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ABSTRACT

Events surrounding the global financial and economic crises of 2008 and 2009 have sparked a renewed interest in discretionary fiscal policy. This paper considers whether private saving in Australia behaves in a manner that is consistent with Ricardian equivalence, thus mitigating the effects of fiscal policy, or conversely, if fiscal policy has some ability to influence real economic activity. A model of private and public saving is estimated using the autoregressive distributed lag approach (ARDL) to cointegration. This estimation procedure is advantageous due to its ability to provide both short- and long-run coefficient estimates, and can accommodate coefficients for structural breaks. Given that the Australian economy has been subject to a substantial amount of structural change over the past 50 years, the estimations attempt to account for these structural effects on long-run savings behaviour. Results indicate that while there is not a full Ricardian response to changes in the fiscal stance, evidence suggests some partial offsetting behaviour – implying that fiscal policy does elicit some (limited) impact on economic activity.

Keywords: Ricardian equivalence, fiscal policy, cointegration, structural breaks.

JEL Classifications: E21, E62, C22, H62.

1. Introduction

Research interest in fiscal policy waned over the 1990s, and for the most part of the 2000s, as the "new consensus" on macroeconomic policy saw monetary policy (inflation targeting) assuming the role of stabilising short-run fluctuations in prices and output in most advanced economies. Fiscal policy over this period was increasingly directed toward the medium-term sustainability of government balance sheets and allowing the automatic stabilisers to freely operate.

However, fiscal policy debates in Australia were reignited in the mid 2000s as the Howard Government undertook a series of personal income tax cuts. At that time, the economy was operating at or near full capacity with unemployment around 30-year lows. Critics argued that this loosening of fiscal policy would only add to aggregate demand – leading to higher inflation and interest rates.

Sharp falls in output associated with the global financial and economic crisis in 2008 and 2009 have seen fiscal stimulus packages enacted in many countries, and a renewed interest in activist fiscal policy. In a number of countries monetary policy had reached the zero bound on nominal interest rates, leaving quantitative easing measures and fiscal policy to support aggregate demand. To prevent a severe and prolonged global downturn, in late 2008 the International Monetary Fund (Spilimbergo et al: 2008) called for a fiscal loosening across the advanced economies amounting to at least 2 per cent of global gross domestic product (GDP). By mid 2009, Australia had implemented fiscal stimulus packages amounting to around 3 per cent of GDP in 2008-09 and 2 per cent of GDP in 2009-10 (Budget: 2009).

Considering the potential efficacy of fiscal policy, Hemming (et al: 2002) provides an excellent survey of the international evidence on fiscal multipliers from simulations using macroeconomic models and reduced-form specifications. In short, Hemming reports that positive fiscal shocks, generated using estimated macroeconomic models, produce positive multipliers, with expenditure multipliers in the range of 0.6 to 1.5 and tax multipliers in the range of 0.3 to 0.8; long-term multipliers are generally smaller and some are negative. More recent estimates have been produced by the Organisation for Economic Co-Operation and Development (2009), and the International Monetary Fund (2009).

However, as Kennedy (et al: 2004) note, there is little empirical evidence on the efficacy of fiscal policy in Australia, or estimates of fiscal multipliers. Perotti (2002) finds a positive short-term impact spending multiplier of 0.6 for Australia, peaking at 0.8 after 14 quarters. Recent estimates from the OECD (2009) suggest that the fiscal multiplier in Australia is 0.2 for tax cuts, and increases up to 1.3 for direct government investment (such as infrastructure).

In contrast to fiscal policy having some impact on aggregate demand, Ricardian equivalence asserts that fiscal deficits merely postpone taxes, and through the actions of altruistically motivated individuals, budget deficits have no real affects on the economy. Barro (1974) considered the effects on bond values and tax capitalisation of finite lives, imperfect capital markets, a government monopoly in the production of bond 'liquidity services' and uncertainty about future tax obligations. Within the context of an overlapping generations model, Barro showed that finite lives will not be relevant for future tax liabilities so long as current generations are connected to future generations by a chain of operative intergenerational transfers (Barro: 1974). This paper gave rise to what is now known as the Ricardian equivalence theorem, or the Barro-Ricardo hypothesis. The key result of Barro's investigation being that so long as there is an operative intergenerational transfer, there will be no net-wealth effect and no effect on aggregate demand; or on interest rates of a marginal change in government debt. Essentially, under the Barro-Ricardo hypothesis deficits do not matter, and do not have any impact on the macroeconomy.

Both Leiderman and Blejer (1988) and Seater (1993) provided in-depth overviews of the Ricardian equivalence theorem. Surveys of previous empirical studies on Ricardian equivalence have been produced by Gale and Orszag (2004), and Ricciuti (2003).

With little (recent) empirical knowledge on the efficacy of fiscal policy in modern economies, fiscal stimulus policies have been enacted without a thorough understanding of the potency of these policy actions – particularly given the marked structural changes in many developed economies over the past two decades (such as the increased integration of global product and financial markets). The analytical model employed in this paper considers the extent to which private saving responds to changes in the total general government (Commonwealth, state and local) fiscal stance. While this framework lends itself towards explaining Ricardian equivalence effects, it can also be considered as a broad measure of the impact of fiscal policy on short- and long-run aggregate demand. The model is estimated using the autoregressive distributed lag approach (ARDL) to cointegration, which provides both short- and long-run coefficient estimates, but also provides the flexibility to accommodate the introduction of coefficients for structural breaks.

However, it is likely that the Australian economy has been subject to a substantial amount of structural change over the past 50 years. From the 1950s through to the early 1980s, the Australian economy was heavily regulated, with markets subject to price controls and tariff protection, a fixed exchange rate, and government controls on bank deposits, interest rates and credit. The 1980s saw a period of rapid reform, with the floating of the dollar, removal of restrictions on credit creation, interest rates, foreign capital inflows and other broader reforms around market pricing and removal (or lowering) of tariffs and subsidies. Not accounting for these changes could lead to spurious results in the econometric analysis. The Lee and Strazicich two-break unit root test is used to test the time series properties of the data.

The following section discusses the analytical model to be estimated in this paper, along with the expected signs of the explanatory variables. Section 3 uses unit root tests that allow for two endogenously determined structural breaks in the individual time series. The analytical model is then estimated through the ARDL approach to cointegration, which provides the flexibility to incorporate structural breaks and both stationary and non-stationary time series. Conclusions are presented in section 4.

2. Analytical framework

The relationship between private and public saving can be estimated through a model with the following functional form:

$$S_t^{priv} = \alpha_0 + \beta_0 S_t^{pub} + \phi_0 Z_t + e_t \quad (1)$$

where S_t^{priv} and S_t^{pub} denotes the ratio of net household plus net corporate saving (which gives total net private saving) to GDP, and the ratio of net general (Commonwealth, local and state) government saving to GDP, while Z_t is a vector of control variables. This reduced-form saving equation allows for the estimation of the private savings offset with a large number of control variables, and is similar to that used in previous empirical studies by Haque (et al: 1999); Masson (et al: 1998); Loayza (et al: 2000); Comley (et al: 2002); de Serres and Pelgrin (2003); and de Mello (et al: 2004). A similar specification of this model was applied to the United States by Cotis (et al: 2006).

The vector Z_t of control variables often includes conventional determinants of private saving, such as the real interest rate, inflation, household income, social assistance payments to households, changes in the terms of trade, and employment. Specifically:

$$Z_t = \{Y_t, AS_t, U_t, R_t, INF_t, TOT_t, FLIB_t, H_t, EQ_t\} \quad (2)$$

Where:

Y_t = Household gross disposable income;

AS_t = Social assistance benefits to household gross disposable income;

U_t = Unemployment rate;

R_t = Real interest rate;

INF_t = Inflation rate;

TOT_t = Terms of trade;

$FLIB_t$ = Net foreign liabilities (proxy for financial openness);

H_t = Australian house price index (proxy for wealth); and

EQ_t = Australian share price index (proxy for wealth).

The hypothesis of a strict private savings offset (Ricardian equivalence) would be supported if the coefficient on public saving in (5.1) above, $\beta_0 = -1$, controlling for the other private saving determinants. A negative coefficient on public savings, but less than 0, that is $(-1 < \beta_0 < 0)$ would indicate a partial savings offset, and that changes in the general government sector's fiscal stance has measurable impacts on the wider economy.

Cotis (et al: 2006) discuss a number of reasons which could give rise to a positive coefficient on public saving, that is, where $\beta_0 > 0$. Sources of changes in the fiscal position arise not only from changes to taxation arrangements, but also from changes in expenditures. For a positive private savings offset, public expenditures need to be considered complimentary, with a clear distinction between expenditures which are permanent, and those which are transitory. Permanent changes will tend to generate negative private savings offsets through the restrictions imposed by the intertemporal budget constraint. Temporary shocks in government spending, however, could generate positive private saving responses, particularly when households see public and private consumption as complements.[1]

[1] Specifically, this arises when the marginal utility of private consumption is positively affected by public spending. Government-subsidised health and education programmes, and government co-payment incentives for first home buyers, could provide examples of public and private complements in consumption.

The coefficient on household disposable income, Y_t, is expected to be positive. As household income may be considered a proxy for labour income in a standard life-cycle model of consumption, an increase in household disposable income is expected to increase private saving. Alternatively, households may suffer from consumption inertia and therefore take time to change their consumption patterns to new levels of income.

Social assistance payments to households, AS_t, are expected to negatively impact private savings. The existence of a welfare safety net in Australia is expected to crowd out precautionary motives for saving, and other privately-run alternatives that would encourage thrift.

Increasing levels of unemployment lowers disposable incomes, and, through a greater incidence of liquidity constraints, lowers saving. However, increases in unemployment may increase the need for precautionary saving. But as noted above the existence of welfare safety nets in Australia may crowd out precautionary motives for saving. Overall the coefficient on the unemployment rate, U_t, is expected to be of negative sign.

The effects of inflation, INF_t, and the real interest rate, R_t, are somewhat ambiguous, and depend largely on the extent of credit constraints and on the relative magnitude of income and substitution effects. Also, higher, and/or accelerating inflation erodes the real value of debt and raises private saving, but may also discourage holdings of assets that are not inflation-indexed.

Terms of trade shocks, TOT_t, are particularly relevant for Australia given a high reliance on commodity-based exports. This coefficient is expected to be positively correlated with private saving to the extent that terms of trade shocks are viewed as being temporary[2] through the Laursen-Harberger-Metzler effect.[3] Permanent shocks should not affect private saving.

As noted earlier, there has been a considerable amount of economic reform undertaken in Australia over the past three decades, most notably the reform of Australia's financial sector. Financial liberalisation in Australia occurred over a decade beginning in the early 1980s, with removals of restrictions on bank deposit rates and lending, and progressed to other significant reforms of which the most notable were the floating of the Australian dollar in December 1983, and deregulation of

[2] This historically has been the case with terms of trade shocks experienced with the Korean War, 1970s oil price shocks, and most recently the rapid industrialisation of China.

[3] According to the Laursen-Harberger-Metzler effect, an adverse (beneficial) transitory movement in the terms of trade results in a decrease (increase) in a country's current level of income which is larger than the decrease (increase) in its permanent income, causing a fall (rise) in aggregate saving.

home mortgage interest rates. This period of financial deregulation lead to a marked structural shift in the Australian economy and the development of sophisticated private markets for credit and financial risk management. More sophisticated private credit markets also enables greater access to personal credit, allowing households to smooth consumption.

As noted by de Mello (et al: 2004), the effect of financial liberalisation on private saving is ambiguous, because improved access to credit may boost consumption but the removal of bank portfolio allocation constraints, which often accompanies financial liberalisation, may result in higher real interest rates, which encourages saving. Given the large increase in foreign capital inflows following financial market deregulation, it may be reasonable to expect that any coefficient representing financial openness in Australia will have a negative sign.

However, adequate proxies for financial openness are difficult to measure, and somewhat subjective in nature. Proxies may include variables such as growth in M2 money and the ratio of household wealth to disposable income (as used by Comley et al: 2002). However, long time series for these variables are generally not available, with most measures only dating back to around the early 1980s at best. Alternative measures of financial openness have been suggested by Lane and Milesi-Ferretti (2001), and include measures based around countries' foreign assets and liabilities. Given this, Australia's level of net foreign liabilities may provide a good proxy for financial openness, particularly as foreign debt has increased substantially since the financial market reforms of the 1980s. Data on Australia's net foreign liabilities is also available back to the late 1950s.

Household wealth is expected to affect consumption/saving decisions based on permanent income considerations. Given that most Australian households have historically tended to hold their wealth through the family home, a house price index is used here as it is expected to provide a good proxy for household wealth in Australia.[4]

A share price index is also considered as an additional measure of private wealth. Historically, the proportion of Australian households participating directly in the sharemarket had been relatively low – until rising markedly over the past two decades. In 2006, approximately 38 per cent of the Australian population owned shares directly (Australian Securities Exchange: 2007),[5] which places Australia as having some of the highest (direct) share ownership rates in the world.

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[4] Around 70 per cent of Australian households owned their home in 2003-04 (Australian Bureau of Statistics: 2006).

[5] Australian households have also been undertaking greater ownership of equities indirectly through their superannuation savings. The Australian Securities Exchange (2007) estimates that in 2006, approximately 46 per cent of

Data has been sourced from the Australian Bureau of Statistics and the Reserve Bank of Australia. The sample size is large in both the number of observations (192) and the time period which is considered: 1959:3 – 2007:2.

3. Econometric methodology and empirical testing

Unit root tests

Previous studies which examine both Ricardian equivalence and fiscal multipliers usually have examined the time series properties of variables by using the Augmented Dickey-Fuller (ADF) (1979, 1981) or Philip-Perron (1988) unit root tests. However, these tests do not allow for the possibility of one or more structural breaks in the time series. Perron (1989) argued that in the presence of a structural break, the standard ADF tests are biased toward the non-rejection of the null hypothesis. The timing of the structural break in Perron’s procedure is assumed to be known *a priori* in accordance with underlying asymptotic distribution theory. Test statistics are constructed by adding dummy variables representing different intercepts and slopes, thereby extending the standard ADF procedure.

However, Perron’s technique was criticised by Christiano (1992) as specific break-dates may be chosen which support the researcher’s results and *a priori* expectations (i.e. data mining). Since then, a number of studies have been developed using different methodologies for endogenising the structural breaks. These studies include Banerjee (et al: 1992), Zivot and Andrews (1992), Perron (1997) Lumsdaine and Papell (1998), and Lee and Strazicich (2003).

Lee and Strazicich (2003) developed a two-break minimum Lagrange Multiplier (LM) unit root test where the alternative hypothesis implies trend stationarity (referred to by the authors as “trend-break stationarity”).[6] This test allows for up to endogenous structural breaks, which may occur in either the level or slope of a series. First consider the following data-generating process:

$$y_t = \delta' Z_t + e_t \tag{3}$$

$$e_t = \beta e_{t-1} + u_t \tag{4}$$

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the Australian population owned shares either directly via shares or indirectly via a managed fund or self managed superannuation fund.

[6] The null hypothesis is a unit root with breaks.

where y_t is the data series in period t, δ is a vector of coefficients, Z_t is a matrix of exogenous variables, and u_t is a standard white noise error term with zero mean and constant variance $u_t \sim \text{iid}\, N\left(0, \sigma^2\right)$, Z_t is described by $\left[1, t, D_{1t}, D_{2t}, DT_{1t}^*, DT_{2t}^*\right]'$, to allow for a constant term, linear time trend, and two structural breaks in level and trend where T_{Bj} denotes the time period of the breaks. Under the trend-break stationary alternative, the D_{jt} terms describe an intercept shift in the deterministic trend, where $D_{jt} = 1$ for $t \geq T_{Bj} + 1$, $j = 1, 2$, and zero otherwise; DT_{jt} describes a change in slope of the deterministic trend, where $DT_{jt} = 1$ for $t \geq T_{Bj} + 1$, $j = 1, 2$, and zero otherwise.

The two-break minimum LM unit root test statistic is obtained from the following regression:

$$\Delta y_t = d' \Delta Z_t + \phi \tilde{S}_{t-1} + \sum y_i \Delta \tilde{S}_{t-i} + \varepsilon_t \tag{5}$$

where $\tilde{S}_t = y_t - \tilde{\psi}_x - Z_t \tilde{\delta}$, $t = 2, ..., T$ and $\tilde{\psi}_t = y_1 - Z_1 \tilde{\delta}$. $\tilde{S}_t$ is a de-trended series of y_t using the coefficients in $\tilde{\delta}_t$, which are estimated from the regression in first differences of Δy_t on $\Delta Z_t = \left[1, \Delta D_{1t}, \Delta D_{2t}, \Delta DT_{1t}, \Delta DT_{2t}\right]$, y_1 and Z_1 are the first observations of y_t and Z_t, respectively, and Δ is the first difference operator. The standard white noise error terms is represented by ε_t. To correct for serial correlation, $\Delta \tilde{S}_{t-1}$, $I = 1, ..., k$ terms are included. The unit root hypothesis in equation (5) is equivalent to $\phi = 0$, and the test statistics are defined as:

$$\tilde{\rho} = T \cdot \tilde{\phi} \tag{6}$$

$\tilde{\tau}$ = t-statistic for the null hypothesis $\phi = 0$. (7)

To determine (endogenously) the location of the two breaks $\left(\lambda_j = T_{Bj} / T, j = 1, 2\right)$, the minimum LM unit root test uses a grid search procedure:

$$\text{LM}_\rho = \text{Inf}_\lambda \tilde{\rho}\left(\lambda\right) \tag{8}$$

$$\text{LM}_\tau = \text{Inf}_\lambda \tilde{\tau}\left(\lambda\right) \tag{9}$$

The LM test is corrected for autocorrelated errors by including lagged augmentation terms $\Delta\tilde{S}t-j, j=1,...,k$ as per the standard Augmented Dickey-Fuller test. The optimal lag length, k, is determined through the general to specific procedure of Perron (1989).

The Lee and Strazicich two-break LM unit root test was conducted in GAUSS using code provided by the authors. Again Models A and C were run, with lag lengths generated automatically through a general to specific procedure.

Critical values for the two-break LM unit root test also vary depending on the location of the breaks $\lambda=(T_{B1}/T,\ T_{B2}/T)$ and are symmetric around λ and $(1-\lambda)$. Critical values for the two-break minimum LM unit root test7 for Model C (intercept and trend break) are shown in Table 1 below, and are drawn from Table 2 in Lee and Strazicich (2003). Critical values for the two-break LM unit root test with change in intercept (Model A) at the 1%, 5% and 10% levels respectively are -4.55, -3.84, and -3.50.

Table 1: Critical values for the two-break LM unit root test (Model C)

Break points λ = (T_{B1}/T, T_{B2}/T)	Critical values		
	1%	5%	10%
λ = (0.2,0.4)	-6.16	-5.59	-5.27
λ = (0.2,0.6)	-6.41	-5.74	-5.32
λ = (0.2,0.8)	-6.33	-5.71	-5.33
λ = (0.4,0.6)	-6.45	-5.67	-5.31
λ = (0.4,0.8)	-6.42	-5.65	-5.32
λ = (0.6,0.8)	-6.32	-5.73	-5.32

Results from Model A (Table 2) suggest that with the exception of private saving (*PS*), all of the variables contain a unit root with at least on statistically significant structural break.

[7] Critical values are provided by Lee and Strazicich for T = 100. Unfortunately the authors do not provide critical values for larger or smaller sample sizes.

Table 2: Results of the two-break LM unit root test (Model A)

Variable	*k*	T_B	$T\ \phi = 0$	Inference
lnPS	0	1997:4#, 2001:1#	-5.4227*	Stationary
GS	4	1976:2#,1999:2#	-3.4204	Non-Stationary
lnY	8	1966:2,1987:3#	-1.7050	Non-Stationary
lnFLIB	7	1971:4#,1976:4#	-3.3650	Non-Stationary
lnU	4	1971:4,1974:4	-2.1289	Non-Stationary
R	4	1977:3#,1983:4#	-3.0836	Non-Stationary
INF	8	1975:3#,1983:2	-2.2589	Non-Stationary
lnAS	7	1992:1#, 1998:3#	-2.8172	Non-Stationary
lnTOT	7	1974:1#,1974:3	-2.4932	Non-Stationary
lnH	2	1973:3, 1980:4#	-1.9984	Non-Stationary
LnEQ	3	1983:2, 1988:1#	-3.3574	Non-Stationary

A maximum of 8 lags was specified in GAUSS. # Denotes significance at the 5% level for the break-point dummy variables. Critical value for $T\ \phi = 0$ is -3.84 at the 5% level.
* Denotes significance at the 5% level.

When allowing for a break in both the level and trend of the series, Model C (Table 3) produces quite different results. In contrast to Model A, the results in Table 3 suggest that household disposable incomes (*Y*), inflation (*INF*), and the terms of trade (*TOT*) are also stationary series.

Table 3: Results of the two-break LM unit root test (Model C)

Variable	*k*	T_B	$T\ \phi = 0$	Critical value break points	Inference
lnPS	0	1997:4#, 2001:1	-6.5213*	λ = (0.8,0.9)	Stationary
GS	7	1974:3#,1997:2#	-4.8116	λ = (0.3,0.8)	Non-Stationary
lnY	6	1973:2#, 1992:3	-6.7481*	λ = (0.3,0.7)	Stationary
lnFLIB	8	1973:1#,1986:1#	-4.3292	λ = (0.2,0.7)	Non-Stationary
lnU	6	1974:2#,1988:1#	-4.5601	λ = (0.3,0.6)	Non-Stationary
R	4	1973:2,1985:3	-4.9872	λ = (0.3,0.6)	Non-Stationary
INF	7	1973:2#,1991:4#	-6.6046*	λ = (0.3,0.7)	Stationary
lnAS	7	1970:1, 1976:1#	-5.4113	λ = (0.2,0.4)	Non-Stationary
lnTOT	4	1969:4#,1995:4#	-6.0485*	λ = (0.2,0.8)	Stationary
lnH	2	1972:2#, 1993:1#	-3.9289	λ = (0.3,0.7)	Non-Stationary
lnEQ	3	1973:2#, 1986:4#	-5.2620	λ = (0.3,0.6)	Non-Stationary

A maximum of 8 lags was specified in GAUSS. # Denotes significance at the 5% level for the break-point dummy variables. Critical values for $T\ \phi = 0$ are contained in Table 1. * Denotes significance at the 5% level.

When interpreting results from the LM unit root tests, the timing of structural breaks could be a useful guide for discerning the reliability and effectiveness of the procedure. Judgement of each model (A or C) based upon economic theory and historical events, such as policy changes and economic shocks (for example), can help to determine the timing of structural breaks, and whether these changes have been sudden or gradual. Results indicate that structural changes have generally coincided with a number of significant events over the past few decades, including:

- the 1960s resources boom;
- the expansion of social welfare programmes (Whitlam Government);
- oil price (terms of trade) and inflation shocks in the 1970s;
- the extensive period of financial deregulation in the 1980s; and
- the 1990-91 recession.

Cointegration

Conventional cointegration procedures (such as that of Johansen (1991, 1995), usually require that all data entering into an equation be non-stationary. As the unit root tests undertaken above suggest that the ratio of private saving to GDP is a stationary time series, conventional cointegration techniques cannot be used to estimate the analytical model. Further, the unit root tests also suggested that each data series contains at least one structural break. This further complicates the use of cointegration techniques as conventional cointegration methods cannot account for endogenous structural breaks. While recent econometric developments allow for cointegration testing in the presence of structural breaks, these techniques are currently in their early stages of development and often can only accommodate one structural break (earlier techniques such as that of Gregory and Hansen (1996) also require all data to be non-stationary). To overcome these difficulties, the analytical model will be estimated through the autoregressive distributed lag (ARDL) approach to cointegration (see Pesaran and Shin 1998; Pesaran et al 1996; and Pesaran et al 2001). This technique allows for a greater degree of flexibility – allowing for both stationary and non-stationary data – and can accommodate additional variables that can represent structural breaks.

Following Pesaran (et al: 2001) the ARDL technique involves two steps for estimating the cointegrating relationship. Under the first step, the existence of a long-run cointegrating relationship is tested. If a long-run cointegrating relationship is found, the second step involves estimating both the long and short-run coefficients. An intercept and trend term will be added to the estimation of the model – particularly as the united root tests considered in the previous section indicated that the dependent variable (*PS*) is stationary – and a visual inspection of the ratio of private saving to GDP indicates a considerable downward trend in the data series. Therefore, the ARDL model is a general ECM with unrestricted intercept and trend:

$$\Delta y_t = a_0 + a_1 t + \pi_{yy} y_{t-1} + \pi_{yx.x} x_{t-1} + \sum_{i=1}^{p-1} \Psi'_i \Delta z_{t-i} + w' \Delta x_t + \varepsilon_t \qquad (10)$$

where $a_0 \neq 0$ and $a_1 \neq 0$. As noted above, the first step of the ARDL procedure involves testing for a cointegrating relationship. This step tests for the *absence of any* level relation between y_t and x_t via the exclusion of the lagged level variables y_{t-1} and x_{t-1} in equation (7.6). Persaran (et al: 2001) define the F-statistic tests for the null hypotheses as $H_0^{\pi_{yy}} : \pi_{yy} = 0$, $H_0^{\pi_{yx.x}} : \pi_{yx.x} = 0'$ and the alternative hypotheses as $H_1^{\pi_{yy}} : \pi_{yy} \neq 0$, $H_1^{\pi_{yx.x}} : \pi_{yx.x} \neq 0'$. The joint null hypothesis for (10) is given by:

$$H_0 = H_0^{\pi_{yy}} \cap H_0^{\pi_{yx.x}} \tag{11}$$

and the alternative hypothesis is correspondingly stated as:

$$H_1 = H_1^{\pi_{yy}} \cup H_1^{\pi_{yx.x}} \tag{12}$$

The asymptotic distribution of the F-statistics are non-standard under the null hypothesis of no cointegrating relationship between the variables, regardless of the order of integration of the variables being considered. The calculated F-statistic is compared with the critical values provided in Pesaran (et al: 2001). The null hypothesis of no cointegration is rejected if the calculated F-statistic is greater than the upper bound critical value. If the calculated F-statistic falls below the lower bound, then the null hypothesis of no cointegration cannot be rejected. The result is inconclusive if the calculated F-statistic lies between the upper and lower bound critical values.

The ARDL specification for equation (1) is as follows:

$$\begin{aligned}
\Delta PS_t &= \alpha_0 + \alpha_1 t + \sum_{i=1}^{p} \delta_i \Delta PS_{t-i} + \sum_{i=1}^{p} \beta_i \Delta GS_{t-i} + \sum_{i=1}^{p} \phi_i \Delta Y_{t-i} + \sum_{i=1}^{p} \varphi_i \Delta AS_{t-i} + \\
&\sum_{i=1}^{p} \gamma_i \Delta U_{t-i} + \sum_{i=1}^{p} \tau_i \Delta R_{t-i} + \sum_{i=1}^{p} \upsilon_i \Delta INF_{t-i} + \sum_{i=1}^{p} \rho_i \Delta TOT_{t-i} + \sum_{i=1}^{p} \psi_i \Delta FLIB_{t-i} + \\
&\sum_{i=1}^{p} \xi_i \Delta H_{t-i} + \sum_{i=1}^{p} \omega_i \Delta EQ_{t-i} + \lambda_1 PS_{t-1} + \lambda_2 GS_{t-1} + \lambda_3 Y_{t-1} + \lambda_4 AS_{t-1} + \lambda_5 U_{t-1} + \\
&\lambda_6 R_{t-1} + \lambda_7 INF_{t-1} + \lambda_8 TOT_{t-1} + \lambda_9 FLIB_{t-1} + \lambda_{10} H_{t-1} + \lambda_{11} EQ_{t-1} + u_t
\end{aligned} \tag{13}$$

where S_t^{priv} and S_t^{pub} have been shortened to *PS* and *GS* respectively. In the ARDL specification above, the summation signs represent the error correction dynamics, while the second section of the equation, denoted by λ_i, represents the long run relationship. The null hypothesis of no cointegration in equation (13) is given by:

$$H_0 : \lambda_1 = \lambda_2 = \lambda_3 = \lambda_4 = \lambda_5 = \lambda_6 = \lambda_7 = \lambda_8 = \lambda_9 = \lambda_{10} = \lambda_{11} = 0$$

or equivalently as:

$$F_{PS}(PS \mid GS, Y, AS, U, R, INF, TOT, FLIB, H, EQ)$$

The corresponding alternative hypothesis is:

$$\lambda_1 \neq 0, \lambda_2 \neq 0, \lambda_3 \neq 0, \lambda_4 \neq 0, \lambda_5 \neq 0, \lambda_6 \neq 0, \lambda_7 \neq 0, \lambda_8 \neq 0, \lambda_9 \neq 0, \lambda_{10} \neq 0, \lambda_{11} \neq 0$$

As noted earlier, the relevant test statistic here is the F-statistic for the joint significance of the coefficients, and as we are dealing with quarterly data, a maximum of 4 lags is included.

Table 4: Results from bounds test on equation (13) – 1959:3 to 2006:2

Dep. Var.	**F-statistic**	**Probability**	**Conclusion**
$F_{PS}(PS \mid GS, Y, AS, U, INF, R, TOT, FLIB, H, EQ)$	3.4906*	0.000	Cointegration
$F_{GS}(GS \mid PS, Y, AS, U, INF, R, TOT, FLIB, H, EQ)$	2.4126	0.009	Inconclusive
$F_{Y}(Y \mid PS, GS, AS, U, INF, R, TOT, FLIB, H, EQ)$	2.2677	0.015	No cointegration
$F_{AS}(AS \mid PS, GS, Y, U, INF, R, TOT, FLIB, H, EQ)$	2. 4465	0.008	Inconclusive
$F_{U}(U \mid PS, GS, Y, AS, INF, R, TOT, FLIB, H, EQ)$	3.0196	0.001	Inconclusive
$F_{R}(R \mid PS, GS, Y, AS, U, INF, TOT, FLIB, H, EQ)$	2.1676	0.020	No cointegration
$F_{INF}(INF \mid PS, GS, Y, AS, U, R, TOT, FLIB, H, EQ)$	2. 0838	0.026	No cointegration
$F_{TOT}(TOT \mid PS, GS, Y, AS, U, INF, R, FLIB, H, EQ)$	3.5018*	0.000	Cointegration
$F_{FLIB}(FLIB \mid PS, GS, Y, AS, U, INF, R, TOT, H, EQ)$	1.7875	0.063	No cointegration
$F_{H}(H \mid PS, GS, Y, AS, U, INF, R, TOT, FLIB, EQ)$	3.1870	0.001	Inconclusive
$F_{EQ}(EQ \mid PS, GS, Y, AS, U, INF, R, TOT, FLIB, H)$	1.8996	0.045	No cointegration

Asymptotic critical value bounds are obtained from Table CI(iii), Case V: unrestricted intercept and unrestricted trends for *k*=10 (Persaran et al: 2001). Lower bound I(0)=2.33 and Upper bound I(1)=3.46 at the 5% significance level. * Denotes significance at the 5% level. ** Denotes significance at the 1% level.

Where private savings is the dependent variable, the calculated F-statistic of 3.4906 is greater than the upper bound critical value at the 5 per cent level, which rejects the null hypothesis of no cointegration – implying a long-run level relationship between the variables (Table 4). Considering the possibility of reverse causation, where government savings is the long-run dependent variable, the calculated F-statistic of 2.4126 falls into the inconclusive region. Consequently, reverse causation cannot be ruled-out. Where the cointegration tests are undertaken with different dependent variables, the results also suggest a long-run relationship between the variables, and that

Y, *R*, *INF*, *FLIB*, and *EQ* act as the long-run forcing variables for private saving. While results in Table 1 show inconclusive results for social assistance payments (*AS*), unemployment (*U*), and house prices (*H*), the subsequent estimations of the short- and long-run parameters may yield further information on the significance of these variables.

The structural breaks identified above may be accounted for by the inclusion of break-point dummy variables in the ARDL model. The structural breaks to be included in the ARDL specification are:

> B1969:1 = 1960s resources boom;
>
> B1973:3 = expansion of social welfare programmes (Whitlam Government); oil price shocks and inflation[8];
>
> B1984:1 = floating of the Australian dollar[9], including broader financial market liberalisation; and
>
> B1990:1 = onset of recession in the early 1990s.

Estimation results

The estimated long-run coefficient estimates for equation (13) are provided in Table 5.[10] With the exception of the unemployment rate (*U*), all variables have the expected sign, although the wealth variables will be discussed in greater detail below. For the level of government savings (*GS*), the results suggest that over the long run, changes in general government saving are offset by changes in private savings by almost half (-0.44). This implies that the behavioural response of households and corporations is not fully Ricardian, and that fiscal policy has a (partial) flow through to the real economy – potentially impacting output, real interest rates, the exchange rate, and subsequently the current account. The value of this coefficient is also similar to the results of Comley (et al: 2002), who estimated a long-run private savings offset coefficient for Australia of -0.5. However, it is important to note here that Comley's estimated long-run coefficient was not statistically significant, possibly due to having a much smaller sample (1981:1-2002:2).

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[8] While two breaks may have been included for each of these effects, the close proximity of both breaks would mean that the inclusion of separate dummy variables for each could increase the likelihood of serial correlation in the regression estimates.

[9] The floating of the Australian dollar is considered to be the most significant of the broader financial market reforms undertaken over the decade from the late 1970s though to the late 1980s.

[10] The appropriate lag length was chosen according to the Schwarz Bayesian Criterion.

The estimated Australian private savings offset of -0.44 is however lower than some estimates derived through international panel studies. Considering private savings across a panel of 21 OECD countries, de Mello (et al: 2004) estimated a long-run private savings offset coefficient of around -0.75; implying that changes in the fiscal stance are almost fully offset by corresponding changes in private saving. Following an analytical model similar to that used here, and to that employed by de Mello (et al: 2004), Cotis (et al: 2006) estimated a long-run private savings offset of around two thirds for a panel of 16 OECD countries. Isolating impacts on the United States, Cotis (et al: 2006) estimated a positive long-run private savings coefficient – implying that US households behave in a non-Ricardian manner.[11]

Table 5: Estimated long-run coefficients for equation (13)

ARDL (1,0,1,2,0,0,0,0,0,0,0) selected lags based on Schwarz Bayesian Criterion

Variable	**Coefficient**	**Standard Error**	**T-Ratio**	**Probability**
Constant	-0.2564	0.1157	-2.2152*	0.028
Trend	0.0003	0.0003	0.9729	0.332
GS	-0.4438	0.1178	-3.7673**	0.000
Y	0.4241	0.1409	3.0100*	0.003
U	0.1571	0.2082	0.7542	0.452
R	0.0301	0.0729	0. 4128	0.680
INF	-0.1460	0.1094	-1.3340	0.184
AS	-0.4579	0.2145	-2.1342*	0.034
TOT	0.0008	0.0002	3.9830**	0.000
FLIB	-0.0364	0.0155	-2.3410*	0.020
H	-0.0066	0.0127	-0.5153	0.607
EQ	0.0179	0.0106	1.6806	0.095
B1969	0.0029	0.0062	0.4685	0.640
B1973	-0.0161	0.0106	-1.5082	0.133
B1984	-0.0035	0.0066	-0.5388	0.591
B1990	-0.0151	0.0078	-1.9209	0.056

* Denotes significance at the 5% level. ** Denotes significance at the 1% level.

For the remaining variables in Table 5, the results indicate that for a 1 per cent rise in household gross disposable income (*Y*), the ratio of private savings to GDP increases by 0.42 per cent. This also implies a marginal propensity to consume of approximately 0.6 – which is consistent with Australian National Accounts data (which indicates that the consumption share of GDP in Australia

[11] Changes in public saving result from both taxation and expenditure. While permanent expenditures will generate an increase in private saving through the intertemporal budget constraint, temporary expenditure shocks can generate positive private saving offsets (particularly when households see public and private consumption as complements; for example, rebates and co-payments).

of 60 per cent). Rising levels of social assistance payments to households (*AS*) are estimated to have a negative impact on private savings over the long-run, with the ratio of private saving to GDP declining by around 0.46 per cent for each one per cent increase in social assistance payments to households. Australia's terms of trade are (*TOT*) is estimated to have a small, although statistically significant, positive impact on private savings over the long run. As expected, financial liberalisation has a negative impact on private savings over the long run. For the unemployment rate (*U*), the real interest rate (*R*), and inflation (*INF*), the results in Table 5 indicate that these variables do not have a statistically significant long-run impact on the level of private saving in Australia.

Both of the wealth variables present some interesting results. Changes in the prices of household assets (and the returns derived from these) will affect household consumption and saving. Additionally, as the dependent variable is private saving (which includes corporate saving), changes in wealth will also affect business borrowing and investment decisions. Results here indicate that wealth from housing does not exert a statistically significant impact on private saving over the long run, although it is of the expected sign. Given that most Australian's hold wealth through the family home, this is somewhat surprising. Equity prices appear to have had a statistically significant (albeit at the 10 per cent level) impact on private saving over the long run. The positive sign of this coefficient is curious, and suggests that for a 1 per cent rise in equity prices, the ratio of private saving to GDP rises by around 0.02 per cent. This positive response may be somewhat indicative of the broad shift toward equity investment, particularly the indirect investment occurring through households' accumulation of assets in superannuation.

Of the dummy variables included in the estimation, only the structural break coinciding with the early 1990s recession (*B1990*) is estimated to have had a statistically significant (at the 10 per cent level) long-run impact on the private savings ratio.

The short-run error correction estimates are presented in Table 6. In the short-run, the error correction equation indicates a private saving offset of one quarter (-0.25) to changes in government saving. The error correction term, $ecm(-1)$, is of the correct sign and statistically significant – indicating that deviations from the long-run rate of private saving are corrected by over 50 per cent in the next period, which is a relatively fast pace of adjustment back to equilibrium. While the unemployment rate (*U*) was statistically insignificant in the long-run relationship, the estimated coefficient here is of the correct sign, and significant at the 10 per cent level, whilst the lagged value of unemployment is significant at the 1 per cent level. This suggests that the unemployment rate negatively impacts private saving in the short-run only, which would be consistent with the impact of temporary shocks to output.

Table: 6 Error correction representation of equation (13)

ARDL (1,0,1,2,0,0,0,0,0,0,0) selected lags based on Schwarz Bayesian Criterion

Variable	**Coefficient**	**Standard Error**	**T-Ratio**	**Probability**
Constant	-0.1469	0.0692	-2.1224*	0.035
Trend	0.0002	0.0002	0.9838	0.327
ΔGS	-0.2544	0.0675	-3.7637**	0.000
ΔY	0.5249	0.0747	7.0231**	0.000
ΔU	-0.3919	0.2228	-1.7593	0.080
$\Delta U(-1)$	-0.7711	0.2184	-3.5302**	0.001
ΔR	0.0172	0.0419	0.4119	0.681
ΔINF	-0.0804	0.0593	-1.3568	0.177
ΔAS	-0.2624	0.1208	-2.1718*	0.031
ΔTOT	0.0004	0.0001	3.8787**	0.000
$\Delta FLIB$	-0.0208	0.0086	-2.4049*	0.017
ΔH	-0.0037	0.0072	-0.5176	0.605
ΔEQ	0.0102	0.0060	1.7059	0.090
$\Delta B1969$	0.0016	0.0036	0.4645	0.643
$\Delta B1973$	-0.0092	0.0061	-1.5230	0.130
$\Delta B1984$	-0.0020	0.0038	-0. 5415	0.589
$\Delta B1990$	-0.0087	0.0047	-1.8481	0.066
$ecm(-1)$	-0.5732	0.0597	-9.6020**	0.000

* Denotes significance at the 5% level. ** Denotes significance at the 1% level.

$ecm = PS + 0.444 * GS - 0.424 * Y - 0.157 * U - 0.03 * R + 0.160 * INF + 0.458 * AS - 0.0007 * TOT + 0.036 * FLIB +$
$0.007 * H - 0.018 * EQ + 0.256 * \text{Constant} - 0.0003 * \text{Trend} - 0.003 * B1969 + 0.016 * B1973 + 0.004 * B1984 + 0.015 * B1990$

$R^2 = 0.6249$ $\quad \bar{R}^2 = 0.5844$ $\quad$ F-stat $F(17,168) = 17.3865\left[0.000\right]$ $\quad$ SER $= 0.0082$

RSS $= 0.011$ $\quad$ DW-statistic $= 2.0817$

Short-run coefficient estimates for household gross disposable income (*Y*), and the terms of trade (*TOT*) are significant at the 1 per cent level, while social assistance payments (*AS*), and financial openness (FLIB) are significant at the 5 per cent level. Similar to the long-run results, the estimated short-run coefficients for the real interest rate (*R*), inflation (*INF*) and break-point dummy variables *B1969*, *B1973*, and *B1984* are statistically insignificant. The short-run results also indicate that housing wealth is not statistically insignificant, while wealth from equities appears to bear a statistically significant influence (at the 10 per cent level) on the ratio of private saving to GDP in Australia.

Diagnostic statistics from the estimations are positive (Table 7), indicating that the error terms do not suffer from serial correlation, and are normally distributed. The model specification also satisfies the RESET test for omitted variables and functional form.

Table 7: Diagnostic tests on equation (13)

LM Test Statistics	χ^2 **statistic**	**Probability**
Serial correlation **a** $\chi^2(4)$	3.3784	0.497
Normality **b** $\chi^2(2)$	1.5196	0.468
Functional form **c** $\chi^2(1)$	0.0038	0.951
Heteroscedasticity **d** $\chi^2(1)$	0.0179	0.893

* Denotes significance at the 5% level. ** Denotes significance at the 1% level.
a Breusch-Godfrey LM test for serial correlation. **b** Jarque-Bera normality test.
c Ramsey RESET test for omitted variables/functional form. **d** White test for heteroscedasticity.

Two subsample estimations for equation (13) will now be undertaken. These cover the period 1959:3 – 1983:4, while the second period is over 1984:1 – 2006:2. This will attempt to account for the effects of financial market liberalisation, and a move toward a greater integration of the Australian economy into the global financial system – particularly as the break-point dummy variable (*B1984*) was not statistically significant in the earlier analysis.[12]

Over the first subsample period, the Australian economy was highly regulated, with a fixed exchange rate, tariff controls, and other regulations over the financial system such as controls on bank lending, deposits, and some interest rates (such as mortgage interest rates, overnight money market rates, and deposit rates). Since the floating of the Australian dollar and associated financial market reforms, foreign capital inflows into Australia have increased markedly, and there has been a commensurate increase in financial market innovation. This integration into global capital markets may have dampened the impact of fiscal policy on the economy. These reforms have also occurred in concert with other reforms in the labour market, tariff reform, the establishment of free trade arrangements with some countries, a national competition policy agenda, fiscal consolidation, privatisation of government business enterprises, and the introduction of inflation targeting.

Private saving offsets – 1959:3 to 1983:4

Cointegration tests where private saving (*PS*) is the dependent variable yield an F-statistic of 3.7095, which is greater than the upper bound critical value at the 5 per cent level – implying that the long-run level relationship between these variables is still observed over the first subsample period (Table 8). However, where government savings is the dependent variable, the calculated F-statistic again falls into the inconclusive zone.

[12] As the financial reforms were phased over the 1980s, with the floating of the Australian dollar one of several major reforms, the insignificance of this dummy variable is not that surprising. This implies that a gradual structural change may have been occurring as opposed to a sudden level shift.

Table 8: Results from bounds test on equation (13) – 1959:3 to 1983:4

Dep. Var.	F-statistic	Probability	Conclusion
$F_{PS}(PS \mid GS, Y, AS, U, INF, R, TOT, FLIB, H, EQ)$	3.7095*	0.001	Cointegration
$F_{GS}(GS \mid PS, Y, AS, U, INF, R, TOT, FLIB, H, EQ)$	2.5843	0.016	Inconclusive
$F_{Y}(Y \mid PS, GS, AS, U, INF, R, TOT, FLIB, H, EQ)$	1.1575	0.349	No cointegration
$F_{AS}(AS \mid PS, GS, Y, U, INF, R, TOT, FLIB, H, EQ)$	3.2765	0.003	No cointegration
$F_{U}(U \mid PS, GS, Y, AS, INF, R, TOT, FLIB, H, EQ)$	2.1103	0.045	No cointegration
$F_{R}(R \mid PS, GS, Y, AS, U, INF, TOT, FLIB, H, EQ)$	2.1373	0.043	No cointegration
$F_{INF}(INF \mid PS, GS, Y, AS, U, R, TOT, FLIB, H, EQ)$	1.6689	0.121	No cointegration
$F_{TOT}(TOT \mid PS, GS, Y, AS, U, INF, R, FLIB, H, EQ)$	2.4355	0.022	Inconclusive
$F_{FLIB}(FLIB \mid PS, GS, Y, AS, U, INF, R, TOT, H, EQ)$	2.2704	0.032	No cointegration
$F_{H}(H \mid PS, GS, Y, AS, U, INF, R, TOT, FLIB, EQ)$	2.7366	0.011	Inconclusive
$F_{EQ}(EQ \mid PS, GS, Y, AS, U, INF, R, TOT, FLIB, H)$	3.7878	0.001	Cointegration

Asymptotic critical value bounds are obtained from Table CI(iii), Case V: unrestricted intercept and unrestricted trends for *k*=10 (Persaran et al: 2001). Lower bound I(0)=2.43 and Upper bound I(1)=3.56 at the 5% significance level. * Denotes significance at the 5% level. ** Denotes significance at the 1% level.

For the ARDL estimation over the period 1959:3-1983:4, initial results for equation (13) were not positive, and indicated that the errors of the estimated ARDL were serially correlated and not normally distributed. Additionally, the estimated trend coefficient was of the wrong sign. The trend coefficient was dropped, along with estimated coefficients for the real interest rate (*R*), inflation (*INF*), financial openness (*FLIB*), and the break-point dummy variables (*B1969*) and (*B1973*) as these variables were all statistically insignificant. Serial correlation was still apparent in the model, and despite theory suggesting that wealth effects may explain some of the variation in private saving behaviour; both the house and equity price series were also dropped from the model. Removing these improved the results markedly, with the Jarque-Bera test indicating that the residuals were normally distributed, while the Breusch-Godfrey LM test suggested that serial correlation had also been alleviated. This left the following specification for the subsample ARDL:

$$\Delta PS_t = \alpha_0 + \sum_{i=1}^{p} \delta_i \Delta PS_{t-i} + \sum_{i=1}^{p} \beta_i \Delta GS_{t-i} + \sum_{i=1}^{p} \phi_i \Delta Y_{t-i} + \sum_{i=1}^{p} \gamma_i \Delta U_{t-i} +$$

$$\sum_{i=1}^{p} \varphi_i \Delta AS_{t-i} + \sum_{i=1}^{p} \rho_i \Delta TOT_{t-i} + \lambda_1 PS_{t-1} + \lambda_2 GS_{t-1} +$$

$$\lambda_3 Y_{t-1} + \lambda_4 U_{t-1} + \lambda_5 AS_{t-1} + \lambda_6 TOT_{t-1} + u_t \tag{14}$$

The estimated long-run coefficient estimates for equation (14) are provided in Table 9. For the ratio of government saving to GDP (*GS*) over the period 1959:3-1983:4, the estimated coefficient is -0.39, which is somewhat lower than the full sample estimation. This potentially suggests that with a lower private saving offset, fiscal policy may have exerted a larger impact on the real economy during this period. Such a result would be consistent with the structure of the economy at that time (markets being subject to a greater degree of regulation, and less exposure to international capital and price movements) and confirms a priori expectations regarding these policy impacts.

Table 9: Estimated long-run coefficients for equation (14)

ARDL (1,0,1,0,2,0) selected lags based on Schwarz Bayesian Criterion				
Variable	**Coefficient**	**Standard Error**	**T-Ratio**	**Probability**
Constant	-0.2085	0.0648	-3.2159**	0.002
GS	-0.3994	0.1861	-2.1455*	0.035
Y	0.3906	0.0700	5.5746**	0.000
U	-0.1998	0.2475	-0.8075	0.422
AS	-0.2438	0.2855	-0.8539	0.395
TOT	0.0007	0.0003	2.6296**	0.010

* Denotes significance at the 5% level. ** Denotes significance at the 1% level.

A one per cent rise in household gross disposable income (*Y*) is estimated to raise the ratio of private saving to GDP by 0.39 per cent over the first subsample, which is slightly higher than for the full sample estimation. The terms of trade (*TOT*) is statistically significant, but is estimated to only exert an extremely small impact on the private saving to GDP ratio. As expected, over this subsample the ratio of social assistance payments to household gross disposable income (*AS*) and the unemployment rate (*U*) are estimated to have had a statistically insignificant long-run impact on private saving.

The short-run error correction estimates are presented in Table 10. In the short-run, the error correction equation indicates a private saving offset of -0.23. The error correction term, $ecm(-1)$, is of the correct sign and statistically significant – indicating that deviations from the long-run rate of private savings are corrected by over 50 per cent in the next period. Household gross disposable income, (*Y*), is statistically significant (at the one per cent level) while the estimated coefficient for social assistance payments (*AS*) is markedly higher in the short-run, and includes an additional lag coefficient for adjustment. The larger sign of this coefficient in the short run may again be explained by the steep rise in the unemployment rate in 1974, then rising again in 1983 (where the unemployment rate reached 10.2 per cent in the September quarter 1983) – suggesting that households were more dependent on the welfare safety net over this period. However, it is interesting that the results indicate that the unemployment rate is statistically insignificant in both the long and short-run estimations. Prior to the large rise in unemployment during the 1970s, the unemployment rate averaged 2 per cent over the 1960s. The introduction of expanded social welfare programmes by the Whitlam government almost coincided with a steep rise in unemployment in 1974, which may explain this curio.[13]

Table 10: Error correction representation of equation (14)

ARDL (1,0,1,0,2,0) selected lags based on Schwarz Bayesian Criterion				
Variable	**Coefficient**	**Standard Error**	**T-Ratio**	**Probability**
Constant	-0.1216	0.0413	-2.9407*	0.004
ΔGS	-0.2329	0.1021	-2.2812*	0.025
ΔY	0.4916	0.0806	6.0980**	0.000
ΔU	-0.1165	0.1462	-0.7968	0.428
ΔAS	-1.4175	0.3407	-4.1602**	0.000
$\Delta AS(-1)$	-0.7800	0.3133	-2.4892*	0.015
ΔTOT	0.0004	0.0002	2.6569*	0.009
$ecm(-1)$	-0.5831	0.0945	-6.1691**	0.000

* Denotes significance at the 5% level. ** Denotes significance at the 1% level.

$ecm = PS + 0.399 * GS - 0.391 * Y + 0.199 * U + 0.244 * AS - 0.0007 * TOT + 0.209 * \text{Constant}$

$R^2 = 0.7104$ $\bar{R}^2 = 0.6800$ F-stat $F(7, 88) = 30.1357\left[0.000\right]$ $\text{SER} = 0.0078$

RSS = 0.0053 DW-statistic = 1.9847

Diagnostic statistics for the error correction mechanism (Table 11) are positive and indicate that the model is correctly specified. The error terms are normally distributed and the Breusch-Godfrey LM test indicates that no serial correlation is present.

Table 11: Diagnostic tests on equation (14)

LM Test Statistics	χ^2 **statistic**	**Probability**
Serial correlation **a** $\chi^2(4)$	2.8417	0.585
Normality **b** $\chi^2(2)$	3.7570	0.153
Functional form **c** $\chi^2(1)$	0.6502	0.420
Heteroscedasticity **d** $\chi^2(1)$	0.4577	0.499

* Denotes significance at the 5% level. ** Denotes significance at the 1% level.
a Breusch-Godfrey LM test for serial correlation. **b** Jarque-Bera normality test.
c Ramsey RESET test for omitted variables/functional form. **d** White test for heteroscedasticity.

[13] In the absence of social welfare arrangements, the coefficient on unemployment could in fact be positive; inferring that a rise in unemployment spurs an increase in precautionary saving.

Private saving offsets – 1984:1 to 2006:2

Cointegration tests where private saving is the dependent variable yield an F-statistic of 2.766, which falls within the inconclusive range of the critical values at the 5 per cent level (Table 12). Results from the bounds test also suggest reverse causation where government savings is the dependent variable. Given the overall sample results presented earlier lend support to cointegration the ARDL estimations will still be undertaken. The inconclusive result (and the suggested reverse causation with government savings as the dependent variable) may in fact suggest that financial liberalisation in Australia, leading to deeper and more open capital markets, has eroded the transmission of changes in the government's fiscal stance.

Table 12: Results from bounds test on equation (13) – 1984:1 to 2006:2

Dep. Var.	**F-statistic**	**Probability**	**Conclusion**
$F_{PS}(PS \mid GS, Y, AS, U, INF, R, TOT, FLIB, H, EQ)$	2.7660	0.012	Inconclusive
$F_{GS}(GS \mid PS, Y, AS, U, INF, R, TOT, FLIB, H, EQ)$	4.7084	0.000	Cointegration
$F_{Y}(Y \mid PS, GS, AS, U, INF, R, TOT, FLIB, H, EQ)$	2.1220	0.047	Inconclusive
$F_{AS}(AS \mid PS, GS, Y, U, INF, R, TOT, FLIB, H, EQ)$	2.6908	0.014	Inconclusive
$F_{U}(U \mid PS, GS, Y, AS, INF, R, TOT, FLIB, H, EQ)$	3.1875	0.005	Inconclusive
$F_{R}(R \mid PS, GS, Y, AS, U, INF, TOT, FLIB, H, EQ)$	2.6692	0.014	Inconclusive
$F_{INF}(INF \mid PS, GS, Y, AS, U, R, TOT, FLIB, H, EQ)$	2.3367	0.029	No cointegration
$F_{TOT}(TOT \mid PS, GS, Y, AS, U, INF, R, FLIB, H, EQ)$	3.6749	0.002	Cointegration
$F_{FLIB}(FLIB \mid PS, GS, Y, AS, U, INF, R, TOT, H, EQ)$	2.7118	0.013	Inconclusive
$F_{H}(H \mid PS, GS, Y, AS, U, INF, R, TOT, FLIB, EQ)$	2.3422	0.029	Inconclusive
$F_{EQ}(EQ \mid PS, GS, Y, AS, U, INF, R, TOT, FLIB, H)$	4.4042	0.000	Cointegration

Asymptotic critical value bounds are obtained from Table CI(iii), Case V: unrestricted intercept and unrestricted trends for *k*=10 (Persaran et al: 2001). Lower bound I(0)=2.43 and Upper bound I(1)=3.56 at the 5% significance level. * Denotes significance at the 5% level. ** Denotes significance at the 1% level.

After initially estimating equation (13), the results suggested that social assistance payments as a proportion of household disposable income (*AS*), inflation (*INF*), the real interest rate (*R*) and the break-point dummy variable coinciding with the early 1990s recession (*B1990*) were statistically insignificant. The following ARDL was estimated:

$$\Delta PS_t = \alpha_0 + \alpha_1 t + \sum_{i=1}^{p} \delta_i \Delta PS_{t-i} + \sum_{i=1}^{p} \beta_i \Delta GS_{t-i} + \sum_{i=1}^{p} \phi_i \Delta Y_{t-i} +$$

$$\sum_{i=1}^{p} \gamma_i \Delta U_{t-i} + \sum_{i=1}^{p} \rho_i \Delta TOT_{t-i} + \sum_{i=1}^{p} \psi_i \Delta FLIB_{t-i} + \sum_{i=1}^{p} \xi_i \Delta H_{t-i} +$$

$$\sum_{i=1}^{p} \omega_i \Delta EQ_{t-i} + \lambda_1 PS_{t-1} + \lambda_2 GS_{t-1} + \lambda_3 Y_{t-1} + \lambda_4 U_{t-1} +$$

$$\lambda_5 TOT_{t-1} + \lambda_6 FLIB_{t-1} + \lambda_7 H_{t-1} + \lambda_8 EQ_{t-1} + u_t \qquad (15)$$

The estimated long-run coefficient estimates are provided in Table 13. For the ratio of government saving to GDP (*GS*) over the period 1984:1-2006:2, the estimated coefficient is -0.39, and statistically significant only at the 10 per cent level. For the other variables, a one per cent rise in household gross disposable income (*Y*) is estimated to raise the ratio of private savings to GDP by 0.43 per cent. Net foreign liabilities (*FLIB*) are also significant at the 1 per cent level – and indicate that Australian financial markets have become more integrated with global capital flows. The long-run coefficient on the terms of trade (*TOT*) is slightly higher than the previous estimations, which possibly indicates that as Australia has become more integrated with the global economy and that international price determination for traded goods may be exerting a greater influence over household incomes, consumption and saving. The house price index is now statistically insignificant, while equity prices remain significant at the 10 per cent level.

Table 13: Estimated long-run coefficients for equation (15)

ARDL (2,1,0,2,0,1,0,0) selected lags based on Schwarz Bayesian Criterion

Variable	**Coefficient**	**Standard Error**	**T-Ratio**	**Probability**
Constant	-0.3901	0.2294	-1.7001	0.093
Trend	-0.0006	0.0004	-1.2942	0.200
GS	-0.3855	0.2386	-1.6160	0.110
Y	0.4338	0.2371	1.8295	0.071
U	0.4296	0.3463	1.2407	0.219
TOT	0.0012	0.0003	3.5862**	0.001
FLIB	-0.0700	0.0227	-3.0776**	0.003
H	0.0202	0.0242	0.8328	0.408
EQ	0.0341	0.0187	1.8232	0.072

* Denotes significance at the 5% level. ** Denotes significance at the 1% level.

The short-run error correction estimates are presented in Table 14. In the short-run, the error correction equation indicates a private savings offset of -0.40 to changes in government saving, which is both statistically significant and roughly equivalent to the estimated long-run coefficient.

The error correction term, $ecm(-1)$, is of the correct sign and statistically significant – indicating that deviations from the long-run rate of private savings are corrected by around 50 per cent in the next period.

Table 14: Error correction representation of equation (15)

ARDL (2,1,0,2,0,1,0,0) selected lags based on Schwarz Bayesian Criterion				
Variable	**Coefficient**	**Standard Error**	**T-Ratio**	**Probability**
Constant	-0.1816	0.1008	-1.8006	0.076
Trend	-0.0003	0.0002	-1.3609	0.177
$\Delta PS(-1)$	-0.1769	0.0805	-2.1976*	0.031
ΔGS	-0.3977	0.1049	-3.7921**	0.000
ΔY	0.2019	0.1110	1.8187	0.073
ΔU	-0.4623	0.3714	-1.2445	0.217
$\Delta U(-1)$	-1.1101	0.3230	-3.4367**	0.001
ΔTOT	0.0006	0.0002	3.4544**	0.000
$\Delta FLIB$	-0.0776	0.0189	-4.0914*	0.000
ΔH	-0.0094	0.0108	0.8707	0.387
ΔEQ	-0.0158	0.0078	2.0123*	0.048
$ecm(-1)$	-0.4654	0.0906	-5.1340**	0.000

* Denotes significance at the 5% level. ** Denotes significance at the 1% level.

$ecm = PS + 0.385 * GS - 0.434 * Y - 0.429 * U - 0.001 * TOT + 0.070 * FLIB - 0.020 * H - 0.034 * EQ +$
$0.390 * INPT + 0.006 * \text{Trend}$

$R^2 = 0.6690$ $\bar{R}^2 = 0.6072$ F-stat $F(11, 78) = 13.7805\,[0.000]$ SER $= 0.0073$

RSS $= 0.0041$ DW-statistic $= 2.0543$

Diagnostic statistics for the error correction mechanism (Table 15) are positive, and indicate that the model is correctly specified.

Table 15: Diagnostic tests on equation (15)

LM Test Statistics	χ^2 **statistic**	**Probability**
Serial correlation **a** $\chi^2(4)$	1.8555	0.762
Normality **b** $\chi^2(2)$	0.4971	0.780
Functional form **c** $\chi^2(1)$	0.4583	0.498
Heteroscedasticity **d** $\chi^2(1)$	0.3776	0.539

* Denotes significance at the 5% level. ** Denotes significance at the 1% level.
a Breusch-Godfrey LM test for serial correlation. **b** Jarque-Bera normality test.
c Ramsey RESET test for omitted variables/functional form. **d** White test for heteroscedasticity.

4. Conclusions

Results from the estimations suggest that while there is no full Ricardian response in Australia to changes in the fiscal stance, fiscal policy has some ability to impact the real economy. Estimates suggest a long-run private saving offset around one half, and between -0.25 and -0.40 in the short run.

While the lower short-run offsets revealed through the error correction mechanisms indicate that nominal and real frictions and/or rigidities prevent some proportion of the offsetting behaviour occurring more quickly, this result is consistent with Keynesian models – suggesting that fiscal policy has a greater ability to influence the real economy over the short term (particularly where some households are liquidity constrained). While full Ricardian equivalence has not been observed in the results, they do suggest that over the longer-term, households and organisations are more forward-looking, and exhibit some partial Ricardian behaviour.

A critical question this paper has also sought to answer is the extent to which the development of the Australian financial sector (and increased integration into global capital markets) may have dampened the impact of fiscal policy on the real economy. Estimates of the long-run coefficient on government saving over the two subsamples (1959:3-1983:4 and 1984:1-2006:2) did not provide any clear indication that this may be occurring (both sets of estimations produced a long-run coefficient on government saving around -0.39). However, the short-run error correction coefficients were markedly different, with the second subsample estimation yielding a short-run private saving offset that was close to the long-run estimate (-0.40).

Results also confirm greater linkages between Australia and the global economy. While the coefficient on net foreign liabilities (*FLIB*), which was taken as a proxy for financial market openness, was statistically insignificant in the first subsample, this coefficient was found to be statistically significant in the second subsample. The negative value of this coefficient (-0.07) suggests that greater access to international capital has lowered private saving. The coefficient on the terms of trade (*TOT*) was also higher in the second subsample, which indicates that Australia may have been deriving higher income from commodities over this period.

While results in this paper suggest that households are not fully Ricardian, fiscal policy can nonetheless exert some impact on real economic activity. However, it is unreasonable to expect that any discretionary fiscal policy actions will have a one-for-one impact on the real economy. To the extent that households anticipate higher (lower) taxes in the future, they will partially offset any policy action through higher (lower) saving. Where policymakers see a need for discretionary policy, it is important to consider the composition of expenditure, as policies directed at particular sectors or households will likely generate different impacts.

While there is a role for activist fiscal policy under extreme economic circumstances, the results also indicate that fiscal policy will only exert a partial impact on activity. It would take substantial movements in the fiscal stance (greater than 1 per cent of GDP) to have a marked impact on the real economy. Such large movements in the fiscal position only exacerbate the risks of poor policy, which includes a risk of excessive debt accumulation, entrenched expenditures and pro-cyclical impacts (arising from poorly timed policy).

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