UK Consumers' Expenditure over 40 Years

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Abstract

Using quarterly data for the last four decades, we test a number of traditional assumptions about aggregate consumer behaviour in the UK, with regard to the order of integration of the time series, the income elasticity of consumption and the stability of the parameters of the consumption function. In all cases, modification of these assumptions now appears to be necessary.

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1. Introduction

It is now two decades since Davidson et al. (1978) introduced the model of consumer expenditure which has become the template for empirical work on aggregate UK consumption, and the quantity of data available is now twice as large as that with which the original paper dealt. In this paper we will use data running up to 1997 to test some of the underlying assumptions of the DHSY model which are usually taken for granted:

• That income and consumption are difference stationary rather than trend stationary, so that the appropriate way to model consumption is by using the error correction isomorphism.
• That the long run income elasticity of consumption is equal to one.
• That there is no trend in the deterministic component of the consumption function, i.e., that the propensity to consume out of income in the steady state is constant, except perhaps for some seasonal variation.

These assumptions are more than just theoretical curiosa: if it turns out that any of them is unjustified then traditional DHSY-type models of consumption used, for example, in macroeconomic forecasts will be misspecified, bringing into question the reliability of their predictions. They also imply strong claims about the stability and simplicity consumer preferences: that the parameters of the representative consumer's intertemporal utility function are constant, and that this function is simple enough to deliver a unit income elasticity. It would be a remarkable finding if consumer preferences were demonstrated to comply with these restrictions over the four decades for which data exist.

Section 2 of the paper presents an overview of the characteristics of the time series data used in standard consumption functions. The focus of attention is particularly on inferences about the order of integration of the series, using both the traditional Dickey-Fuller methodology and more recent tests which allow for more complexity in the deterministic components of the series. Section 3 then presents a structural econometric model which is consistent with the findings in section 2 and which permits tests of assumptions about the structure of the aggregate consumption function.

2. Long-Run Trends in Income and Consumption

Visual inspection of income and consumption time series for the last four decades suggests that they are stationary around a deterministic trend. Figure 1 plots two of the series which will feature in the econometric model: the logarithms of real unadjusted nondurable consumption (c_t) and real unadjusted personal disposable income (y_t), taken from Economic Trends.[1] This suspicion is confirmed by standard Augmented Dickey-Fuller tests (Tables 1-2), which indicate that the null that the series are I(1) can be rejected at the 5% level. However, there are substantial departures from this trend in the middle of the sample period. In the early 1970s (from about the time of the first oil price hike) there is a deterioration in the rate of growth of income and consumption; and in the mid 1980s there is an increase in their growth rate which ends in about 1988. It is therefore not surprising that papers modeling income and consumption from the late 1960s to the late 1980s (for example, Carruth and Henley, 1990, and the papers they cite) should infer or assume that these variables are I(1): this is precisely the period in which the series have deviated substantially from their long run trends. The standard deviations of the two variables around their fitted whole-sample linear trends are as follows:[2]

	1955-1969	1970-1989	1990-1997
c_t	1.58%	4.08%	2.20%
y_t	3.24%	3.95%	2.21%

In other words, the two decades which have been intensively studied in previous papers represent a period of unusually high macroeconomic instability, giving the (false) impression that macroeconomic time series are I(1); the increase in sample variance in the middle column above does not represent a permanent feature of the data. Estimation over longer sample periods indicates the stationarity of these series.

[1] There is no adjustment to y_t for perceived capital gains or losses. The results of using an income measure in which such an adjustment is made are available on request, and are very similar to the ones reported below.

[2] The deterministic trends here include a seasonal component.

[Figure 1 and Tables 1-2 here]

There remains the possibility that the deviations from the linear trend are partly deterministic, i.e., that the trend is really non-linear. Following the logic of Perron (1988), this possibility does not necessitate the estimation of stationarity test statistics for c_t and y_t around a non-linear trend (if we could not reject the null of non-stationarity when the non-linear components were included in the autoregression, we would end up testing down to the linear form anyway). However, there are other macroeconomic time series for which the inclusion of a non-linear trend does make a difference to inferences about the order of integration. One example is consumer price inflation qua the rate of growth of the personal consumption deflator ($\Delta p_t = \pi_t$), depicted in Figure 2.[3] An ordinary ADF test does not lead to rejection of the null that inflation is I(1), but a stationarity test based on the methodology of Leybourne et al. (1995) does lead to such a rejection. The test employed (reported in Table 3) allows for a smooth transition from a low inflation rate to a higher one in the 1970s, and a smooth transition back down to a lower rate in the 1980s, by testing for the stationarity of the residual u_t from the regression:

$$\pi_t = \alpha_0 + \alpha_1 \cdot t + [\beta_0 + \beta_1 \cdot t] \cdot S_A(t) + [\gamma_0 + \gamma_1 \cdot t] \cdot S_B(t) + \Sigma \mu_j \cdot Q(j)_t + u_t \qquad (1)$$

$$S_A(t) = [1 + \exp(-\theta \cdot (t-\phi))]^{-1}; S_B(t) = [1 + \exp(-\eta \cdot (t-\iota))]^{-1}$$

where the $Q(j)_t$ are quarterly dummies. This test allows for instantaneous structural breaks (as the limit when $\theta \rightarrow \infty$ or $\eta \rightarrow \infty$), but does not impose instantaneity; for finite θ and η the parameters ϕ and ι represent not breakpoints as such but midpoints in smooth transitions between two linear trends. For π_t neither of the transitions is instantaneous: $\theta = 0.04$ and $\eta = 0.98$. The midpoints are at 1970(1) and 1986(1), implying a long upward transition beginning in the late 1960s and continuing until the mid 1970s, and a slightly shorter downward transition in the late 1980s. The ADF t-ratio for u_t is -9.90, which entails rejection of the null that π_t is I(1) at the 5% level.[4]

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[3] N.b. in order to remove the seasonal component of inflation in the figure, it depicts the annual moving average of the series used in the regression analysis: i.e., the figure plots $\Delta_4 p_t/4$ and the regressions use Δp_t.

[4] N.b. the critical values of the test (computed by Monte-Carlo methods) are higher in absolute value than the ones in Leybourne et al. (1995) because we are allowing for two smooth transitions, not just one.

[Figure 2 and Table 3 here]

Two other macroeconomic time series will be used in our consumption model: a measure of real personal sector net financial wealth (w_t) and real lending to the personal sector (b_t). In order to have a consistent definition of wealth for as long a period as possible, w_t is based on a national accounting definition: it is the stock to date of gross personal sector saving net of gross personal sector physical investment, deflated by the consumer price deflator (so $\Delta [p_t \cdot w_t]$ is equal to "personal sector financial surplus" in Economic Trends). It is this variable which constrains the length of the sample on which the final model is estimated, observations on w_t beginning only in 1963. The trend in the w_t series is similar to those in the income and consumption series, and the null that it is I(1) can be rejected without recourse to a smooth transitions model (Table 4). For b_t, a single smooth transition is required, to capture the collapse of real lending to the personal sector in the 1990s (largely associated with the house price collapse and emergence of negative equity), which is evident in Figure 3. The smooth transition parameters are reported in Table 5, along with the stationarity test results. Here the t-ratio is only -4.64, and the null of nonstationarity can be rejected at only the 10% level (see Leybourne et al., op. cit., Table 1); so t-ratios on this variable in later regressions should be treated with some caution.

[Figure 3 and Tables 4-5 here]

3. The Consumption Model

(i) Model structure

The DHSY consumption function (Davidson et al., op. cit.), and its many subsequent modifications, are designed to model consumption in a world of I(1) variables. The basic form of this function is:

$$\Delta_n c_t = \beta_0 + \beta_1 \cdot \Delta_n y_t - \beta_2 \cdot [c_{t-n} - y_{t-n}] + \ldots + u_t \qquad (2)$$

When n = 1 this is an autoregressive distributed lag model of c_t on y_t (plus various extras), reparameterised in error-correction format with the imposition of a unit elasticity in the cointegrating vector. When n > 1 the equation represents a restricted n^{th} order ADL model. There are various theoretical interpretations of such a model, but a common thread to all is that consumption can in the long run be described by aggregating over households behaving according to a lifecycle model with a C.E.S. utility function, which ensures that planned consumption is proportional to permanent income; the coefficient β_2 measures the speed at which consumers adjust towards an equilibrium in which steady-state c = y. More recent versions of equation (2) include an error correction term in financial wealth, $[w_{t-n} - y_{t-n}]$, allowing consumers to adjust spending so that in the steady-state financial assets are a fixed fraction of income (Price, 1989, Whitley, 1989, Carruth and Henley, op. cit.). Some also include a measure of real lending to the personal sector, to allow for the possibility that some consumers are credit-constrained, and terms in the real interest rate (to capture changes in the slope of the intertemporal budget constraint) and inflation.

The great variety of model specifications which arose in the late 1980s was driven by the failure of successive vintages to forecast changes in consumer spending in this period. Interest focused on extending the set of explanatory variables, and fine-tuning their definitions, in order to produce a model which correctly anticipated the consumer boom. In this paper, we wish to explore two conjectures about the traditional model: (i) that the long-run unit elasticity restriction is inappropriate, because preferences are not characterized by a C.E.S. utility function (and something more general, for example an L.E.S. function, is needed);[5] (ii) that the deterministic component of the model has not been constant: some of the changes in consumer spending reflect changes in aggregate behavioural patterns which need to be incorporated into the model.

Visual inspection of the consumption time series in Figure 1 suggests that the most likely periods for such changes to have taken place are the early 1970s and the mid 1980s, especially since these are also periods of marked structural change in the evolution of some of the key potential explanatory variables, such as inflation. However, the existence of non-linear trends in the explanatory variables does not entail the existence of a non-linear trend in the model of consumption conditional on these variables. For example, it might be the case that the deviations in consumption from its linear trend are due entirely to deviations in income and inflation, the

[5] See for example Deaton and Muellbauer (1980, page 324) on intertemporal utility functions.

parameters of the consumption function remaining constant throughout. In order to determine whether the structure of the function has changed, we need to estimate an equation which allows for non-linearities in the time trend.

The form of the function which we shall estimate is as follows:

$$c_t = f(t) + \sum_{i=1}^{i=NC} \delta_i \cdot c_{t-i} + \sum_{i=0}^{i=NY} \kappa_i \cdot y_{t-i} + \sum_{i=0}^{i=N\Pi} \lambda_i \cdot \pi_{t-i} + \sum_{i=1}^{i=NW} \mu_i \cdot w_{t-i} + \sum_{i=1}^{i=NB} \nu_i \cdot b_{t-i} + u_t \qquad (3)$$

where the variables are defined as above, and f(t) is some deterministic, but not necessarily linear, function of time. For each variable x, the lag order NX is chosen so as to optimize standard model selection criteria (Schwartz, Hannon-Quinn and Akaike: the lag order chosen does not depend on which criterion is used).[6] We anticipate that the long run elasticities estimated as $\Sigma_i\kappa_i/[1 - \Sigma_i\delta_i]$, $\Sigma_i\mu_i/[1 - \Sigma_i\delta_i]$ and $\Sigma_i\nu_i/[1 - \Sigma_i\delta_i]$ will be positive, but there is no a priori assumption that the first two will be equal to unity. In the standard DHSY specification the short run inflation elasticity (λ_0) is negative and the long run elasticity ($\Sigma_i\lambda_i/[1 - \Sigma_i\delta_i]$) is zero, which is often interpreted as evidence that consumers find it difficult in the short run to distinguish between real and nominal shocks. However, we will not impose any long run restriction on the inflation elasticity.[7]

In order to capture the possibility of two deterministic changes in the consumption function over the sample period, f(t) was allowed to take the form:

$$f(t) = \alpha_0 + \alpha_1 \cdot t + [\beta_0 + \beta_1 \cdot t] \cdot S_A(t) + [\gamma_0 + \gamma_1 \cdot t] \cdot S_B(t) + \Sigma_j \mu_j \cdot Q(j)_t \qquad (4)$$

$$S_A(t) = [1 + \exp(-\theta \cdot (t-\phi))]^{-1}; S_B(t) = [1 + \exp(-\eta \cdot (t-\iota))]^{-1}$$

i.e., the time trend can have up to two transitions with midpoints at ϕ and ι. t = 1 in 1955(1). (It is possible to build more than two transitions into the model, but no more than two were found to be

[6] Contemporaneous values of w and b are excluded from the regression because of their likely endogeneity. The regression can be seen as a reduced form in which w_t and b_t are modeled as AR(NW) and AR(NB) processes. y and π are treated as exogenous; the magnitude of long run coefficients in a model which excludes contemporaneous y and π are very similar to the ones reported below, which suggests that the exogeneity assumption is not biasing the results.

[7] Some papers (for example, Carruth and Henley, 1990) include nominal or real interest rates in the consumption model. When lags of the nominal interest rate (qua the t-bill rate) are added to the regression equations reported below, they are jointly and individually insignificant.

significant.) The deterministic component of the final consumption equation estimated, which is reported in Table 6, is slightly more restrictive than this: γ_1 was found to be insignificantly different from zero, and is omitted, and the transition in $S_A(t)$ turned out to be instantaneous: fitted $S_A(t)$ is within 10^{-15} of zero before 1975(1) and within 10^{-15} of unity from this date onwards. The parameters θ and ϕ are not reported in the table: θ is set at infinity and ϕ is set at 81.

In the equation reported in Table 6 the lag orders are, respectively, NC = 6, NY = 4, NΠ = 1, NW = 1 and NB = 1. With this specification the equation passes standard diagnostic tests, which are reported at he bottom of the table. Moreover, the parameters appear to be stable over time. Recursive estimation with final observations ranging from 1991(4) to 1997(4) (well after the second transition has worked itself out) does not produce significant forecast Chow Test statistics (Figure 4), and one-step forecast residuals are well within the two standard error bar (Figure 5). The parameter instability statistics of Hansen (1992) are all insignificant; the joint parameter mean (H_1) and joint parameter variance (H_2) statistics are reported in the table.

[Figures 4-6 and Table 6 here]

(ii) The deterministic component of the model

All seven of the deterministic parameters ($\alpha_0, \alpha_1, \beta_0, \beta_1, \gamma_0, \eta, \iota$) are significant. Their interpretation is best discussed in the context of Figure 6, which plots the sum of the deterministic components of the model, excluding the seasonals; that is, it shows how consumption would have evolved had other variables remained constant. First, there is a sudden drop in the exogenous rate of growth of consumption at the beginning of 1975, from around 0.070% per quarter to around 0.018% per quarter. Second, there is a gradual increase in consumption between 1983 and 1988, which does not however correspond to any increase in the long run exogenous growth rate ($\gamma_1 = 0$). The midpoint of this transition is at the beginning of 1986. Consumption in any one period after the transition is around 4.38% higher than it would have been without the transition. These changes are estimated after having controlled for income, inflation, financial wealth and the supply of credit. In other words, there have been substantial changes in consumer behaviour over the sample period which are not explained by changes in the variables contained in a standard aggregate consumption function.

(iii) Consumption elasticities

The slope coefficients of the consumption function are most readily interpreted if they are presented in "error correction" format; Table 6 also shows these figures. These represent the same regression as the autoregressive distributed lag equation, but with Δc_t as the dependent variable, only the first lag of each explanatory variable in levels, and [NX - 1] lags of each variable in differences. Also of interest are the long run coefficients on each explanatory variable; these too are reported in the table.

Current consumption growth depends positively on current income growth, and on income growth over the last year; the coefficient on Δy_t (0.26) is greater than that on Δy_{t-1} (0.16), which is greater than that on Δy_{t-2} and Δy_{t-3} (0.10). Other things being equal, a recent history of high income growth will encourage more growth in consumption. The steady state elasticity is 0.52; this is significantly different from both zero and unity, indicating that the traditional unit elasticity restriction is invalid. (Though in a model without a deterministic trend this might not be apparent: the common trend to income and consumption will tend to push up the income elasticity.) This result suggests that the standard assumption that intertemporal preferences can be represented within a C.E.S. framework is incorrect.

As in traditional consumption models there is a negative coefficient on inflation growth, and this is of a similar order of magnitude to figures reported in previous papers; here the estimate is -0.18. However, there is also a long run coefficient of -0.97, significantly different from zero but not unity: higher inflation permanently depresses consumption. It is somewhat implausible to attribute the long run effect to a signal extraction problem; one alternative explanation is that consumers associate high inflation with economic instability (for example it may be seen as a precursor of fiscal contraction), and so tend to engage in more precautionary saving.

Higher levels of financial wealth are associated with higher consumption, as in previous studies, but the long run elasticity (0.06) is much smaller than, and significantly different from unity. Similarly, the coefficient on lending to the personal sector, though significantly positive, is very small. The estimated long run elasticity is only 0.01. It is perhaps worth emphasizing again that these coefficients appear to be time-invariant: there is no significant change in the sensitivity of consumption to lending over the sample period, i.e., no period in which credit constraints were more important.

Current Δc_t depends negatively on Δc_{t-1}, Δc_{t-2} and Δc_{t-3}, and positively on Δc_{t-4} and Δc_{t-5}; one interpretation of these coefficients is that current consumption growth depends negatively on both past consumption growth and the change in past consumption growth. Consumers are more reluctant to increase their spending if consumption has been growing recently, and are even more reluctant if this growth has been accelerating. This introduces an intuitively appealing conservatism into the model. If the "error correction" term is normalized on consumption, the error correction coefficient is -0.34, a figure somewhat larger in absolute value than previously estimated adjustment coefficients; adjustment to the steady state may be rather more rapid than is implied in the DHSY model.

(iv) Overview

There is a significant nonlinear trend in the intercept aggregate consumption function, whilst the slope parameters of the function have remained stable. The trend is upward-sloping: had aggregate income remained constant, aggregate consumption would still have risen. The simplest explanation for the overall upward trend is the increase in the average age of the population in the postwar period: as the fraction of the population past retirement age (and therefore in a period of dissaving in the lifecycle) increases, aggregate consumption increases for any given aggregate income level. The reason that the savings ratio has not collapsed entirely is that real income has risen, and the marginal propensity to consume out of income is less than unity.

However, there have been changes in this trend path: a sharp fall in the rate of exogenous consumption growth after 1975, and a temporary increase in the growth rate in the mid-late 1980s.

One interpretation of these transitions is that they represent changes in consumer attitudes: first of all an increase in precautionary saving by younger generations as the oil crisis led to an increase in perceived macroeconomic instability; then a temporary "feel good" boom in the 1980s.

4. Summary and Conclusion

The availability of time-series income and consumption data from the 1990s has facilitated the estimation of a consumption model which overturns some of the conventional wisdom about aggregate consumer behaviour. First, when a long enough series is used, the null hypothesis of non-stationarity can be rejected for both income and consumption. Second, traditional time-series consumption functions have assumed a long run unit income elasticity, and have not allowed for any deterministic trend in the aggregate propensity to consume. Common trends in income and consumption data have given the impression that these assumptions are correct, but a time-series model based on data from the early 1960s to the late 1990s suggests that the income elasticity is far less than unity, and that there is a strong trend in the propensity to consume.

Moreover, there is some evidence that the growth rate of the propensity to consume has not been constant: the 1970s and 1980s saw marked deviations away from the long run trend, which might be interpreted as changes in consumers' preferences in response to the perceived degree of stability and growth in the UK economy. There appear to have been marked upturns and downturns in consumer confidence.

The existence of such deviations, even when controlling for wealth and credit constraint effects, suggests that any forecasts based on a time series model need to be treated with extreme caution. Whilst the model presented here forecasts reasonably well over the mid-late 1990s, there is every reason to suspect that future decades will see changes in consumer confidence to match the swings of the 1970s and 80s. Until there is a way of predicting such changes, any forecast based on a model of aggregate consumption should be regarded as provisional.

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Table 1: Stationarity Test for c_t

Sample: 1957(2)-1997(4)

Regression: $\Delta c_t = \alpha_0 + \alpha_1 \cdot t + \Sigma_i \delta_i \cdot \Delta c_{t-i} + \kappa \cdot c_{t-1} + u_t$ + seasonal[8]

variable	coeff.	std. err.	t ratio
Δc_{t-1}	-0.09112	0.07831	-1.164
Δc_{t-2}	-0.02715	0.07116	-0.381
Δc_{t-3}	-0.03133	0.07773	-0.403
Δc_{t-4}	0.41836	0.07680	5.448
Δc_{t-5}	0.18456	0.07848	2.352
Δc_{t-7}	0.07852	0.07938	0.989
Δc_{t-8}	0.24856	0.07883	3.153
α_0	1.04730	0.28039	3.735
α_1/1000	0.60543	0.16555	3.657
$Q(1)_t$	-0.05327	0.01516	-3.514
$Q(2)_t$	-0.00239	0.01132	-0.211
$Q(3)_t$	-0.02156	0.01459	-1.478
c_{t-1}	-0.09933	0.02714	-3.660

$R^2 = 0.97413$

Residual autocorrelation (order 1): $F(1,149) = 1.31210$ [0.2539]

Residual autocorrelation (order 4): $F(4,146) = 1.21860$ [0.3055]

[8] u_t is the regression residual; the seasonal for the j^{th} quarter is denoted $Q(j)_t$. t = 1 in 1955(1).

Table 2: Stationarity Test for y_t

Sample: 1958(3)-1997(4)

Regression: $\Delta y_t = \alpha_0 + \alpha_1 \cdot t + \Sigma_i \delta_i \cdot \Delta y_{t-i} + \kappa \cdot y_{t-1} + u_t$ + seasonal

variable	coeff.	std. err.	t ratio
Δy_{t-1}	-0.15404	0.09044	-1.703
Δy_{t-2}	0.18065	0.09037	1.999
Δy_{t-3}	-0.08531	0.09156	-0.932
Δy_{t-4}	0.26588	0.08755	3.037
Δy_{t-5}	0.20084	0.08895	2.258
Δy_{t-6}	0.20946	0.09061	2.312
Δy_{t-7}	-0.02022	0.09021	-0.224
Δy_{t-8}	0.15239	0.09026	1.688
Δy_{t-9}	0.08739	0.09125	0.958
Δy_{t-10}	-0.14590	0.09104	-1.603
Δy_{t-11}	-0.01589	0.08918	-0.178
Δy_{t-12}	0.04428	0.08903	0.497
Δy_{t-13}	0.18772	0.08413	2.231
α_0	2.34090	0.64642	3.621
α_1	0.14271	0.04000	3.567
$Q(1)_t$	-0.01341	0.00697	-1.924
$Q(2)_t$	0.01703	0.00673	2.531
$Q(3)_t$	0.00325	0.00696	0.467
y_{t-1}	-0.22298	0.06181	-3.607

$R^2 = 0.69441$

Residual autocorrelation (order 1): $F(1,138) = 0.04098$ [0.8399]

Residual autocorrelation (order 4): $F(4,135) = 0.28194$ [0.8893]

Table 3: Smooth Transition for π_t

Sample: 1955(2)-1997(4)

Regression: $\pi_t = \alpha_0 + \alpha_1 \cdot t + [\beta_0 + \beta_1 \cdot t] \cdot S_A(t) + [\gamma_0 + \gamma_1 \cdot t] \cdot S_B(t) + u_t$ + seasonal

$$S_A(t) = [1 + \exp(-\theta \cdot (t-\phi))]^{-1};\ S_B(t) = [1 + \exp(-\eta \cdot (t-\iota))]^{-1}$$

variable	coeff.	std. err.	t ratio
α_0	0.00450	0.00696	0.646
$\alpha_1/100$	-0.15197	0.00806	-18.856
β_0	0.23332	0.01087	21.470
$\beta_1/100$	-0.02405	0.01019	-2.361
γ_0	-0.15850	0.00520	-30.493
$\gamma_1/100$	0.13376	0.00438	30.553
θ	0.04477	0.00359	12.481
ϕ	60.4580	1.72330	35.083
η	0.98180		
ι	124.910		
$Q(1)_t$	-0.00422	0.00223	-1.892
$Q(2)_t$	0.00402	0.00222	1.811
$Q(3)_t$	-0.00754	0.00222	-3.399

$R^2 = 0.53923$

Stationarity Test

Sample: 1957(3)-1997(4)

variable	coeff.	std. err.	t ratio
Δu_{t-4}	0.15484	0.06131	2.526
Δu_{t-8}	0.14313	0.06093	2.349
u_{t-1}	-0.73718	0.07446	-9.900

$R^2 = 0.44781$

Residual autocorrelation (order 1): $F(1,158) = 2.11650$ [0.1477]

Residual autocorrelation (order 4): $F(4,155) = 0.66728$ [0.6156]

Table 4: Stationarity Test for w_t

Sample: 1964(2)-1997(4)

Regression: $\Delta w_t = \alpha_0 + \alpha_1 \cdot t + \Sigma_i \delta_i \cdot \Delta w_{t-i} + \kappa \cdot w_{t-1} + u_t$ + seasonal

variable	coeff.	std. err.	t ratio
Δw_{t-1}	0.29993	0.08274	3.625
Δw_{t-2}	0.23425	0.08682	2.698
Δw_{t-3}	-0.01539	0.08927	-0.172
Δw_{t-4}	0.36411	0.08497	4.285
α_0	0.39637	0.10813	3.666
$\alpha_1/1000$	0.36527	0.09813	3.722
$Q(1)_t$	0.02575	0.00490	5.255
$Q(2)_t$	0.01421	0.00543	2.617
$Q(3)_t$	0.00177	0.00506	0.349
w_{t-1}	-0.03810	0.01011	-3.770

$R^2 = 0.67438$

Residual autocorrelation (order 1): $F(1,124) = 0.61440$ [0.4346]

Residual autocorrelation (order 4): $F(4,121) = 0.94712$ [0.4393]

Table 5: Smooth Transition for b_t

Sample: 1963(1)-1997(4)

Regression: $b_t = \alpha_0 + \alpha_1 \cdot t + [\beta_0 + \beta_1 \cdot t] \cdot S_A(t) + u_t$ + seasonal

$$S_A(t) = [1 + \exp(-\theta \cdot (t-\phi))]^{-1}$$

variable	coeff.	std. err.	t ratio
α_0	6.56430	0.13250	49.543
α_1	0.02122	0.00129	16.512
β_0	-5.99780	2.50770	-2.392
β_1	0.02826	0.01546	1.828
θ	0.53849	0.18916	2.847
ϕ	146.790	1.03950	141.221
$Q(1)_t$	-0.24362	0.09800	-2.486
$Q(2)_t$	-0.11394	0.09796	-1.163
$Q(3)_t$	-0.12256	0.09792	-1.252

R^2 = 0.71499

Stationarity Test

Sample: 1963(3)-1997(4)

variable	coeff.	std. err.	t ratio
Δu_{t-1}	-0.29963	0.08171	-3.667
u_{t-1}	-0.40684	0.08770	-4.639

R^2 = 0.352676

Residual autocorrelation (order 1): F(1,135) = 1.81680 [0.1800]

Residual autocorrelation (order 4): F(4,132) = 0.69358 [0.5977]

Table 6: The Consumption Equation

Dependent variable: c_t; Sample: 1963(2)-1997(4)

variable	coeff.	std. err.	t ratio	prob.	Ins.*
c_{t-1}	0.35429	0.08976	3.947	0.0001	0.03
c_{t-2}	-0.01355	0.09140	-0.148	0.8824	0.03
c_{t-3}	0.06370	0.07889	0.807	0.4211	0.03
c_{t-4}	0.50879	0.07793	6.529	0.0000	0.03
c_{t-5}	-0.14692	0.09018	-1.629	0.1060	0.03
c_{t-6}	-0.11036	0.07082	-1.558	0.1219	0.03
y_t	0.25635	0.04482	5.719	0.0000	0.03
y_{t-1}	0.08956	0.05091	1.759	0.0812	0.03
y_{t-2}	-0.06731	0.05118	-1.315	0.1911	0.03
y_{t-3}	0.01304	0.05355	0.244	0.8080	0.03
y_{t-4}	-0.11386	0.05095	-2.235	0.0274	0.03
π_t	-0.17900	0.07588	-2.359	0.0200	0.08
π_{t-1}	-0.15376	0.07968	-1.930	0.0561	0.05
w_{t-1}	0.02036	0.00580	3.507	0.0006	0.03
b_{t-1}	0.00384	0.00158	2.425	0.0169	0.03
$Q(1)_t$	-0.05501	0.01108	-4.966	0.0000	0.23
$Q(2)_t$	-0.01036	0.00965	-1.074	0.2852	0.32
$Q(3)_t$	-0.02644	0.01103	-2.398	0.0181	0.05
α_0	1.48220	0.03232	45.863	0.0000	0.03
$\alpha_1/1000$	0.69542	0.03296	21.098	0.0000	0.03
β_0	0.03993	0.00306	13.044	0.0000	0.03
$\beta_1/1000$	-0.51856	0.03312	-15.657	0.0000	0.03
γ_0	0.04287	0.00323	13.286	0.0000	0.04
η	0.36512	0.10652	3.428	0.0008	___
ι	124.760	1.04630	119.240	0.0000	___

$R^2 = 0.99915$ $\sigma = 0.00802$ RSS = 0.00733

Joint significance: $F(24,114) = 5548.6$ [0.0000]

Residual normality: $\chi^2(2) = 0.86007$ [0.6505]

Residual autocorrelation (order 1): $F(1,113) = 1.36550$ [0.2450]

Residual autocorrelation (order 4): $F(4,110) = 0.81972$ [0.5153]

Heteroscedasticity:* $F(40,75) = 0.93171$ [0.5891]

ARCH (order 1): $F(1,112) = 0.01164$ [0.9143]

ARCH (order 4): $F(4,106) = 0.82356$ [0.5130]

RESET test:* $F(1,115) = 0.88372$ [0.3492]

Ins.: Hansen (1992) parameter instability statistic

Parameter instability test:* $H_1 = 5.15296$

Variance instability test:* $H_2 = 0.33837$

Table 6 (continued)

Long Run Equation

variable	coeff.	std. err.	t ratio	prob.
y	0.51670	0.12360	4.180	0.0000
π	-0.96720	0.45600	-2.121	0.0339
w	0.05917	0.03777	1.567	0.1171
b	0.01115	0.00565	1.973	0.0485

Regression Coefficients in "Error Correction" Format

variable	coeff.	std. err.	t ratio	prob.
Δc_{t-1}	-0.30166	0.10610	-2.843	0.0053
Δc_{t-2}	-0.31521	0.09094	-3.466	0.0007
Δc_{t-3}	-0.25151	0.08454	-2.975	0.0036
Δc_{t-4}	0.25728	0.07213	3.567	0.0005
Δc_{t-5}	0.11036	0.07046	1.566	0.1200
Δy_t	0.25635	0.04475	5.728	0.0000
Δy_{t-1}	0.16813	0.06591	2.551	0.0120
Δy_{t-2}	0.10082	0.05882	1.714	0.0892
Δy_{t-3}	0.11386	0.04950	2.300	0.0232
$\Delta \pi_t$	-0.17900	0.08779	-2.039	0.0437
c_{t-1}	-0.34405	0.09500	-3.621	0.0004
y_{t-1}	0.17779	0.06430	2.765	0.0066
π_{t-1}	-0.33276	0.12295	-2.706	0.0078
w_{t-1}	0.02036	0.00993	2.051	0.0425
b_{t-1}	0.00384	0.00179	2.147	0.0339

Statistics indicated by * are calculated from a regression holding η and ι constant at the values estimated above.

Figure 1: Logarithms of Real Personal Disposable Income (Upper Series) and Nondurable Consumption (Lower Series)

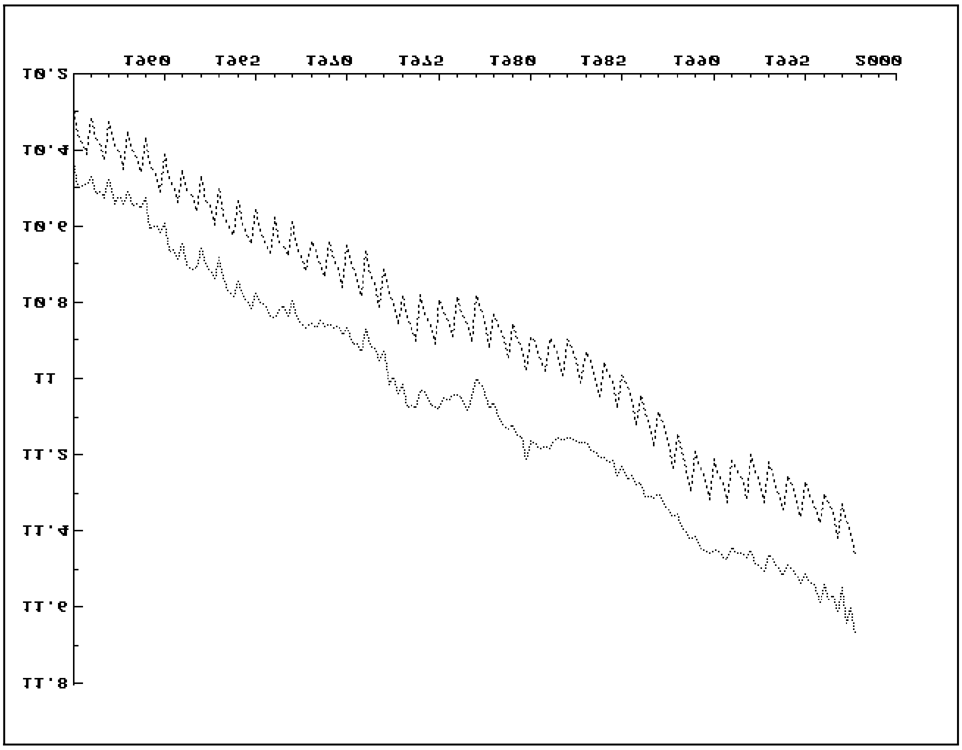

Figure 2: Annual Moving Average of the Consumer Price Quarterly Growth Rate

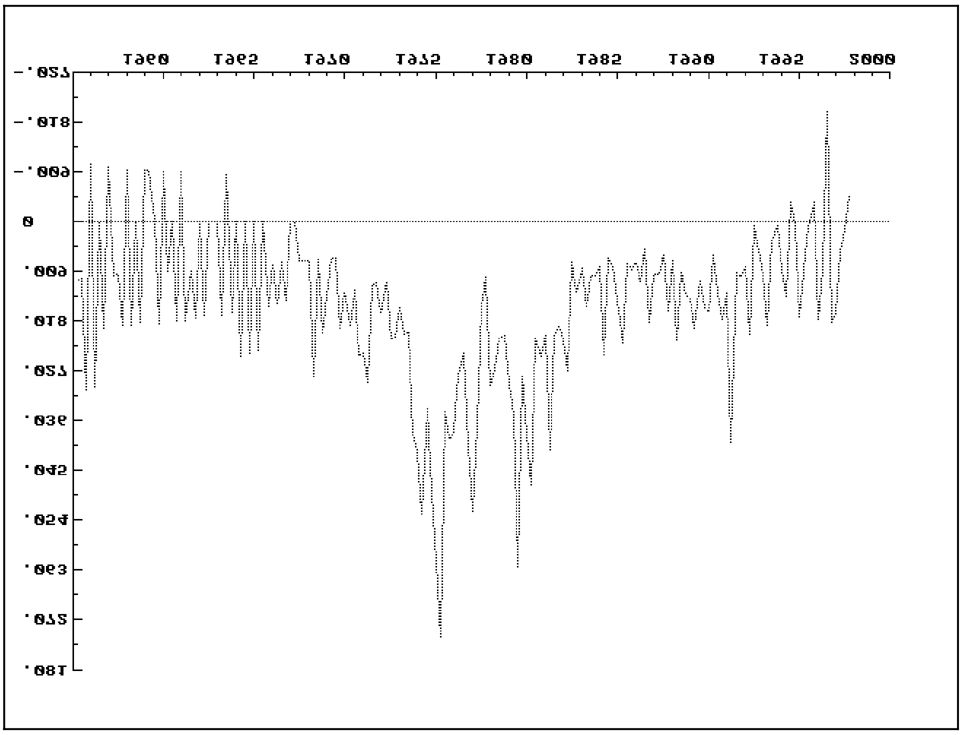

Figure 3: Logarithm of Real Lending to the Personal Sector

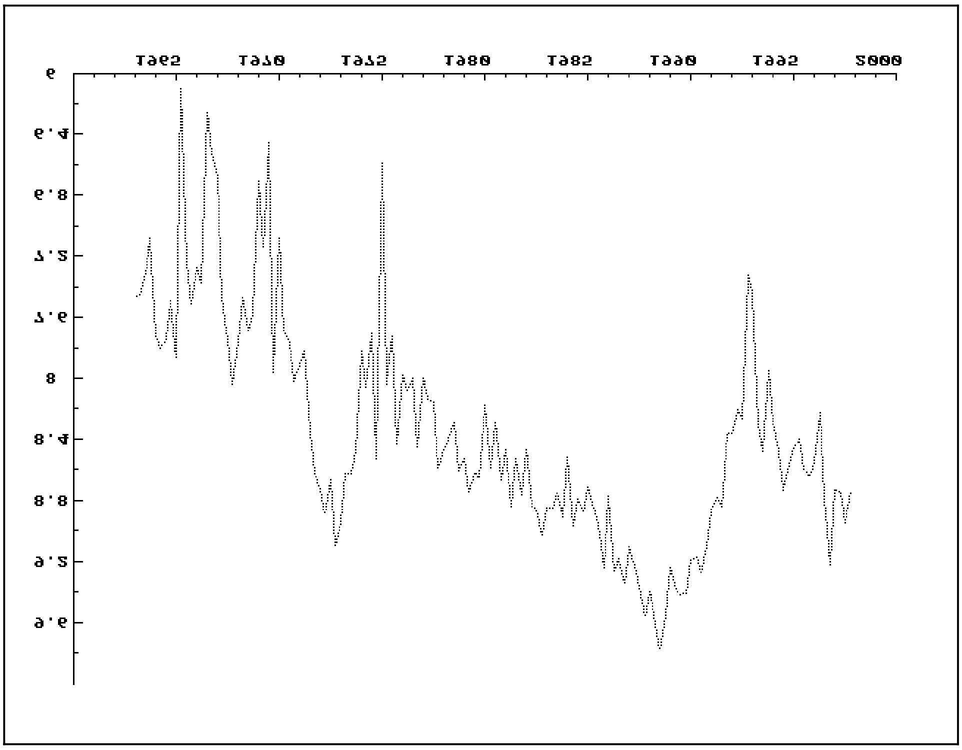

Figure 4: Forecast Chow Test Statistics Relative to the 5% Critical Value

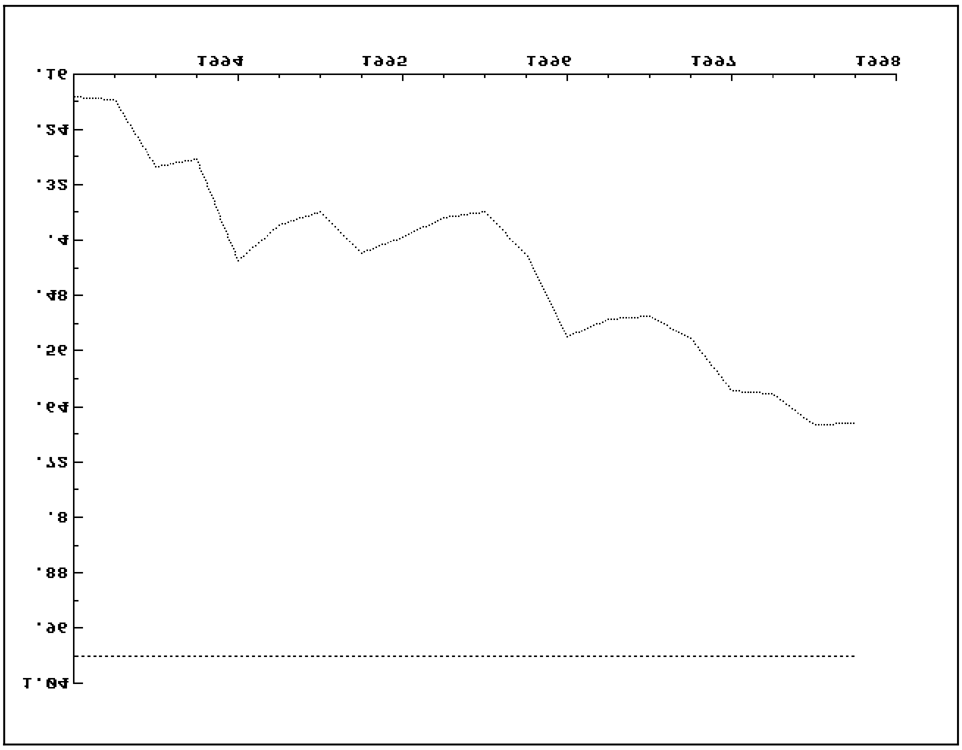

Figure 5: One-Step Forecast Residuals ± Two Standard Errors

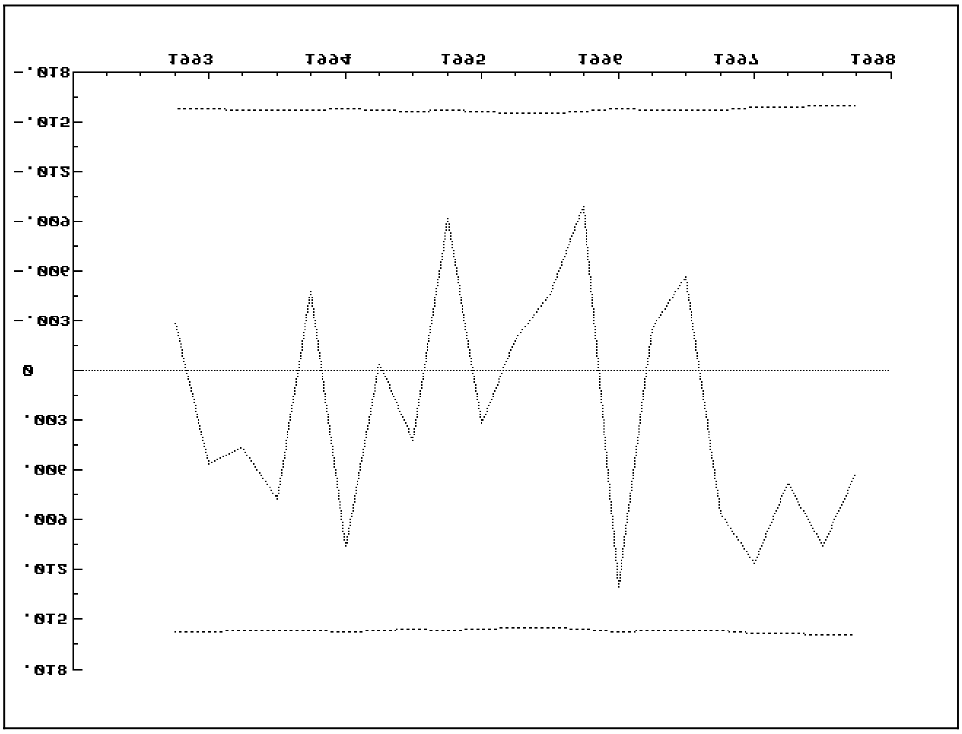

Figure 6: Steady-State Deterministic Component of Consumption