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An International Comparison of Long-Run Consumer Behabiour

Chris Stewart*
London Guildhall University

Abstract

Using the Johansen procedure I test for cointegration between consumption, private disposable income and inflation for 20 OECD countries over the period 1955–1994. There is evidence of cointegration for all countries. Plausible long–run consumption functions are obtained for 18 countries, and feature heterogeneous parameter estimates across countries. Evidence against a unit–income elasticity is obtained for 11 countries suggesting that one would be unwise to assume consumption is homogenous of degree one in income. Inflation is statistically significant and negative for only 7 countries indicating that it is not a fundamental explanatory factor of consumption for many countries. Cross–country regressions for the income elasticity reveal a negative association with income growth, the log–level of income and income inequality and a positive correlation with the fiscal surplus/deficit and the availability of credit. The cross–country regressions of the inflation elasticity are consistent with inflation acting as a proxy for asset effects.

Keywords: Cointegration, cross-country variations, private consumer behaviour, OECD countries.

JEL Classification: C51, C52, D12.

1. Introduction

* Address for correspondence: Department of Economics, London Guildhall University, 31 Jewry Street, London, EC3N 2EY; telephone: 020-7320-1350; email: stewart@lgu.ac.uk. I gratefully acknowledge the helpful comments of George Hadjimatheou, Costas Milas, Joe Pearlman, Nick Sarantis, Ron Smith and two anonymous referees. I am responsible for any errors.

This paper tests for cointegration between consumption, income and inflation and estimates long-run consumption functions for 20 OECD countries. These variables form the long-run relationship utilised in Davidson *et al*'s (1978) [DHSY hereafter] pioneering work, and may be interpreted as approximating Ando and Modigliani's (1963) Life-cycle Hypothesis (LCH) formulation with naive income expectations and inflation proxying wealth effects. Inflation is used to proxy asset effects because data of reasonable coverage on wealth is unavailable for the majority of countries.[1] Since accepted consumer theory strongly suggests that factors beyond current income, especially wealth, influence consumption I do hold a strong prior belief that consumption, income and inflation will form a unique equilibrium relationship and use it to inform the empirical analysis. These estimated consumption functions are used to draw comparisons of consumer behaviour across the OECD.

There are other important influences on aggregate consumption that it would be desirable to investigate, such as, demography, income uncertainty, interest rates and liquidity constraints – see Muellbauer and Lattimore (1995). These factors are not considered because I apply the Johansen (1988) procedure, taking account of subsequent extensions, to each country using 35 time-series observations. Since this method is based upon a vector autoregression (VAR), degrees of freedom become increasingly scarce with the proliferation of variables entered endogenously in the equilibrium relation. Thus, the efficiency of parameter estimates and the reliability of inference can be undermined.

One novelty of this investigation is the use of private disposable income over the estimation period 1960-1994 for all 20 OECD

[1] [1] I am only aware of reliable time series on broadly defined wealth being used in consumption functions for 4 OECD countries (Australia, Japan, the UK and the USA). Studies using financial wealth have been used for more countries see, for example, Sefton and In't Veld (1999) for applications to Canada, France, Germany, the UK and the USA.

countries. The few previous studies of OECD countries' consumer behaviour use national disposable income or GDP. Another novelty is the development of country-specific models, which are free from evident misspecification, to identify each economy's long-run consumption function. The country-specific models allow heterogeneous dynamics across countries, consider whether an intercept should be included in the cointegrating vector or not and, where appropriate, examine whether inflation should be omitted. Both the use of private disposable income and development of country-specific models should yield superior estimates and inference relative to previous studies A third novelty is the examination of whether the estimated income and inflation elasticities vary systematically across countries using cross-section regressions.

The next section discusses the theoretical underpinnings and recent empirical literature on the specified long-run consumption function. Section 3 tests for cointegration using the Johansen procedure. Discussion of the favoured long-run consumption functions is given in section 4. Section 5 conducts a cross-country analysis of the estimated parameters and section 6 draws conclusions.

2. Review of Recent International Empirical Studies of the DHSY Model

The empirical analysis is based upon the dynamic long-run solution of DHSY's model, relaxing the unit-income elasticity and ignoring the income growth term:

$$\ln C_t = b_1 + b_2 \ln Y_t + b_3 \Delta \ln P_t, \;\; 0 \leq b_2 \leq 1 \text{ and } b_3 < 0 \qquad (1)$$

where C_t, Y_t and $\Delta\ln P_t$, denote consumption, income and inflation respectively. The difference and natural logarithm operators are Δ and ln.

The use of current, rather than expected, income may be justified by the presence of liquidity constraints, income uncertainty and information constraints limiting expectation formation. Additionally, one could assume that expected income is proportional to current income following Ando and Modigliani (1963). Given data constraints and the importance of wealth, I employ inflation as a proxy for various asset effects in the long-run consumption function.[2] Hadjimatheou (1987) points out that studies generally find inflation to be negatively related to consumption, which is consistent with its use as a proxy for wealth effects. Any other direct inflation effects will also be captured.[3]

I am aware of only three recent analyses of this model for a number of OECD countries. Carruth *et al* (1996) estimate the dynamic DHSY model, which *implies* the equilibrium **(1)**, for a panel of the 15 European Union (EU) countries. They find implicit evidence favouring cointegration for 8 countries – the rejection of cointegration for 7 countries manifests itself in the imposition of the long-run unit-income elasticity. They also find inflation effects are statistically significant and negative for only 7 countries. Pesaran *et al* (1997) investigate **(1)** for 24 OECD countries using time-series regressions of the same dynamic autoregressive distributed lag model and find implicit evidence of cointegration for 20 countries. The estimated long-run income elasticity is significantly less than one in 9 countries, greater than unity in 3 and insignificantly different from one in 12. The long-run inflation coefficient is statistically significant and negative in only 10 countries. Estimating the model in a panel they reject the hypothesis of common long-run coefficients across countries. Larsson *et al* (1998) apply their panel cointegration test to **(1)** for 23 OECD countries. Time-series tests suggest 1 cointegrating vector for 17 countries, 2 cointegrating vectors for 4 countries and 3 cointegrating vectors for 2 economies. Their panel cointegration test indicates that the *largest* number of common cointegrating vectors across the panel is 2. They find that the overidentification restriction that consumption and income

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2 [2] For example, high inflation affects consumption by eroding the *real* value of money-fixed assets.

3 [3] See, for example, Deaton (1977) for a justification of (unanticipated) inflation effects.

constitute one cointegrating vector and inflation is a second, distinct, stationary vector cannot be rejected.

These recent investigations indicate, explicitly or implicitly, that there exist one or two cointegrating vectors between consumption, income and inflation for OECD countries and that the coefficients vary substantially across countries, suggesting the need to develop country-specific models. At present valid panel estimation methods that allow both the specification of short run dynamics and estimates of long-run elasticities to be different from country to country this means that are not available, thus the most flexible country-specific models will be secured through time-series estimation. I will explicitly test for cointegration and estimate long-run consumption functions using time-series regressions to allow for as much heterogeneity as possible. Both Carruth *et al*'s (1996) and Pesaran *et al*'s (1997) studies feature models that suffer from evident misspecification. I aim to choose country-specific models that are free from evident misspecification. All three studies use income measures that incorporate government income (GDP or national disposable income) and employ shorter time series than I use here. I seek to obtain superior parameter estimates by using income measures solely based upon the private sector and have data which facilitates the estimation of equation **(1)** for twenty OECD countries using 1960-1994 as the estimation period. To my knowledge, there is no previous study that estimates consumption functions for so many countries, using such a long time-series of data based solely upon the private sector. These models should help to clarify whether the long-run unit-income elasticity postulate is valid for the majority of OECD countries and indicate whether inflation constitutes a part of the long-run consumption function.

3. Cointegration Analysis

The empirical analysis uses annual data, available over the period 1955-1994 for 20 OECD countries,[4] on the natural logarithms of real (1990) per-capita total private consumers' expenditure, lnC_t, real (1990) private disposable income, lnY_t, and the log of the consumers implied price deflator, lnP_t, (1990~100).[5] All equations are estimated over the sample 1960-1994 (35 observations), allowing 5 observations for lags and transformations. Horioka (1996) applied ADF tests and the Johansen procedure to a Japanese consumption function with 3 variables using a *maximum* of 38 observations. My specification features almost identical degrees of freedom so should provide valid inference.

3.1 Integration Tests

ADF tests are used to assess whether the data are second-order stationary. The number of lagged dependent variables in each country's test equation is chosen to minimise the SBIC whilst ensuring non-autocorrelated residuals.[6] Referring to Tables 1 and 2 one observes some heterogeneity of inference across countries, however, the following generalisations can be drawn. The logarithms of consumption and income generally appear to be I(1). Prices are probably I(2), which is what I infer, but could be I(1). Some of the results deviate from the general inference stated above, which may be due to factors such as the low power of the ADF test, these anomalies may be explained in multivariate modelling. Nevertheless, this general inference, which allows consistency

[4] 4 The 20 OECD countries considered are: Australia (AUL), Austria (AUT), Belgium (BEL), Canada (CAN), Denmark (DEN), Finland (FIN), France (FRA), Germany (GER), Greece (GRE), Iceland (ICE), Ireland (IRE), Italy (ITA), Japan (JAP), Netherlands (NET), Norway (NOR), Spain (SPA), Sweden (SWE), Switzerland (SWZ), the UK the USA

[5] 5 A discussion of data definitions, construction, sources, coverage and transformations is provided in the data appendix.

[6] 6 The autocorrelation test statistic and SBIC are not reported to save space.

of model specification across countries, is the starting point for the multivariate cointegration analysis.

3.2 Cointegration Tests

I employ the standard Johansen procedure to test for cointegration. The general specification of the vector error correction model (VECM) is:

$$\Delta lnC_t = \gamma_{10} + \Sigma\gamma_{11i}\Delta lnC_{t-i} + \Sigma\gamma_{12i}\Delta lnY_{t-i} + \Sigma\gamma_{13i}\Delta\Delta lnP_{t-i} + \pi_{11}lnC_{t-1} + \pi_{12}lnY_{t-1} + \pi_{13}\Delta lnP_{t-1} + \Sigma b_{1j}D_{jt} + u_{1t}$$

$$\Delta lnY_t = \gamma_{20} + \Sigma\gamma_{21}\Delta lnC_{t-i} + \Sigma\gamma_{22}\Delta lnY_{t-i} + \Sigma\gamma_{23}\Delta\Delta lnP_{t-i} + \pi_{21}lnC_{t-1} + \pi_{22}lnY_{t-1} + \pi_{23}\Delta lnP_{t-1} + \Sigma b_{2j}D_{jt} + u_{2t} \qquad (2)$$

$$\Delta\Delta lnP_t = \gamma_{30} + \Sigma\gamma_{31}\Delta lnC_{t-i} + \Sigma\gamma_{32}\Delta lnY_{t-i} + \Sigma\gamma_{33}\Delta\Delta lnP_{t-i} + \pi_{31}lnC_{t-1} + \pi_{32}lnY_{t-1} + \pi_{33}\Delta lnP_{t-1} + \Sigma b_{3j}D_{jt} + u_{3t}$$

where i=1,...,L–1, π_{11}, π_{12}..., π_{33} are the reduced-form long-run coefficients, D_{jt} denotes the J contemporaneous exogenous variables (dummy variables in this case), and u_{ht} are the equations' error terms (h=1,2,3). The number of cointegrating vectors, r, is determined by the rank of the matrix containing reduced-form long-run coefficients (π_{11}, π_{12}..., π_{33}) using the standard maximum eigenvalue and trace test statistics.

Dummy variables may be included to remove evident misspecification (primarily departures from normality) which may arise due to many factors, including omitted variables. Omitted country specific events include, German reunification in 1990/1991, the dramatic slowdown in Japan's remarkable post-war growth in 1973/1974 and the financial deregulation that occurred in the UK and the Nordic countries during the 1980s. Further, inflation may not fully approximate asset effects for all countries. Further, outliers in non-consumption equations of the VECM could also cause misspecification. A parsimonious means of removing residual autocorrelation *and* departures from normality is desirable because the Johansen procedure is very sensitive to the independent normal errors assumption (see, Huang and Yang, 1996). Using dummy variables to remove misspecification rather than increasing the VECM's dimension is advocated by Clements and Mizon (1991).

The VECM used for testing cointegration for each country is determined by estimating (2) for L=1,2,3 and 4 and selecting the model with the lowest SBIC from those which are free from evident autocorrelation and non-normality (using both system and unreported individual equation tests). Table 3 summarises the model selection results for each country. A lag length of 3 is favoured for Australia, Denmark, Sweden and the UK.[7] For 7 of the 20 countries the favoured lag length is 1 (Austria, France, Greece, Ireland, the Netherlands, Norway and Spain). The remaining 9 countries favour a lag length of 2.

Inference from the Johansen test can be unreliable, especially when using small samples. This is due to its low power, the possibility of spurious cointegration (see Gonzalo and Lee, 1998 & Maddala and Kim, 1998), its sensitivity to how restricted the VECM is and the chosen VAR lag length (see Hall, 1991). Such potential biases can cause one to infer too much or too little cointegration. However, it is difficult to discern the overall impact of any such biases on inference. To the extent that such biases may exist in this analysis, I show pragmatism when interpreting the statistical results. Indeed, given the possible sensitivity of inference to specification and that, theoretically speaking, I have a strong prior belief that only one cointegrating vector exists, the aim is to see whether I can uncover statistical support for a unique cointegrating relation. Thus, I consider whether an evident unique cointegrating vector can be secured at the 1%, 5% *or* 10% levels of significance.

Table 4 reports the cointegration test results. If r=1 can be inferred by *either* the trace or maximum eigenvalue statistics then I will infer the presence of one cointegrating vector, as suggested by

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[7] [7] I favour L=3 for for Denmark and the UK because this is the only specification which yields a plausible, unique cointegrating vector. For Sweden this is the only specification that does not reject cointegration.

economic priors. Three cointegrating vectors will only be inferred if the tests for the null hypotheses of r=0, r=1 and r=2 are *all* rejected. This is because r=3 suggests all the variables are stationary, which is inconsistent with the order of integration results.

I only report cointegration results for the preferred model for each country.[8] For all 20 countries evidence of *at least* one cointegrating vector is found. For Sweden I had to search for a specification to secure cointegration. For 17 countries one can infer *exactly* 1 cointegrating vector.[9] The exceptions are Finland, Ireland and Spain.

4. Favoured Cointegrating Vectors

4.1 Selection Criteria

Economic as well as statistical criteria are employed to select favoured long-run consumption functions. The statistical criteria are hypothesis tests placed on the identified cointegrating relations and corresponding adjustment coefficients. Two potential forms of cointegrating vectors are nested within the general equation **(3)**. The first restricts the intercept into the equilibrium relation ($\beta_{r1} \neq 0$). The second, the unrestricted intercept specification, excludes the constant term from **(3)** ($\beta_{r1}=0$).

$$Z_{rt} = -\beta_{r0}\ln C_t + \beta_{r1} + \beta_{r2}\ln Y_t + \beta_{r3}\Delta\ln P_t. \tag{3}$$

where the subscript, r, denotes the first (r=1) or second (r=2) cointegrating vector and, Z_{rt}, the error correction term.

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[8] [8] The selection criteria are outlined in section 4.

[9] [9] From Table 4 the Canadian model appears to unambiguously suggest r=2. However, if one uses the trace statistic adjusted for degrees of freedom, which is 58.40 for the null of r=0 and 23.92 for the null of r=1, one cannot reject r=1 at the 1% level. Although Doornik and Hendry (1995, p. 222) note that it is not clear whether this is the preferred small sample correction this result is utilised to provide statistical support for the strong prior belief of a single cointegrating vector. Further, the overidentification restrictions (applied assuming r=2) are rejected (the test statistic is 6.158).

The favoured model for each country is selected using four criteria. The first two are based upon zero restrictions on the parameters of the cointegrating vector(s) and are applied using the standard likelihood ratio (LR) statistic. When r=1 the single hypothesis is, $\beta_{1k}=0$, and when r=2 one must test the significance of a single variable on *both* cointegrating vectors using the joint hypothesis, $\beta_{1k}=\beta_{2k}=0$.[10] The first criterion is that $\ln C_t$ and $\ln Y_t$ must be statistically significant, because I am interested in uncovering a consumption function and income is postulated as the main explanatory factor. The second criterion is that *when* inflation is significant it should have a negative coefficient to proxy wealth effects. The third criterion is that the long-run average propensity to consume (APC) should be less than one to reflect the persistence of positive aggregate saving observed for OECD countries and because consumers cannot spend more than they earn over the long term. This implies that there is a below unit-income elasticity or, if the income elasticity is not significantly different from one, there should be a statistically significant and negative intercept or inflation term. To test for a unit-income elasticity, when r=1, the restriction, $\beta_{10}+\beta_{12}= 0$ is applied.[11] The fourth criterion is that the adjustment coefficient(s) in the consumption growth equation must be positive and statistically significant. That is, for valid error correction behaviour the coefficient on $\ln C_t$ must be negative. Since consumption is normalised upon by setting $-\beta_{r0}=-1$ in **(3)** and given the coefficient on $\ln C_t$ is, $\pi_{r1} = (\alpha_{r1})(-\beta_{r0})$, this implies that the adjustment coefficient, α_{r1}, must be *positive*. That is, $\alpha_{11}>0$ (for r=1)

[10] [10] When r=2 one tests whether the variable is jointly significant in both cointegrating vectors. From tests on the first cointegrating vector one can determine the statistical significance of a variable in that first vector, however, one cannot always deduce whether such a variable is significant in the second.

[11] [11] This test is not conducted when r=2.

in the restricted VECM, equation **(4)**.[12] This statistical significance of the adjustment coefficient is tested with the restriction, $\alpha_{11}=0$.[13]

$$\Delta lnC_t = \delta_{10} + \Sigma\delta_{11i}\Delta lnC_{t-i} + \Sigma\delta_{12i}\Delta lnY_{t-i} + \Sigma\delta_{13i}\Delta\Delta lnP_{t-i} + \Sigma\phi_{1j}D_{1jt} + \alpha_{11}Z_{1t-1} + \alpha_{21}Z_{2t-1} + u'_{1t}$$

$$\Delta lnY_t = \delta_{20} + \Sigma\delta_{21i}\Delta lnC_{t-i} + \Sigma\delta_{22i}\Delta lnY_{t-i} + \Sigma\delta_{23i}\Delta\Delta lnP_{t-i} + \Sigma\phi_{2j}D_{2jt} + \alpha_{12}Z_{1t-1} + \alpha_{22}Z_{2t-1} + u'_{2t} \quad (4)$$

$$\Delta\Delta lnP_t = \delta_{30} + \Sigma\delta_{31i}\Delta lnC_{t-i} + \Sigma\delta_{32i}\Delta lnY_{t-i} + \Sigma\delta_{33i}\Delta\Delta lnP_{t-i} + \Sigma\phi_{3j}D_{3jt} + \alpha_{13}Z_{1t-1} + \alpha_{23}Z_{3t-1} + u'_{3t}$$

4.2 Consistency of Long–run Consumption Functions with Selection Criteria when r=1

For all countries, except Ireland, Japan and Spain, the favoured model, selected using the four criteria outlined above, has statistical support as a unique cointegrating relation. The consistency of these favoured models with the specified criteria is discussed with reference to Table 5.

Four countries' (Austria, Canada, Greece and the UK) favoured models satisfy all four specified criteria. Therefore, they provide good approximations to these countries' long-run consumption functions.

For six countries (France, Germany, Iceland, the Netherlands, Norway, and the UK) only one of the desirable conditions is not satisfied by the favoured long-run consumption functions. The criteria *not* satisfied are as follows. In the case of France the income term is statistically insignificant while both consumption and income are insignificant for the Netherlands. For Germany and Norway the unit-income elasticity hypothesis cannot be rejected and both the intercept and inflation terms are statistically insignificant,

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[12] [12] When the intercept is restricted into Z_{rt}, $\delta_{10}=\delta_{20}=\delta_{30}=0$ and when r=1, $\alpha_{21}=\alpha_{22}=\alpha_{23}=0$.

[13] [13] When r=2 the joint hypothesis, $\alpha_{11}=\alpha_{21}=0$ is tested.

suggesting that the APC is unity in the long-run.[14] The adjustment coefficient in the consumption equation is statistically insignificant for Iceland and the UK (if positive for both countries). Although not satisfying all the specified criteria these six countries' cointegrating vectors are plausible in many senses, and are presented as reasonable approximations of their countries' long-run consumption functions.

For four countries (Australia, Belgium, Finland and Italy) two of these *plausibility* criteria are not met. In the case of Australia both consumption and income terms are statistically insignificant and there is no significant and negative intercept or inflation term to compensate for the evidence against the presence of a below unit-income elasticity. However, the adjustment coefficient in the consumption equation is positive and statistically significant and the cointegrating vector's estimated parameters are plausible, if not well determined. The favoured cointegrating vectors for Belgium and Italy are comprised of statistically insignificant coefficients (including consumption and income). Further, although the coefficient on income is less than one for both countries, it is not significantly less than one, implying a unit long-run APC because both intercept and inflation terms are statistically insignificant. However, the adjustment coefficient in the consumption equation is positive and statistically significant and the estimated coefficients are theoretically plausible for both countries, if the income elasticity is quite low for Italy (being 0.569).[15] For Finland the adjustment coefficient in the consumption equation is statistically insignificant, if positive.[16] Although consumption and income are both statistically significant in the cointegrating vector, there is evidence that the

[14] [14] The coefficients on income for Germany and Norway are both below one, if not statistically different from unity, so *may* be considered completely plausible.

[15] [15] This low income-elasticity is consistent with Italy's historically low APC (see Guiso *et al* 1991), but may also be due to this parameter's poor determination.

income elasticity is significantly greater than one.[17] The cointegrating vectors for Australia, Belgium, Finland and Italy are presented as usefully plausible because they exhibit many desirable features for credible long-run consumption functions and their departures from the specified criteria do not seem too severe.

The favoured cointegrating vector for Denmark fails to satisfy three of the desired criteria. There is evidence of an above unit-income elasticity, the adjustment coefficient is statistically insignificant and the coefficient on inflation has the *incorrect* positive sign. However, all the estimated coefficients in the cointegrating vector are statistically significant and the adjustment coefficient is positive. Therefore, this vector represents an approximate long-run consumption function with some desirable features.

For Sweden and Switzerland, the most plausible cointegrating vector fails to satisfy many of the specified criteria including all variables being statistically insignificant with rather large estimated income elasticities. Further, and somewhat crucially, the adjustment coefficients are negative, which is inconsistent with continually forcing consumption towards its equilibrium, suggesting that these vectors provide poor approximations to credible long-run consumption functions.[18]

16 16 There is *support* for zero or three cointegrating vectors for Finland's favoured model (see Table 4). Thus, the statistics seem completely unhelpful regarding the choice of r so the prior of r=1 is imposed. Pesaran and Pesaran (1997, p. 297) similarly impose theoretical priors when the Johansen procedure is uninformative on the choice of r.

17 17 The estimated income elasticity being greater than unity for Finland may be due to the omission of explanatory factors capturing the effects of financial deregulation.

18 18 For Switzerland the possibility that r=2 was also investigated, given the results reported in Table 4, however, a plausible cointegrating vector could not be found.

4.3 Consistency of Long–run Consumption Functions with Selection Criteria when r=2

When there is no statistical *support* for a *unique* cointegrating vector the possibility that r=2 is examined. In this case one must apply at least 4 overidentification restrictions, with two on each vector - see Pesaran and Shin (1994). Within the context of the DHSY model, Larsson *et al* (1998) suggest that consumption and income may constitute one cointegrating vector and that inflation, on its own, forms a second. This involves imposing two normalisation restrictions and three exclusion restrictions, $\beta_{13}=0$, $\beta_{21}=0$ and $\beta_{22}=0$, on **(3)**.[19] This produces an overidentified long-run matrix which, following Pesaran and Shin (1994), can be tested with an LR statistic that has a χ^2 distribution with, in this case, one degree of freedom. If these overidentification retrictions cannot be rejected *and* the estimated parameters on the long-run consumption equation are plausible, the overidentified vector will represent the favoured specification.

For Ireland, Japan and Spain there was no statistical support for a *plausible* unique cointegrating relation, however, there was evidence for two cointegrating vectors. (There was evidence that r=1 for Japan, however, the adjustment coefficient in the favoured model is negative). The results of the overidentification restrictions are reported in Table 6. The overidentification restrictions are rejected for Japan and Spain but not Ireland. The first cointegrating vector for Ireland is plausible as a long-run consumption function in the sense that the adjustment coefficient is positive and the income elasticity is very close to unity (1.010) with a negative intercept (allowing the long-run APC to be below one).[20] This overiden-

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[19] 19 It is not obvious that any other form of economically sensible overidentification restrictions exist, so no other form is considered.

[20] 20 The statistical significance of the parameters is not tested because they involve joint (overidentification) restrictions, so do not specifically refer to the hypothesis of interest.

tified consumption vector therefore represents the favoured long-run consumption function for Ireland.

The economic prior of a unique cointegrating vector is imposed for Spain because of the rejection of the overidentification restrictions and the possibility of spurious cointegration when testing for r>1 (see Maddala and Kim 1998, pp. 173 and 220). The favoured Spanish vector, reported in Table 5, provides a reasonable approximation to a long-run consumption function because only one of the four desirable criteria, outlined above, is not satisfied, being evidence of an above unit-income elasticity.

For Japan the overidentification restriction is rejected so the overidentified consumption function is not favoured. However, I do not assume r=1 because the first cointegrating vector reported in Table 5 features a significant and negative adjustment coefficient. Since the second vector in Table 5 satisfies all four of the specified criteria for a plausible long-run consumption function it represents the favoured model for Japan.

4.4 General Characteristics of Long–run Consumption Functions

The favoured long-run consumption functions of 6 countries (Austria, Canada, France, Greece, Japan and the Netherlands) exhibit a below unit-income elasticity, in the sense that the unit-income elasticity hypothesis is rejected and the coefficient on income is less than one. The unit-income elasticity cannot be rejected for 9 countries (Australia, Belgium, Germany, Iceland, Ireland, Italy, Norway, Switzerland and the USA).[21] An above unit-income elasticity is inferred for 5 countries (Denmark, Finland, Spain, Sweden and the UK).[22] The rejection of the unit-income elasticity postulate for 11 of

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[21] 21 This homogeneity postulate is not tested for Ireland because the favoured consumption function is an overidentified vector. However, the estimated income elasticity (1.010) is so close to unity I believe it is safe to assume a unit-income-elasticity.

[22] 22 The evidence of an above unit long-run income elasticity may reflect the omission of explanatory factors such as wealth and credit.

the 20 countries suggests that one should not automatically assume consumption is homogeneous of degree one in income for any particular country. This is consistent with the findings of Carruth *et al* (1996) and Pesaran *et al* (1997).

For only 7 countries (Canada, France, Greece, Japan, Spain the UK and the USA) is inflation negative and significant in the long-run consumption function. This is consistent with Carruth *et al* (1996) and Pesaran *et al* (1997), both of whom find that inflation is negative and significant in less than half of the countries' consumption functions that they analyse. Thus, inflation does not appear to be a fundamental determinant of many OECD countries' consumer behaviour.

For cointegration to imply that an equilibrium consumption function has been found the adjustment coefficient in the consumption growth equation should be positive and statistically significant. This condition is satisfied for all countries except Sweden and Switzerland, suggesting that a valid long-run consumption function has not been uncovered for these 2 countries.

5. Explaining Cross–Country Differences In Consumer Behaviour

This section employs cross-country regressions to explain the variation in the estimated long-run elasticity of consumption with respect to income and inflation.[23] I am not aware of any previous attempt to do this. These estimated coefficients vary considerably across countries. Figure 1 plots the estimated long-run income elasticity, which ranges in value from 0.569 for Italy to 1.464 for Denmark relative to an average value of 1.014 (the standard deviation is 0.205). Figure 2 plots the estimated long-run inflation elasticity, with values ranging from -3.645 for Italy to 1.926 for Denmark relative to an average value of -0.394 (the standard devi-

[23] [23] This analysis does raise questions about the nature of an OECD-wide data generation process. I am grateful to an anonymous referee for drawing this issue to my attention.

ation is 1.135). The Italian value is extremely low and is regarded as an outlier.

5.1 Explaining Cross–Country Differences in the Income Elasticity

I am not aware of any theories that directly rationalise variations in estimated income elasticities. The potential explanatory factors considered here are based upon reasons why different responses of consumption to income may occur under the assumption that such factors will also be relevant for explaining the variation in income elasticities.

Modigliani's (1986) LCH and Brown's (1952) Habit Persistence version of the Relative Income Hypothesis (RIH) suggest a negative relationship between an economy's APC and its income growth (denoted GRTH). The LCH also suggests that the length of retirement (LRET) is positively (negatively) related to the saving rate (APC) and that the proportion of dependents in the population (DEP) is negatively (positively) related to the saving rate (APC). Miles and Patel (1996) suggest a parsimonious way of capturing the demographic effects of the LCH. They argue that the support ratio (SUPT), the number of working age to the number of pensionable age, is positively (negatively) related to the saving rate (APC). They also suggest that, due to the needs of children, only the proportion of the population aged 50 to 64 (RSAV) accrue substantial saving for retirement, implying a negative relationship between pre–retirement savers and the APC. Modigliani (1990) extends the LCH specification to consider Ricardian equivalence. In the present context, some degree of Ricardian offset suggests a positive association between the income elasticity parameter and the fiscal surplus / deficit to GDP ratio (GDEF). Jappelli and Pagano (1994) extended Modigliani's (1990) model to include liquidity constraints with the implication that the availability of credit (CRED) is positively related to the APC.

Keynes (1936) has been attributed with the suggestion that the marginal propensity to consume (MPC) falls as the level of income rises. However, a linear relationship would imply that a continual rise in the *level* of per-capita income would cause an unbounded fall in the APC, eventually making it negative, which is implausible. Therefore, various nonlinear relationships are considered [the natural logarithm of income (lnINC) is favoured as a regressor], allowing consumption out of income to decrease at a decreasing rate as income rises – I am not aware of previous attempts to investigate such a nonlinear relation. In contrast, Modigliani's (1986) LCH implies that a country's APC is independent of its income level.

A negative relationship between the real interest rate (r) and APC arises due to intertemporal substitution. However, with an offsetting income effect, the overall impact is ambiguous, possibly being positive or yielding a small unstable relationship – see Muellbauer (1994). Deaton (1992) suggests that increased income uncertainty (UNCT) will generate greater precautionary savings implying a potential negative relationship between UNCT and the income elasticity. While Duesenberry's (1949) RIH suggests that the degree of income inequality (INEQ) *within* a country will be negatively associated with the proportion of income consumed.

The discussion above suggests the following general eclectic model for the estimated income elasticity, β_Y.[24] Expected signs of coefficients are given beneath the variables.

β_Y = f(GRTH, LRET, DEP, RSAV, SUPT, GDEF, CRED, lnINC, r, UNCT, INEQ) (5)

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24 24 Detail on variables employed in the *favoured* models is provided in the data appendix. Detail on further variables considered but not reported is given in Stewart C (1999) *Modelling and Comparing OECD Countries' Consumer Behaviour*, PhD thesis, London Guildhall University.

All regressions use 20 observations except those including income inequality, which use 13 observations (due to data constraints on this variable). The general-to-specific methodology is employed to search for parsimonious forms of (5). Table 7 reports the OLS coefficient estimates, with White's t-ratios in parentheses, for 5 models nested within (5). All reported models exhibit statistically significant explanatory power and are free from evident misspecification at the 5% level, except equation **5c** which features significant nonlinearity at the 5% (but not 1%) level. Thus, the inference from these models is legitimate.

The reported models contain various combinations of the five main explanatory factors: GRTH, GDEF, CRED, lnINC and INEQ. Equation **5a** includes GRTH, GDEF, CRED and lnINC whose coefficients exhibit the expected sign and are statistically significant, except CRED, which is insignificant. Excluding CRED from **5a** yields equation **5b** and causes the adjusted R^2 to drop marginally from 0.606 to 0.587. All remaining variables are statistically significant and *correctly* signed. These two regressions indicate that income growth is negatively associated with the income elasticity and GDEF is positively related to it. Further, the level of income exhibits a negative nonlinear correlation with the income elasticity. Exclusion of lnINC from **5b**, yielding equation **5c**, causes a large fall in the adjusted R^2, from 0.587 to 0.517, and induces evident nonlinearity, suggesting that this is an important explanatory factor and should not be excluded.

Equation **5d** includes income inequality, which is negative if insignificant. GRTH, CRED and lnINC are also statistically significant and *correctly* signed, while GDEF is highly insignificant. Excluding GDEF gives equation **5e**. All retained variables, including income equality, are statistically significant. This model confirms the inferences drawn from the previous regressions regarding GRTH and lnINC whilst suggesting an additional role for income inequality.

However, unlike previous regressions it indicates that CRED is an important explanatory factor, and that there is no role for GDEF.

Overall, the results suggest that income growth negatively determines the income elasticity, consistent with Modigliani's LCH and Brown's version of the RIH. The log of per-capita income has a nonlinear negative influence on the income elasticity such that the elasticity decreases at a decreasing rate as the level of income rises, which does not necessitate that the elasticity eventually becomes negative. This supports a suggestion often attributed to Keynes (1936), if it contradicts an implication of Modigliani's LCH. There is also some evidence indicating that increased income inequality reduces the income elasticity, which is consistent with Duesenberry's RIH. The fiscal surplus/deficit exerts a positive and statistically significant influence on the income elasticity for some models suggesting some evidence of a Ricardian offset, consistent with the majority of empirical work. There is some tentative evidence that the amount of credit available to the private sector has a positive impact upon the long-run income elasticity, which is consistent with Jappelli and Pagano (1994).

5.2 Explaining Cross–Country Differences in the Inflation Elasticity

Since inflation is primarily used to approximate wealth effects I consider whether the variation in the inflation elasticity is related to factors that affect the MPC out of assets, assuming an inverse relation between inflation and asset effects.

Within the context of the LCH, Muellbauer and Lattimore (1995) suggest that the MPC out of assets increase with age. This implies a negative (positive) relationship between the proportion of the population who are young and economically active (YNG) and wealth (the inflation elasticity) and a positive (negative) correlation between the retired proportion of the population (RET) and the elasticity out of assets (inflation).[25] In early middle age the household with dependents reduces savings (borrows) suggesting a positive (negative) relationship between the dependency ratio (DEP) and the MPC out of wealth (inflation). In later middle age, once dependents have left home, the household will save for its retirement, suggesting that the proportion of the population comprised of pre–retirement savers (RSAV) is negatively (positively) related to wealth (the inflation elasticity).

Additional potential explanatory factors include the following. The LCH suggests that the expected length of retirement (LRET) is negatively (positively) related to the wealth (inflation) elasticity of consumption. The precautionary saving motive suggests that income uncertainty (UNCT) is negatively (positively) correlated with expenditure out of assets (inflation). Since less binding credit constraints suggests greater fungibility of wealth the availability of credit (CRED) may be positively (negatively) associated with the MPC out of wealth (inflation). The general model for the estimated *inflation* elasticity, β_I, is:

$$\beta_I = f(\underset{-}{\text{CRED}}, \underset{+}{\text{UNCT}}, \underset{+}{\text{LRET}}, \underset{-(+)}{\text{DEP}}, \underset{+}{\text{YNG}}, \underset{+}{\text{RSAV}}, \underset{-}{\text{RET}}) \quad \textbf{(6)}$$

Table 8 presents the only satisfactory model that could be secured. The outlying Italian observation (see Figure 2) is excluded from the regressions because it causes severe non–normality, restricting the sample to 19 observations. There is no evident misspecification according to the reported diagnostics suggesting inference is valid. UNCT exhibits a positive and statistically significant impact upon the inflation elasticity while CRED and DEP feature negative and significant correlations. The model provides significant explanatory power with a 55.1% fit. The estimated coefficients' signs are consistent with cross–country variations expected if in–

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25 25 The presence of a bequest motive may reduce or eliminate this effect.

flation were approximating wealth effects in the long-run consumption function.

6. Conclusions

The Johansen procedure has been employed to test whether the logs of consumption and disposable income and inflation cointegrate for 20 OECD countries. The use of disposable income and the heterogeneity of model specification across countries should provide superior inference relative to previous studies of OECD countries' consumer behaviour. Statistical evidence *supports* cointegration for all countries, however, for only 18 countries do the favoured cointegrating vectors represent plausible long-run consumption functions – the exceptions are Sweden and Switzerland.

The estimated elasticities of the favoured models are heterogeneous across countries. There is evidence of a below unit-income elasticity for 6 countries, a unit-income elasticity for 9 countries and an above unit-income elasticity for 5 countries. The above unit long-run income elasticity possibly reflects omitted variable bias. The impact of omitted variables, the poor determination of some countries' income elasticities and the evidence of a below unit-income elasticity for 6 countries suggests that one should not automatically assume that consumption is homogenous of degree one in income for any particular OECD country. Inflation is statistically significant and negative for only 7 countries, suggesting that inflation is not a fundamental explanatory factor of consumption for all countries.

The long-run consumption elasticities with respect to income and inflation have been modelled using cross-section regressions. I am not aware of any previous attempt to model the variation in consumption elasticities.

The long-run income elasticity is negatively correlated with income growth and features a plausible nonlinear negative relationship with the log of per-capita income. The latter finding represents an innovation of the current study and supports a proposition often attributed to Keynes. The implication of these two correlations is that policies that raise development will reduce the proportion of income consumed and so raise savings. There is some evidence that the fiscal surplus/deficit is positively associated with the income elasticity suggesting some degree of Ricardian offset without eliminating the possibility that fiscal policy can influence consumption. There is also some evidence that holdings of private sector credit positively influences the income-elasticity, suggesting that policies increasing financial liberalisation and integration can raise consumption for a given income level. There is also some evidence that increased income inequality reduces the income elasticity. Thus, policies that redistribute income, such as taxation policy, may affect consumer demand.

The availability of credit and the dependency ratio are found to have negative (implicitly positive) impacts upon the long-run inflation (wealth) elasticity while income uncertainty exhibits a positive (negative) association. These results are consistent with inflation approximating wealth effects (through a negative correlation) in the long-run consumption function.

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Data Appendix

The 20 OECD countries considered are: Australia (AUL), Austria (AUT), Belgium (BEL), Canada (CAN), Denmark (DEN), Finland (FIN), France (FRA), Germany (GER), Greece (GRE), Iceland (ICE), Ireland (IRE), Italy (ITA), Japan (JAP), Netherlands (NET), Norway (NOR), Spain (SPA), Sweden (SWE), Switzerland (SWZ), the UK the USA.[26] The data are annual observations (1955-1994) on the following series:

[26] [26] The Federal Republic of Germany and the German Democratic Republic were united monetarily, economically and socially on July 1, 1990 and unified on October 3, 1990. Data refers to West-Germany over the period 1955-1990 (inclusive) and to *unified* Germany from 1991-1994.

AC_t Aggregate total private consumers' expenditure in current (AC_t) and 1990 (ARC_t) prices. The primary source is OECD National Accounts, volume 1 (Main Aggregates) with other observations obtained from UN National Accounts, OECD Quarterly National Accounts 96/1, and OECD Economic Outlook 6/96.

AY_t Aggregate private disposable income in current prices. The main source is UN National Accounts with other data derived/obtained from OECD National Accounts, OECD Economic Outlook, Ireland's National Accounts 1975-1981 and National Economic Institute: *Historical Statistics*, Reykjavik, September 1995 (very kindly provided by Thorarinn Petursson of the Bank of Iceland). Kari H Eika (Bank of Norway) kindly provided a time-series for the Norwegian real (1970) private disposable income series, over the period 1962-1978. GDP was spliced to the disposable income series for Iceland and Norway to obtain data prior to 1962.

AG_t Current (AG_t) and 1990 (ARG_t) price GDP was taken from OECD National Accounts, OECD Quarterly National Accounts, and OECD Economic Outlook.

POP_t Population (millions of people) was obtained from International Monetary Fund International Financial Statistics (IMFIFS) line 99z. Most of the data, to 1990, was obtained using Manchester's Data Archive. For Belgium the 1990-1994 observations were obtained from Eurostatistics, July 1996.

Data transformations employed are:
$P_t=[(C_t/RC_t)x100]$; $C_t=[(ARC_t)/(POP_t x1000000)]$;
$Y_t=\{[AY_t/(P_t/100)]/[POP_t x1000000]\}$; $G_t=[(ARG_t)/(POP_t x1000000)]$.

The variables used to explain the cross-country variations in the estimated parameters are averages, over the period 1960-1994 (unless otherwise stated), of the variables listed below. To save space only details on the variables used in the favoured models is reported.

$CRED_t$ is proxied by the private sector domestic credit to GDP ratio (**$PSDC_t/AG_t$**). Private sector domestic credit, **$PSDC_t$**, is obtained from line 32d of IMFIFS.

$GDEF_t$ $= D_t/AG_t$. The Central Government fiscal surplus/deficit, **D_t**, is obtained from line 80 of IMFIFS.

$lnINC_t$ is the natural logarithm of per-capita income in a common currency, $lnINC_t$. The measure of income, **INC_t**, is real (1990) per-capita GDP in thousands of Geary-Khamis US dollars. This is available for all countries, except Iceland, from Maddison (1995). For Iceland I use real GDP in Kronur (from UN/OECD National Accounts) and convert it into dollars by dividing by the Kronur/dollar exchange rate (line rf from IMFIFS) and apply an adjustment to compensate, to some degree, for purchasing power parity.

$GRTH_t$ $= lnY_t - lnY_{t-1}$.

$INEQ_t$ Gini coefficient of income inequality. This is reported in Atkinson (1996, p. 21). It is only available for 13 countries: Belgium, Finland, France, Germany, Ireland, Italy, the Netherlands, Norway, Spain, Sweden, Switzerland, the UK and the USA.

$UNCT_t$ is income uncertainty proxied with the absolute deviation of income growth, defined as: $|\Delta lnY - \text{"trend"}|$, where "trend" is, following Muellbauer and Lattimore (1995), an MA(5) of income growth. The unemployment rate was tried but was not retained in the reported models.

DEP_t The dependency ratio. The proportion of the *total* population aged between 0 and 14 years. This variable is based upon the age distribution of the population (of both sexes) available in *The Sex and Age Distribution of the World Populations the 1994 Revision* UN 1994.