THE DYNAMIC RELATIONSHIP BETWEEN THE STATE PENSION SCHEME AND HOUSEHOLD SAVING IN NEW ZEALAND

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ABSTRACT

The paper employs the cointegrating vector autoregression and autoregressive distributed lag approaches to estimate the long-run equilibrium model and the short-run dynamics of household saving incorporating certainty of retirement income in New Zealand. Of particular concern is the persistent negative saving by households. There seems to be a need to reform the funding mode of the universal tax-funded public pension scheme because of the projected increases in the cost of providing superannuation for an ageing population. The long-run parameters from the two approaches are comparable but the short-run ones show some variation. The long run results indicate that whereas the trend in the household saving rate has been negative, increases in disposable income and gross social wealth boost saving; the introduction of the government-run Super Fund in 2001 has elicited a slight positive response in the saving rate; there is significant propensity to consume out of household net wealth; and inflation and unemployment engender significant precautionary saving. The error correction term takes the expected negative sign in all the models. Cuts to the pension benefits rate would reduce the social security wealth and exacerbate an already abysmal saving rate.

Key words: error correction model, household saving, superannuation, New Zealand.
JEL Classifications: C20, D91, E21, I38, J14, J26.

1. INTRODUCTION

For the provision of retirement[2] income, New Zealand has a universal tax-funded state pension scheme for all residents over 65 years of age who have lived in the country for ten years since age 20, and five of those years must have been since age 50 (St John, 1999; Kritzer, 2007). The scheme is known as the New Zealand Superannuation (NZS). The proportion of the population aged 65 and over has risen from 8.5% in 1972 to 12.6% in 2008 and is projected to reach 25% by 2030. Spending on pension benefits, as in other OECD/high-income countries, is projected to increase on account of a relatively long life expectancy and a low birth rate. From the current level of 3.5% of GDP in 2008/09, pension benefits payments are projected to be 5.6% in 2030 and 6.6% by 2050 (NZSF, 2009). To pay for this, the government will either have to increase taxes or reduce spending. This is a widely recognised fiscal challenge with enormous implications for both economic growth and business cycles and the welfare

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[2] In this paper the terms retirement, pension, superannuation and social security are synonymous and will be used interchangeably.

of different generations. Increased taxation would mean heavier tax burden on those currently working (those in the 15-64 age bracket) and it could also have distortionary effects on the economy. Reducing spending would mean either (1) cutting benefits or (2) making people take a greater role in caring for themselves by either increasing the retirement age or increasing incentives for people to save more during their working years. When it is considered that for 50 to 60 percent of retirees the superannuation is their only income (St John, 1999) and that since 1994 saving by households out of their disposable income has been negative, the issue of retirees' welfare takes on increasing importance over time. It is inevitable that the 'pay-as-you-go' (PAYGO) system would be reformed. The key macroeconomic issue is not so much whether the state will pay pensions, but how the pensions will be funded. Ultimately, however, it is an issue about saving.

In response to the concerns about the impending increases in the cost of running the NZS, the New Zealand Superannuation and Retirement Act was enacted in 2001, establishing the New Zealand Superannuation Fund (also known as the 'Super Fund'). The NZS Fund is an investment fund that accumulates and invests government contributions paid out of general taxes which, in later years, will progressively be drawn on to supplement the superannuation expenses. The Super Fund is meant to provide a smoothing mechanism (or 'buffer fund') for the current PAYGO system and inject some certainty in the durability of the NZS. In addition, to encourage workers to save more, the government introduced the 'KiwiSaver' scheme in July 2007. The KiwiSaver is a defined contribution retirement savings plan created to supplement NZS and help increase an individual's retirement income (Kritzer, 2007).

The foregoing discussion leads to a revisiting of the relationship between public pension schemes and private saving. Whereas some empirical work accounting for certainty of retirement income has been done in countries such as Australia, Canada and the US, a comparable study has not been done on New Zealand. The paper seeks to fill that void by estimating both a long-run equilibrium model and a short-run dynamic model of the household saving function that takes account of the certainty derived from the presence of the NZS and the creation of the buffer fund for New Zealand. After reviewing the theoretical and empirical studies on saving, a number of variables were identified. The final set of variables was dictated by data availability in the New Zealand context. For robustness, the two alternative approaches to estimate error correction models (the cointegrating vector autoregression, CVAR, and autoregressive distributed lag, ARDL, approaches) were used to estimate the long-run and short-run parameters. The long-run results from the two approaches are comparable but there is some variation in the short-run results. Increases in the gross social wealth and the establishment of the Super Fund have had positive influences on household saving although the saving rate continues to trend downwards. The rest of the paper is organised as follows: Section 2 reviews the theoretical and empirical literature; Section 3 gives a short history of the NZS; Section 4 describes the data and the analytical methods employed; Section 5 covers the analytical results; and Section 6 concludes the paper.

2. LITERATURE REVIEW

Most models of saving behaviour are based on the hybrid Life Cycle-Permanent Income Model (LCPIM) of consumption[3] duly extended to capture other important factors such as uncertainty and liquidity constraints. The underlying theme of the LCPIM is inter-temporal choice and the micro-foundations depict an economic agent who maximises lifetime utility from consumption subject to lifetime budget constraint.

It is useful to comment here on the Life Cycle Hypothesis and the Permanent Income Hypothesis separately. The Life Cycle Hypothesis asserts that individuals save and accumulate wealth during the working phase of their life cycle and dis-save or run down their wealth during their retirement period. Current consumption (conversely, saving) is a function of current real wealth (W) and the present value of labour income (Y); *certainty* about future income is assumed. The Permanent Income Hypothesis partitions current income (Y_t) into transitory (Y^T) and permanent (Y^P) components and asserts that consumption is a function of permanent income which is the individual's expectation of their lifetime incomes. Current income can be equal to, greater than or less than permanent income. The time pattern of current income is not important to consumption but it is critical to saving. People save when current income exceeds permanent income ($Y_t > Y^P$) and dis-save when current income falls below permanent income ($Y_t < Y^P$), thus they use saving and borrowing to smooth the path of consumption. Individuals can borrow at the same interest rate at which they can save, as long as they eventually repay their loans. In other words, there is perfect substitutability among all forms of saving and there are no liquidity constraints. According to the Permanent Income Hypothesis, 'saving is future consumption' and the instantaneous utility function is usually assumed to be *quadratic*. The hybrid LCPIM retains the wealth variable of the life-cycle model and utilises permanent income as the income variable.

The certainty assumption and quadratic utility function underlying the LCPIM have been found to be inconsistent with consumption or saving behaviour and give incorrect predictions. Leland (1968) has demonstrated that the combination of a positive third derivative of the utility function and uncertainty about future income reduces current consumption and thus raises saving. This is known as precautionary saving (Romer, 2006 p. 372). It comes under the heading of buffer stock saving behaviour that Deaton (1991) asserts is exhibited by most households. When households face borrowing constraints, they are forced to consume less than they otherwise would, imposing a forced saving of sorts. Extensions of the LCPIM to address those weaknesses have therefore attempted to incorporate variables to capture uncertainty, precautionary motives, liquidity constraints and different forms of saving. For our modelling purposes, we would suggest that the determinants of household saving considered in the extensive literature can be divided into three broad categories. Firstly, there are those core variables suggested by the LCPIM: disposable income, household wealth, interest rates, and demographic characteristics. Secondly, there are the variables used to proxy precautionary motives (e.g., variation in income,

[3] The seminal expositions on the life cycle model and the permanent income hypothesis are credited, respectively, to Modigliani and Brumberg (1954) and Friedman (1957).

inflation and unemployment) and liquidity constraints. Thirdly, there are other forms of saving (e.g., retirement schemes).

Despite the potential importance of social security (i.e., pensions) on saving behaviour having been noted earlier by Friedman (1957), it was largely ignored in both theoretical and empirical analysis for almost two decades. Feldstein (1974) was the first to incorporate pensions into an empirical life-cycle model based on the specification by Ando and Modigliani (1963) that relates current consumption to current income and a lagged wealth variable. He argued that the existence of a public pension has two countervailing effects on saving: the [positive] *retirement effect* and the [negative] *benefit effect*. For people who would have preferred to work longer and retire later than the pension qualifying age, the decision to retire at the eligibility age and take up the pension extends the implicit post-retirement period that they have to save for during the working years, other things being equal. To the extent that this phenomenon elicits an increase in saving, it is argued that the existence of a public pension, by lengthening the period of retirement over which accumulated assets will be spread, stimulates saving. This is the explanation underlying the retirement effect. The benefit effect is when the need to save during the working years is reduced because the retirement benefit is seen as a substitute for household assets. The net effect will therefore depend on the relative strengths of these two opposing forces.

Feldstein (1974) created a social security wealth (SSW) variable that has two specifications: *gross social security wealth* which is defined as the present value in year t of the retirement benefits which could eventually be claimed by all those who are either in the labour force or already retired in year t; and *net social security wealth* that is equal to gross social security wealth minus the present value of the social security taxes payable by the current labour force. His technique, that made use of the gross social security wealth, proved popular, as it was adopted by later researchers such as Munnell (1974), Barro (1978), Boyle and Murray (1979), Morling and Subbaraman (1995) and Connolly and Kohler (2004). The findings from these studies, however, are mixed. Whereas Feldstein (1974, 1995) and Munnell (1974) analysing United States data, and Morling and Subbaraman (1995) and Connolly and Kohler (2004) analysing Australia data find that the existence of public pension plans depress household savings, Barro (1978) extending Feldstein's (1974) study and Boyle and Murray (1979) analysing Canadian data find no such relationship. The differences in the findings may be attributed to the differences in the mix of additional variables, estimation techniques, countries and sample periods studied. For the dependent variable some studies used consumption expenditure (e.g., Feldstein 1974, 1995) and some others used the saving:income ratio (e.g., Morling and Subbaraman, 1995).

From this review of the literature, two main issues need resolution: variable selection and estimation techniques. The selection of variables is relatively straightforward. For our purposes, the dependent variable is the ratio of aggregate saving to disposable income for households. In addition to the social security wealth variable, the three broad categories of household saving determinants identified earlier in this section of the paper would constitute the set of independent variables to be considered. Thus, the generic aggregate household saving model may be specified as:

$$S/YD = f(SSW^{(+/-)}; YD^{(+)}, W^{(-)}, r^{(?)}, \mathbf{D}^{(?)}; \Delta YD^{(+)}, \pi^{(+)}, \Delta U^{(+)}, L^{(+)}; R^{(-)}) \qquad (1)$$

where S is household saving, YD is household disposable income, SSW is the constructed variable social security wealth, W is household wealth, r is interest rate, **D** is a vector of demographic characteristics, Δ is the first difference operator used to represent change in a variable, π represents the inflation rate, U is the unemployment rate, L represents liquidity constraints, and R represents other retirement saving schemes. The superscripted signs indicate the expected directions of influence of increases in the variables on the saving rate. Depending on which of the two countervailing effects of public pensions on saving dominates the other, SSW can take a positive or negative sign. The received knowledge about the positive-intercept short-run consumption function and the zero-intercept long-run consumption function (Mankiw, 2007, Chapter 16) points to an inverse relationship between the consumption:income ratio (the average propensity to consume, APC) and the saving:income ratio (the average propensity to save, APS). The corollary is that the constant APC expected from a very long data set would imply a constant saving rate, whilst a diminishing APC expected from a relatively short or cross-section data set would imply an increasing saving rate as income increases. Hence, YD can be expected to take a zero coefficient if the time series data set is sufficiently long and a positive coefficient otherwise. Since most macroeconomic data sets tend to be small to moderate in size, YD is expected to take a positive sign. An increase in household wealth pushes up the intercept in the 'short-run' consumption function, so that at any given income level the APC increases and, by implication, the saving rate falls. It can also be argued that, when people feel wealthier, they tend to increase consumption expenditure and decrease saving out of current income. For these reasons, W is expected to take a negative sign. Because the effect of demographic characteristics would depend on the specifications of the variables incorporated and because the effect of interest rate on saving is ambiguous, the signs of the variables **D** and r are undetermined. As explained above, uncertainty about future income induced by variation in income, inflation and unemployment would instigate precautionary saving and credit-constrained households may be forced to defer consumption and increase savings. Hence, the variables ΔYD, π, ΔU and L are expected to take positive signs. Finally, if part of the finite amount that can be saved has to be partitioned to cover other retirement schemes, then naturally the non-superannuation household saving rate would fall, hence, the negative sign of the variable R.

A brief comment on methodology is warranted. Advances in econometric knowledge of the analysis of time series data have put tests for stationarity and use of cointegration techniques at the centre stage. The techniques are essentially aimed at ameliorating the spurious regression problem.[4] Among the empirical studies reviewed here, that by Morling and Subbaraman (1995) was the first to explicitly check for stationarity and apply the error correction methodology; Connolly and Kohler (2004) implemented similar modern techniques. The earlier studies all used short-run dynamic models that were estimated using Ordinary Least Squares (OLS) that ignores the spurious regression problem. In Section 4, a detailed explanation is given of the modern time series techniques employed to preclude spurious regressions and incorporate lagged responses in the models utilised in this study.

[4] More is said about the spurious regression problem in Section 4.

3. A SHORT HISTORY OF THE NEW ZEALAND SUPERANNUATION

New Zealand's old-age pension scheme dates back to 1893 when the state started providing an 'age benefit' of 18 English pounds a year for people aged 65 or older who had 'good moral character and sober habits' (St John, 1999). Since that time the scheme has undergone many changes. The Social Security Act of 1938 lowered the age of eligibility from 65 to 60 for the income-tested age benefit and introduced a universal pension that was taxable but not income tested for those over 65 years of age. In 1977 the two-tiered provisions were consolidated into the National Superannuation with a single eligibility age of 60. The revamped scheme offered a very generous pension by setting the benefit rate for a married/partnered couple at 80% of the national net average wage, with a higher rate for a single person. It also eliminated the means-testing element (St John, 2003). As a result, the incidence of poverty among the elderly dropped significantly but the cost soon proved unsustainable. In less than a decade, expenditure on public pensions as a proportion of GDP rose from about 3% to nearly 8%. To reduce the spending on pensions, in 1985 a tax surcharge of 25% on other income of superannuitants was introduced, and in 1990 a program was initiated to gradually increase the eligibility age to 65 by the year 2001. These succeeded in reducing the expenditure on pensions as a proportion of GDP to about 5% by the late 1990s.

An accord reached among the major political parties in 1993 led to a flat-rate taxable pension that was adjusted for price increases so that it would move within a band of 65% to 72.5% of the national net average wage (NAW). The Retirement Income Act of 1993 saw the National Superannuation renamed New Zealand Superannuation (NZS) but other provisions remained unchanged. The surcharge, which turned out to be quite unpopular, was discarded in 1998 and the government also lowered the pension-NAW band 'floor' from 65% to 60%. The following year a new coalition government restored the 65% floor. Since 2005 the pension-NAW band floor has been set at 66%.

The instability in superannuation policy had continually raised widespread concerns about the future of public pensions in the country. To reduce the uncertainties about the viability of the NZS, the New Zealand Superannuation and Retirement Income Act that established the New Zealand Superannuation Fund was passed in October 2001. And as broached in the introductory part of this paper, the Fund is a dedicated investment fund aimed at building up resources to help reduce the net fiscal cost of NZS in the future. Starting with NZ$2.4 billion (US$1.9 billion) cash in September 2003, the funding framework originally provided for fiscal transfers of an average of NZ$1.95 billion (US$1.6 billion) annually until about 2027. [5] The mandate to the Fund's governing body, the 'Guardians', is to invest the money in a way that maximises returns without undue risk over the long term while avoiding prejudice to

[5] Owing to fiscal difficulties and economic slowdown brought on largely by the current world-wide financial crisis the government, in its 2009 Budget, announced a drastic change in the contributions to the Fund: a reduction to NZ$250 million in the 2009/10 fiscal year and suspension of contributions thereafter until 2020/21. Depending on changes in economic and fiscal conditions in the near future the non-contribution period may be reviewed. The main impact of the policy change seems to be a delay in the date of cessation of capital contributions and the date the Fund's assets are expected to peak.

New Zealand's reputation as a responsible member of the world community. By law, the Government may not make any capital withdrawals from the Fund before 1 July 2020. By 31 May 2009 the total assets of the Fund stood at NZ$13.1 billion (US$8.2 billion) and the annualised rate of return on its investments since inception was reported to be 3.83% per annum.[6] Currently the New Zealand Treasury estimates that capital contributions will cease in 2031, at which time the Government will start to take money out of the Fund to finance between 15% and 20% of the superannuation expenditure. In spite of the capital withdrawals, the Fund is expected to continue growing over time in nominal dollar terms, because the capital withdrawals are projected to be less than the Fund's after-tax income. Under the present scenario, the Fund's assets are projected to peak at about 23% of GDP in 2056 and then gradually decrease in subsequent decades.

The effect of the Fund on household saving is an empirical issue, since theory cannot give a clear indication as to what the net impact might be. An important part of the objective of the superannuation initiative is to create a more certain and stable environment that will make planning for one's retirement easier. On the one hand, the clearer picture that emerges can make people realise the inadequacy of current saving and so increase their saving efforts. On the other hand, the increased certainty can lessen the need for precautionary saving against adverse events (McCulloch, 2000). To provide an answer, this study used a dummy variable in the empirical models for the sub-period that the Fund has been in existence.

4. DATA AND ANALYTICAL METHODS

4.1 Data and Sources

In the empirical models some of the variables identified in the last section were side-stepped because of either non-availability of data in the New Zealand context or questionable usefulness in other empirical studies. Variables that were not considered included interest rate, demographic characteristics and liquidity constraints. Theory is ambiguous on the role of interest rate on saving and many studies find it to be an insignificant determinant. Among the array of demographic variables that could be utilised, perhaps the most relevant for our circumstance might be the labour market participation rates of the different age groups and the proportion of the over-65s in the population. Owing to the relatively small size of the sample and the tendency of demographic characteristics to change with time, demographic factors were subsumed in a trend variable to conserve on degrees of freedom. The absence of a reliable proxy to capture economy-wide liquidity constraints militated against directly controlling for liquidity constraints in the model. Of the two specifications of social security wealth, the more tractable 'gross' variety has been utilised more often in empirical studies than the 'net' variety. In the New Zealand context the absence of a specific social security tax that is required to calculate the net social security wealth variable

[6] The total assets and annual rate of return have fluctuated over the years reflecting the performance of the Fund's investment portfolio. For instance, total assets peaked at NZ$14.7 billion in May 2008 and the annual rate of return peaked at 19.2% for the July 2005-June 2006 period. The latest figures report a return of negative 22.45% for the July 2008-May 2009 period. For details on the profile, governance arrangements and performance of the New Zealand Superannuation Fund see the relevant website: http://www.nzsuperfund.co.nz.

precluded the employment of that variable. Hence, this study will join that genre of studies that have used the gross social security wealth variable.

Data on household net wealth were obtained from the Reserve Bank of New Zealand, and those on GDP, benefits, household disposable income, saving, inflation and unemployment were sourced from Claus and Scobie (2002) and from Statistics New Zealand. Construction of values for the gross social security wealth variable followed closely the formulation by Feldstein (1974) and Munnell (1974). The required survival probabilities were calculated using data in the cohort life tables provided by Statistics New Zealand.[7] Because the earliest date for which data could be obtained on household disposable income and saving was 1972, the sample period was restricted to cover 1972 to 2008. Monetary variables were deflated to constant 2000 New Zealand dollars (NZ$). Summary statistics of the raw data on the six variables employed in this study are presented in Table 1. There it can be observed that over the 37-year period the household saving rate averaged negative 1.71%, with a maximum of 4.44% (posted in 1988) and a minimum of negative 17.18% (posted in 2008). The declining household saving rate, which has been negative since 1994 when the sustained government budget surpluses began, is ascribed to: (i) New Zealanders' penchant to hold wealth in the form of real estate rather than financial assets[8]; (ii) financial deregulation in the 1980s that made it easier to acquire home loans; and (iii) strong capital gains on housing in recent times[9] making households feel that they do not need to save as much from current income (Briggs *et al*., 2006; OECD, 2007).

Table 1
Summary Statistics of the Raw Data for the Sample Period

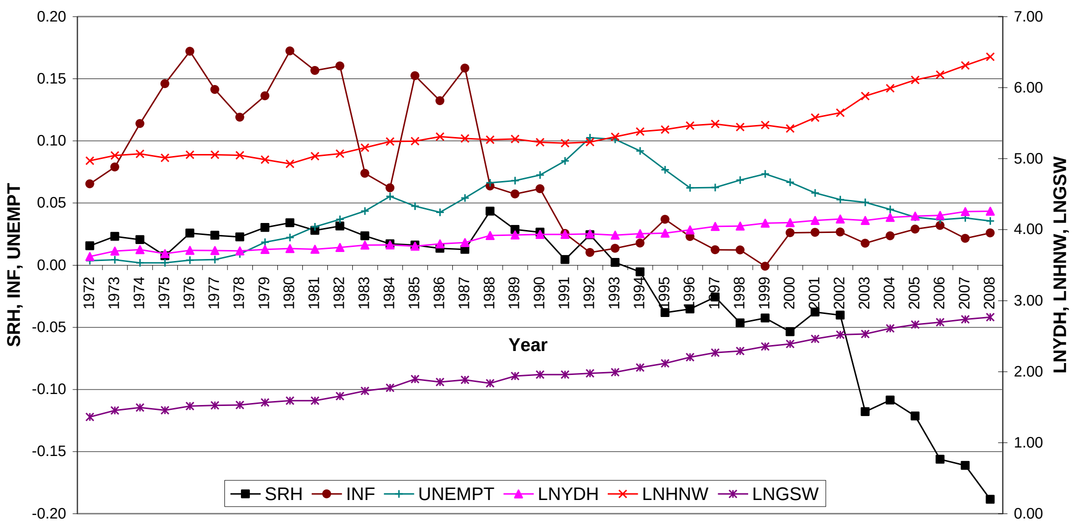
Statistic	Household Saving-Disposable Income Ratio (%)	Annual CPI Inflation Rate (%)	Annual Unemploy-ment Rate (%)	Household Disposable Income (constant NZ$ bn)	Household Net Worth (constant NZ$ bn)	Gross Social Wealth (constant NZ$ bn)
Average	-1.71	7.03	4.68	51.1230	237.8155	8.1223
Minimum	-17.18	-0.10	0.18	37.4065	137.9425	3.9014
Maximum	4.44	17.23	10.25	70.7036	622.4164	15.8919

For modelling purposes, however, the variables needed to be transformed and renamed. Inflation and unemployment rates (INF and UNEMPT, respectively) were expressed in decimals. Natural logs of disposable income (LNYDH), gross social security wealth (LNGSW) and household net worth (LNHNW) were taken. Because the saving rate took negative values in some years, that variable could not be logged. Rather, the natural log of the sum of 1 and the saving rate in decimal (SRH, suggestive of saving rate of households) was used. This allowed the regression

[7] Details about the computations are available from the authors upon request.
[8] Between 1979 and 2008 housing value as a proportion of household net wealth rose steadily from 67.6% to 98.8%, averaging 78.5% for the period. During the same period household net financial wealth expressed as a proportion of household net wealth fell steadily from 32.4% to 1.2%, averaging about 21.5%.
[9] From a value of 100 in 2000Q4, the housing price index rose progressively to peak at a value of 2,415 in 2008Q1 but by 2009Q1 had slipped to 1,782.

coefficients to be interpreted as the percentage-point change in the saving rate following one percent relative change (in the case of the logged variables) and one percent absolute change in the case of the non-logged variables (see Wooldridge 2009, pp. 192, 701). The graphs of the transformed variables are shown in Figure 1. Two vertical axes are used in order to allow the trends of the lower-magnitude variables to be discernible without being swamped by the higher-magnitude variables. The trends of the different variables seem distinct. LNGSW, LNYDH and LNHNW can be described as having risen steadily over the sample period. INF fluctuated widely at relatively high levels in the first half of the period but became more settled at lower levels in the second half of the period. UNEMPT exhibited a humped trend curve. The downward sloping trend curve of SRH is consistent with a household saving rate that fluctuated mildly at low positive values in the first half of the period but dropped into the negative region in the second half of the period and has been plummeting continually.

Figure 1
Graphs of the Transformed Data: 1972-2008



4.2 The Model and Specification Issues

4.2.1 Preamble

Since economic time series were analysed in this research, the choice of methodology was driven by two considerations: the need to preclude spurious regressions and the need to incorporate lagged responses in the model. Following the seminal work of Engle and Granger (1987), it has become customary when specifying regression models involving time series to check that the different variables are integrated of the same order, otherwise the regression might not make sense. A variable is said to be integrated of order d (i.e., I(d)) if it must be differenced d-times before it can be rendered stationary or weakly dependent (Wooldridge, 2009). Stationary variables are integrated of order zero (i.e., I(0)) and nonstationary variables are integrated of order equal to or greater than one (i.e., $d \geq 1$). A regression of one nonstationary variable on other nonstationary variables is deemed spurious unless the variables are cointegrated. A set of I(1) variables are said to be cointegrated if there exists a linear combination of them which is I(0).

For instance, if each of the time series in a standard OLS regression such as Equation (2) is I(1) and they are cointegrated, that static regression, known as the cointegrating regression, is meaningful and represents the long-run equilibrium relationships between the dependent variable and the independent variables.

$$Y_t = \beta_0 + \sum_{j=1}^{m} \beta_j X_{jt} + e_t \qquad (2)$$

The coefficients (the β_j's) represent the independent variables' long-run impacts on the dependent variable. The concept of equilibrium here is that of no tendency to change. Because of lagged responses among economic variables, however, it is very likely that the short-run impacts of the independent variables will be different from their long-run impacts and therefore the short-run value or behaviour of the dependent variable may be different from its long-run value or behaviour resulting in a short-run disequilibrium captured by the error term e_t. How the disequilibrium is eliminated from the short run to the long run needs to be modelled. According to the Granger representation theorem (Harvey 1993, p. 260), if I(1) variables are cointegrated the short-run dynamics corresponding to the long-run equilibrium can be described by the error correction model (ECM). This presents two challenges: (a) how to ascertain the existence of cointegration, and (b) how to estimate the ECM.

On cointegration, Greene (2003, Chapter 20) proffers that two broad approaches for testing for cointegration have been developed: (i) the Engle and Granger (1987) method based on testing for a unit root in the residuals of the cointegrating regression, and (ii) the VAR (vector autoregression) approach due to Johansen (1988, 1991) and Stock and Watson (1988). As noted by Pesaran and Pesaran (1997, p. 291), "The residual-based cointegration tests are inefficient and can lead to contradictory results, especially when there are more than two I(1) variables under consideration. A more satisfactory approach would be to employ Johansen's ML procedure." The most popular test for cointegrating rank based on the VAR approach (CVAR) is the specification by Johansen and its highlights will be outlined shortly.

The estimation of the ECM involves a regression of the first difference of the dependent variable on its own lags, the distributed lags of the first differences of the independent variables plus the lagged residuals from the cointegrating regression used as the error correction term. At the most general level the ECM[10] may be represented as

$$\Delta Y_t = a + \sum_{k=1}^{p} c_k \Delta Y_{t-k} + \sum_{j=1}^{m} \sum_{i=0}^{q_j} b_{ji} \Delta X_{j,t-i} + \lambda \hat{e}_{t-1} + u_t \qquad (3)$$

[10] Some variants of the ECM may not incorporate lags of the (first difference of the) dependent variable. Here, we ignore the possibility of apposite deterministic variables such as time trend and seasonal variables in order to simplify the exposition and concentrate on the key variables.

where p is the optimal lag of the dependent variable; q_j is the optimal lag of the j^{th} independent variable; $\hat{e}_{t-1}$ is the error correction term and its coefficient, λ, is the speed-of-adjustment coefficient. It must be noted that λ, which gives the proportion of the disequilibrium eliminated in one period, has the range $-1 \leq \lambda < 0$. The optimal lag structure may be selected based on the scores from one or more of the conventional model selection criteria such as the Akaike Information Criterion (AIC) or the Schwarz Bayesian Criterion (SBC). Empirically, the ECM may be estimated from either a single-equation approach or from a multiple-equation (vector) approach. In the single-equation approach there are two alternative methods: the two-step procedure suggested by Engle and Granger (1987) and the simultaneous estimation of the short-run and long-run parameters using the ARDL (autoregressive distributed lag) model. Bewley (1979), Banerjee *et al.* (1986), Wickens and Breusch (1988) and Maddala (1992) have shown that more efficient parameters can be obtained from the ARDL approach. The vector approach to estimating the ECM (VECM) is intertwined with the vector approach to testing for the cointegrating rank (CVAR). It can be inferred from the review so far that the relevant methods are CVAR (or VECM) and ARDL approaches to estimating the short-run and long-run parameters.

4.2.2 VAR-based Cointegration Test and ECM Estimation

In the VAR methodology, each of the variables in the system is regressed on its own lags and the lags of the other variables. The optimal number of lags can be decided based on statistical selection criteria such as the Akaike Information Criterion (AIC), the Schwartz Criterion (SC) or the Hannon-Quinn (HQ) criterion. Assume $\mathbf{Z_t}$ is a vector of k-jointly determined endogenous variables[11] and $\mathbf{W_t}$ is a vector of m exogenous variables. A p^{th} order VAR model of the inter-related time series, VAR(p), can be written as:

$$Z_t = \sum_{i=1}^{p} \Phi_i Z_{t-i} + \Psi W_t + \varepsilon_t \tag{4}$$

where Φ_i and Ψ are matrices of coefficients to be estimated, and ε_t is a vector of independent and identically distributed disturbances. This version of the model may be referred to as unrestricted VAR (UVAR). If the endogenous variables are each I(1) we can write the VAR(p) model as a vector error correction model (VECM):

$$\Delta Z_t = \Pi Z_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta Z_{t-i} + \Psi W_t + \varepsilon_t \tag{5}$$

where $\Pi = \sum_{i=1}^{p} \Phi_i - I$, and $\Gamma_i = -\sum_{j=i+1}^{p} \Phi_j$

Granger's representation theorem asserts that if the coefficient matrix Π has reduced rank (i.e., rank(Π) = r < k) then there exist k-by-r matrices α and β each with rank r such that Π is equal to $\alpha\beta'$ and $\beta'\mathbf{Z_t}$ is integrated of order zero, I(0). The rank r is the number of cointegrating or long-run relations among the variables and each column of β is a cointegrating vector. The Johansen maximum likelihood estimation procedure (Johansen, 1988, 1991, 1995a; Johansen and Juselius, 1990) can be used to estimate

[11] It may be assumed that Z is composed of the set of Y and X's considered in the static model earlier.

the two matrices α and β and to test for the number of distinct cointegrating vectors. Restrictions on the elements of β help to determine which variables are relevant in the long-run relations; economic theory may have to be invoked to decide on the restrictions to impose on each cointegrating vector (Johansen, 1995b). The elements of α are known as the adjustment parameters in the VECM. When appropriate and binding restrictions are imposed on the identified cointegrating vectors, the VECM becomes a restricted VAR and is also called cointegrating VAR (CVAR).

4.2.3 ARDL-based Cointegration Test and ECM Estimation

The ARDL may be formed by augmenting the static long-run model with lags of the dependent and independent variables on the right hand side to yield an equation such as Equation (6) with the optimal lag structure ARDL(p, q_1, …. q_m).

$$Y_t = \alpha + \sum_{k=1}^{p} \gamma_k Y_{t-k} + \sum_{j=1}^{m} \sum_{i=0}^{q_j} \delta_{ji} X_{j,t-i} + \varepsilon_t \qquad (6)$$

The selection of the optimal lag structure of the ARDL model can also be based on model selection criteria such as the AIC or the SBC. The coefficients of the ARDL give the short-run parameters and the contingent long-run coefficients and adjustment coefficient can be calculated from them. It may also be noted that the ARDL can equivalently be re-specified as the ECM or formulated as the Bardsen transformation or the Bewley transformation each of which incorporates both the short-run and long-run impacts.[12]

Using the Bardsen transformation, Pesaran *et al*. (1996, 2001) have demonstrated that in addition to estimating the short-run and long-run multipliers, the ARDL can be used to check for the existence of cointegration (i.e., long-run relationship) among a set of variables without needing to know the order(s) of integration of the variables even when the variables are a mixture of I(0)'s and I(1)'s. Their procedure involves two stages. At the first stage an F-test is done to ensure that a long-run relationship exists between the variables. This is effected by estimating an unrestricted error correction version (the Bardsen transformation) of the ARDL model:

$$\Delta Y_t = a + \sum_{k=1}^{p} c_k \Delta Y_{t-k} + \sum_{j=1}^{m} \sum_{i=0}^{q_j} b_{ji} \Delta X_{j,t-i} + \lambda_1 Y_{t-1} + \sum_{j=1}^{m} \lambda_{j+1} X_{j,t-1} + u_t \qquad (7)$$

If a long relationship exists between Y and the Xs, their lagged levels would belong in Equation (7), otherwise they would not. Thus the test for the existence of long-run relationship is a test of the joint significance of the coefficients of the one-period lagged levels of the variables in the model. The null hypothesis of 'no long-run relationship' or 'no cointegration' is the joint hypotheses test of H_0: $\lambda_1 = \lambda_2 = \ldots = \lambda_{m+1} = 0$ against the alternative hypothesis that H_0: $\lambda_1 \neq \lambda_2 \neq \ldots \neq \lambda_{m+1} \neq 0$. This variable inclusion/exclusion test yields a non-standard F-statistic and to operationalise it, Pesaran *et al*. (2001) have tabulated the appropriate critical values for different numbers of regressors when all the variables are I(1) and when all the variables are

[12] For these specifications see, for example, Maddala and Kim (1998, Ch. 2), Banerjee *et al*. (1993, Ch. 2), Hendry (1995, Ch. 6) and Patterson (2000, Ch. 8).

I(0), thus catering for fractionally integrated variables as well. If the computed F-statistic is greater (smaller) than the upper (lower) critical value, it can be inferred that the variables are cointegrated (not cointegrated) irrespective of whether the variables are I(0) or I(1). If the computed F-statistic falls within the critical value band, the "inconclusive zone", the order(s) of integration of the variables need to be checked before the correct inference can be made. After ascertaining that a genuine long-run relationship exists between the variables the procedure moves to the second stage where the auxiliary ARDL regression is run and the contingent short-run and long-run parameters are estimated. An econometric/time series software that can implement all this is Microfit.

5. EMPIRICAL RESULTS

In the pre-estimation tests on the variables, the F-statistic in the ARDL bounds test was estimated to be 4.0414 with a p-value of 0.037. For five regressors, the critical values for I(0) and I(1) variables are 3.189 and 4.329, respectively, at the 5% significance level.[13] Since the estimated statistic falls within the inconclusive zone it became imperative to ascertain the order of integration of each of the variables. The result also made the VAR-based cointegration test very compelling. The ADF and PP unit root test results reported in Table 2 indicate each variable is I(1). In the process to select the order of the underlying VAR model, SC chose VAR(1) whilst AIC, FPE and HQ chose VAR(3). Because of the small sample size, however, VAR(2) was chosen. To check the cointegrating rank of the VAR model, the Johansen cointegration test was implemented. Both the trace and maximum-eigenvalue statistics indicated that there was one cointegrating equation at the 5% level of significance (see Table 3). Thus, the way was cleared for the single-equation estimation of the saving function.

Table 2
Unit Root Tests

Variable	ADF				PP			
	Levels		1st Differences		Levels		1st Differences	
	Stat	p-value	Stat	p-value	Stat	p-value	Stat	p-value
SRH	-0.8782	0.9477	-8.3144	0.0000	-0.8782	0.9477	-9.8062	0.0000
LNYDH	-2.5908	0.2864	-7.4389	0.0000	-2.5735	0.2938	-7.7207	0.0000
LNHNW	0.8996	0.9997	-3.2073	0.1010	0.5953	0.9992	-4.9326	0.0017
LNGSW	-1.8012	0.6833	-7.4321	0.0000	-1.6883	0.7358	-7.6846	0.0000
INF	-3.2614	0.0891	-5.3628	0.0007	-3.2541	0.0904	-9.3178	0.0000
UNEMPT	-0.4413	0.9816	-4.7502	0.0029	-0.5712	0.9748	-3.2061	0.0997

Note: In levels, the critical values at the different significance levels are: 1%, -4.2350; 5%, -3.5403; and 10%, -2.2025. In first differences, the critical values at the different significance levels are: 1%, -4.2436; 5%, -3.5443; and 10%, -3.2047.

[13] The corresponding critical values at the 10% level are 2.782 and 3.827; at the 1% level they are 4.011 and 5.331, respectively. This means SRH is cointegrated with other variables at the 10% level but the test is inconclusive at the 1% and 5% levels. When the other variables are sequentially made the dependent variables, the conclusions were: LNYDH: inconclusive at the 10% level, not cointegrated at the 1% and 5% levels; LNHNW: not cointegrated at all the conventional levels; LNGSW: inconclusive at the 5% and 10% levels, not cointegrated at the 1%; INF: cointegrated at all the conventional levels; UNEMPT: not cointegrated at all the conventional levels.

Table 3
Results of the Johansen Cointegration Tests

Eigenvalue	Null hypothesis	Maximal Eigenvalue		Trace Test	
		Test stat	5% CV	Test stat	5% CV
0.759658	r = 0	49.89925	43.41977	114.4931	107.3466
0.571243	r ≤ 1	29.64026	37.16359	64.59385	79.34145
0.383221	r ≤ 2	16.91355	30.81507	34.95359	55.24578
0.311376	r ≤ 3	13.05711	24.25202	18.04004	35.01090
0.110583	r ≤ 4	4.101606	17.14769	4.982926	18.39771
0.024866	r ≤ 5	0.881320	3.841466	0.881320	3.841466

Before estimating the ECM via the CVAR and ARDL approaches, the set of explanatory variables was augmented with a pre-funding dummy variable (PFDUM) to capture the influence of the presence of the NZS Fund and a trend variable (TREND) to capture independent, time-driven factors such as an ageing population with those in retirement living longer[14], increasing income inequality[15], decreasing housing affordability[16] and average household size, and distinct labour market developments such as increasing female and decreasing male labour force participation rates[17]. It will be noticed in Figure 1 that the dependent variable, household saving rate (SRH), has a pronounced quadratic trend curve. Whereas the VECM results from the Johansen specification do not depend on the model selection criteria, the ARDL approach leads to different sets of results depending on the model selection criterion used. SBC selected ARDL(1,1,0,0,0,0) and AIC selected ARDL(0,1,1,0,2,2). For circumspection, we report the results from the ARDL models selected by SBC and AIC as well as the results obtained for the VECM. The long-run results are reported in Table 4 and the error correction (short-run) results are reported in Table 5.

[14] Between 1972 and 2008 life expectancy at birth for males increased from 69 to 78 years and that for females increased from 75 to 82 years, meaning expected life after age 65 increased from 4 to 13 years for males and from 10 to 17 years for females. Additionally, the frequency of retirement among the over-65s *decreased* over that period: the percentage for men dropped from 89% to 80% and that for women dropped from 96% to 90% implying the labour force participation rates of the over-65s increased.

[15] Martin (2002) reports that as measured by the Gini coefficient, income inequality increased by 10.7 percentage points between 1976 and 1996, and Hyslop and Yahanpath (2005) estimate that between 1998 and 2004 the equivalised household income inequality increased 2-3 percent. In its latest Social Report, the Ministry of Social Development states that, for the 2004-2007 period, the Gini coefficient for the country has remained steady at 0.34 (MSD, 2008, p. 61).

[16] As measured by the proportion of households that spent more than 30% of their disposable income on housing costs, housing affordability worsened from 10.6% in 1988 to 26% in 2007 (MSD, 2008, p. 64).

[17] Between 1972 and 2008, the labour force participation rate for males dropped from 80.1% to 75.3% whilst that for females rose from 33.1% to 61.6%

Table 4
Long-run Estimated Parameters (Dependent Variable is SRH)

Regressor	CVAR Approach	ARDL Approach	
		SBC Results	AIC Results
C	–0.0040 (none)	0.1839 (0.42)	–0.0489 (-0.14)
LNYDH	0.1087 (1.66)	0.0901 (0.78)	0.1202 (1.27)
LNHNW	–0.1068 (-8.48)	–0.1317 (-5.93)	–0.1178 (-7.92)
LNGSW	0.0892 (2.00)	0.1216 (1.63)	0.1504 (2.66)
INF	0.3633 (6.51)	0.1836 (1.70)	0.3333 (3.47)
UNEMPT	1.3917 (11.22)	1.2456 (5.15)	1.4210 (7.57)
TREND	–0.0071 (none)	–0.0094 (-2.79)	–0.0108 (-4.40)
PFDUM	…	0.0332 (2.62)	0.0334 (3.13)

Note: t-ratios are in parentheses below the coefficients.

As shown in Table 4, there is consistent signage in the long-run models although the magnitudes of the corresponding coefficients from the different models are not exactly equal. The trend variable takes a negative sign in all of the models and is shown to be significant in the ARDL models. This suggests that, apart from the 'customary' explanatory variables considered in the saving model, there are important independent factors that have changed systematically with time and whose net effect on the household saving rate has been negative. Anecdotal evidence suggests that a population that is becoming older and living longer and experiencing worsening income distribution and housing affordability over time is not likely to be able to increase its saving. Our conjecture is that, because New Zealand households have experienced these phenomena over the study period, the declining aggregate saving rate is not enigmatic.

Among the economic variables explicitly considered, only household net wealth (LNHNW) depresses the saving rate, and quite significantly: a 10% increase in household net worth decreases the saving rate by about 1 percentage point; the other variables boost the saving rate. From this result it may be inferred that increases in household net worth, emanating primarily from rising housing values, engender increases in consumption and thus reductions in the saving rate. The finding of a significant propensity to dis-save out of household wealth is consistent with what many other commentators have noted about consumer behaviour in New Zealand (e.g., Orr and Purdue, 2001; Scobie *et al*., 2004; NZIER, 2007).

Table 5

Short-Run Estimated Parameters (Dependent Variable is ΔSRH)

Regressor	CVAR Approach	ARDL Approach: SBC Results	ARDL Approach: AIC Results
C	0.0123 (1.49)	0.1296 (0.42)	–0.0489 (-0.14)
ΔSRH_{-1}	–0.2188 (-0.88)		
ΔLNYDH		0.4125 (5.13)	0.3013 (3.51)
$\Delta LNYDH_{-1}$	0.0678 (0.44)		
ΔLNHNW		–0.0928 (-4.97)	–0.1684 (-4.38)
$\Delta LNHNW_{-1}$	–0.0384 (-0.63)		
ΔLNGSW		0.0857 (1.69)	0.1504 (2.66)
$\Delta LNGSW_{-1}$	–0.0026 (-0.03)		
ΔINF		0.1294 (1.76)	0.0449 (0.51)
ΔINF_{-1}	–0.1728 (-1.10)		–0.1445 (-2.39)
ΔUNEMPT		0.8783 (4.48)	0.7980 (2.56)
$\Delta UNEMPT_{-1}$	–0.2049 (-0.42)		–0.3477 (-1.27)
TREND	–0.0012 (-2.28)	–0.0066 (-2.93)	–0.0108 (-4.40)
PFDUM	0.0127 (0.96)	0.0234 (2.49)	0.0334 (3.13)
$ECT_{(-1)}$	–0.7978 (-2.60)	–0.7051 (-5.56)	–1.0000 (none)

Note: t-ratios are in parentheses below the coefficients.

Concerning gross social security wealth (LNGSW), the variable of key interest, the positive sign in all the models indicates that in New Zealand, after controlling for the nominated variables and time-driven factors such as demographic changes and labour market developments, an increase in the social security wealth does not offset private saving as found elsewhere. The absence of both means-testing and a compulsory element in retirement saving in New Zealand may be the main reason for this finding. The regression results suggest that an increase of 10% in the social security wealth would induce between 0.9 and 1.5 percentage points increase in the household saving rate. The implication is that certainty from the existence and enhanced viability of the NZS scheme engenders greater saving. And indeed this finding is buttressed by the positive significant coefficients of PFDUM (the dummy variable for the sub-period the NZS Fund has been in existence) in the ARDL models.[18] It can therefore be concluded that in New Zealand, the [positive] induced-retirement effect of public pensions on saving trumps the [negative] asset-substitution effect.

The positive effect of household disposable income (LNYDH) on the saving rate (albeit of low statistical significance) contrasts with the negative effect of household net wealth. The life-cycle/permanent-income hypothesis asserts that over long periods of time one should expect a constant consumption:income ratio (average propensity to consume) or, equivalently, a constant saving:income ratio (average propensity to save, the saving rate) because wealth and income are expected to grow at such rates as to keep the wealth:income ratio constant and the variation in income is assumed to come from the permanent component. If the hypothesis is true, then LNHYD should have a zero or insignificant coefficient in the long-run models but a positive coefficient in the short-run models. The hypothesis of zero-coefficient for LNHYD cannot be rejected at the 5% level of significance in all the long-run models in Table 4 but is rejected in the ARDL short-run models in Table 5.

The positive and significant coefficients of INF and UNEMPT indicate that increases in inflation and unemployment lead to increases in the saving rate which may be termed precautionary saving. An increase in the rate of inflation leads to less than proportional increase in the saving rate, but an increase in the unemployment rate leads to a more than proportional increase in the saving rate. High inflation may be associated with financial or real shocks to the economy and the response to the resulting increased variability of real income flows may be the precautionary saving by households. When unemployment goes up, risk-averse households, one or more of whose members have become newly unemployed, and who may be liquidity constrained, may be compelled to increase saving.

With reference to Table 5, the short-run parameters obtained from the CVAR are all insignificant except those for TREND and $ECT_{(-1)}$. On the other hand, most of the short-run parameters obtained from the ARDL models are significant. It will also be noted that different mixtures of regressors were selected by the different models. When the ARDL results are viewed as complementing the CVAR results, a much more harmonious picture emerges than the configuration of estimates might at first suggest.

Most importantly, the error correction term $ECT_{(-1)}$ is correctly [negatively] signed in all the models. This supports the hypothesis that the respective long-run models are stable, or that the variables identified as significant in the long-run models are cointegrated. The range of estimates from negative 0.71 to negative 1.00 suggests that, after a shock, the speed of adjustment of the saving rate towards the long-run equilibrium rate is reasonably fast; at least, 71% of the disequilibrium is eliminated in the subsequent year after a shock. Another implication is that, depending on the

[18] Being classified as an exogenous variable the time dummy, PFDUM, was not selected in the long-run model by CVAR but it was selected by both ARDL models.

model, it takes anywhere from one year to four years for the effects of a shock to dissipate. TREND is consistently negatively signed and PFDUM is consistently positively signed, which accord with expectations.

Concerning the variables with first differences and lagged differences, it will be seen in the second column of Table 5 that CVAR selected only lagged differences which were all insignificant but the ARDL results encompassed both the first differences and lagged differences. Ignoring the CVAR results and considering the ARDL results, it will be realised unambiguously that, just as in the long term, household disposable income and gross social security wealth boost the change in household saving rate and household net wealth dampen the change in household saving rate in the short term. When the rate of inflation or unemployment picks up from one year to the next, the corresponding change in the saving rate also becomes larger. The model selection criterion AIC, that tends to select more variables than the parsimonious SBC, however, is suggesting that the increment in the saving rate tapers off in successive years.

6. SUMMARY AND CONCLUSION

The projected increases in fiscal outlays on the universal pay-as-you-go pension scheme in New Zealand, in a climate of global recession, has reignited the debate on the affordability of New Zealand Superannuation (NZS). Reforms seem inevitable and they are more likely to involve expenditure reduction rather than tax increases. To reduce the fiscal burden, the government can either cut benefits or make people take a greater role in caring for themselves by either increasing the retirement age or increasing incentives for people to save more during their working years. A reduction in pension benefits would give greater prominence to private savings in determining retirees' welfare. This prospect occasions a revisit of the relationship between public pension schemes and private saving and the empirical question of whether increases in the former encourage or depress the latter. The distinguishing feature of the literature on that topic is the incorporation of certainty of retirement income in 'traditional' consumption/saving models. In empirical studies, the preeminent methodology is the employment of the constructed social security wealth variable pioneered by Feldstein (1974) to capture the certainty of retirement income. Whereas such studies have been done on countries such as Australia, Canada and the US, a comparable one has not been done on New Zealand. The paper set out to fill that void by estimating both a long-run equilibrium model and short-run dynamic model of the household saving function that takes account of the certainty derived from the presence of the NZS and the creation of the NZS buffer fund in 2001.

For robust results, the two alternative approaches to estimating error correction models, the cointegrating vector autoregression (CVAR) and autoregressive distributed lag (ARDL), were taken to estimate the long-run and short-run parameters. The two approaches yielded comparable long-run parameters but their short-run parameters showed some variation. The long-run results indicate that: whereas the trend in the household saving rate has been negative, increases in disposable income and gross social wealth boost saving; the introduction of the NZS Fund in 2001 has elicited a slight positive response in the saving rate; there is a significant propensity to dis-save out of household net wealth; and inflation and unemployment engender significant precautionary saving. The short-run results are consistent with the long-run results. In all the short-run models the error correction term takes the expected negative sign; estimates of the adjustment coefficient suggest that, following a shock, no less than 71% of the deviation of the saving rate from the long-run level is corrected in one year. From these results it can be concluded that in New Zealand the certainty of retirement income is auspicious to household saving and so cuts to the benefits rate that would reduce the social security wealth are likely to reduce the household saving rate and, by implication, the welfare of superannuitants.

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