AN AGGREGATE CONSUMPTION FUNCTION FOR CANADA: A COINTEGRATION APPROACH

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February 1993

ACKNOWLEGDEMENTS: This paper has benefitted from comments provided by Professor Yatchew of the University of Toronto. We also thank Professor Sawyer of the Institute for Policy Analysis (IPA) at the University of Toronto for initial discussions and for providing the data used in the IPA's FOCUS model. However, whatever errors remain are attributable only to the authors.

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ABSTRACT

This study utilises cointegration theory and error correction models (*ECMs*) to specify a candidate dynamic formulation of the aggregate consumption function using quarterly Canadian data. Following Engle and Granger (1987), the variables entering the long-run equilibrium relationship are tested for unit roots using a variety of techniques such as the Dickey-Fuller, the Augmented Dickey-Fuller and the Phillips-Perron test procedures. Johansen's maximum likelihood approach is also used to determine the rank of the cointegrating matrix. The error correction model, which appears to be a tentatively adequate conditional characterisation of the data generating process, reveals that disposable income, wealth, government expenditures, relative prices and liquidity constraints (proxied by the unemployment rate) are important variables. Non-nested tests are carried out against a recent dynamic specification found in Sawyer (1992). Encompassing tests show that our model can encompass the Sawyer model.

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INTRODUCTION:

The aggregate consumption function has been a key component of macro-models since the days of Keynes (Keynes,1936). However, the clear theoretical and empirical inadequacies of the simple Keynesian Absolute Income Hypothesis have led economists to develop new hypotheses about consumer behaviour. Doubtlessly, the most influential of these theories have been the life cycle (*LCH*) and permanent income (*PIH*) hypotheses (Modigliani and Ando,1957; Friedman, 1957). These theories, explicitly based in the microeconomics of rational utility maximising behaviour, have enjoyed widespread appeal. In a seminal paper, Hall (1978) incorporated rational expectations into the permanent income framework (*REPIH*) to derive striking empirical predictions: under a stringent set of assumptions, Hall showed that consumption follows a random walk. Since then, much attention has focussed on testing the Euler Equation form of the *REPIH-LCH* model derived from first principles of economic theory and testing for the implied restrictions. However, this line of research has had mixed results with the majority of studies rejecting the strict over-identifying restrictions imposed by the theoretical model (e.g, Bean, 1986). As a consequence, the precise empirical formulation of the aggregate consumption function cannot be considered settled (Blinder and Deaton, 1985).

An alternative, pragmatic framework for studying economic relationships has been recently developed, largely by British econometricians. This approach, the cointegration-error correction model methodology, while still supporting the need for a theoretical basi for the variables entering the consumption function, loosens the bonds of rigid adherence t theory and permits the data to tell a substantial part of the story. Consequently, these model are not usually constructed from explicit reference to underlying micro behaviour an typically do not involve expectational mechanisms, but concentrate on specifying long-ru 'equilibrium' properties while allowing a rich enough dynamic specificaton to explain shor run activity (Davidson *et al*, 1978). However, this does not mean that a cointegration-erro correction model is inconsistent with microeconomic theory: as Muellbauer and Bover (198 demonstrate, an error correction model (*ECM*) can be derived from a multi-period utilit maximising model with some agents facing liquidity constraints.

Cointegration theory focusses attention on the temporal properties of economic tir series. Essentially, cointegration theory asserts that despite the empirical fact that mo economic variables exhibit non-stationarities (thus rendering much of classical inferen invalid), if there exists a linear combination of these non-stationary series that is stationa (cointegrated), then valid inference is still possible and the cointegrating variables can regarded as defining a 'long-run equilibrium subspace' by which it is simply meant that variables comprising the cointegration vector 'move together closely' in the long-run. I set of variables are indeed related in a long-run or equilibrium sense as postulated by theo then they should be cointegrated and a levels equation is valid. Consequently, cointegrat tests offer an initial check for the data consistency of (long-run) theoretical relationship:

This paper estimates an aggregate consumption function for Canada, usin cointegration approach. Section II of the paper motivates the variables to be included in empirical formulation. Section III describes the econometric methodology while Section IV analyses the data and empirical results. The model found in Sawyer (1992) is also re-estimated, and subjected to a wider range of tests than reported by Sawyer; non-nested model tests are used to evaluate the Sawyer formulation against our alternative specification. Lastly, conclusions and thoughts about future work are presented in Section V.

II. The Empirical Consumption Function

Sawyer (1992) proposes an empirical consumption function for Canada. His function allows constant elasticities of consumption with respect to income, wealth and prices while permitting the more realistic assumption of changing elasticities with respect to interest rates. His results would seem to indicate that (a subset of) these variables form a cointegrated vector and that the resulting *ECM* is an adequate initialisation point for the study of the data generating process underlying aggregate consumption behaviour. In this study, we make use of the Sawyer equation as the point of departure, augmenting it in the empirical tests by the addition of theoretically plausible variables and by investigating whether the results of cointegration are sensitive to such additions and to some alternative variable specifications.

The Sawyer equation is given by

$$C_t = Y_t^{\alpha_1} P_t^{\alpha_2} W_t^{\alpha_3} S_t^{\alpha_4} \exp^{\alpha_0 + \alpha_5 R_t + \alpha_6 T + a_t} \qquad (1)$$

where C_t is real consumer expenditures, Y_t is real disposable income, P_t is a relative price variable, W_t is a wealth construct, S_t is the ratio of labour income to total personal income, and R_t is a real interest rate variable. a_t is an error term with assumed classical properties and T is a time trend. Upper case letters represent untransformed levels of variables.

Taking logarithms of both sides of equation (1) yields

$$c_t = \alpha_0 + \alpha_1 y_t + \alpha_2 p_t + \alpha_3 w_t + \alpha_4 s_t + \alpha_5 R_t + \alpha_6 T + a_t \qquad (2)$$

where lower case letters represent natural logarithms. The regressor set in equations (1) and (2) is rather standard. A positive relationship between c_t and y_t has been postulated from earliest times; the *LCH* would suggest the inclusion of a labour income variable such as *s* and a wealth variable; as Molana (1991) shows, the *LCH* can be modified to specifically support an error correction relationship between consumption and wealth. If real interest rates or relative prices are not assumed *ex ante* to be constant as in Hall (1978), then they too should enter the aggregate consumption function (see Blinder and Deaton, 1985).

An Alternative Specification:

Following Blinder and Deaton (1985), an 'agnostic' approach to the specification of an alternative model is adopted. The usual regressor set is expanded by considering additional variables normally omitted from consideration such as government expenditure and the unemployment rate which may act as a proxy for liquidity constraints. The unemployment rate can conceivably enter on its own merit. Since it is itself a measure of constraints in the labour market, it could have an independent effect on savings and consumption. Prospects of future unemployment may rationally cause agents to reduce current consumption. Flavin (1985), Cuddington (1982), Carroll and Summers (1987) have all argued for the inclusion of the unemployment rate as a proxy for liquidity constraints in an aggregate consumption/savings function. While recognising that a study of liquidity constraints may well be best done using disaggregated data on individuals or households as in Hall and Miskin (1982), we nevertheless incorporate the unemployment rate as an additional regressor to proxy for such liquidity or employment constraints.[1]

Regarding the inclusion of the government spending variable, as Aschauer (1985) and Bean (1986) show, if government expenditures substitute for private expenditures in utility, then this variable rightly enters as an argument in the consumption function.[2] Other candidate variables included in the cointegration tests are income, wealth, interest rates, inflation and relative prices.

This specification is employed as an alternative model to that proposed by Sawyer (1992). To ascertain whether the conditional representation of the *DGP* developed herein is robust, a wide range of model diagnostic criteria is utilised. One further check of our model is to test it against a rival data coherent model such as the Sawyer model. This is done in

[1] Sawyer (1992) only allows the unemployment rate to enter the *ECM* stage of his model, assuming that liquidity constraints only influence consumption in the short-run. This would suggest that in the long-run, market imperfections, rationing etc. are eliminated and liquidity constraints vanish. However, if one interprets 'long-run' in a 'cointegration sense', an argument can be made for allowing liquidity constraints to directly enter the levels equation. All that is required is that the relationship between consumption and the liquidity variable has been in existence for a long time. There is, in our opinion, no reason why liquidity or unemployment constraints are not pervasive enough (Hayashi, 1985) or necessarily vanishing.

[2] Substitutability would imply a negative coefficient. However, as government spending may provide productive services, the net effect may be one of complementarity.

Section IV, using non-nested testing procedures.

III. The Econometric Methodology

Much of classical inference is predicated on the assumption that economic time seri are covariance stationary; i.e. the series have finite second moments, and the mean a covariance structure of the data do not change across observations. However, as is w known, most economic time series of interest exhibit stochastic non-stationarities, th undermining the foundations of traditional inference. Regressing one trended variable another could lead to the erroneous inference of a significant relationship where none exis a problem Granger and Newbold (1974) refer to as the 'spurious' regression problem.[3]

Following Granger (1981), define d as the order of integrability of a time se where d represents the number of times the series must be differenced to induce a station *ARMA* representation. Cointegration can be defined formally as follows: all compo variables of a vector x_t are said to be cointegrated of order d,b i.e. $(x_t \sim CI(d,b)$ (i) all components of $x_t \sim I(d)$ and (ii) there exists a vector $\alpha \neq 0$ so $z_t = \alpha' x_t \sim I(d-b)$, $b > 0$; the vector α is then called the cointegrating vector (E

[3] The early work on the consumption function fell afoul of 'spurious' regression critique: this work focussed on regressing the leve consumption on the level of income, two trended variables. Spanos (1989) reviewed the early empirical consumption literature and argues that muc the conclusions drawn from such models reflect *statistical* biases induce insufficient attention to the dynamic specification of such models. Sp views the subsequent literature on the *PIH*, *LCH* and relative income hypoth as attempts to derive more adequate statistical models (i.e, theoret reparameterisations of a statistically adequate specification).

and Granger, 1987).[4,5,6]

Cointegration theory states that if a linear combination of non-stationary time series is itself stationary, these series can be viewed as cointegrating and forming a 'long-run equilibrium' subspace. One can conceptualise that in the long-run, market forces or government intervention act to bring these series together although in the short-run they may drift apart. Cointegration theory therefore allows a separation of the long-run information contained in the data from the complicated short-run dynamics.

The cointegrating vector describes a static, long-run relation. The problem of modelling short-run dynamics is solved by the following theorem called the Granger Representation Theorem which states that if a set of variables are cointegrated *CI(1,1)*, then there is a corresponding, valid *ECM* form of those variables (Engle and Granger, 1987). This makes it possible to incorporate the notion of a steady state in a dynamic model. More formally, for N linearly independent cointegrating vectors, $d = b = 1$, there exists an error correction representation of N stationary random variables x_t if one can write

$$B(L)\,(1-L)\,x_t = -\gamma' z_{t-1} + v(L)\,\epsilon_t \qquad (3)$$

where $B(L)$ is a finite order polynomial with $B(0) = I_N$ and $v(L)$ is a finite order lag polynomial; ϵ_t is a stationary multivariate disturbance and $\gamma \neq 0$. As (3) contains only stationary variables, there is no problem with employing the usual inference. The Granger Representation theorem thus offers a sound theoretical rationale for employing *ECM*'s when the level terms form a cointegrating set; it also shows that if the data generating process is an equation such as (3), then x_t must be a cointegrating set of varaibles.[7]

Engle and Granger (1987) demonstrate that if *OLS* is employed to estimate the parameters of the cointegrating equation, then the parameters of the *ECM* can be consistently estimated if the first stage estimates of the cointegrating equation are imposed on a second stage *ECM*. This 'control' is implemented by including the lagged error terms from the cointegrating regression in a general *ECM*.[8] This procedure is referred to as the Granger-Engle two-step procedure. Stock (1987) shows that the order of convergence of these first stage *OLS* estimates is $O(T^1)$ compared with $O(T^{0.5})$ in the standard case; this faster convergence in the non-stationary case is sometimes referred to as 'super-consistency'. This fast rate of convergence of the estimator of the cointegrating parameter vector means that the

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[4] When the dimension of $x_t > 2$, α is not necessarily unique as there are possibly several cointegrating vectors, some of which may be linearly dependent.

[5] The above asserts that two variables with different orders of integrability cannot be cointegrated although when there are more than three series, mixtures of different order series are possible (Hall and Henry, 1988)

[6] In general, a linear combination of I(d) variables produces an I(d) random variable. Thus, a cointegration result is not general as it requires the same relationship to hold between variables for an extended time period.

[7] This theorem has been criticised by Wickens and Breusch (1988) who argue that there is no strong relationship between cointegration and the conventional *ECM* but Engle and Granger's version of the *ECM* can be interpreted as just another transformation of the original dynamic model. Wickens and Breusch also argue that the two-step procedure is unnecessary and that instrumental variables estimation is more convenient.

[8] The lagged residuals represent a levels term which acts to dampen short-run deviations and return the system towards the steady state equilibrium values: an 'error correction'. An *ECM* can be developed from such economic considerations as incomplete information or costs of adjustment (Engle and Granger, 1987).

estimators of the short-run parameters are asymptotically independent of the cointegrating vector. The second stage standard errors are 'correct', i.e. consistent for the true standard errors and the standard t-statistics can then be used for the short-run estimates (Stock, 1987; Engle and Granger, 1987). The full *ECM* model thus derived escapes the charge of a 'spurious' regression.

The actual test for a cointegrated vector proceeds in the following way (for $d > 0$): (i) investigate the temporal characteristics of every variable in the static 'long-run' equilibrium equation of interest; (ii) test for a cointegrated vector; (iii) formulate and estimate a general *ECM*; (iv) allow data-based diagnostic statistics to determine the *ECM*'s adequacy as a conditional characterisation of the data generating process.

IV. Data and Empirical Results

Data

For purposes of comparison with Sawyer (1992) we make use of the same data drawn from the National Accounts and the IPA's *Focus* model. Consumption, c_t, is defined as real consumption expenditures on goods (1986 = 100)[9,10]. The wealth-income ratio, w_t, is from the *Focus* model (see Sawyer (1992) for a precise explanation of its constructio Disposable income, y_t, is used as the income variable; the implicit price of goods relative the implicit price of services, p_t, is the price variable. g_t is the level of federal expenditur We experimented with splitting government spending into real aggregate defence expenditu and *per capita* real non-defence expenditures.[11] Based on our initial tests, it was deci to work only with the non-defence component. The (nominal) interest rate, I_t, is the yi on 1 - 3 year Government of Canada bonds. The interest rate variable should be the after rate of return. Some researchers like Carroll and Summers (1987) and Mankiw (19 simply impose an arbitrary value for the marginal tax rate. Rather than follow suit, simply use the before-tax rate. The unemployment rate, UR_t, is the rate of involun unemployment [12].

The data are seasonally adjusted values [13,14] and cover the 1958.1 - 1990.3 pe

[9] Sawyer argues that changes in classification of some service expenditures in the 1960's probably suggest separate treatment of consumption of services. Strictly speaking, this is valid only if consumption of goods is separable from consumption of other categories.

[10] The consumption measure should ideally capture only the service flows from durables during the observation period. In practice, a precise measure of such service flows is almost impossible to obtain so that in practice, either consumption of nondurables and services or total consumption is used. The consumption measure in the IPA's study was therefore expenditures on durables plus nondurables goods. In the regression analysis that followed, we attempted to test the sensitivity of results to using an alternate measur consumption, namely nondurable goods expenditures. Results proved extre sensitive to this change, with the equilibrium sub-space being far diffe from that estimated in Sawyer (1992) and that implied by our later resu These results are available on request. See also Lee and Siklos (1993).

[11] That is, defence spending is treated as a public good unlike defence spending.

[12] I.e., if *RU* is the total unemployment rate, *UR* = *RU* - *RUNAT* w *RUNAT* is the IPA's estimate of the natural rate of unemployment.

[13] The pre-filtering involved may conceivably overly smooth the c altering the time series properties of the raw series and hence cointegration results. With raw data, one must consider the possibilit potential unit roots at seasonal frequencies. However, Lee and Siklos (have shown that, considering a wide menu of standard macroeconomic vari for Canada, only the unemployment rate gives a different result in the unit roots tests when raw data is used as opposed to adjusted data. Th course does not imply that cointegration tests would be invariant wit choice of raw versus adjusted data. One paper investigating the sensit of cointegration results to this choice in Canada is Lee and Siklos (1993 found that " cointegration between consumption and income may be due, at in part, to seasonal adjustment at the source since a finding of cointegr between these same two series is easily rejected for seasonally unadj data." Our rationale for using seasonally adjusted data is a pragmatic

Consumption, income and government spending are all real, *per capita* magnitudes. Our empirical consumption function is of the form

$$c_t = \beta_0 + \beta_1 y_t + \beta_2 g_t + \beta_3 w_t + \beta_4 UR_t + \beta_5 I_t + \beta_6 p_t + \beta_7 INF + \eta_t \tag{4}$$

where *INF* is the rate of inflation (i.e., we test the restriction that the absolute values of β_5 and β_7 are equal). β_1 and β_3 are expected to be positive while β_4 is expected to be negative. β_2 can be negative or positive; a negative sign would suggest that c and g are substitutes; however, as g may be a productive input, c and g may be complements. As the interest rate and relative price variable may have both income and substitution effects, *a priori* the signs of β_5 and β_6 are uncertain although the usual expectation is for a common negative sign. The sign of β_7 may be ambiguous if one allows for Deaton-type confusion effects.

Empirical Results

Table 1 contains the results of the tests for the order of integrability of several series proposed for inclusion in the long-run static regression or later in the *ECM*. The test for stationarity in the levels of variables is given by

$$\Delta x_t = \alpha + \beta t + \delta_0 x_{t-1} + \sum_{j=1}^{J} \delta_j \Delta x_{t-j} + \epsilon_t \tag{5}$$

J is chosen to be sufficiently large to ensure that the error term is free of significant serial dependence. When $J = 0$, the Dickey-Fuller test obtains and $J \neq 0$ defines the Augmented Dickey-Fuller test. The null hypothesis that x_t follows a random walk is rejected if the coefficient on x_{t-1} is significantly negative. [15] Correspondingly, the test for an *I(1)* variable is that δ_0 is not significantly negative in (5) but γ_0 must be significantly negative in (6) below:

$$\Delta^2 x_t = \alpha + \beta t + \gamma_0 \Delta x_{t-1} + \sum_{j=1}^{J} \gamma_j \Delta^2 x_{t-j} + \epsilon_t \tag{6}$$

Both the *DF* and *ADF* statistics are presented in Table 1. Since the *DF* tests may lose power when the i.i.d assumption is invalid - see Phillips (1987) - the residuals (ϵ_t) are tested for serial correlation using a Lagrange Multiplier test and a variant of White's 1980 test for heteroscedasticity.[16] For the *ADF* tests, we employed different lag structures as was necessary to eliminate excess serial correlation. The results of the unit root tests indicate that

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to match Sawyer (1992) and because some of the non-standard variables used in our regressions such as *UR* and *W* are constructed (by the IPA) using adjusted data.

[14] Note that c, y and g are in billions of dollars, while population is measured in billions.

[15] The asymptotic distribution of the t-statistic on δ_0 is independent of the lagged first difference terms. The t-statistic on δ_0 does not however follow the standard t-distribution, but rather what is known as the Dickey-Fuller distribution (Dickey and Fuller, 1979). The presence of a trend and/or constant in (5) and (6) alters the distribution and different tables must be used for each case.

[16] Engle and Granger (1987) show that the *DF* test is likely to be more powerful when it is valid. It should be noted that the *ADF* and *DF* statistics have the same limiting distribution.

all variables considered are integrated of order 1. Consequently, first differencing is required for stationarity. [17] The critical values for the *DF* test statistics with $T = 127$, $N = 1$ for a regression with both constant and trend can be computed using the algorithm in Mackinnon (1990). These are: 1% (-4.03), 5% (-3.45), 10% (-3.15).

Since some of the DF results imply an error term that is not i.i.d, the Phillips (1987) and Phillips and Perron (1988) tests for weak dependence and heteroscedasticity in the residuals are also constructed. These tests were derived allowing for up to twelve lags using a Bartlett window to ensure positive definiteness of the variance-covariance matrix (Newey and West, 1987). Critical values are found in Fuller (1976) and Dickey and Fuller (1981). Only a summary of these results are reported here in order to conserve space: *c, p, y, g* are all I(1) with possible drift while *I, R, UR, w, s, INF and R* are zero mean I(1) variables. Thus these tests confirm the findings of the *DF* and *ADF* tests.[18]

The next step in the analysis is to investigate whether equation (4) represents a cointegrating set (or possesses at least one cointegrating subset). To ascertain whether the null hypothesis of no-cointegration is rejected, we check to see whether the *OLS* residuals from equation (4) are *I(0)*. A finding of *I(0)* implies cointegration. Estimating equation (4) over the period 1959.1 - 1990.3, the following results are obtained:

[17] We also tested those variables in the Sawyer model for completeness.

[18] Details of the Phillips and Phillips-Perron tests can be obtained from the authors on request.

$$\underset{(3.61)}{c = 0.813} \; \underset{(31.30)}{+0.795y} \; \underset{(2.05)}{+0.063g} \; \underset{(2.36)}{-0.188p} \; \underset{(9.58)}{-0.011UR} \; \underset{(2.37)}{+0.056w} \; \underset{(8.17)}{-0.008I}$$

$$R^2 = 0.9967 \qquad RSS = 0.028152 \qquad D.W = 1.09$$

$$DF = -6.69 \qquad ADF = -5.61 \qquad (7)$$

RSS is the residual sum of squares, *D.W* is the cointegrating regression Durbin-Wats (*CRDW*) statistic due to Sargan and Bhargava (1983). This latter statistic can be used to gi a rough indication as to whether there is cointegration. The value of 1.09 compared with critical value of about 0.51 would seem to indicate cointegration. [19]

The more formal tests are the *DF* and *ADF* tests. The *t*-ratios from these tests longer have the Dickey-Fuller distribution given the need to estimate the cointegrati parameters. The distribution of the test, first tabulated by Engle and Granger is referred as the Engle-Granger distribution. The form of the tests, however, remains unaltered. T Mackinnon (1990) tables allow one to compute these critical values for up to six variabl As equation (4) contains seven variables, some informal extrapolation of these results inevitable. For six variables and no trend, the 1% critical value of the test is -5.45 and critical value is -4.841. Both tests clearly support a cointegrating result.

[19] This is at the 1% level. See Engle and Granger (1987, Table 2). should be noted that the *CRDW* test has low power to reject the null of cointegration (a unit root) against alternatives close to the unit cir although an argument can be made for its use on the grounds that distribution is invariant to nuisance parameters such as the const (Banerjee *et.al.*, 1986).

Although t-statistics are reported in parentheses, these are to be interpreted very carefully as the standard errors are not correct, being biased downward. However, this allows us to 'confidently' eliminate insignificant variables from the cointegrating vector. Based on this, equation (7) differs from equation (4) by the exclusion of the insignificant inflation variable.[21]

Banerjee *et.al.* (1986) have also shown that there could be substantial small sample bias in the cointegrating vector estimators; this bias declines more slowly than theoretically expected. Their Theorem 2 shows, however, that $(1-R^2)$ is an indicator of the bias in the *OLS* estimator: the bias goes to zero as R^2 goes to 1. Since our reported R^2 is 0.9967, the bias may be small in our case, although it is not possible to assert that the bias for individual parameters is negligible (Banerjee *et.al.*, 1986, p.262). Thus, a high R^2 is a necessary condition for adopting the two-step procedure.

The results of the cointegration tests would suggest that a cointegrated set of variables have been obtained. However, before going on to the *ECM* formulation, there is a further complication to consider. Although equation (7) is a valid cointegrating regression, it may not be unique. There can be several 'equilibrium' relationships linking $N > 2$ variables (Granger, 1986). In fact, any one of the regressors in equation (7) could have been used as the regressand, producing r distinct cointegrating vectors, where $r \leq N-1$.[22] However, given the properties of OLS, the vector of cointegrating parameters implied by the different inversions will differ from equation (7). Inspection of Table 2 would indicate that the rank of the cointegrating matrix, r, is greater than one and the different inversions imply sometimes vastly different though consistent estimates of the long-run static parameters.[22] Based on the diagnostic statistics reported in Table 2, the rank of the cointegrating vector is 4. This non-uniqueness and the different implied estimates compound the problems of meaningfully interpreting the cointegrating parameters derived using *OLS*.

The literature provides two ways of handling this problem: (i) Hall and Henry (1988) suggest normalisation on the dependent variable in the cointegrating equation with the highest R^2. In our case, this would point to *both* c and y! At this point an arbitrary decision must be made to normalise on consumption, the variable of interest, rather than income.[23] (ii) The maximum likelihood approach of Johansen which provides consistent estimates of the cointegrating matrix as well as statistical tests for the size of r (critical values are found in Johansen (1988) and Osterwald and Lenum (1992)). In our case, based on the maximal eigenvalue criterion, r was found to be one when three lags were employed for the vector autoregression in the Johansen procedure.[24]

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[21] Thus, as in Blinder and Deaton (1985), it is found that the nominal rate of interest appears relevant; in fact, the inflation rate proved insignificant and was subsequently dropped from the regression. This finding is quite puzzling and difficult to rationalise unless one argues, say, that the nominal rate already incorporates an expectation of the rate of inflation.

[22] As Granger(1986) points out, this is a problem which parallels the usual identification problem of classical simultaneous equations models.

[22] Thus the endogeneity of a right hand side variable does not affect the consistency of the OLS estimates (Hall and Henry, 1988, p58-59)

[23] Engle and Yoo (1987, p.156) suggests simply picking a particular (natural) normalisation

[24] The VAR length was chosen by testing down from a general VAR of length 5 until reducing the order by one lag could be rejected using a likelihood ratio test.

As the Johansen procedure is more formal and powerful we proceed on the belief that $r = 1$. The above complications demonstrate why cointegration theory is somewhat weak in practice and why a test for cointegration must be considered as only an initial procedure (a pre-test) aiding in formulating the dynamic (*ECM*) model (Granger, 1986).

Let's proceed to the second stage of the Granger-Engle procedure: the *ECM*. An initially over-parameterised model with 4 lags on the dependent and independent variables was sequentially simplified by use of F-tests of restrictions until a parsimonious representation of the data generating process (*DGP*) was achieved. The lagged residual term from equation (7), the cointegrating equation is included in the *ECM* and in effect incorporates a set of cross-equation restrictions derived from the static relationship. It is also possible to test for non-linear long-run adjustment. To do this, squared and cubed lagged error correction terms are added to the *ECM*, following Escribano (1987).

The general *ECM* estimated was

$$\Delta c_t = constant + \sum_{i=1}^{4} \gamma_i \Delta c_{t-i} + \sum_{j=0}^{4} \delta_j \Delta q_{t-j} + \sum_{k=1}^{3} \beta_k z_{t-1}^{k}$$

where q_t are the regressor variables from equation (7). F-tests again indicated the non-inclusion of an inflation measure.

The *ECM* model was estimated over the 1959.2 - 1987.3 period to save 12 data observations for forecast tests. A battery of diagnostic checks was conducted to ascertain the chosen *ECM* specification:

$$\begin{aligned}\Delta c = {} & \underset{(0.61)}{0.001} + \underset{(4.69)}{0.356\Delta y} + \underset{(3.44)}{0.255\Delta y_{-3}} + \underset{(2.96)}{0.055\Delta g} - \underset{(3.26)}{0.523\Delta p} \\ & \underset{(4.19)}{-0.014\Delta UR} - \underset{(2.59)}{0.009\Delta UR_{-1}} + \underset{(2.70)}{0.009\Delta UR_{-3}} \qquad (8) \\ & \underset{(3.09)}{-0.010\Delta UR_{-4}} + \underset{(2.66)}{0.118\Delta w_{-1}} - \underset{(7.45)}{0.575\Delta z_{-1}}\end{aligned}$$

$R^2 = 0.6104$ $\quad RSS = 0.0110$ $\quad D.W. = 2.08$ $\quad F(10,103) = 16.14$
$PC1\,[\chi^2(12)/12] = 1.91$ $\quad PC2\,[F(12,103)] = 1.58$
$LMN\,[\chi^2(2)] = 3.98$ $\quad SC1\,[F(8,95)] = 1.93$
$ARCH\,[F(4,95)] = 0.68$ $\quad HET\,[F(20,82)] = 0.68$
$RESET\,[F(1,102)] = 0.33$ $\quad SC2\,[\chi^2(12)] = 18.67$
$MIS\,[F(27,78)] = 1.28$ $\quad WUT\,[F(4,99)] = 1.80$
$F_R(27,76) = 1.24$

where Δ is the first difference operator; z_{t-1} is the lagged error correction term from equatic (7). t-statistics are in parentheses. Correcting the standard errors using White's (198 adjustment for general heteroscedasticity produced little change and hence we do not repc the adjusted standard errors. F_R is an F-test of restrictions for going from the general EC to equation (8) and supports the reduction.[25] *PC1* is a parameter constancy test as is *P*

[25] To test the robustness of the general-to-simple modelling, tho variables excluded in the simplification search were added, one at a tim The t - statistic showed in every case but one that the exclusions we warranted. However, including this other variable - the third lag on t government variable - caused the model to fail predictive accuracy test strengthening our confidence in the parsimonious specification (8).

which is Chow's (1960) test for structural change over the forecast period. *LMN* is the Jarque and Bera (1980) normality test; *SC1* is a lagrange multiplier test for serial correlation (up to the 8th lag); *ARCH* is an autoregressive conditional heteroscedasticity test for order 4; *HET* is a test for heteroscedasticity based on White (1980); *RESET* is a general test for misspecification due to Ramsey (1969); *SC2* is the Ljung-Box residual correlogram statistic for lags of order 12 and *MIS* is White's heteroscedasticity/functional form misspecification test. *WUT* is Wu's (1973) T^2 statistic for testing the independence of regressors. The terms χ^2 or F denote the distribution of the test statistics under the null. Where applicable, the F-versions of the tests are used in light of Kiviet's (1985) Monte Carlo evidence that the F variants are more powerful in small samples.[27]

Equation (8) suggests further restrictions that could be imposed in an effort to achieve even greater parsimony. This leads to equation (8a) below, where the data have accepted an equality restriction on the (absolute) values of the coefficients of the lagged unemployment rate terms. As seen, coefficient values are identical in equations (8) and (8a). We focus on equation (8a) in the analysis that follows.

[27] Only in one case did the χ^2 variant contradict the F test.

$$
\begin{aligned}
\Delta c = {} & \underset{(0.59)}{0.001} + \underset{(4.80)}{0.356}\,\Delta y + \underset{(3.63)}{0.258}\,\Delta y_{-3} \\
& + \underset{(2.98)}{0.055}\,\Delta g - \underset{(3.35)}{0.524}\,\Delta p - \underset{(4.81)}{0.014}\,\Delta UR \\
& - \underset{(4.18)}{0.009}\,[\Delta UR_{-1} - \Delta UR_{-3} + \Delta UR_{-4}] \\
& + \underset{(2.82)}{0.186}\,\Delta w_{-1} - \underset{(7.78)}{0.574}\,z_{-1}
\end{aligned}
\qquad (8a)
$$

$R^2 = 0.6100$ $RSS = 0.0111$ $D.W. = 2.00$ $F(8,105) = 20.53$

$PC1\,[\chi^2(12)/12] = 1.94$ $PC2\,[F(12,105)] = 1.65$

$LMN\,[\chi^2(2)] = 3.50$ $SC1\,[F(8,97)] = 1.92$

$ARCH\,[F(4,97)] = 0.63$ $HET\,[F(16,88)] = 0.77$

$RESET\,[F(1,104)] = 0.24$ $SC2\,[\chi^2(12)] = 18.83$

$MIS\,[F(27,78)] = 1.14$ $WUT\,[F(4,101)] = 1.66$

The various tests for autocorrelation are not significant as are the tests for heteroscedasticity, including *ARCH*. These results, plus the indirect evidence from the White standard errors referred to above would indicate that the residuals from equation (8a) are empirical white noise.

The correlation matrix (Table A2) reveals that the explanatory variables constitute near-orthogonal regressors suggesting that multicollinearity is not a problem. The normality test of Jarque and Bera indicates that the residuals are normal. Additional data coherence checks are provided by *MIS*[28] and *RESET*. *RESET* has power against many forms of misspecification but is particularly useful when the maintained model has under-represented the curvature of the true regression function. *MIS* is also a very general test for functional form misspecification. Both tests accept the null of no misspecification.

R^2 for the *ECM* is 0.6100 which is not particularly high for time series data. R^2, however, is not the only or indeed best criterion for judging the fit of a model. A more illuminating check is the out of sample forecasting/parameter constancy tests (Hendry, 1980). Forecasts were done for the period 1987.4 - 1990.3 post-sample period. The statistics *PC1* and *PC2* indicate that parameter constancy was maintained over this period. However, plots of fitted and unfitted and forecast versus actual values showed that the model understated the sharp falls in consumption in 1974, 1983 and in 1990 during the recent recession.[29] Overall, however, the model picks up turning points rather well. Tests for structural breaks in 1974.1 and 1983.2 were carried out using the standard Chow test. Estimated test values were 1.79 and 1.34 for breaks at 1974.1 and 1983.2, respectively, well below the 5% critical value of about 1.96 for F(9,108). Informal tests such as the Cusum and Cusum-Squared tests also suggested parameter constancy. No structural break was indicated and the model appears stable.

[28]For squares and cross-products of variables. This limits the number of variables that can be simultaneously included in the test. The reported *MIS* statistic is based on the first seven variables although results based on the next seven and other combinations were similar.

[29] This in itself is not surprising, given that the model is only an estimate. Such a result hints at the possible exclusion of some form of information probably associated with expectations.

In a macroeconomic context, it is quite plausible to speculate that y, g, UR, p etc. a jointly determined along with consumption. The Wu-Hausman statistic (WUT) is used to te the independence between the regressor set and the error term. The implementation of t test requires the regression of each of the current first differenced regressor terms in equati (8a) on a set of instruments such as lagged regressors from the *ECM* as well as lagged lev of variables, say, from the static regression. The residuals from these regressions are sav and an F-test for their inclusion in the *ECM* is conducted. A significant test statistic impl rejection of the null hypotheses of weak exogeneity.[30] Consequently, the null of w exogeneity is accepted as the estimated value of 1.66 lies well below the critical value 2.46 for F(4,101).

As required, if equation (7) is a valid cointegrating vector, the coefficient on z significant and negative. It should be noted that while the standard t-tests can be used to hypotheses about the other coefficients, the coefficient on z_{t-1} has a limiting non-nor distribution. Critical values for this coefficient are much higher, with a rule of thum about 3 (Hendry, 1986).[31] All variables are significant at the 1% level and have expe signs. Note that I while part of the equilibrium sub-space, does not significantly affec short-run dynamics of the model. Given that the *ECM* model has satisfied such a wide a of diagnostic tests, equation (8a) is accepted as a well specified model in a statis

[30] As is well known, weak exogeneity is a sufficient condition conducting conditional inference without loss of relevant sample informa (see Engle *et.al.*, 1983).

[31] Banerjee *et.al.* (1986) provide simulation results to show that i *ECM* with restrictions imposed by the presence of the error correction t testing the significance of z_{t-1} at the usual 5% level proves surprising rob

sense[31]. . A behavioural interpretation of the *ECM* results can now be attempted.

Of particular note is the negative coefficient on the unemployment variable. A similar finding for Canada was reported by Cuddington (1982). If *UR* is interpreted as reflecting liquidity or employment constraints, then equation (8a) implies that such constraints may be present and significant in both the long- and short-run although the direct effect on *per capita* consumption is relatively small. It appears that government non-defence expenditures complement private consumption expenditures, augmenting rather than substituting for private spending. This is not a novel result as Blinder and Deaton (1985) report a similar finding for the USA. As income and the wealth to income ratio grow, consumption expands but it contracts with increases in relative prices as the substitution effect outweighs the income effect and expenditure switches to the purchase of services.

Comparison with Sawyer(1992)

Another test of the adequacy of equation (8a) is to compare it with other models 'explaining' agggregate consumption. A natural candidate for the rival model is one of the models found in Sawyer (1992). Sawyer's model provides a case for non-nested testing. The Granger-Engle two-step procedure applied to the Sawyer formulation - equation (IV), Table 2. p.13 - yields the following results for the static equation for the 1959.1 - 1990.3 period.

[31] An unrestricted *ECM*, where z_{-1} was replaced by one period lags of the variables in equation (7), was ran as one further check of our model. If the cointegration finding is valid, the implied steady-state relationship should be similar to equation (7). Inspection of Table (A3) in the Appendix shows this to be the case. However, the unrestricted *ECM* exhibits some instability around 1973Q4 as evidenced by a Cusum-squared plot. This lends support to imposing the 'control' implied by inclusion of z_{-1} in equation (8a).

$$\underset{}{c} = \underset{(11.32)}{3.240} + \underset{(17.15)}{0.562\,y} - \underset{(11.42)}{0.009\,R} - \underset{(6.30)}{0.297\,p} - \underset{(5.92)}{0.002\,TREND} \qquad (9)$$

$$R^2 = 0.9959 \quad D.W = 0.83 \quad DF = -5.70 \quad ADF = -5.46$$

Critical values for the above case with a trend included are: 1% (-5.15); 5% (-4.55). Sawyer does not report the results of any inversions of the cointegrating relationship so we estimate the inversions as a further check of the Sawyer model. These results are reported below in Table 3.

As can be seen on inspection of Table 3, the rank of the cointegrating matrix is at least 2. However, in this case, Hall and Henry's (1988) suggestion of normalising on the cointegrating equation with the highest R^2 leads unambiguously to the choice of normalising on consumption. However, for the case of four lags, the more formal Johansen approach suggest that $r = 1$, where the choice of 4 lags was determined using a likelihood ratio test as before.

The *ECM* formulation used in the non-nested tests is presented below. This is essentially Model I in Sawyer (1992, p.17), his preferred model.[32] However, the estimation period is now 1959.2 - 1987.3, saving 12 observations for an out of sample forecast test.

[32] Sawyer includes variables not considered in the cointegration tests in his *ECM*, such as *UR*. We have added a dummy variable to the Sawyer equation as all of his equations show signs of instability, even over the longer period, 1959.2 -1990.3. His models "marginally" but uniformly fail Chow tests for any period between 1972.1 - 1974.4. The dummy takes on a value of unity during this period; zero otherwise. Its inclusion cause parameter values to change little; they are quite close to those reported by Sawyer.

$$\Delta c = 0.0003 + 0.406\Delta y + 0.310\Delta y_{-3} - 0.004\Delta R - 0.608\Delta p$$
$$(0.21) \quad (5.23) \quad (3.99) \quad (3.21) \quad (3.56) \qquad (10)$$
$$-0.016\Delta UR_{-1} - 0.497z_{-1} - 0.007DUMMY$$
$$(5.47) \quad (7.56) \quad (1.84)$$

$R^2 = 0.5524$ $\quad F(7,106) = 18.69$ $\quad RSS = 0.01269$

$D.W = 1.99$ $\quad PC2[F(12,106)] = 0.65$ $\quad LMN[\chi^2(2)] = 2.43$

$PC1[\chi^2(12)/12] = 0.68$ $\quad SC1[F(8,98)] = 0.72$

$HET[F(13,92)] = 1.22$ $\quad ARCH[F(4,98)] = 1.08$

$RESET[F(1,105)] = 0.80$ $\quad SC2[\chi^2(12)] = 15.27$

$MIS[F(27,78)] = 0.90$ $\quad WUT[F(3,103)] = 0.48$

The cointegration tests and diagnostics reported above suggest that equation (10) is a valid *ECM*.

The non-nested tests employed are the F-test, calculated by embedding the null and alternative models in a comprehensive model and the adjusted Cox-type tests, *NT* and *W*, which incorporate small sample corrections designed to bring actual significance levels close to the normal values (see Godfrey and Pesaran, 1983). Godfrey and Pesaran present simulation evidence demonstrating that the *NT* and *W* tests have finite sample significance levels which are quite close to the normal values over a wide range of *DGP*'s. Moreover, these tests are more likely to lead to the correct decision of accepting the true model and rejecting the false model than the familiar F tests based upon the comprehensive model or regression parameter encompassing approaches. Note that these tests do not exclude t possibility of both models being rejected against each other.

The results of the three tests show that equation (8a) and equation (10) reject ea other using *NT* and *W* (Table 4 below). The encompassing F-test indicates that the Sawy model cannot encompass equation (8a) at conventional levels of significance; howev equation (8a) can be treated as a valid restricted form of the embedding model which inclu them both if we use the 1% level (critical value 3.51). In addition, the Akaike's Informat Criterion and the Schwarz's Bayesian Criterion both favour equation (8a) over equation (1

V. Conclusion

This study applied a cointegration approach to the estimation of an aggreg consumption function using quarterly Canadian data. A dynamic *ECM* model was develo which satisfies a number of model diagnostics including tests for predictive accuracy out the sample period. The results of the model suggest that income, interest rates, governm spending, wealth and liquidity constraints may well be relevant 'explanatory' variables aggregate consumption. Unlike Sawyer (1992), we find that our data based model approach supports a role for the rate of unemployment, in both the short-run and long formulations. The procedure suggested by a cointegration approach is not without its pit and the results contained herein would indicate several *caveats* in interpreting the resul

Non-nested and encompassing tests using a model suggested by Sawyer (1992) a rival model hinted at the 'superiority' of our specification. Further investigation is how warranted on at least two counts before either equations (8a) or (10) can be accepted as more than initial, tentatively adequate conditional characterisations of the *DGP*: (i) both models fail to fully track the most significant movements in the data - although equation (8a) passes a variety of parameter stability tests - and the original Sawyer model fails a formal Chow test for structural break; (ii) results may be compromised by the use of seasonally adjusted data. An interesting focus of future research work would be the use of raw data to test for the added presence of seasonal unit roots and to see whether results are significantly altered when such data are employed. Nonetheless, the *ECM* results (equation (8a)) survive a battery of diagnostic tests and indicate that future work along the general lines of this paper may be profitable.

APPENDIX

Table A.1 Cointegration Results

Dependent variable is c_t

Regressors	I	II	III
constant	0.816	0.724	0.813
	(3.63)	(2.85)	(3.61)
y	0.795	0.802	0.795
	(31.30)	(29.63)	(31.30)
g	0.063	0.069	0.063
	(2.04)	(2.16)	(2.05)
p	-0.189	-0.154	-0.188
	(2.37)	(1.68)	(2.36)
UR	-0.011	-0.011	-0.011
	(9.57)	(9.28)	(9.58)
w	0.057	0.047	0.056
	(2.38)	(1.73)	(2.37)
R	-0.009	-----	----
	(8.16)		
INF	-----	-0.319	----
		(-0.77)	
I	-----	-0.008	-0.008
		(7.70)	(8.17)
R^2	0.9967	0.9967	0.9967
DW	1.09	1.10	1.09
DF	-6.69	-6.76	-6.69
ADF	-5.61	-5.66	-5.61

Table A2

Estimated Correlation Matrix of Variables for Equation (8a)

	Δc	Δy	Δy_{-3}	Δg	Δp	ΔUR	Δw_{-1}	z_{-1}	ADUR
Δc	1.00	0.36	0.21	0.18	-0.29	-0.27	-0.01	-0.35	-0.28
Δy		1.00	0.00	0.01	-0.06	-0.27	0.15	0.24	-0.10
Δy_{-3}			1.00	0.15	-0.06	0.05	-0.23	-0.03	0.05
Δg				1.00	-0.02	0.03	-0.08	0.02	0.01
Δp					1.00	-0.16	-0.04	0.23	-0.13
ΔUR						1.00	0.19	-0.28	0.29
Δw_{-1}							1.00	0.16	0.15
z_{-1}								1.00	-0.10
ADUR									1.00

*ADUR is $[\Delta UR_{-1}-\Delta UR_{-3}+\Delta UR_{-4}]$ as in equation (8a).

Table A3

Unrestricted *ECM*

Period 1959.2 - 1987.3

Dependent variable is Δc

constant	0.446	(2.46)	Δy	0.385	(4.86)
Δy_{-3}	0.276	(3.80)	Δg	0.054	(2.33)
Δp	-0.548	(3.26)	ΔUR	-0.009	(2.69)
ADUR*	-0.010	(4.19)	Δw_{-1}	0.129	(1.71)
p_{-1}	-0.132	(1.92)	c_{-1}	-0.529	(7.05)
I_{-1}	-0.005	(5.71)	UR_{-1}	-0.004	(2.95)
w_{-1}	-0.021	(1.00)	g_{-1}	0.024	(0.85)
y_{-1}	0.427	(6.78)			

$R^2 = 0.6421$ $F[14,99] = 12.68$ D.W $= 2.26$ $SC1[F(4,95)] = 1.69$

$HET[F(1,112)] = 0.38$ $LMN[\chi^2(2)] = 1.95$

$RESET[F(1,98)] = 0.09$

The implied long-run equation from the above regression is obtained by setting $\Delta x = 0$, all variables x. This gives

$$c = 0.842 + 0.807y + 0.046g - 0.250p - 0.008UR + 0.041w - 0.01$$

which is quite similar to equation (7).

*ADUR is $[\Delta UR_{-1}-\Delta UR_{-3}+\Delta UR_{-4}]$ as in equation (8a).

Table 1

DF and *ADF* test statistics+

Variable	Level	First Difference
c	-0.73 (-0.54)	-12.33 (- 7.67)
y	-0.81 (-2.00)	-13.43 (- 7.09)
p	-0.73*(-0.93)	- 8.96 (- 6.01)
R[$]	-1.70 (-1.89)	-10.42*(- 8.25)
INF	-3.27*(-2.19)	-16.50 (-10.93)
I	-2.66*(-2.46)	- 9.95*(- 9.45)
w	-0.22*(-1.37)	-12.54*(- 6.77)
s	-2.39*(-2.35)	-12.21 (- 7.94)
UR	-1.88*(-2.84)	- 5.59 (- 4.72)
g	-1.41*(-0.02)	-19.42*(- 5.38)

Period: 1959.1 - 1990.3

* denotes serial correlation and/or heteroscedasticity in the *DF* formulation at the 5% level.

$ the real interest using the *IPA*'s estimate of expected inflation as used in Sawyer(1992).

+ ADF test statistics are in parentheses.

Table 2

Inversions of equation (7)

Dependent variable	constant	y	g	p	UR	w	I	R^2	DW	DF	ADF
c	0.813	0.795	0.063	-0.188	-0.011	0.056	-0.008	0.9967	1.09	-6.69	-5.61
y	0.309	0.892	0.009	-0.123	-0.102	0.002	-0.009	0.9967	1.16	-7.02	-5.99
g	-5.146	0.112	1.869	3.277	-0.043	1.133	-0.028	0.9886	1.63	-9.22	-6.27
p	10.932	0.523	-1.100	-4.255	0.009	-0.336	0.021	0.9777	0.93	-6.13*	-4.03
UR	-0.075	0.745	0.249	0.136	-0.025	0.253	-0.009	0.7813	0.71	-5.00*	-3.60
w	-0.799	0.031	1.267	0.031	-0.049	1.255	-0.013	0.8543	0.99	-6.28*	-3.75
I	-1.587	0.952	0.224	0.479	-0.012	0.092	-0.022	0.8736	1.07	-6.70*	-6.24

* denotes serial correlation/heteroscedasticity in the *DF* formulation.

NB: In interpreting the above table, note that the variables in the first column represent the dependent variable in that particular inversion. However, to facilitate ease of comparison with equation (7) - i.e, the first row - all equations have been rewritten so that the coefficient on *c* is always unity. Coefficient values can therefore be compared directly across equations.

Table 3. Inversions of Equations (9)

Dependent Variable	constant	y	R	p	Trend	R^2	DW	DF	ADF
c	3.240	0.562	-0.009	-0.297	0.002	0.9959	0.83	-5.70	-5.46
y	1.207	0.795	-0.007	-0.207	4×10^{-5}	0.9936	0.61	-4.73	-4.04
R	4.181	0.444	-0.018	-0.040	0.004	0.8049	0.72	-5.28	-5.73
p	4.976	0.391	-0.001	-1.208	2×10^{-4}	0.9390	0.17	-2.55	-2.44

NB: In interpreting the above table, note that the variables in the first column represent the dependent variable in that particular inversion. However, to facilitate ease of comparison with equation (7) - i.e, the first row - all equations have been rewritten so that the coefficient on *c* is always unity. Coefficient values can therefore be compared directly across equations.

Note: the inclusion of a time trend in the cointegrating equation is equivalent to detrending the data before the cointegration test; this raises the critical value of the *ADF* and *DF* tests.

Table 4

Non-nested and Encompassing Tests for

Equation (8a) and Equation (10): 1959.2 - 1987.3

Test Statistic	Distribution	(8a) vs (10)	Distribution	(10) Vs (8
NT-test	N(0,1)	-2.59	N(0.1)	-5.47
W-test	N(0,1)	-2.38	N(0,1)	-4.62
Encompassing				
F-test	F(4,101)	2.63	F(5,101)	5.40

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