(c/y)_{t-2}$	11
	with $i = 2, 3, 4$.	

Notes:
a. With some results we added dummy variables to the equation and therefore added them as instruments as well.
Symbols c, y, r, i, s denote real per capita consumption and income, real and nominal interest rates, and a real stock price index (see data appendix).

Our choice of the instruments reflects two other considerations. First, there is the need for an appropriate selection of instruments to be used for r_{t+1}. Second, to test the LCPIH-RE restrictions on lagged information variables, we want to include instruments that are commonly regarded as closely related to the consumption decision. Besides these, more or less general, conditions there is no obvious rule to choose a single set of instruments. Therefore, we used several sets of instruments with different combinations of lagged variables that can be regarded as relevant to our problem. We always include at least one lagged value of the real interest rate. In our choice of the other instruments we closely followed the selections made in earlier studies. The remainder of each instrument set consists of lagged values of consumption growth and income growth, changes in the nominal interest rate, changes in share prices as a proxy for changes in wealth and/or interest rates, and a lagged value of the consumption-income ratio representing an error correction mechanism. The LCPIH-RE tests examine the predictive power of the instrument sets in Table 2 which goes beyond the information these instruments may contain about r_{t+1}.

4.2 *The Basic Model*

We estimated the basic regression model for Canada (CN), France (FR), Germany (GE), Japan (JP), The Netherlands (NL), Sweden (SW), the United Kingdom (UK), and the United States (US). Table 3 presents a summary of the estimation results. Allowing for the use of first differences and the maximum lag of instrument values, our estimation period is 1971 : 2–1989 : 4.

The estimated coefficient for the real rate of interest, here also interpreted as the elasticity of intertemporal substitution, is not particularly large. Positive values range from 0.1321 to 0.2562, and only for the United Kingdom do we find a significantly positive intertemporal-substitution effect. These magnitudes are broadly consistent with other estimates. Values found in other studies include for the United States Mankiw (1981) 0.17–0.19 (data 1948 : 1–1980 : 4), Mankiw (1985) 0.19–0.38 (fourth quarter data 1950–1981), Bean (1986) 0.40 (data 1949 : 2–1979 : 4), and Campbell and Mankiw (1989) 0.27–1.3 (data 1953 : 1–1986 : 4). For the United Kingdom Muellbauer's (1983) estimates are in the range 0.38–1.08 (data 1956 : 2–1979 : 3), Wickens and Molana (1984) estimate 0.3 (data 1964 : 1–1979 : 4). The estimates imply a coefficient of relative risk aversion, the inverse of the interest rate elasticity, ranging from 3.9 to 7.6. For Canada and The Netherlands the interest rate coefficient is significantly negative at the 10 percent level.This implies not risk-averse but risk-seeking agents. Another possible explanation is that the negative coefficient indicates significant independent – rather than operating through the real interest rate – effects of inflation on consumption through involuntary savings, as was suggested by Deaton (1977). We find that for seven out of eight samples we do not estimate significant coefficients for the real interest rate that have the expected sign. Either the elasticity of intertemporal substitution is really close or equal to zero, or else the true elasticity somehow cannot be accurately measured due to data problems (Hall (1988), Campbell and Mankiw (1989)).

The LCPIH-RE restrictions are tested for the different subsets of instruments specified in Table 2. The combined evidence of the residual autocorrelation Q-statistics and the LM test for the six sets of instruments is that, formally, the rational expectations restrictions must be rejected for all eight countries.[10]

10 In an earlier version of this paper we were inclined to take a more flexible position. The source of the residual autocorrelation is left unexplained. It may very well reflect mechanisms that are not related to (ir)rational expectations, *i.e.* habit persistence, unexplained seasonal patterns, *etc.* For five countries the LM tests provided no or just a single scattered and sometimes only marginally significant statistic. This could be a statistical quirck within a large set of instruments. Rejection of the restrictions is most obvious for the Japanese and United Kingdom data. Alternative estimates with instrument variables starting at lag one, a longer sample period (UK and US), and (personal) disposable income series to replace GNP/GDP increased the significance and the number of rejections of the overidentifying restrictions. Qualitatively, the results remained very similar.

TABLE 3 – THE BASIC LCPIH CONSUMPTION EQUATION

$$\Delta log\ c_t = a_0 + a_1\ log(1 + r_t) + \eta_t$$

	CN	FR	GE	JP	NL	SW	UK	US
a_0	2.9372[a]	3.2544[a]	2.8950[a]	2.7695[a]	2.4236[a]	1.2381[b]	2.3807[a]	1.5060[a]
	(7.10)	(3.54)	(2.76)	(5.01)	(4.37)	(2.11)	(5.89)	(5.59)
a_1	−0.1837[c]	−0.0952	−0.2990	0.1460	−0.2781[c]	0.1786	0.2562[a]	0.1321
	(1.70)	(0.95)	(0.90)	(1.34)	(1.85)	(1.17)	(3.07)	(1.58)
SE	2.862	2.808	3.509	4.500	4.019	4.832	3.472	2.053
DW	1.79	2.49	2.47	2.33	2.48	2.34	1.75	1.34
$Q(8)$	6.51	16.36[b]	14.03[c]	18.46[b]	14.93[c]	6.97	14.80[b]	25.59[a]
Lagrange Multiplier tests for the overidentifying restrictions with instrument list (degrees of freedom):								
I (3)	3.74	6.36[c]	6.59[c]	12.58[a]	4.49	1.47	9.56[b]	8.41[b]
II (3)	3.19	1.67	2.54	6.74[c]	3.69	5.32	4.86	3.37
III (5)	8.80	6.76	7.57	15.94[a]	5.53	8.93	12.41[b]	8.54
IV (8)	9.44	10.71	10.06	19.17[b]	9.21	13.43[c]	15.46[c]	22.81[a]
V (11)	19.74[b]	13.62	14.47	17.74[c]	10.28	12.77	15.38	12.12
VI (9)	10.81	10.18	11.30	16.61[c]	11.04	12.69	14.26	13.58

Notes:

Estimation period is 1971:2–1989:4. Regression results are for IV/2SLS estimation using instrument list V (see Table 2). $Q(8)$ is the Ljung-Box Q-statistic. The LM test for the rational expectations restrictions equals $n * R^2$, with n the number of observations and R^2 the coefficient of determination for the CLS regression of the residuals on the instruments. LM is $\chi^2(k - l)$ distributed with k the number of instruments and l the number of coefficients estimated Superscripts a, b, and c indicate significant test statistics at 1, 5, and 10 percent levels, respectively.

4.3 *Alternative Approaches: Business Cycle Effects and Regime Shifts*

Recently, it has been argued in the literature that refinements of the marginal substitution model should focus on the less-than-perfect integration of capital markets and goods markets, on the existence of liquidity constraints, or on regime shifts as possible causes for rejection of the simple model. For example, Ferson and Merrick (1987) found that allowing for stage-of-the-business-cycle effects reduces the evidence against the model with U.S. data. Their tests show, that, compared to the full business cycle data, there is 'virtually no evidence against the model when the economy is not in recession,' thus suggesting the operation of liquidity constraints. Kandel and Stambaugh (1990) found that the conditional means and variances of consumption growth change with the business cycle. This would violate the classical assumptions underlying our regression analysis and test statistics, becuse it implies non-constant coefficients in our regressions.

There are several ways to introduce the possibility of time-varying coefficients in the basic model. Equation (3) shows that consumption growth rates vary for different levels of (expected) real interest rates. But equation (3) also shows that average consumption growth rates depend on the variance or uncertainty of the joint distribution of consumption and interest rates, represented by σ^2. If economic conditions change so as to increase or decrease uncertainty about future changes in interest rates and income, consumption growth rates will change to reflect the changed demand for precautionary savings. A similar effect on consumption growth could occur when the risk aversion index γ or the subjective discount rate of expected utility β changes. This, however, would contradict a more fundamental theoretical notion of constant structural parameters. Alternatively, we could introduce taste-shift variables in the utility function. This, however, is empirically not an easily tractable solution because we do not have the data to measure these taste shifts. Still other interpretations of changing coefficients could focus on changes in liquidity constraints and/or distribution effects. Presence of these effects would invalidate the representative consumer assumption in our basic model. In the end, we could reclassify these alternative approaches into two categories representing different implications for our basic theory. For instance, if the cause of changing coefficients is to be found in the adaptation of consumer behavior to changes in risk,our theory remains valid, but the classical – constant coefficient – regression analysis is inappropriate. If, however, the cause is found to be the presence of taste shifts or liquidity constraints, then our simple theory must be exchanged for more complex considerations, or, as an ultimate consequence, the life cycle-permanent income hypothesis must be discarded entirely.[11]

11 Some examples of extended Euler equations are described in a separate appendix available from the authors.

4.3.1 Business cycle effects

In this section we test whether time-varying moments of the joint behavior of consumption and interest rates can account for the rejection of the over-identifying restrictions. Others, *i.e.* Flavin (1985), Wilcox (1989), Campbell and Mankiw (1989, 1991), have attempted to capture the effects of liquidity/borrowing constraints by adding current income, unemployment or nominal interest rates to the basic equation. We follow Ferson and Merrick (1987) and employ dummy variables to account for the effects of business cycles on the joint behavior of consumption and interest rates. From the OECD (1987) study into member countries' business cycles and subsequent OECD publications we compiled GNP/GDP business cycle chronologies for the countries in our sample. Ignoring the smaller business cycles, we identified 2 to 4 major recession periods for each country. Next, we constructed recession-period dummy variables that cover the first quarter after the business cycle peak up to and including the subsequent business cycle through. The results of the equations where we allow for a business cycle shift in the intercept of the equation are summarized in Table 4.

In most samples, except Japan and the UK, the business cycle shifts reduce the overidentifying test statistic. The combined evidence from residual correlation and overidentifying restrictions is that the rational expectations restrictions must be formally rejected in five countries. Rejection remains very strong in the JP and UK samples. The coefficients on the dummies indicate a major role for the early-1980s recessions, and much less for the 1973–75 recessions. The results suggest that the business cycle dummies and the real interest rate are correlated, lowering the t-statistics for the interest rate coefficient.[12] In none of the samples do we find a significant independent effect of the real interest rate when the consumption process is allowed to change with the business cycle. That business cycles, consumption growth rates and real interest rates relate in this way does not, of course, invalidate the LCPIH model. It reflects the required positive correlation between consumption growth and real interest rates. We must conclude that our business cycle dummies do not improve the specification of our consumption equation, neither with respect to estimates of the intertemporal substitution nor with respect to our findings on the RE restrictions.

4.3.2 Regime shifts after 1979

As an alternative to business cycle effects we examine the effect of a permanent change in the joint behavior of consumption and interest rates after 1979.

12 Campbell and Mankiw (1991) interpret the evidence for the interest rate-income relationship in favor of their λ-model. They introduce current income in the consumption equation to capture effects of liquidity constraints. We prefer to maintain the interest rate effect that features explicitly in our theoretical model and avoid current income (or OECD cyclical indicators) because we cannot distinguish between transitory and permanent changes in income.

TABLE 4 – ESTIMATION RESULTS WITH BUSINESS CYCLE EFFECTS

$$\Delta log\ c_t = (a_0 + \sum a_0^i D_i) + a_1\ log(1 + r_t) + \eta_t$$

	CN	FR	GE	JP	NL	SW	UK	US
a_0	3.0505[a] (7.33)	3.8885[a] (3.58)	2.3233[a] (2.89)	3.1492[a] (4.02)	2.6382[a] (4.15)	1.6391[b] (2.15)	3.3976[a] (6.72)	2.4330[a] (8.37)
a_0^1	−0.4602 (0.34)	−0.7959 (0.70)	−0.0188 (0.01)	1.5334 (0.59)	−0.7091 (0.39)	2.7384 (1.23)	−3.1138[b] (2.08)	−2.5504[a] (3.39)
a_0^2	−4.9846[a] (4.05)	−1.3998[c] (1.93)	−2.2107 (1.47)	0.0616 (0.02)	3.4024 (1.47)	−0.0911 (0.06)	−3.8068[a] (3.12)	−2.2233[a] (4.37)
a_0^3	0.3231 (0.20)		−1.9422 (1.34)	−2.1358 (1.59)	−2.5236[b] (2.11)	−3.7513[b] (2.26)	−0.7570 (0.48)	−1.7493[c] (1.66)
a_0^4				−2.3506 (1.18)		−1.6804 (0.48)		
a_1	−0.0515 (0.42)	−0.1101 (0.99)	0.0776 (0.27)	0.2436 (1.40)	−0.1741 (1.05)	0.1601 (0.98)	0.0991 (0.95)	0.0594 (0.75)
SE	2.557	2.781	3.541	4.383	3.875	4.710	3.288	1.756
DW	2.10	2.67	2.37	2.39	2.68	2.52	1.94	1.94
$Q(8)$	3.48	20.50[a]	11.42	18.06[b]	18.39[b]	9.57	16.57[b]	23.82[a]
Lagrange Multiplier tests for the overidentifying restrictions with instrument list (degrees of freedom):								
I (3)	0.45	4.65	1.99	15.54[a]	1.84	1.38	8.93[b]	7.93[b]
II (3)	0.77	1.11	2.54	8.68[b]	2.07	3.98	5.08	2.59
III (5)	4.81	4.80	7.91	16.70[a]	2.86	7.85	10.65[c]	8.47
IV (8)	5.77	8.72	9.60	19.25[b]	10.00	11.01	15.87[b]	12.81
V (11)	15.66	11.74	12.38	18.60[c]	6.43	13.69	15.43	12.47
VI (9)	6.14	9.90	9.95	17.18[b]	6.45	12.10	14.23	9.81

Notes:

Estimation period is 1971:2–1989:4. Regression results are for IV/2SLS estimation using instrument list V (see Table 2). The dummy variables D^i have value one during each individual major business cycle recession period. $Q(8)$ is the Ljung-Box Q-statistic. The LM test for the rational expectations restrictions equals $n * R^2$, with n the number of observations and R^2 the coefficient of determination for the CLS regression of the residuals on the instruments. LM is $\chi^2(k - l)$ distributed with k the number of instruments and l the number of coefficients estimated. Superscripts a, b, and c indicate significant test statistics at 1, 5, and 10 percent levels, respectively.

Elsewhere, it has been suggested that the financial deregulation in industrialized countries after 1979 affected the operation of liquidity constraints (*i.e.* Bayoumi and Koujianou (1989), Blundell and Browne (1991)). However, it is difficult to identify specific institutional developments, their timing or their effects on household liquidity constraints. Financial deregulation is a slow-moving, continuous process embedded in a more general pattern of financial markets innovations. Ferson and Merrick (1987) refer to the 1979 change in Federal Reserve operating procedures as a major cause for changes in the underlying behavior of US real interest rates. Person (1989) examines stochastic nonstationarity in macroeconomic time series. He identifies major regime shifts in the GNP time-series behavior.[13] In a European context we refer to the establishment of the exchange rate mechanism of the European Monetary System as a probable cause of changes in the behavior of European interest rates. Winder and Palm (1989) and Palm and Winder (1990) examine models in which the trend rate of growth of consumption is related to changes in the trend rate of growth of income. Multiple dummy variables, based on visual inspection of the income series, improve the time-series characteristics of consumption. However, they assume a constant real interest rate and therefore make no inferences about intertemporal substitution.

Our approach is to include a single dummy variable in our variable interest rate regression to capture the effect of a change in the joint behavior of consumption and interest rates after 1979.[14] In the context of Palm and Winder, this could relate to the shift towards more stable (less uncertain) but reduced income growth in the 1980s. The results are displayed in Table 5.

Allowing a shift after 1979 improves our results favorably in terms of estimated coefficients. Real interest rate coefficients are now significant and positive for Japan, Sweden, the U.K., and U.S. All coefficients, except in the German equation, now have the expected positive sign and the point estimates are similar in magnitude across countries. Just as with our business cycle dummies, allowing a shift in the intercept of the equation reduces the overidentifying restrictions test statistics. However, formally, the LCPIH-RE restrictions are still rejected, and most clearly rejected in the case of the United Kingdom and Japan.

An intriguing question that arises from the previous analysis is why the rejection of the life cycle-permanent income hypothesis is especially prominent in the case of the United Kingdom and Japan. Our results corroborate those of Japelli and Pagano (1989). Their estimates of the presence and size of liquidity constraints ranked Sweden and the U.S. below the U.K. and Japan (with even

13 Perron (1989) identifies the 1929 crash and 1973 oil crisis as major regime shifts in GNP time-series behavior for the US, not, however, 1979. In contrast, Balke and Fomby (1991) identified shifts in quarterly real GNP in 1980/1981, not, however, in 1973.

14 The intercept is a function of σ^2, β, and γ. One referee has pointed out that allowing a shift in the intercept, but at the same time assuming that the coefficients for the real interest rate and the residual variance are unchanged, must imply that it is the rate of time preference β that has changed.

TABLE 5 – ESTIMATION RESULTS WITH A SHIFT IN THE INTERCEPT TERM AFTER 1979:4

$$\Delta log\ c_t = (a_0 + a_0' D_{79}) + a_1\ log(1 + r_t) + \eta_t$$

	CN	FR	GE	JP	NL	SW	UK	US
a_0	3.4166[a]	1.8197	3.0564[a]	3.9946[a]	2.4127[a]	2.9692[a]	3.5833[a]	2.6485[a]
	(5.83)	(1.63)	(3.46)	(5.59)	(4.83	(3.11)	(4.68)	(6.64)
a_0'	−2.7591	−2.6822[b]	−0.8804	−2.8616[b]	−3.4765[c]	−3.8625[b]	−2.3783[c]	−3.3536[a]
	(1.34)	(2.21)	(0.68)	(2.46)	(1.92)	(2.08)	(1.76)	(3.50)
a_1	0.2488	0.2392	−0.1929	0.3208[a]	0.1488	0.4602[c]	0.3657[a]	0.5564[a]
	(0.79)	(1.30)	(0.47)	(2.63)	(0.52)	(1.95)	(2.80)	(3.43)
SE	2.897	2.640	3.495	4.149	3.891	4.704	3.404	2.051
DW	1.57	2.74	2.47	2.45	2.59	2.50	1.90	1.35
$Q(8)$	8.70	21.97[a]	14.40[c]	19.64[b]	16.08[b]	8.06	9.34	16.68[b]
Lagrange Multiplier tests for the overidentifying restrictions with instrument list (degrees of freedom):								
I (3)	4.35	4.92	1.55	12.69[a]	1.42	1.53	4.01	4.36
II (3)	1.76	1.45	2.25	6.87[c]	3.99	5.39	8.60[b]	0.66
III (5)	6.18	4.95	8.24	17.36[a]	2.67	6.16	15.06[b]	4.81
IV (8)	9.45	8.79	9.96	18.26[b]	10.13	9.72	17.59[b]	18.72[b]
V (11)	18.48[c]	11.80	14.61	19.12[c]	9.63	14.08	20.30[b]	8.34
VI (9)	9.50	9.55	11.27	19.31[b]	10.39	11.82	18.45[b]	9.41

Notes:

Estimation period is 1971:2–1989:4. Regression results are for IV/2SLS estimation using instrument list V (see Table 2). The dummy variable D_{79} has value one, starting 1980:1. $Q(8)$ is the Ljung-Box Q-statistic. The LM test for the rational expectations restrictions equals $n * R^2$, with n the number of observations and R^2 the coefficient of determination for the CLS regression of the residuals on the instruments. LM is $\chi^2(k - l)$ distributed with k the number of instruments and l the number of coefficients estimated. Superscripts a, b, and c indicate significant test statistics at 1, 5, and 10 percent levels, respectively.

higher estimates of liquidity restrictions for Italy, Spain and Greece, countries that are not included in our sample).

5 CONCLUSIONS

In this paper we provide international evidence on the life cycle-permanent income hypothesis. We test the rational expectations restrictions and examine equilibrium relationship between consumption and the real interest rate relationship in a unified framework and with a consistent data set for eight countries – Canada, France, Germany, Japan, The Neteherlands, Sweden, the United Kingdom, and the United States. Cross-country comparisons can tell us whether rejections of the LCPIH are general or country specific. Formally, we must reject the rational expectations restrictions. For different countries the standard test results depend on the information set used and residual autocorrelation must be examined in more detail. Rejection of the RE restrictions is most evident for Japan and the United Kingdom, and perhaps the United States.

We proposed to examine non-stationarity in the consumption-interest rate relationship as a possible source for rejection of the RE restrictions. We tested the effects of business cycles and a regime shift. The interest rate effect closely follows the business cycle. We concluded that incorporating independent business cycle effects does not improve our specification, neither in terms of reducing the RE restrictions nor in terms of more reliable estimates of the intertemporal substitution effect. Introducing a post-1979 regime shift affected the measurement of intertemporal substitution effects. We found positive real interest rate coefficients (except for Germany) in the range of 0.15–0.56 and a five percent significance range of 0.46–0.56 for Japan, Sweden, the United Kingdom, and the United States. However, the rejection of the RE restrictions remains, and is prominent for two of the eight countries – the United Kingdom and Japan. An intriguing question is why this is the case for precisely these countries. This, however, as well as a more adequate examination of serial correlations, is left for future research.

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Summary

THE PERMANENT INCOME HYPOTHESIS, BUSINESS CYCLES, AND REGIME SHIFTS: EVIDENCE FROM EIGHT COUNTRIES

We provide international evidence on the joint behavior of consumption and the real rate of interest and examine the rational expectations restrictions of the permanent income hypothesis. We extend the basic model to allow for independent effects of the stage of the business cycle or a regime shift after 1979. In our eight-country sample (using 1970s–1980s data) we find a small but internationally similar rate of intertemporal substitution once we allow for a regime shift affecting the average growth of consumption after 1979. The rational expectations restrictions are formally rejected, most prominently for the United Kingdom and Japan.