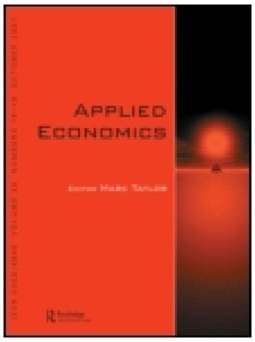

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Testing the validity of the Feldstein–Horioka puzzle for Australia

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This article presents the details of an investigation into the relationship between investment and savings in Australia over the period 1960 to 2007. Using five time series techniques our results reveal that the Feldstein–Horioka puzzle exists in a weak form, with a lower saving retention coefficient. Granger causality tests illustrate that savings Granger causes investment, both in the short and long runs. Our results suggest Australia could effectively adopt policies that focus on increasing investment through increasing domestic savings.

I. Introduction

Low capital mobility among Organization for Economic Co-operation and Development (OECD) countries, signified by a high Saving–Investment (S–I) correlation and known as the Feldstein–Horioka (henceforth F–H) puzzle, has set off a vigorous debate in the empirical literature. The S–I relationship is important, because it has implications for taxation rates and for policies that seek to affect international capital mobility; at the same time there is no doubt that capital accumulation, savings and investment are important stimuli for growth (Coakley *et al*., 2004; Rao *et al*., 2010).

Feldstein and Horioka (1980) present estimates of the following equation for 16 OECD countries over the sample period 1960 to 1974

$$ITY_t = \alpha + \beta STY_t + \varepsilon_t \qquad (1)$$

where ITY is the investment divided by income (hereafter termed the 'investment ratio'), STY the savings divided by income (hereafter termed the 'savings ratio') and ε the error term. Their saving retention estimate (β) is within the range 0.85–0.95, indicating low capital mobility in the sample of countries.

A comprehensive review of the relevant literature is presented by Apergis and Tsoumas (2009), who conclude that the majority of the empirical studies oppose the original strong results of F–H although the correlation exists in a weaker form.

To shed light on country-specific time-series studies, Felmingham and Cooray (2006) utilize an Error Correction Model (ECM) to examine Australian data and found a long-run S–I relationship and imperfect capital mobility. Estimations of the S–I relation by Cooray and Felmingham (2008) suggest that Australia could successfully adopt policies that focus on increasing investment through increasing domestic savings, but this finding contradicts Schmidt (2003), who found that investment in Australia is strongly exogenous and therefore policies that aim to increase investment through savings are unlikely to be successful.

The objective of this article is to re-examine the F–H hypothesis using a comprehensive set of five

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powerful, time-series econometric techniques in an attempt to identify the relative stability of results and the causal relationship between S and I for Australia using data for the period 1960 to 2007. Section II provides a brief history of the F–H puzzle. Section III presents details of the methodologies employed and Section IV details the empirical results. The implications of the results are considered in a brief concluding section.

II. Brief Survey on F–H Puzzle

There is a vast literature that presents investigations into the F–H puzzle using cross-section, time-series or panel data estimation methods; however the results are not consistent.[1] For instance, Hussein (1998) employs Dynamic Ordinary Least Squares (DOLS) techniques to estimate the S–I relationship for a sample of 23 OECD countries using data for the period 1960 to 1993 and obtain results which imply that capital mobility is remarkably high in the majority of the sample countries.[2] Amirkhalkhali *et al.* (2003) examine the S–I deficit relationship within the context of a random coefficients model for 19 OECD countries over the period 1971 to 1999 and find that the S–I correlation is present, but the crowding out effect appears to weaken in the 1990s at the same time that the degree of capital mobility appears to increase. Sinha and Sinha (2004) examine the short- and long-run relationships between S and I for 123 countries using an error correction framework and show that there is evidence for capital mobility for only 16 countries, and most of these are developing countries. Hoffmann (2004) uses a bivariate cointegrated Vector Autoregressive (VAR) model and finds that the long-run capital mobility in UK and US is remarkably stable and high over the mid nineteenth century up to the early 1990s.

Pelgrin and Schich (2008) apply panel error correction techniques to data for 20 OECD countries from 1960 to 1999 and find a long-run S–I relationship, with an increase in the persistency of the deviations from this long-run relation, which suggests that capital mobility has increased. Grier *et al.* (2008) use data from 1947 to 2007 to examine the relationship between S and I in the USA, using Bai and Perron (1998, 2003) techniques to test for structural breaks. They found a positive relationship in the short run that has weakened considerably over time in terms of magnitude and statistical significance but find no cointegration between S and I in the long run.

Early studies on the F–H puzzle with an Australian focus include Obstfeld (1986) who pools time-series observations to estimate the F–H relationship for seven developed countries and finds β to be as low as 0.2 for Australia for the period 1960 to 1984. However, when Coakley *et al.* (1996) extend the sample period from 1960 to 1992, they obtain a value for β of about 0.6. Georgopoulos and Hejazi (2005) emphasize, however, that there should be a home bias, and that this bias should be falling through time.

For the purpose of analysing the cyclical and trend behaviour of the S–I relationship for Australia, Felmingham and Cooray (2006) utilize spectral analysis and find a long-run relationship between the two variables. In relation to capital mobility, their results based on an ECM, show that imperfect capital mobility exists, and their more recent work (Cooray and Felmingham, 2008) argues that Australia could successfully adopt policies that focus on increasing investment through increasing domestic savings. However, these results are in contrast to Schmidt (2003), who finds that investment is strongly exogenous and so policies that aim to increase investment through domestic savings are unlikely to be successful for Australia.

It is unlikely that complete stability will exist in empirical tests of the F–H puzzle in a changing and less than perfectly competitive dynamic international economic environment. Nevertheless given the uncertainty caused by contradictory results generated using a variety of econometric techniques, time periods and samples, this article seeks to identify the stability of causal relationship between S and I for Australia across econometric techniques, using annual data for the period 1960 to 2007.

III. Methodologies

Hendry's well known General to Specific (GETS) technique consists of a broad dynamic lag structure between the dependent and explanatory variables, with the cointegrating equation comprising lagged levels and first differences of the variables. Using standard variable deletion tests, the general unrestricted model is reduced to a parsimonious

[1] See Apergis and Tsoumas (2009) for a comprehensive survey.
[2] This finding is supported by Coakley *et al.* (2004) who used mean group estimator techniques to estimate F–H for 12 OECD countries over the period 1980 to 2000. Their results show that high capital mobility exists in these countries.

dynamic adjustment model, while ensuring that the residuals satisfy the underlying classical assumptions.

Engle and Granger (EG, 1987) developed a two-step technique to estimate long- and short-run equations. The first stage is to estimate the cointegrating equation and the residuals from the cointegrating equation are then used to estimate the short-run dynamic model.

The Phillips and Hansen's (1990) Fully Modified Ordinary Least Squares (FMOLS) is also a single equation estimation technique. We used the Parzen lag window to estimate the cointegrating equations. Following Rao (2007), we start with smaller lag lengths and increase the size while keeping an eye on the estimated elasticities. The specific lag length is confirmed when there are no significant changes in the implied coefficients.

Johansen's (1988) contribution to multivariate cointegration tests, frequently referred to as the Johansen Maximum Likelihood (JML) test,[3] can be applied to justify that the regressors are exogenous. The first stage is to determine the order of the VAR, and then to test for the existence of cointegrating vector(s). Rejection of the null of no cointegration can be obtained through the eigenvalue and trace test statistics. Identification is tested by regressing the lagged ECM normalized on respective variables.

JML offers a more unified framework for estimating and testing cointegrating relationships in the context of ECMs. Granger causality tests can be applied to both long- and short-run situations, with the one period lagged error correction term being derived from the long-run cointegrating equation. In each case, the dependent variable is regressed against past values of itself and past values of other variables together with the lagged error correction term. Significance of the lagged error correction term in respective equations will determine the causality relationship.

To consider the finite sample properties, we apply the Autoregressive Distributed Lag (ARDL) bounds tests approach developed by Pesaran *et al.* (2001). Pesaran and Shin (1999) argued that the ARDL technique can be reliably used in small samples to test hypotheses on the long-run coefficient in both cases, where the underlying regressors are $I(1)$ or $I(0)$. In other words, the ARDL has good small sample properties as compared to alternative techniques. To examine the small sample performance, Pesaran and Shin (1999) used Monte Carlo experiments and showed that the ARDL approach has the advantage of yielding consistent estimates of long-run coefficients. As stated in Pesaran and Pesaran (1997, p. 304), the ARDL procedure contains two steps. First, the existence of the long-run relation between the variables in the system is tested. The F tests are used for testing the existence of long-run relationships. Second, the lag order is selected with the Schwarz Bayesian Criterion (SBC), and thus the ARDL cointegrating equation is estimated.

Table 1. Descriptive statistics of ITY and STY in the period 1960 to 2007

	ITY	STY
Mean	26.591	23.880
SE	0.182	0.157
Median	26.588	25.829
SD	2.648	4.412
Range	9.005	15.122

IV. Results

We use annual data from the International Financial Statistics (IFS), which is published by the International Monetary Fund. The sample period is from 1960 to 2007. This sample period is chosen due to the availability of the data across all variables in the model. The two variables of interest are ITY, which is gross domestic investment as a share of Gross Domestic Product (GDP; IFS, 2008), and STY, which is gross domestic savings as a share of GDP (IFS, 2008). Descriptive statistics related to ITY and STY for the period 1960 to 2007 are provided in Table 1. The averages of ITY and STY are 26.59 and 23.88, respectively. At the same time, the standard deviations for the two variables are 2.65 and 4.41, respectively. For both variables, the standard errors are low at around 0.2. In addition, the STY is the highest (lowest) in 1960 (1991) and ITY is the highest (lowest) in 1966 (1992). This yield ranges from 15.12 to 9.01, respectively. The median lies in the mid-twenties for both the variables.

We test for the presence of a unit root in our variables using the Augmented Dickey–Fuller (ADF) and Phillips–Perron (PP) tests. These results are reported in Table 2. The ADF tests are applied to both levels and their first differences, with an intercept and trend. The ADF and PP statistics

[3] This technique is sometimes also referred as the Vector Error Correction Method (VECM) or the ECM.

Table 2. Results of ADF and PP unit root tests

Variables	ITY	ΔITY	STY	ΔSTY
ADF statistic	1.100 (2)	5.816 (1)	1.669 (0)	5.821 (1)
PP statistic	2.387 (4)	7.230 (5)	1.638 (2)	4.276 (4)

Notes: The ADF and PP critical values at 5% are 3.516 and 3.519, respectively. The lag lengths for ADF and PP are in parenthesis.

for the level variables do not exceed the critical values (in absolute terms), but when we take the first difference of each of the variables, the ADF and PP statistics are higher than the respective critical values (in absolute terms), suggesting that the level variables are $I(1)$ and their first differences are stationary.

Alternative estimates

The alternative time-series techniques are employed to examine the validity of F–H puzzle for Australia. We estimate Equation 1 with GETS, EG, FMOLS, JML and ARDL techniques for the whole sample 1960 to 2007 and for two sub-samples, 1960 to 1980 and 1981 to 2007. We selected the sample break date of 1980 because of the substantial financial reforms and liberalization undertaken in the early 1980s in Australia and many other developed economies. The financial reforms have resulted in substantial gains such as increases in bank profits, new financial products and services and a decline in average costs. For a review of financial reforms in Australia, see Edirisuriya and O'Brien (2001), Sathye (2001) and Sturm and Williams (2002).

GETS, EG, FMOLS, JML and ARDL estimates of Equation 1 are presented in Table 4. The null hypothesis of the savings retention coefficient, β, is that it should be equal to zero for complete capital mobility. The dummy variable (DUM) captures the effects of financial sector reforms and liberalization and it is expected that the sign of DUM should be positive because financial reforms improve the quantity and quality of financial services and trigger more opportunities for investment and savings.

We estimate the general dynamic equation in GETS with a lag length of four periods for the entire sample and with three lags for sub-periods. Using standard variable deletion tests, these were later reduced to manageable parsimonious versions. Note that GETS estimates both levels and their first differences of the variables together. Since the dynamic estimates do not have any economic meaning, it is the long-run estimates that are of interest. The EG cointegrating equation is estimated with Ordinary Least Squares (OLS) and provides estimates of the cointegrating equation in the first stage. FMOLS estimates are obtained with the parzen lag window and a lag length of zero for the whole period and zero and four, for the two sub-periods, respectively. Unlike JML and ARDL, there are no formal tests for cointegration in GETS, EG and FMOLS. In JML, the order of the VAR is five for the entire period and one for the sub-periods. We estimate the JML cointegrating equation using (un)restricted intercepts and no trends.[4] For the whole period and the first sub-period, we employ the restricted intercept and no trend option; for the second sub-period, we use the unrestricted intercept and no trend option. The eigenvalues and trace tests reject the hypothesis of no cointegration at the 95% level, thereby accepting the null hypothesis of one long-run relationship at the 95% level. Details of the eigenvalues and trace statistics are presented in Table 3. Thus the JML results of the tests for cointegration indicate that there is cointegration between savings and investment.

When we applied the ARDL tests for cointegration, we used the SBC to select the optimal lag length. The SBC criterion indicated an optimal lag length of three and two periods for whole and sub-sample periods, respectively. For the whole sample, the computed F statistic (6.126) is greater than the upper bound of the 95% critical value (4.378). Similarly, the computed F statistics are 8.924 (4.378) and 5.490 (4.378) for the first and second sub-periods, respectively, with corresponding 95% critical values in parentheses. In all cases, there is rejection of the null hypothesis of no long-run relationship.

The results show similarity in the estimates of β across all five techniques and are consistent and slightly above 0.5 for the whole period.[5] Sub-sample results could be used to draw inferences about international capital mobility; explicitly, the estimates of β are between 0.6 and 0.8 and 0.5 and 0.6 in the first and second sub-periods, respectively. To this end, it is observed that the magnitude of β is lower for the later sub-period, suggesting that international capital mobility has increased slightly. DUM has the expected positive sign and is significant in most tests. The consistent estimates of β across the five techniques confirm robustness in our results.

In developing the appropriate ECMs for the short run, we adopted the GETS approach in the second stage. The short-run equations are only estimated for

[4] Only these options gave us meaningful results.
[5] For brevity, Table 4 contains the estimates of β and DUM.

Table 3. JML cointegration tests

	Eigenvalue			Trace		
	Test statistic	95%	90%	Test statistic	95%	90%
1960 to 1980						
$r=0$	16.2132	15.8700	13.8100	26.3480	20.1800	17.8800
$r \leq 1$	7.5300	9.1600	10.1348	9.7906	19.2200	17.1800
1981 to 2007						
$r=0$	24.380	14.8800	12.9800	28.7316	17.8600	15.7500
$r \leq 1$	4.3423	8.0700	6.5000	4.3423	8.0700	6.5000
1960 to 2007						
$r=0$	15.8700	14.678	13.8100	20.9633	20.1800	17.8800
$r \leq 1$	6.2852	9.1600	7.5300	6.2852	9.1600	7.5300

Note: r is the number of cointegrating vectors.

Table 4. Alternative estimates of β

	GETS	EG	FMOLS	JML	ARDL
1960 to 2007					
β	0.568	0.531	0.544	0.503	0.516
	(5.56)*	(12.30)*	(12.80)*	(3.27)*	(4.48)*
DUM	0.811	0.808	1.303	0.737	0.870
	(1.75)**	(1.18)	(2.00)*	(2.15)*	(2.53)*
1960 to 1980					
β	0.827	0.579	0.691	0.790	0.695
	(2.88)*	(5.95)*	(6.02)*	(2.10)*	(4.26)*
1981 to 2007					
β	0.625	0.533	0.649	0.641	0.627
	(3.02)*	(2.73)*	(3.47)*	(4.43)*	(2.39)*
DUM	0.852	0.323	0.132	1.628	0.357
	(1.84)**	(2.34)*	(2.45)*	(1.89)**	(1.97)*

Notes: β is the savings retention coefficient. DUM captures the effects of financial reforms and is constructed 1 in the period 1980 to 1985 and zero otherwise. Absolute t-ratios are in the parentheses. The t-ratio is calculated by dividing the coefficient with its SE.
* and ** denote the significance at 5 and 10% levels, respectively.

whole period. The cointegrating equations from Table 4 are used to formulate the respective Error Correction Terms (*ECT*). The second-stage equations are estimated with OLS in which ΔITY_t is regressed on its lagged values, the current and lagged values of ΔSTY_t and the lagged *ECT* from the cointegrating vectors of EG, FMOLS, JML and ARDL. We use lags up to four periods and after employing variable deletion tests, we obtain the parsimonious versions given in Table 5.

It is worth noting that all the estimated coefficients are significant at conventional levels, except the intercept in EG, FMOLS, JML and ARDL. The coefficient of the lagged *ECT* is significant at 5% level with the expected negative sign, and serves as a negative feedback mechanism in the equations. The X^2 statistics indicate that there is no serial correlation (X^2_{sc1}), functional form misspecification (X^2_{ff}), non-normality (X^2_n) or heteroskedasticity (X^2_{hs}) in the residuals. Since our ARDL estimates are indifferent from GETS, EG, FMOLS and JML, we argue that there are no signs of sample biases or endogeneity problems.

The Granger causality tests

The existence of cointegration implies Granger causality, but it does not indicate the direction of causality. We employ JML modelling to assess the direction of causation. If the variables are cointegrated, then equations should be estimated with JML rather than a VAR as in a standard Granger causality test and we follow EG (1987) by estimating a JML

Table 5. Short-run adjustment equations: 1960 to 2007

	GETS	EG	FMOLS	JML	ARDL
Intercept	12.864	0.014	0.011	0.010	0.286
	(5.26)*	(0.11)	(0.08)	(0.08)	(1.24)
λ	−0.306				
	(2.71)*				
STY_{t-1}	0.568				
	(5.56)*				
ECT_{t-1}		−0.432	−0.340	−0.326	−0.335
		(3.42)*	(3.03)*	(2.95)*	(4.17)*
ΔITY_{t-1}	0.331	0.411	0.381	0.371	
	(2.73)*	(3.42)*	(3.19)*	(2.95)*	
ΔSTY	0.373	0.393	0.343	0.362	0.401
	(4.01)*	(4.27)*	(3.84)*	(3.96)*	(2.16)*
ΔSTY_{t-3}	−0.191	−0.198	−0.209	−0.210	
	(2.05)*	(2.13)*	(2.23)*	(2.24)*	
DUM	0.811				
	(1.76)**				
Adjusted R^2	0.574	0.674	0.560	0.654	0.601
SEE	0.129	0.090	0.080	0.096	0.098
X^2_{sc}	0.171	0.015	0.010	0.024	1.509
	(0.68)	(0.90)	(0.92)	(0.88)	(0.22)
X^2_{ff}	1.272	1.485	1.668	0.841	0.445
	(0.26)	(0.22)	(0.20)	(0.36)	(0.51)
X^2_{n}	0.913	0.845	0.361	0.208	0.998
	(0.63)	(0.66)	(0.84)	(0.90)	(0.32)
X^2_{hs}	0.064	0.055	0.074	0.134	0.050
	(0.80)	(0.81)	(0.79)	(0.71)	(0.82)

Notes: Absolute *t*-ratios are in the parentheses. The *t*-ratio is calculated by dividing the coefficient with its standard error. λ is the speed of adjustment. The DUM is part of the cointegrating equations in EG, FMOLS, JML and ARDL and therefore it is not included again in the short-run equations.
* and ** denote statistical significance at the 5 and 10% confidence levels, respectively.

model for Granger causality as follows:

$$\Delta ITY_t = \upsilon + \sum_{i=1}^{p} \lambda_i \Delta ITY_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta STY_{t-1} + \pi_1 ECT_{t-1} + \varepsilon_{1t} \tag{2}$$

$$\Delta STY_t = \upsilon + \sum_{i=1}^{p} \gamma_i \Delta STY_{t-1} + \sum_{i=1}^{p} \alpha_i \Delta ITY_{t-1} + \pi_2 ECT_{t-1} + \varepsilon_{2t} \tag{3}$$

where ECT_{t-1} is the lagged ECT, derived from the long-run cointegrating relationship and ε_{1t} and ε_{2t} are the serially independent random errors. In each case, the dependent variable is regressed against past values of itself and past values of other variables. The JML-based Granger causality test is applied in both the short run and long run. The results are presented in Table 6.

In the short run, the investment ratio is insignificant at the 5% level in the savings ratio equation,

Table 6. Results of Granger causality tests

Dependent variable	ΔSTY_t	ΔITY_t	ECM_{t-1}
ΔSTY_t	–	0.663	0.277
		(1.11)	(1.06)
ΔITY_t	0.315	–	−0.430
	(2.91)*		(2.58)*

Notes: Absolute *t*-ratios are in the parentheses. The *t*-ratio is calculated by dividing the coefficient with its SE.
* Denotes statistical significance at the 5% level.

implying that the investment ratio does not Granger cause the savings ratio in the short run. However, the savings ratio is significant at the 5% level in the investment ratio equation, implying that there is a bi-directional causality running from savings ratio to investment ratio. The long-run results suggest that the coefficient of ECM_{t-1} is significant at the 5% level, with the expected negative sign in the investment ratio equation, which implies that in the long run, the savings ratio Granger causes the investment ratio.

Our results suggest that the endogeneity problem is limited, because the savings ratio is only weakly exogenous, and this is in contrast to the results of Schmidt (2003), who found that investment in Australia is strongly exogenous. Therefore, we reach the opposite conclusion and suggest that policies, which aim to increase investment through domestic savings, are likely to be successful.

V. Conclusion

We have attempted to estimate the savings retention coefficient (β) and determine the causality relationship between S and I for Australia over the 1960 to 2007 period using GETS, EG, FMOLS, JML and ARDL techniques. The β coefficient is identified to be slightly larger than 0.5 and statistically significant over the entire time period and appears to be falling when we compare sub-periods. This implies that the F–H puzzle exists in a weaker form with a lower saving retention coefficient. The effects of the financial reforms of 1980s and liberalization on investment are also significant, but only temporary.

Further improvements in capital mobility may be difficult, as Australia is a highly open economy with a stable ratio of current account balances to GDP over longer periods. Financial reforms may play an influential role in improving international capital mobility. Our Granger causality tests reveal that both in short and long runs, the savings ratio Granger causes investment ratio. With this finding, and in line with Cooray and Felmingham (2008), we argue Australia could effectively adopt policies that focus on increasing investment through increasing domestic savings.

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