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HOUSEHOLD SAVING AND PERMANENT INCOME IN CANADA AND THE UNITED KINGDOM

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ABSTRACT

Recent theoretical research in open-economy macroeconomics has emphasized the connection between a country's current account and the intertemporal savings and investment choices of its households, firms, and governments. In this paper, we assess the empirical relevance of the permanent income theory of household saving, a key building block of recent theoretical models of the current account. Using the econometric approach of Campbell (1987), we are able to reject the theory on quarterly aggregate data in Canada and the United Kingdom. However, we also assess the economic significance of these statistical rejections by comparing the behavior of saving with that of an unrestricted vector autoregressive (VAR) forecast of future changes in disposable labor income. If the theory is true, saving should be the best available predictor of future changes in disposable labor income. We find the correlation between saving and the unrestricted VAR forecast to be extremely high in both countries. The results suggest that the theory provides a useful description of the dynamic behavior of household saving in Canada and Britain.

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1. Introduction

Recent theoretical research in open-economy macroeconomics has emphasized the connection between a country's current account and the intertemporal savings and investment choices of its households, firms, and governments (Buiter (1981), Sachs (1981, 1982), Dornbusch (1983), Frenkel and Razin (1985), Persson and Svensson (1983)). To the extent that national savings and investment patterns reflect forward looking behavior on the part of households and firms which optimize in the absence of liquidity constraints, expectations about future economic variables can significantly influence the magnitude and persistence of current account deficits and surpluses observed in the present.

The objective of this paper is to assess the empirical relevance of the permanent income theory of household saving, a key building block of recent theoretical models of the current account. According to the permanent income hypothesis, consumption is proportional to permanent income; it thus tends to be above current income when current income is relatively low and expected to rise, and to be below current income when current income is relatively high and expected to fall. This intuitive observation has the striking and until recently ignored econometric implication that household saving should be the best available predictor of future changes in disposable labor income. Using the econometric methodology developed in Campbell (1987), we test this implication of the permanent income hypothesis on Canadian and British quarterly aggregate data and compare our findings to those obtained by Campbell for U.S. data.

Our approach is to test cross-equation restrictions on the coefficients of a bivariate vector autoregression (VAR) comprised of saving and changes in disposable labor income. Thus, by contrast with most of the recent literature, we do not directly test the consumption martingale property emphasized by Hall (1978).

In both the Canadian and British data, we find that saving Granger causes changes in disposable labor income and that it is significantly negatively correlated with the subsequent change in disposable labor income. This is exactly what one would expect if the permanent income hypothesis is true, since saving occurs because labor income is expected to decline in the future. However, we are able to reject statistically the cross-equation restrictions implied by the permanent income hypothesis. In the case of Canada, the statistical rejection of the theory appears to result from the fact that the mean of the unrestricted forecast of future changes in labor income differs substantially from the mean of Canadian saving. In the case of Britain, violations of the cross-equation restrictions implied by the theory are more pervasive.

To assess the economic significance of these statistical rejections of the theory, we compare the behavior of saving in each country with that of the unrestricted VAR forecast of future changes in disposable income. If the permanent income hypothesis is true, saving and the unrestricted VAR forecast should have identical standard deviations and should be perfectly correlated. In fact, we find that the correlation between saving and the unrestricted VAR forecast of future changes in disposable labor income is extremely high in both countries, exceeding .89 in the British data and .99 in the Canadian data. While the sample standard deviation of saving is somewhat less than the standard deviation of the VAR forecast, time series plots of the two series for each country reinforce the impression conveyed by the correlation results that a substantial fraction of the forecastable variation in disposable labor income is incorporated in Canadian and British household saving behavior. Thus, even though it is possible to reject the permanent income hypothesis at conventional levels of statistical significance, our results suggest that the theory provides a useful description of the dynamic behavior of household saving in Canada and Britain. Our findings correspond quite closely to those obtained by Campbell in his test of the permanent income hypothesis on U.S. data.

2. The Permanent Income Hypothesis

Following Flavin (1981), we write the permanent income model as

$$(1) \qquad c_t = \gamma y_{Pt} = \gamma[y_{kt} + (\frac{r}{1+r}) \sum_{i=0}^{\infty} (\frac{1}{1+r})^i E_t y_{\ell,t+i}],$$

where c_t is real per capita consumption, y_{Pt} is permanent income, y_{kt} is real per-capita capital income, $y_{\ell t}$ is real per capita labor income, r is the expected real interest rate, and γ is the propensity to consume out of permanent income. The model assumes that γ and r are constant and that $\gamma \leqslant 1$.

Permanent income is defined as the Hicksian income generated by nonhuman and human wealth. The Hicksian income from human wealth is r times the present discounted value of expected labor income. The Hicksian income from nonhuman wealth, W_t, is just $y_{kt} = rW_t$. Wealth evolves according to $W_t = (1+r)W_{t-1} + y\ell_{t-1} - c_{t-1} + \eta_t$ so that capital income obeys

$$(2) \qquad yk_t - (1+r)yk_{t-1} - r[y\ell_{t-1} - c_{t-1}] = r\eta_t;$$

where η_t represents unanticipated capital gains and is unforecastable as of time t-1. Note that in general the conditional variance of η_t will be positively related to the level of wealth W_{t-1}.

Our interest is in ascertaining and testing the restrictions which the permanent income hypothesis places on saving behavior. Define transitory income y_{Tt} as the difference between total disposable income, $y_t \equiv y_{kt} + y_{\ell t}$, and permanent income y_{Pt} which is defined in equation (1). Define saving, s_t, as the difference between total disposable income and consumption: $s_t \equiv y_t - c_t$. If $\gamma = 1$, so that consumption is equal to permanent income, equation (1) can be rearranged so that it becomes a statement about saving.

$$(3) \qquad s_t = y_t - y_{Pt} = y_{Tt} = -(\frac{r}{1+r}) \sum_{i=0}^{\infty} (\frac{1}{1+r})^i [E_t y_{\ell,t+i} - y_{\ell t}]$$

$$= - \sum_{i=1}^{\infty} (\frac{1}{1+r})^i E_t \Delta y_{\ell,t+i} ,$$

where Δ denotes a standard backward difference.

Equation (3) says that saving equals transitory income, which in turn can be expressed as the expected present value of future declines in labor income. It follows from (3) that

$$(4) \qquad s_t - \Delta y_{\ell t} - (1+r)s_{t-1} = -r\epsilon_t ,$$

where $\epsilon_t = [1/(1+r)] \sum_{i=0}^{\infty} [1/(1+r)]^i [E_t y_{\ell,t+i} - E_{t-1} y_{\ell,t+i}]$ is the unforecastable revision from t-1 to t in the expected value of human wealth. Equation (4) is not immediately intuitive, stating that a linear combination of the change in labor income, current, and lagged saving is unforecastable. However, subtracting (4) from (2) and using the definition of s_t, we obtain Hall's (1978) celebrated result that consumption is a random walk:

$$(5) \qquad c_t - c_{t-1} = r[\eta_t + \epsilon_t].$$

If $\gamma < 1$, the above analysis needs to be modified. We can define a new variable $\tilde{s}_t$ as $\tilde{s}_t \equiv y_t - c_t/\gamma$. Equations (3) and (4) now apply to $\tilde{s}_t$ rather than s_t and we have

$$(6) \qquad s_t = \tilde{s}_t + (1-\gamma)y_{Pt} = y_{Tt} + (1-\gamma)y_{Pt}.$$

Equation (6) says that people save their transitory income, and a fraction $(1-\gamma)$ of their permanent income. Our approach is to evaluate the model by examining $\tilde{s}_t$.

Although tests of the permanent income hypothesis usually focus on the consumption martingale implication given by equation (5), we shall adopt the econometric methodology developed by Campbell (1987) which uses the restriction on savings behavior given by equations (3) and (4) to assess the empirical relevance of the theory. There are two main reasons for this choice.

First, the random walk behavior of consumption is only one implication of the permanent income hypothesis. A time series can follow a random walk and yet not be determined by permanent income (although such a series will not obey the intertemporal budget constraint). Equations (3) and (4) directly examine the relation between consumption and income.

Secondly, the restriction on household saving behavior given by equation (3) can be used to characterize the fit of the permanent income model. In the context of the recent theoretical work in open-economy macroeconomics, it is worthwhile to investigate the extent to which the movements of household saving incorporate forecastable variations in income, even -- or perhaps especially -- if the theory is statistically rejected.

3. Econometric Methodology

In this section we explain the econometric concepts and techniques which we use to evaluate the permanent income model. We begin by discussing the stationarity of the different variables in the model; then we show that the model implies an intuitive set of restrictions on a stationary VAR; and finally we discuss the estimation of nuisance parameters.

If the permanent income hypothesis holds with $\gamma = 1$ and changes in labor income are stationary, then equations (2), (3) and (5) imply that consumption and capital income are also stationary in first differences but that saving is stationary in its level. This is because the theory restricts saving to equal the discounted present value of expected changes in labor income; these changes are stationary and thus so is saving.

More formally, define the vector $x_t = [y_{kt}, y_{\ell t}, c_t]'$. Each of the elements of x_t is stationary in first differences but a linear combination of the elements, $s_t = [1\ 1\ -1]x_t$ is stationary in its level. The vector x_t is said to be cointegrated.

Definition (Granger and Engle, 1987). A vector x_t is said to be cointegrated of order d,b denoted x_t CI(d,b) if (i) all components of x_t are integrated of order d (stationary in d'th differences), and (ii) there exists at least one vector $\alpha(\neq 0)$ such that $z_t = \alpha' x_t$ is integrated of order d - b, b > 0.

The vector α is called the cointegrating vector; it is unique up to a scalar normalization and, in the present example, is proportional to [1 1 -1]'. Stock (1987) proves that if there is a single unknown element of α, a variety of methods provide estimates with a standard error which goes to zero at a rate proportional to the sample size T (rather than $\sqrt{T}$). Intuitively, all linear combinations of the elements of x_t other than $\alpha' x_t$ have infinite variance because variables in a CI(1,1) vector share a common stochastic trend (a unit root) while exhibiting stationary deviations from one another in the short run. The practical implication of this result is that an unknown element of α may be estimated in a first-stage regression and then treated as known in second stage procedures, whose asymptotic standard errors will still be correct.

If the propensity to consume is known -- or assumed -- to equal unity and the economically relevant measure of consumption is observable, all elements of α are known _a priori_. However, in the context of the permanent income hypothesis, the relevant measure of consumption is consumption of nondurables, services, and the services yielded by the existing stock of durable goods. The latter is unobservable, and in what follows we shall postulate that consumption, c_t, is proportional to the observed consumption of nondurables and services, c_{nt}, so that $c_t = \lambda c_{nt}$. In this case, the vector $x_{nt} = [y_{kt}\ y_{\ell t}\ c_{nt}]'$ is CI(1,1) and the scale factor λ can be estimated from the cointegrating vector $[1\ 1\ -\lambda]$ using Stock's theorem.

If the permanent income hypothesis holds with $\gamma < 1$, then from equations (2) and (5), both y_{kt} and c_t are explosive rather than stationary in first differences and the vector x_{nt} no longer satisfies the formal definition of cointegration. However, x_{nt} still exhibits the key property that a linear combination of its elements, $s_{nt} = [1\ 1\ -\lambda/\gamma]x_{nt}$, is stationary. This follows from equations (1) and (2) and the assumption that $c_t = \lambda c_{nt}$. The linear combination s_{nt} can still be estimated precisely in a first stage regression since it is the only linear combination with asymptotically finite variance. Note, however, that with data only on c_{nt}, the cointegrating vector identifies only the ratio λ/γ. In the discussion which follows, we shall refer to $s_{nt} = y_t - (\lambda/\gamma)c_{nt}$ as saving. If in fact the theory holds with $\gamma < 1$, s_{nt} equals the difference between current saving and the fraction $(1-\gamma)$ of permanent income; i.e., $s_{nt} = \tilde{s}_t = s_t - (1-\gamma)y_{pt}$.

As we have seen the permanent income hypothesis implies that saving is equal to the discounted present value of expected future declines in labor income. Campbell (1987) shows that this can be tested as a set of cross-equation restrictions on a bivariate VAR comprised of saving and changes in disposable labor income. The test of restrictions on the VAR is, in fact, equivalent to a single-equation regression test of equation (4), but the VAR easily generates an optimal unrestricted forecast of future changes in labor income. Because the VAR includes saving as a variable, the unrestricted forecast should equal saving if the model is true; this can be used to characterize informally the fit of the permanent income hypothesis.

Consider the following VAR:

$$(7) \qquad \begin{bmatrix} \Delta y_{\ell t} \\ s_{nt} \end{bmatrix} = \begin{bmatrix} a(L) & b(L) \\ c(L) & d(L) \end{bmatrix} \begin{bmatrix} \Delta y_{\ell,t-1} \\ s_{n,t-1} \end{bmatrix} + \begin{bmatrix} u_{1t} \\ u_{2t} \end{bmatrix},$$

which can be rewritten in "companion" form as $z_t = Az_{t-1} + v_t$. Note that, for all i, $E[z_{t+1}|H_t] = A^i z_t$, where H_t is the information set $\{z_t, z_{t-1}, \ldots\}$, a proper subset of agents' information set I_t. As is common in empirical research, we take conditional expectations to be linear projections on information.

If the theory is correct, s_{nt} is an optimal forecast of future changes in income conditional on the full information set I_t. A weak implication is that s_{nt} will have incremental explanatory power for future labor income changes if agents have information useful for forecasting labor income beyond the history of that variable. s_{nt} must Granger cause $\Delta y_{\ell t}$, unless agents have no useful information beyond the history of labor income, in which case s_{nt} is an exact linear function of current and lagged changes in labor income.

Projecting equation (3) onto the information set H_t, and noting that the left hand side is unchanged because s_{nt} is in H_t, we obtain the following set of cross-equation restrictions on the matrix A.

$$(8) \qquad g' = -\sum_{i=1}^{\infty} [1/(1+r)^i h'A^i]$$

where g' and h' are row vectors with 2p elements, all of which are zero except for the $(p+1')^{st}$ element of g' and the first element of h'. These non-linear cross-equation restrictions are in fact equivalent to the restriction that the linear combination of the change in disposable labor income, current, and lagged saving derived in equation (4) is orthogonal to lagged $\Delta y_{\ell t}$ and s_{nt}.

To see this, note that the right hand side of equation (8) can be expressed as $-h'[1/(1+r)]A\{I - [1/(1+r)]A\}^{-1}$. Postmultiplying equation (8) by $\{I - [1/(1+r)]A\}$, we obtain

$$(9) \qquad g'\{I - [1/(1+r)]A\} = -h'[1/(1+r)]A.$$

Using the structure of the A matrix, the cross-equation restrictions defined by equation (9) can be written as follows: $a_1 = c_1, \ldots, a_p = c_p$, $d_1 - b_1 = (1+r)$, $b_2 = d_2, \ldots, b_p = d_p$. Subtracting the $\Delta y_{\ell t}$ equation of the VAR from the s_{nt} equation, we obtain $s_{nt} - \Delta y_{\ell t} = (c_1-a_1)\Delta y_{\ell,t-1} + \ldots + (c_p-a_p)\Delta y_{\ell,t-p} + (d_1-b_1)s_{n,t-1} + (d_2-b_2)s_{n,t-2} + \ldots + (d_p-b_p)s_{n,t-p} + u_{1t} - u_{2t}$. Thus, if the theory is true, $s_{nt} - \Delta y_{\ell t} - (1+r)s_{n,t-1}$ is orthogonal to lagged $\Delta y_{\ell t}$ and s_{nt}. The implication is that a single-equation regression test of equation (4) is equivalent to a test of (9). However, the VAR can be used to generate optimal unrestricted forecasts (conditional on lagged $\Delta y_{\ell t}$ and s_{nt}) of the discounted present value of future changes in labor income. These forecasts can be compared with s_{nt} to characterize the fit of the permanent income model, since the theory predicts that the unrestricted forecast and s_{nt} should have identical standard deviations and should be perfectly correlated.

A weaker version of the permanent income hypothesis which allows for a "transitory consumption" error in equation (2) can also be tested in this framework, provided that the error is assumed to be orthogonal to all lagged information. In this case, the theory restricts the conditional expectation of saving, one period ahead, to equal the discounted present value of expected future declines in labor income

$$(3') \qquad E_t s_{n,t+1} = -\sum_{i=1}^{\infty} [1/(1+r)]^i E_t \Delta y_{\ell,t+1+i},$$

since by definition, the conditional expectation of next period's transitory consumption is zero. Equation (3') can be tested by regressing $s_{nt} - \Delta y_{\ell t} - (1+r)s_{n,t-1}$ on twice lagged $\Delta y_{\ell t}$ and s_{nt}.

There are at least two methods available for estimating the cointegrating vector $[1\ 1\ -\lambda/\gamma]$. The first is a "levels regression" of total income y_t on non-durables and services consumption c_{nt}, while the second is an "error-correction regression" of Δy_t on lagged changes in and levels of y_t and c_{nt}. In the levels regression, the estimate of λ/γ is the coefficient on c_{nt}, while in the error-correction regression the estimate is given by the ratio of the coefficient on lagged consumption to that on lagged income. The residual from the levels regression can be used to test the hypothesis that c_{nt} and y_t are not cointegrated. In particular, Granger and Engle (1987) develop an "Augmented Dickey-Fuller" test in which the change in the residual is regressed on one lagged level of itself and at least one lagged change. Based upon the Monte Carlo work of Granger and Engle, a t-statistic on the coefficient of the lagged level exceeding 2.84 is sufficient to reject the hypothesis of no cointegration at the 10 percent level; a t-statistic exceeding 3.17 is sufficient to reject at the 5 percent level.

4. Data and Empirical Results

Quarterly data on real per capita consumption, disposable income, and disposable labor income in Canada and Britain are used in the empirical work. All data are seasonally adjusted quarterly series for the period 1955:1-1984:4. Canadian data are from Statistics Canada, while the British data are from the Central Statistical Office. Following Blinder and Deaton (1985), the disposable labor income series are constructed as follows. Proprietor's income and personal income taxes are attributed to labor and capital according to their overall factor shares; social income contributions are deducted from labor income; obviously, income from dividends, rent, and interest is not included in the disposable labor income series. Nominal, aggregate magnitudes are converted to a real, per capita basis by dividing by total population and the consumer spending deflator for each country. Annual population figures are from International Financial Statistics; quarterly population series are constructed by interpolation.

As a preliminary diagnostic, in Table 1 we test for a unit root in the disposable labor income process in Canada and the U.K. The VAR methodology described in the previous section relies on the presence of a unit root in labor income. We also test for unit roots in total income y_t, nondurables and services consumption c_{nt}, and "saving" $s_{nt} = y_t - (\lambda/\gamma)c_{nt}$ where the parameter λ/γ is estimated in Table 2. We expect to be unable to reject the unit root hypothesis for y_t and c_{nt} (although strictly speaking these variables have explosive rather than unit roots if $\gamma < 1$), and to reject strongly for s_{nt} (in fact, the test in Table 1 is biased towards rejection because it ignores the fact that λ/γ must be estimated -- a test which takes this into account is presented in Table 2).

The test statistics in Table 1 have recently been proposed by Phillips (1986) and Phillips and Perron (1986). To test the null hypothesis that a series X_t has a unit root (perhaps with drift), against the alternative that it is stationary around a linear trend, one runs the regression $\Delta X_t = \mu + \beta t + \alpha X_{t-1} + \epsilon_{Xt}$. Fuller (1976) and Dickey and Fuller (1981) tabulated critical values for the t statistic on α, $t\tilde{\alpha}$, and the F statistic testing (H_0: α=0, β=0), Φ_3, but these are correct only if ϵ_{Xt} is serially uncorrelated. We present Phillips and Perron's modified statistics, $Zt\tilde{\alpha}$ and $Z\Phi_3$, which make a nonparametric correction for serial correlation in ϵ_{Xt}.

The results in Table 1 are generally consistent with our prior expectations. For Canada, there is no evidence at even the 10 percent level against the hypothesis that $y_{\ell t}$, y_t and c_{nt} have unit roots; but the hypothesis that s_{nt} has a unit root is rejected at the 1 percent level. Results for the U.K. are similar except that the unit root hypothesis for y_t is rejected at the 10 percent level by the statistic $Z\Phi_3$.

Table 2 presents estimates of λ/γ and the tests for no cointegration. In the Canadian data, the parameter λ/γ is estimated at 1.698 by the levels regression and 1.720 by the error-correction regression; in the British data, λ/γ is estimated at 1.801 by the levels regression and 1.802 in the error-correction regression. Recall that, since the consumption of services given by durables is unobservable, we postulate that $c_t = \lambda c_{nt}$. While the cointegrating regressions do not identify the λ parameter, we note that, if the propensity to consume out of permanent income is assumed to be unity, the estimates of λ/γ imply a share of nondurables and services in total consumption of 59 percent in Canada and 55 percent in Britain. These implied shares seem somewhat low, suggesting that the value of γ is less than unity. Evidence on the value of γ, along with an assessment of the robustness of our findings, is presented in the next section. Finally, we note that the hypothesis of no cointegration is rejected at the 5 percent level in the Canadian data, while in the British data it can only be rejected at the 10 percent level.

We next construct time series for saving, $s_{nt} = y_t - (\lambda/\gamma)c_{nt}$, in each country, estimate the bivariate VARS given by equation (7), and use the estimates to evaluate the permanent income hypothesis. The Akaike Information

Criterion is used to select the VAR lag length: four lags are selected for the British VAR, and one lag is selected for the Canadian VAR. (The robustness of the results to alternative choices of lag length is discussed in the next section.) In computing standard errors, we allow for conditional heteroskedasticity by using White's (1984) heteroskedasticity-consistent estimate of the variance-covariance matrix of the VAR coefficient estimates. This is given by $(X'X)^{-1}X'VX(X'X)^{-1}$, where V is a diagonal matrix with squared residuals on the diagonal.

Tables 3 and 4 present the empirical results for Canada and Britain respectively. In each table there are four columns of regression coefficients (sums of coefficients in the British case, to reduce the complexity of the table). Column (1) reports the regression of $\Delta y_{\ell t}$ on the lagged change in labor income and lagged saving. Column (2) reports the regression of s_{nt} on these variables. The two columns together make up the VAR system. Column (3) is the regression of $s_{nt} - \Delta y_{\ell t} - (1+r)s_{n,t-1}$ on the VAR explanatory variables; recall that, if the permanent income hypothesis is true, all coefficients should be zero in this regression. Column (4) is the regression with $s_{n,t+1} - \Delta y_{\ell,t+1} - (1+r)s_{nt}$ as the dependent variable; this tests the hypothesis that the permanent income model holds except for serially uncorrelated transitory consumption. A fifth column reports the coefficients of the optimal unrestricted VAR forecast of the present value of labor income declines, which we write $s'_{n,t-1}$. If the permanent income hypothesis is true, we should have $s'_{n,t-1} = s_{n,t-1}$.

In both Canada and the United Kingdom, the full set of restrictions of the permanent income hypothesis is strongly rejected. In each country, the coefficients in column (3) are jointly significant at the 0.004 percent level or better. For Canada, the strength of the rejection is due largely to the restriction that the mean of s_{nt} should equal (-1/r) times the mean change in labor income, a result which follows immediately from equation (3). When $\gamma < 1$, $s_{nt} = s_t - (1-\gamma)y_t^P$ so it is possible to have s_{nt} be negative even though s_t is positive; nevertheless, the mean of s_{nt} is too high to satisfy the model's restrictions. When the mean restriction is dropped, the other restrictions of the model are rejected at only the 1.7 percent level. In Britain, however, the dynamic restrictions of the model are rejected just as strongly as the mean restriction.

Allowing serially uncorrelated transitory consumption also helps the model fit the Canadian data; the coefficients in column (4), excluding the intercept, are not even significant at the 60 percent level for Canada. But for the U.K., the model with transitory consumption is rejected as strongly as the model without.

As discussed in the Introduction, our goal is to do more than simply conduct formal tests of the permanent income model. We are interested in characterizing the fit of the model, bringing out its strengths and its weaknesses. Tables 3 and 4 also present summary statistics for the VAR systems which help us to do this.

In both Canadian and British data, saving Granger causes changes in labor income at extremely high levels of statistical significance. Furthermore, the estimated coefficients on $s_{n,t-1}$ are negative in both countries and statistically significant. These findings are exactly those implied by the permanent income hypothesis since saving is forward looking and rises in anticipation of future declines in labor income. As shown in detail in Section 3, the single-equation regression tests already discussed are equivalent to the test of the cross-equation restrictions on the VARs. In fact, the deviations of the estimated VAR coefficients from these restrictions are just the coefficients reported in the regressions of column (3).

We also report, for each country, the standard deviation of saving, the ratio of the standard deviation of the unrestricted VAR forecast of the discounted present value of future changes in labor income to the standard deviation of saving, and the correlation between saving and the unrestricted forecast. In the Canadian data, the correlation is .997 and the standard deviation ratio is 1.079. If the theory were exactly correct, the correlation and the standard deviation ratio would both equal one. We have seen that these differences are statistically significant, but Canadian household saving appears to incorporate virtually all of the forecastable variation in future labor income.

In the British data, the correlation between saving and the unrestricted forecast is also quite high at .896. However s_{nt} is less than half as volatile as the unrestricted forecast of the present value of the future labor income declines. This may be taken as evidence of excess sensitivity of consumption to current income in Britain since, if consumption were in fact determined by current as opposed to permanent income, saving -- the difference between income and consumption -- would be expected to have substantially less volatility than the unrestricted optimal forecast. It should be noted, though, that the asymptotic standard error on the excess sensitivity statistic is high at 0.679.

We conclude this section by plotting s_{nt} and the unrestricted optimal forecast of labor income declines. These plots are given in Figures 1 and 2. The figures convincingly support the inferences drawn from the summary statistics. In particular, the statistical rejection of the permanent income hypothesis in Canadian data does not appear to have substantial economic significance. Forecastable variations in the present value of future labor income declines are incorporated virtually one-for-one in Canadian household saving behavior (Figure 1). The noteworthy feature of the British plot (Figure 2) is that, while the magnitude of swings in saving does not in general match that of swings in forecastable labor income declines, the former tracks every turning point in the latter. Despite the statistical rejections of the present income hypothesis reported in Table 2, the theory seems to be a reasonable first approximation to the behavior of household saving in Canada and Britain.

5. How Robust are the Findings?

In this section we check to see whether our results are robust to changes in the econometric specification and the measure of consumption. We are particularly concerned with the possibility that our results are sensitive to lag length, since Campbell (1987) found this to be the case for U.S. data.

Fortunately, in British and Canadian data we obtain rather similar results to those reported for all lag lengths between 1 and 5. In Canadian data, the permanent income hypothesis (without transitory consumption) is rejected at the 2.1 percent level for lag length 2, the 1.1 percent level for lag length 3, the 0.3 percent level for lag length 4, and the 0.4 percent level for lag length 5. The correlation of s_{nt} and s'_{nt} falls slightly from 0.996 with lag length 1 to 0.908 with lag length 5, and the standard deviation ratio rises to 1.518 with lag length 5, but the standard errors on these numbers also rise so that individually they remain insignificantly different from one. In British data the model is strongly rejected but the correlation of s_{nt} and s'_{nt} remains above 0.85 for all lag lengths. We find strong evidence of excess sensitivity only at the lag lengths above 3.

The assumption that the unobservable consumption of services yielded by the existing stock of durable goods is proportional to c_{nt}, the consumption of non-durables and services, involves a potential specification error. To evaluate the robustness of our findings, we test and evaluate the fit of the permanent income model using data on total consumption expenditures, c_t, to construct a series for saving $\tilde{s}_t = y_t - c_t/\gamma$ in each country.

In this case, the cointegrating regression of y_t on a constant and c_t uniquely identifies the parameter $1/\gamma$. The estimate of $1/\gamma$ for Canada is 1.289 with a standard error of .013; the estimate for Britain is 1.299 with a standard error of .012. The null hypothesis of no cointegration is rejected at the 10 percent level in Britain, and at the 5 percent level in Canada, as in Table 2.

For Britain, the results of the single equation regression tests are unchanged: the restriction that $\tilde{s}_t - \Delta y_{\ell t} - (1+r)\tilde{s}_{t-1}$ or $\tilde{s}_{t+1} - \Delta y_{\ell,t+1} - (1+r)\tilde{s}_t$ be orthogonal to lagged $\Delta y_{\ell t}$ and $\tilde{s}_t$ is resoundingly rejected. However, for Canada, we cannot reject at the 5 percent level the hypothesis that the source of the rejection is solely a significant constant term in both the strict and transitory-consumption regression tests.

The only substantive differences in the VAR summary statistics relative to those reported in Tables 3 and 4 are that the standard deviations of Canadian saving and the unrestricted forecast increase from .661 and .669, respectively, to .900 and 1.408. However, the correlation between Canadian saving and the unrestricted forecasts falls only slightly to .996 from .997. The corresponding British correlation rises to .913 (from .896) while the standard deviation of the unrestricted British forecast declines to .154 from .221.

Figures 3 and 4 give a visual impression of the results using total consumption expenditure. The conclusions drawn in Section 4 are not weakened when total consumption is used. If anything, the British plots in Figure 4 provide stronger support for the proposition that the permanent income hypothesis provides a parsimonious and empirically relevant account of the cyclical dynamics of household saving behavior.

6. Concluding Remarks

In this paper, we have found substantial support for the prediction of the permanent income hypothesis that forecastable variations in disposable labor income are incorporated in household saving behavior in Canada and Britain. The tight formulation of the permanent income hypothesis tested in this paper can be statistically rejected, but we conclude that the theory has surprising empirical content. Because it abstracts from the demographic considerations of the life-cycle hypothesis, the permanent income hypothesis is not successful at explaining average saving rates, or differences in these rates across countries. However, the theory's predictions about the dynamics of saving and income appear to be worth taking seriously.

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Table 1: Univariate Tests for Unit Roots

	Country/Test Statistic			
	Canada		U.K.	
Variable	$Zt\tilde{\alpha}$	$Z\Phi_3$	$Zt\tilde{\alpha}$	$Z\Phi_3$
$y_{\ell t}$	-1.452	1.057	-2.553	3.271
y_t	-1.816	1.848	-3.303 (10%)	5.472 (10%)
c_{nt}	-1.407	1.409	-2.453	3.199
s_{nt}	-4.357 (1%)	9.505 (1%)	-6.182 (1%)	19.145 (1%)

Notes: These test statistics are from Phillips and Perron (1986) and Perron (1986). $Zt\tilde{\alpha}$ is formed from the t statistic on α in the regression $\Delta X_t = \mu + \beta t + \alpha X_{t-1}$. $Z\Phi_3$ is formed from the F statistic for H_0: (β=0, α=0) in this regression. All statistics are corrected for serial correlation in the equation error using a 4th-order Newey-West (1987) correction. Asymptotic critical values, from Fuller (1976) and Dickey and Fuller (1981), are as follows: $Zt\tilde{\alpha}$ 1% -3.96, 2.5% -3.66, 5% -3.41, 10% -3.12; $Z\Phi_3$, 1% 8.27, 2.5% 7.16, 5% 6.25, 10% 5.34.

Table 2: Estimation of λ/γ and Cointegration Tests

- Canada -

(1) $y_t = \underset{(0.223)}{-3.241} + \underset{(0.012)}{1.698}\, c_{nt}$

$R^2 = 0.995$
estimate of λ/γ = 1.698

(2) $\Delta y_t = \underset{(0.283)}{-0.758} + \underset{(0.301)}{0.764}\, \Delta y_{t-1} - \underset{(0.097)}{0.455}\, \Delta c_{n,t-1}$

$R^2 = 0.197$
estimate of λ/γ = 1.720

$\underset{(0.072)}{-0.263}\, y_{t-1} + \underset{(0.122)}{0.452}\, c_{n,t-1}$

Augmented Dickey-Fuller test with 1 lag 3.217*; with 5 lags 3.365*

- U.K. -

(1) $y_t = \underset{(0.069)}{-2.107} + \underset{(0.016)}{1.801}\, c_{nt}$

$R^2 = 0.991$
estimate of λ/γ = 1.801

(2) $\Delta y_t = \underset{(0.221)}{-0.952} + \underset{(0.268)}{1.105}\, \Delta y_t - \underset{(0.102)}{0.536}\, \Delta c_{n,t-1}$

$R^2 = 0.249$
estimate of λ/γ = 1.802

$\underset{(0.101)}{-0.463}\, y_{t-1} + \underset{(0.182)}{0.835}\, c_{n,t-1}$

Augmented Dickey-Fuller test with 1 lag 3.812*; with 5 lags 2.073

Note: The Granger and Engle (1986) critical values for the null hypothesis of no cointegration are 3.17 at the 5 percent level (*) and 2.84 at the 10 percent level (**).

Table 3: Tests of the Permanent Income Hypothesis -- Canada

	Regression coefficients of column variables on row variables:				
	(1)	(2)	(3)	(4)	(5)
	$\Delta y_{\ell t}$	s_{nt}	$s_{nt}-\Delta y_{\ell t}$ $-(1+r)s_{n,t-1}$	$s_{n,t+1}-\Delta y_{\ell,t+1}$ $-(1+r)s_{nt}$	$s'_{n,t-1}$
Constant	--	--	-0.139 (0.158)	-0.044 (0.163)	
$\Delta y_{\ell,t-1}$	-0.207 (0.143)	-0.339 (0.143)	-0.132 (0.049)	-0.052 (0.054)	-0.135 (0.115)
$s_{n,t-1}$	-0.236 (0.064)	0.765 (0.070)	-0.008 (0.047)	0.024 (0.048)	1.096 (0.313)
R^2	0.196	0.565	0.042	0.005	

s_{nt} Granger causes $\Delta y_{\ell t}$ at 0.02 percent level in column (1).
$\Delta y_{\ell t}$ Granger causes s_{nt} at 1.8 percent level in column (2).

Coefficients are jointly significant at 1.6×10^{-7} percent level in column (3). Coefficients, excluding the intercept, are jointly significant at 1.7 percent level in column (3).

Coefficients are jointly significant at 2.9×10^{-4} percent level in column (4). Coefficients, excluding the intercept, are jointly significant at 60.8 percent level in column (4).

Summary statistics:

$\sigma(s_{nt}) = 0.619$

$\sigma(s'_{nt})/\sigma(s_{nt}) = 1.079$ (0.302)

$\rho(s'_{nt}, s_{nt}) = 0.997$ (0.004)

Table 4: Tests of the Permanent Income Hypothesis -- U.K.

Sums of regression coefficients of column variables on row variables:

	(1) $\Delta y_{\ell t}$	(2) s_{nt}	(3) $s_{nt}-\Delta y_{\ell t}$ $-(1+r)s_{n,t-1}$	(4) $s_{n,t+1}-\Delta y_{\ell,t+1}$ $-(1+r)s_{nt}$	(5) $s'_{n,t-1}$
Constant	--	--	0.260 (0.148)	0.377 (0.178)	--
$\sum_{i=1}^{4} \Delta y_{\ell,t-i}$	0.256 (0.226)	0.445 (0.256)	0.189 (0.195)	-0.023 (0.235)	0.513 (0.555)
$\sum_{i=1}^{4} s_{n,t-i}$	-0.274 (0.097)	0.863 (0.097)	0.128 (0.070)	0.181 (0.085)	2.704 (0.898)
R^2	0.201	0.448	0.370	0.184	

s_{nt} Granger causes $\Delta y_{\ell t}$ at 0.007 percent level in column (1).
$\Delta y_{\ell t}$ Granger causes s_{nt} at 4.9 percent level in column (2).

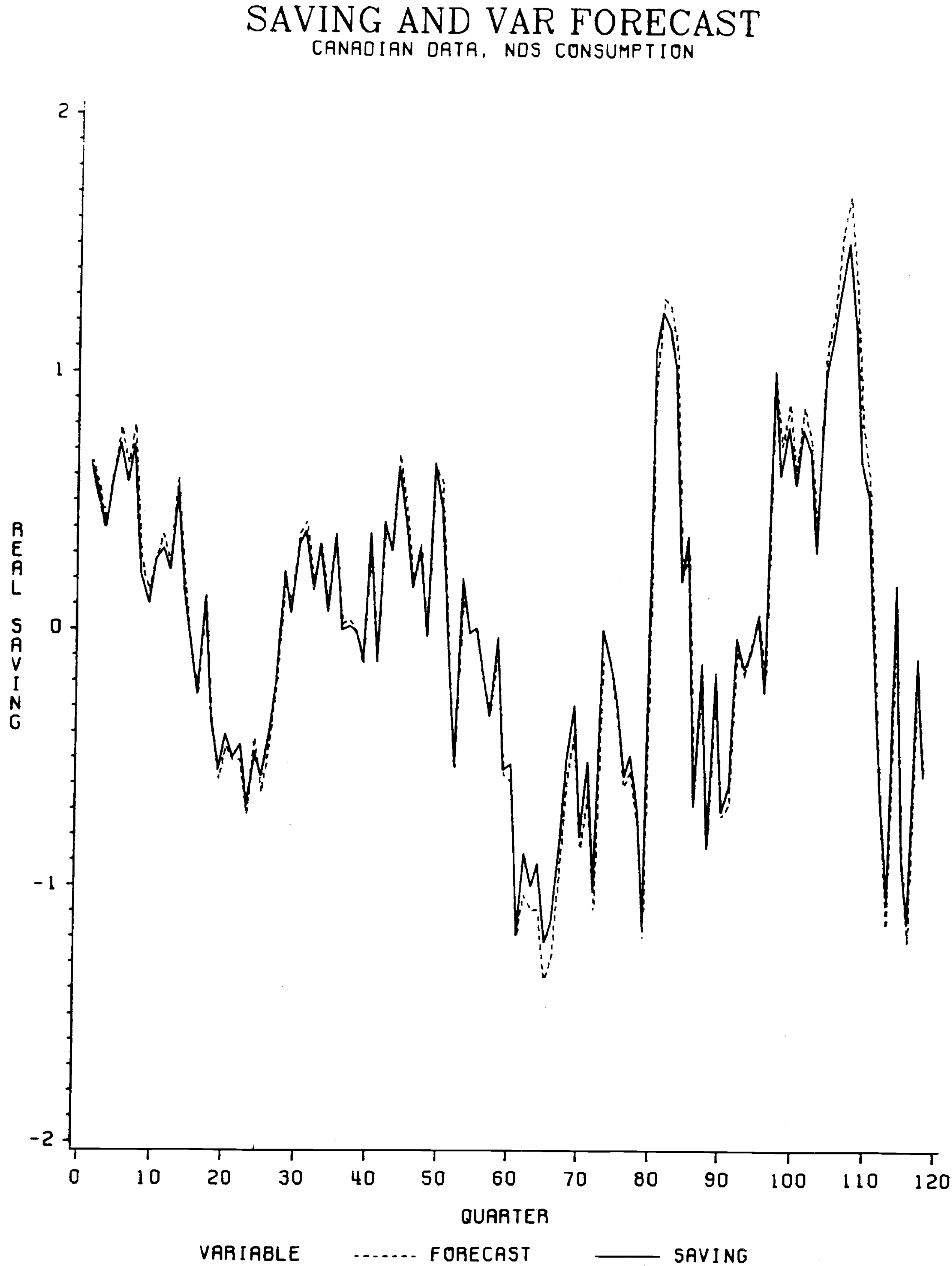
Coefficients are jointly significant at 0.004 percent level in column (3).
Coefficients, excluding the intercept, are jointly significant at 0.002 percent level in column (3).

Coefficients are jointly significant at 4.5×10^{-4} percent level in column (4).
Coefficients, excluding the intercept, are jointly significant at 2.0×10^{-4} percent level in column (4).

Summary statistics:

$$\sigma(s_{nt}) = 0.105$$
$$\sigma(s'_{nt})/\sigma(s_{nt}) = 2.114 \quad (0.679)$$
$$\rho(s'_{nt}, s_{nt}) = 0.896 \quad (0.023)$$

Figure 1



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-2
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0
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20
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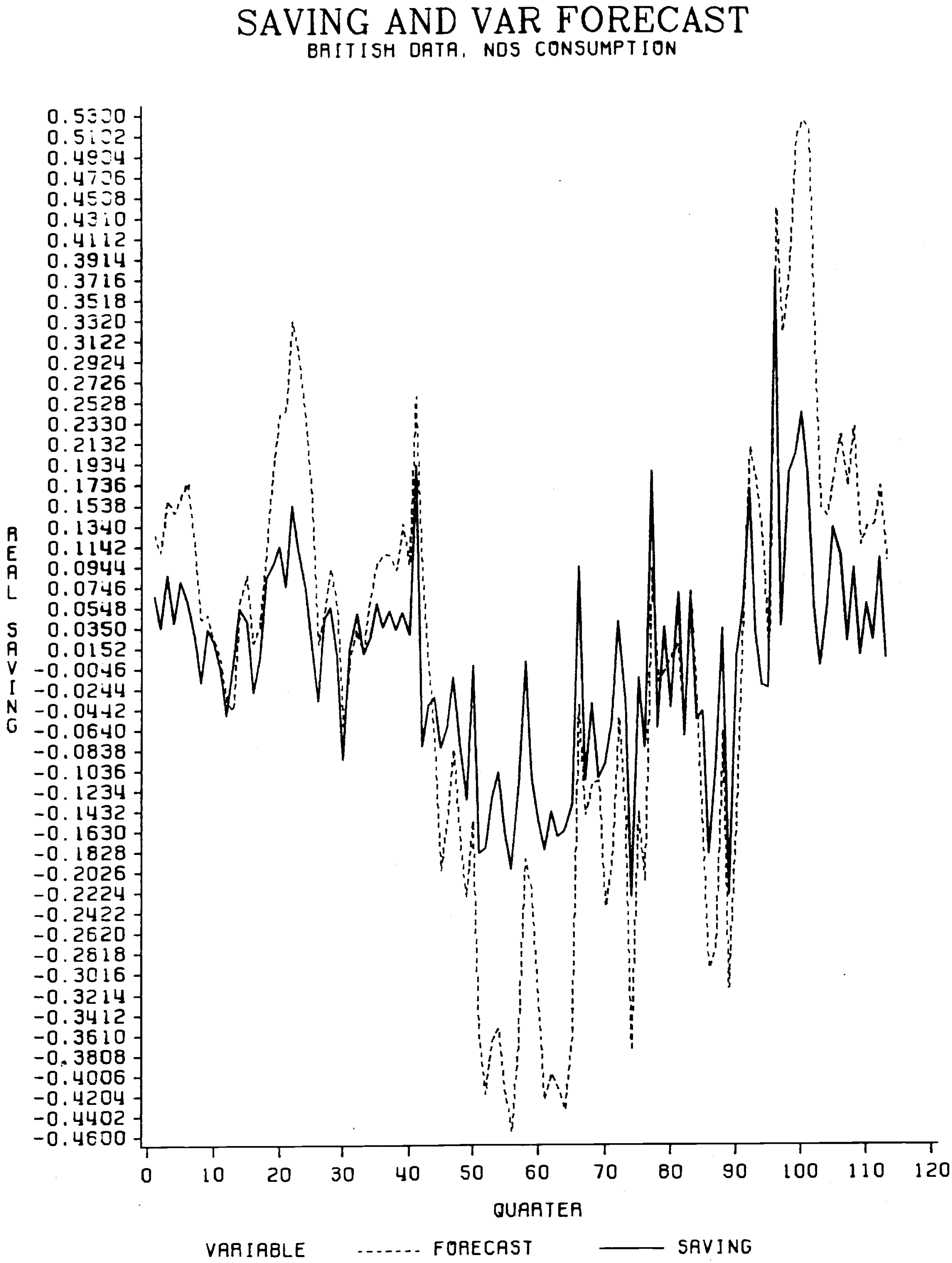
40
50
60
70
80
90
100
110
120
QUARTER
VARIABLE
FORECAST
SAVING

Figure 2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0.3716
0.3518
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0.3122
0.2924
0.2726
0.2528
0.2330
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0.1142
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0.0746
0.0548
0.0350
0.0152
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-0.0244
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-0.1234
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-0.1630
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-0.4600
0
10
20
30
40
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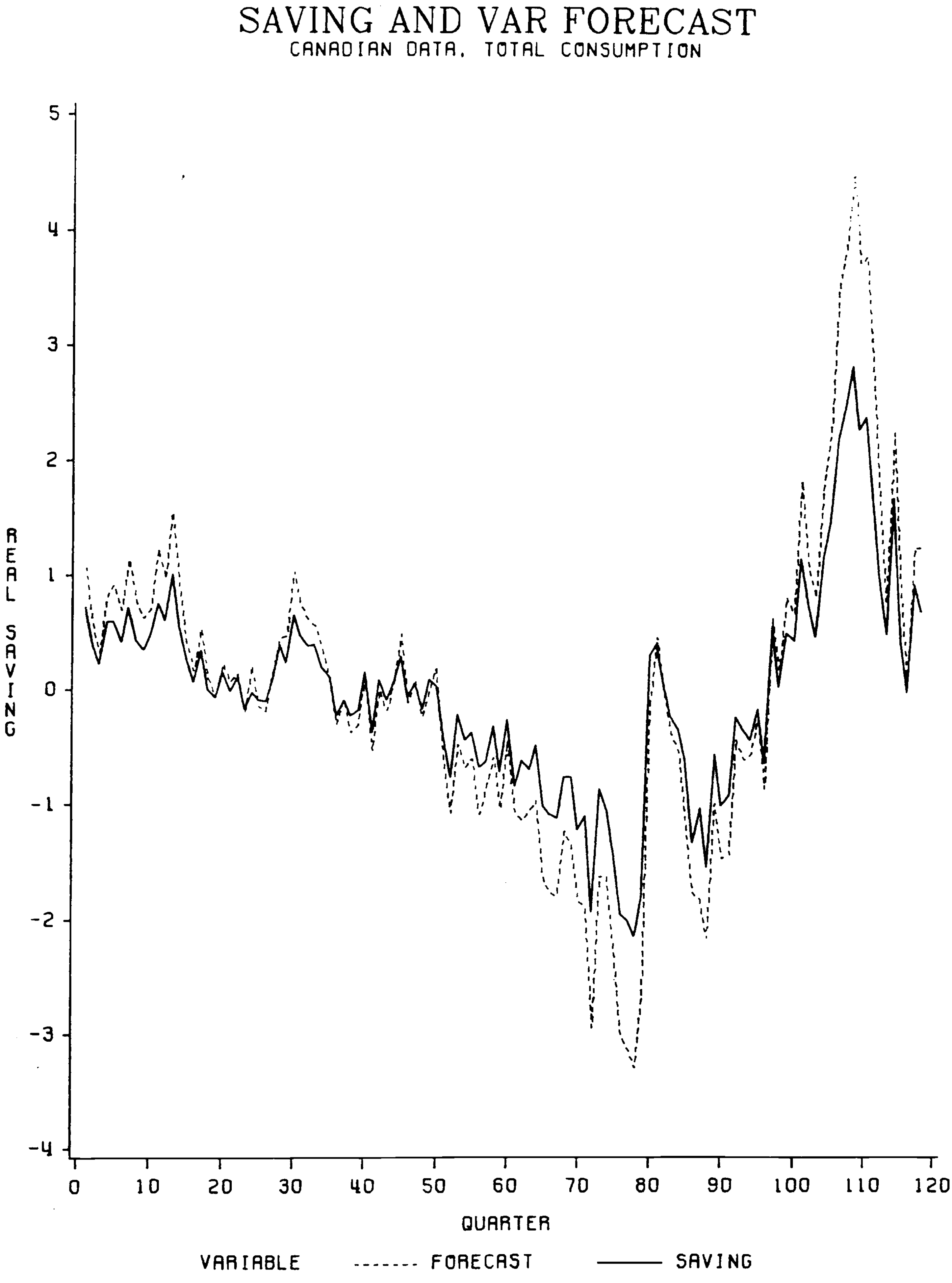
60
70
80
90
100
110
120
QUARTER
VARIABLE
FORECAST
SAVING

Figure 3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-4
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60
70
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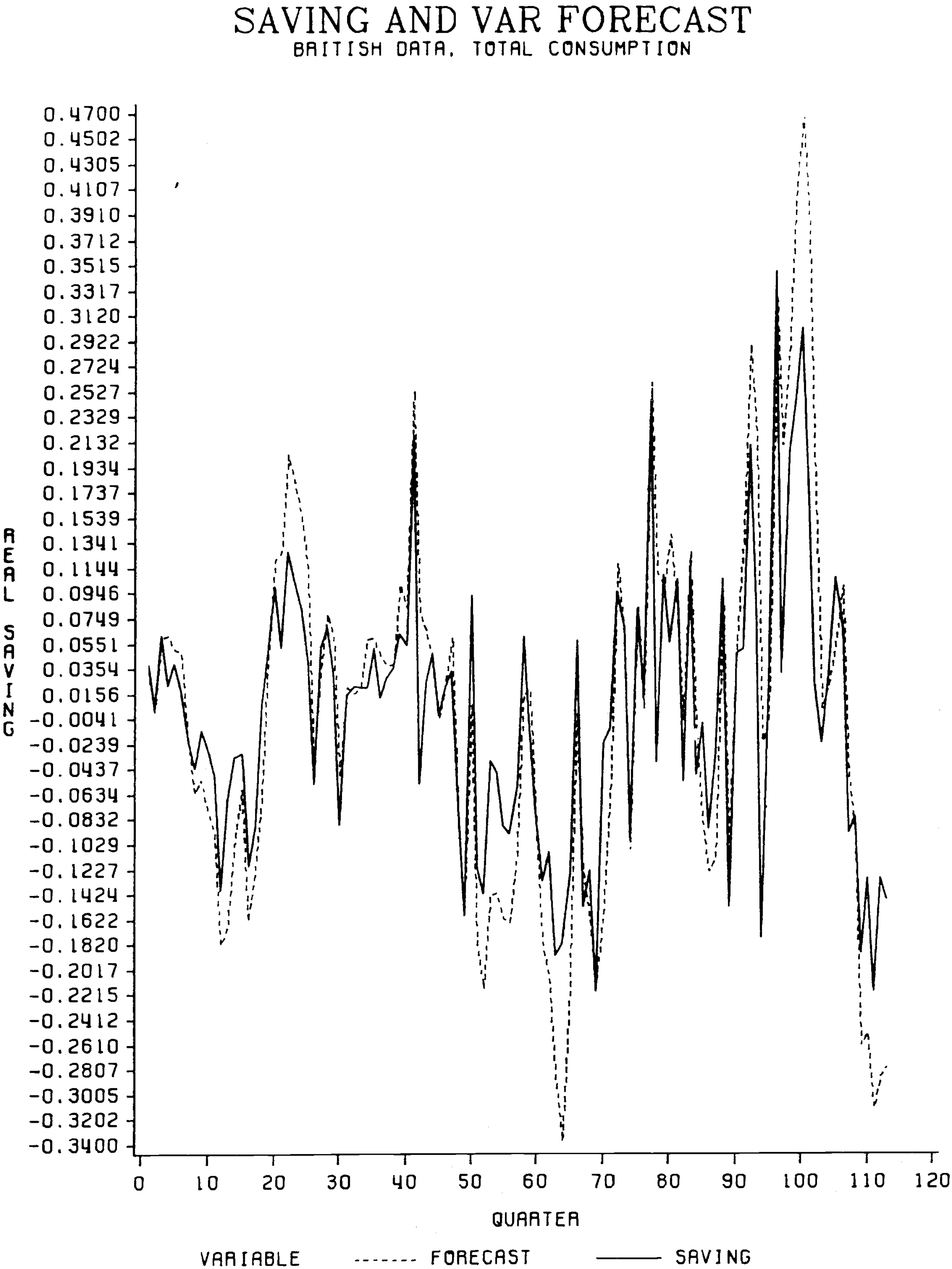
120
QUARTER
VARIABLE
FORECAST
SAVING

Figure 4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120
QUARTER
VARIABLE
FORECAST
SAVING