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The relationship between saving and investment for Japan

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Abstract

In this paper, we revisit the saving and investment nexus as postulated by Feldstein and Horioka (FH) [Econ. J. 90 (1980) 314]. We test for cointegration between saving and investment using the recently developed bounds testing approach to cointegration and derive the long-run elasticities using the autoregressive distributed lag modeling approach for Japan over the period 1960–1999. We establish the unit root properties of the data in the presence of structural break(s) using the Zivot and Andrews (ZA) [J. Business Econ. Stat. 10 (1992) 251] and the Lumsdaine and Papell (LP) [Rev. Econ. Stat. 79 (1997) 212] tests. Finally, we ascertain the direction of causation between saving and investment by using the bootstrap approach. Amongst our key results we find that saving and investment are cointegrated for Japan; investment causes saving and saving causes investment; shocks to saving and investment have a permanent effect; and the long-run coefficient on saving is 0.68, implying a moderate rate of correlation. From the latter finding, we believe that there is no puzzle between saving and investment in the case of Japan, a result contrary to FH (1980).

JEL classification: C32; E21; F41

Keywords: Japan; Saving–investment; Cointegration; Causation

1. Introduction

In their seminal contribution to international economics, Feldstein and Horioka (FH) (1980) contend that, in the presence of a high correlation between saving and investment

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(*S–I*), capital mobility should be low. In the aftermath of this thesis a plethora of studies has emerged, essentially taking two competing lines of inquiry. One strand of literature can be perceived as extending upon the Feldstein and Horioka theory. To this end, studies have attempted to re-examine the *S–I* nexus using a longer sample period (Schmidt, 2003; De Vita and Abbott, 2003; De Haan and Siermann, 1994); several studies (Miller, 1988; Sarno and Taylor, 1998; De Vita and Abbott, 2003; Alexakis and Apergis, 1994) have attempted to divide the sample period into two based on the date of change in the exchange rate regime, while others (Bodman, 1995; Özmen and Parmaksiz, 2003) have revisited the theory by simulating exogenous and endogenous shocks to saving and investment.

The second strand of literature proffers alternative hypotheses to explain the co-movements between investment and saving. A common and fitting hypothesis in this regard is the intertemporal budget constraint or solvency condition. Saving and investment are cointegrated if a country satisfies the solvency constraint (see, Ballabriga et al., 1991; Coakly et al., 1996; Jansen, 1996, 1998; Shibata and Shintani, 1998).[1] Further, Coakly et al. (1996) and Coakly and Kulasi (1997) advance the hypothesis that a high degree of capital mobility and strong correlation between saving and investment can coexist with current account targeting by the authorities. In particular, in the event of imbalances in the current account, monetary policy and/or fiscal policy is/are used to stabilise the external position of the country, which thus strengthens the saving–investment (*S–I*) correlation.

A key feature of the extant literature on the saving and investment nexus is that most are cross-sectional or based on panel data (Jansen, 1998, 2000; Dooley et al., 1987; Bayoumi, 1990; Tesar, 1991; Feldstein, 1983; Dar et al., 1994; Coakly et al., 1999; Wong, 1990; Kim, 2001; Coiteux and Olivier, 2000; Tsung-wu, 2002). There are several limitations of cross-sectional/panel studies which may bear on the results. For instance, given that it is clear that the *S–I* correlation is contingent on the nature of shocks and the structure of an economy, it is incorrect to expect the *S–I* relationship to be the same for countries in the sample—a result implicit in cross-sectional regressions. Further, when the *S–I* relationship is modelled within a cross-sectional framework, it exposes the results to possible distortions since it may be significantly affected by outliers, i.e. countries that are economically large and contain a high correlation coefficient because of capital controls (for more limitations see Caporale et al., 2001).

Our study differs from the extant literature in four novel ways. First, we undertake a thorough investigation of the unit root properties of the data. To this end, apart from using the conventional unit root tests—Augmented Dickey and Fuller (ADF, 1979, 1981) and Phillips and Perron (PP, 1988) tests—we also employ the Zivot and Andrews (ZA, 1992) and Lumsdaine and Papell (LP, 1997) tests. The appealing aspect of the ZA

[1] The short run correlation, however, impinges on the size and the nature of shocks that affect the economy over time and the structure of the economy as it captures the contemporaneous co-movements in saving and investment. Hence, the short-run correlation between saving and investment reflects adjustments in demand and supply conditions rather than anything else (see Jansen, 1996, 1998).

(1992) and LP (1997) tests is that they allow one to establish the unit root properties of the data in the presence of one and two structural breaks respectively. An associated advantage of these techniques is that they search for breaks endogenously. Given that Christiano (1992) and Zivot and Andrews (1992) argue that the break points should be viewed as being correlated with the data, selecting the break exogenously could lead to an over rejection of the unit root hypothesis. Testing the unit root hypothesis in the presence of structural break(s) also allows us to gauge whether policies implemented to boost saving and investment have a permanent or a transitory effect on saving and investment.

Second, we use the bounds testing approach to cointegration. Our study is an advance over existing studies (see, for instance, Abbott and De Vita, 2003) using the bounds testing approach because we calculate bounds F-statistic critical values (CVs) specific to our sample size. This is an important exercise given that existing critical values are based on sample sizes of 500 observations and 1000 observations. By calculating critical values specific to our sample size we ensure that our inference regarding cointegration is correct.

Third, we use the bootstrap approach to testing for causality. It is now widely known that the bootstrap approach to causality produces robust critical values (see, inter alia, Horowitz, 1994; Mantalos and Sukur, 1998; and Mantalos, 2000). The bootstrap approach has not been previously used in the saving–investment literature.

Fourth, our study differs from a recent study (Sinha, 2002) on Japan. Sinha (2002) uses the Johansen (1988) approach and finds investment and saving to be cointegrated. However, his results are normalised on saving which is a priori determined. In contrast, we use the bounds testing approach that tells which variable should be normalised. Our result from the bounds test suggests that there is a cointegration relationship between saving and investment only when investment is the dependent variable. Secondly, Sinha (2002) uses the Ganger F-test for causality and finds no evidence of any causality between saving and investment. In this paper, we use the more powerful bootstrap approach to Granger-causality and find that there is causality running from saving to investment and investment to saving.

The rest of the paper is organized as follows. In the next section we present the model followed by the methodology. In the penultimate section we present the empirical results while in the final section we conclude.

2. Model

To examine the S–I correlation for Japan, we apply the generic long-run model which has the following form:

$$I_t = \alpha_0 + \beta_1 S_t + \varepsilon_t \tag{1}$$

Here, I_t is gross national investment as a proportion of gross domestic product (GDP); S_t the gross national saving as a proportion of GDP; α the constant; and ε_t the disturbance term. The S–I correlation is determined by the size of β.

3. Methodology

3.1. Bounds testing approach to cointegration

To examine the long-run relationship between investment and saving for Japan, as explained earlier, we employ the bounds testing approach to cointegration, developed by Pesaran and Shin (1999) and Pesaran et al. (2001).[2]

This bounds procedure can be applied to models irrespective of whether the variables are integrated of order zero ($I(0)$) or integrated of order one ($I(1)$) or mutually cointegrated. The bounds testing procedure involves two stages. The first stage is to establish the existence of a long-run relationship in Eq. (1). Given that the bounds testing procedure to cointegration is a recent development in the econometric time-series literature, we here present a brief outline of this procedure. To implement the bounds test let us define a vector of two variables, z_t, where $z_t = (y_t, x_t{}')'$, y_t is the dependent variable and x_t a vector of regressors. The data generating process of z_t is a p-order vector autoregression. For cointegration analysis it is essential that Δy_t be modelled as a conditional error correction model (ECM)

$$\Delta y_t = \beta_0 + \pi_{yy} y_{t-1} + \pi_{yx.x} x_{t-1} + \sum_{i=1}^{p} \vartheta_i \, \Delta y_{t-i} + \sum_{j=0}^{q} \phi_j{}' \, \Delta x_{t-j} + \theta w_t + \mu_t \tag{2}$$

Here, π_{yy} and π_{yx} are long-run multipliers. β_0 is the drift and w_t is a vector of exogenous components, e.g. dummy variables. Lagged values of Δy_t and current and lagged values of Δx_t are used to model the short-run dynamic structure. The bounds testing procedure for the absence of any level relationship between y_t and x_t is through exclusion of the lagged levels variables y_{t-1} and x_{t-1} in Eq. (2). It follows, then, that our test for the absence of a conditional level relationship between y_t and x_t entails the following null and alternative hypotheses:

$$H_0 : \pi_{yy} = 0, \quad \pi_{yx.x} = 0', \tag{3}$$

$$H_1 : \pi_{yy} \neq 0, \quad \pi_{yx.x} \neq 0' \quad \text{or} \quad \pi_{yy} \neq 0, \quad \pi_{yx.x} = 0' \quad \text{or} \quad \pi_{yy} = 0, \quad \pi_{yx.x} \neq 0'. \tag{4}$$

These hypotheses can be examined using the standard Wald or F-statistics. We use the F-test which has a non-standard distribution which depends upon; (i) whether variables included in the ARDL model are $I(0)$ or $I(1)$, (ii) the number of regressors and (iii) whether the ARDL model contains an intercept and/or a trend. Critical values are reported by Pesaran and Pesaran (1997) and Pesaran et al. (2001). However, these CVs are generated for sample sizes of 500 observations and 1000 observations and 20,000 and 40,000 replications, respectively. Given the relatively small sample size in our study (40 observations) we calculate critical values specific to our sample size. To this end, we use the same GAUSS™ code used to generate the original set of CVs.

[2] Tang (2003) applied the bounds test approach to the analysis of import demand behaviour of Japan. Contrary to Mah (1994) and Hamori and Matsubasyashi (2001), the bounds test confirms a long-run equilibrium relationship between the quantity of imports and its determinants.

The critical value bounds are calculated using stochastic simulations for $T = 40$ and 40,000 replications for the F-statistic. The F-statistic is used for testing the null hypothesis of $\psi = \delta_1 = \delta_2 = \cdots = \delta_k = 0$ in a model with an intercept but no trend. In the Pesaran et al. (2001: T3) terminology, a model with an intercept and no trend is referred to as Case II, and has the following form:

$$\Delta y_t = \beta_0 + \psi y_{t-1} + \sum_{i=1}^{k} \delta_i x_{i,t-1} + \varepsilon_t \tag{5}$$

Here, $t = 1, \ldots, T$; $x_t = (x_{1t}, \ldots, x_{kt})'$ and $z_{t-1} = (y_{t-1}, x_{t-1}{}', 1)'$, $w_t = 0$. The variables y_t and x_t are generated from $y_t = y_{t-1} - \varepsilon_{1t}$ and $x_t = Px_{t-1} - \varepsilon_{2t}$, $t = 1, \ldots, T$, $y_0 = 0$, $x_0 = 0$ and $\varepsilon_t = (\varepsilon_{1t}, \varepsilon_{2t}{}')$ is drawn as $(k - 1)$ independent standard normal variables. If x_t is purely $I(1)$, i.e. integrated of order one, $P = I_k$. On the other hand, $P = 0$ if x_t is purely $I(0)$, i.e. if it is integrated of order zero. The critical values for $k = 0$ correspond to those for the Dickey and Fuller (1981) unit root F-statistics. Two sets of critical values are generated and presented. One set refers to the $I(1)$ series and the other for the $I(0)$ series. Here, critical values for the $I(1)$ series are referred to as the upper bound critical values while the critical values for the $I(0)$ series are referred to as the lower bound critical values.

If the computed F-statistics fall outside the critical bounds, a conclusive decision can be made regarding cointegration without knowing the order of integration of the regressors. For instance, if the empirical analysis shows that the estimated F-statistic is higher than the upper bound of the critical values then the null hypothesis of no cointegration is rejected.

Once cointegration is ascertained, the second stage involves estimating the long-run and short-run coefficients of the cointegrated equation. The mathematical derivation of the long-run and short-run parameters can be found in Pesaran and Pesaran (1997).

3.2. Unit roots

To ascertain the order of integration we use the Zivot and Andrews (1992) one break test and Lumsdaine and Papell (1997) two break test. In conducting the unit root tests in the presence of structural break(s) we seek to establish new information: whether policies designed to stimulate saving and investment have a permanent or a transitory effect.

We use two versions of the Zivot and Andrews (1992) sequential trend break model to investigate the unit root hypothesis for investment and saving. Model A allows for a change in intercept, while model C allows for a change in both the intercept and slope. Previous studies of the unit root hypothesis which have employed the Zivot and Andrews (1992) sequential trend break model have tended to use model C (e.g. Raj, 1992; Perron, 1994; Ben-David and Papell, 1995); however, there is no consensus on which model is superior.

Model A has the following form:

$$\Delta y_t = \kappa + \alpha y_{t-1} + \beta t + \theta_1 DU_t + \sum_{j=1}^{k} d_j \, \Delta y_{t-j} + \varepsilon_t \tag{6}$$

Model C takes the following form:

$$\Delta y_t = \kappa + \alpha y_{t-1} + \beta t + \theta_1 DU_t + \gamma_1 DT_t + \sum_{j=1}^{k} d_j \, \Delta y_{t-j} + \varepsilon_t \tag{7}$$

Here, Δ is the first difference operator, ε_t is a white noise disturbance term with variance σ^2, and $t = 1, \ldots, T$ is an index of time. The Δy_{t-j} terms on the right-hand-side of Eqs. (6) and (7) allow for serial correlation and ensure that the disturbance term is white noise. Finally, DU_t is an indicator dummy variable for a mean shift occurring at time TB and DT_t is the corresponding trend shift variable, where

$$DU_t = \begin{cases} 1 & \text{if } t > TB \\ 0 & \text{otherwise} \end{cases}$$

and

$$DT_t = \begin{cases} t - TB & \text{if } t > TB \\ 0 & \text{otherwise} \end{cases}$$

In applying the Zivot and Andrews (1992) test, some region must be chosen such that the end points of the sample are not included, for in the presence of the end points the asymptotic distribution of the statistics diverges to infinity.[3] Zivot and Andrews (1992) suggest the 'trimming region' [0.15, 0.85], which in our case is the sample sub-period 1959–1991. The break points are selected by choosing the value of TB for which the ADF t-statistic (the absolute value of the t-statistic for α) is maximized. The test essentially amounts to testing the null hypothesis that the series $[y_t]$ is an integrated process without a structural break, against the alternative hypothesis that $[y_t]$ is trend stationary with a structural break in the trend function which occurs at an unknown time.

Following Hall (1994), we select the lag length (k) using the "t-sig" approach. Hall recommends starting with a predetermined upper bound k. If the last included lag is significant, the upper bound k is chosen. However, if k is insignificant, it is reduced by one lag until the last lag becomes significant. If no lags are significant k is set equal to zero. One of the key advantages of using the 't-sig' approach is that it has been shown to produce test statistics with better properties (stable size and higher power) than information-based methods, such as the Akaike Information Criterion (e.g. Ng and Perron, 1995; Perron, 1997, pp. 359–360).

Given that our sample size is relatively small (40 observations) we set $k_{\max} = 5$ and following the norm in recent studies (see Perron, 1989; Zivot and Andrews, 1992; Ben-David and Papell, 1995; Lumsdaine and Papell, 1997) we use a critical value of 1.60 to determine the significance of the t-statistic on the last lag and we do not increase the upper bound when the procedure selects $k = k_{\max}$.

Whilst asymptotic critical values are available for this test, Zivot and Andrews (1992) warn that with small sample sizes the distribution of the test statistic can deviate substantially from this asymptotic distribution. To circumvent this distortion, we calculate 'exact' critical values for the test following the methodology recommended in Zivot and Andrews (1992: 262). In assuming that the errors driving the data series are normal ARMA (p,q) processes we estimate an ARMA (p,q) model for each Δy_t, with p and q selected according to the Akaike information criterion (AIC). The AIC minimizes $2 \ln L + 2(p + q)$, where L denotes the likelihood function. The implied ARMA process is then used as the data generating process for generation of 5000 replications of the series under the null of a unit root. We then follow

[3] For a detailed proof, see Corollary 1 of Andrews (1993)

Zivot and Andrews (1992) in determining k, and obtain a minimum ADF statistic for each of the 5000 series. The critical values are then constructed from this empirical distribution.

Lumsdaine and Papell (1997) propose a model that tests endogenously for two structural breaks. In essence, they extend Zivot and Andrews (1992) models A and C and call these model AA and model CC, respectively.

Model AA takes the following form:

$$\Delta y_t = \kappa + \alpha y_{t-1} + \beta t + \theta_1 DU1_t + \psi_1 DU2_t + \sum_{j=1}^{k} d_j \, \Delta y_{t-j} + \varepsilon_t \tag{8}$$

Model CC can be represented as follows:

$$\Delta y_t = \kappa + \alpha y_{t-1} + \beta t + \theta_1 DU1_t + \gamma_1 DT1_t + \psi_1 DU2_t + \omega_1 DT2_t + \sum_{j=1}^{k} d_j \, \Delta y_{t-j} + \varepsilon_t \tag{9}$$

Here, $DU1_t$ and $DU2_t$ are indicator dummy variables for a mean shift occurring at times $TB1$ and $TB2$, respectively, where $TB2 > TB1 + 2$ and $DT1_t$ and $DT2_t$ are the corresponding trend shift variables. That is,

$$DU1_t = \begin{cases} 1 & \text{if } t > TB1 \\ 0 & \text{otherwise} \end{cases}$$

$$DU2_t = \begin{cases} 1 & \text{if } t > TB2 \\ 0 & \text{otherwise} \end{cases}$$

and

$$DT1_t = \begin{cases} t - TB1 & \text{if } t > TB1 \\ 0 & \text{otherwise} \end{cases}$$

$$DT2_t = \begin{cases} t - TB2 & \text{if } t > TB2 \\ 0 & \text{otherwise} \end{cases}$$

The lag length is selected using the ‘‘t-sig’’ method and the break points are chosen using the same approach as in the one break case.

3.3. Granger-causality test

In this section we will present the bootstrap approach to testing for causality between investment and saving in Japan using the LR-test. We begin by considering the Granger-causality test by using the LR-test. Following Dolado and Lutkepohl (1996) the following data generating process is considered. It entails k-dimensional multiple time series generated by a VAR(p) process:

$$y_t = \alpha + A_1 y_{t-1} + \cdots + A_p y_{t-p} + \varepsilon_t \tag{10}$$

Here $\varepsilon_t = (\varepsilon_{1t}, \ldots, \varepsilon_{kt})'$ is a zero mean independent white noise process with nonsingular covariance matrix $\sum_\varepsilon$ and, for $j = 1, \ldots, k$, $E|\varepsilon_{jt}|^{2+\tau} < \infty$ for some $\tau > 0$. The order of p is assumed to be know. Here, following Hatemi (2002) we choose the lag length of the VAR model based on an amalgam of techniques—the Akaike (1969) information criterion, Schwarz (1978) Bayesian Criterion, LR-test and diagnostic tests. In the next step we partition y_t in two subvectors y_t^1 and y_t^2 and A_i matrices partitioned conformably we can write the Eq. (10) in the following form:

$$y_t = \begin{bmatrix} y_t^1 \\ y_t^2 \end{bmatrix} = \begin{bmatrix} \eta_1 \\ \eta_2 \end{bmatrix} + \begin{bmatrix} A_{11,1}A_{12,1} \\ A_{21,1}A_{22,1} \end{bmatrix} \times \begin{bmatrix} y_{t-1}^1 \\ y_{t-1}^2 \end{bmatrix} + \cdots + \begin{bmatrix} A_{11,p}A_{12,p} \\ A_{21,p}A_{22,p} \end{bmatrix} \times \begin{bmatrix} y_{t-p}^1 \\ y_{t-p}^2 \end{bmatrix} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{bmatrix} \tag{11}$$

In this setting, it follows that y_t^2 does not Granger-cause y_t^1 if the following hypothesis is true:

$$H_0 : A_{12,i} = 0 \quad \text{for } i = 1, \ldots, p$$

In investigating whether y_t^1 Granger-causes y_t^2 one needs to test the test the following hypothesis:

$$H_0 : A_{21,i} = 0 \quad \text{for } i = 1, \ldots, p$$

and so forth.

Before defining the multivariate LR-test for the Granger-causality analysis, we define:

$$Y := (y_1, \ldots, y_T) \quad (k \times T) \quad \text{matrix},$$

$$B := (\eta, A_1, \ldots, A_p) \quad (k \times (kp+1)) \quad \text{matrix},$$

$$Z_t := \begin{bmatrix} 1 \\ y_t \\ \vdots \\ y_{t-p+1} \end{bmatrix} \quad ((kp+1) \times 1) \quad \text{matrix},$$

$$Z := (Z_0, \ldots, Z_{T-1}) \quad ((kp+1) \times T) \quad \text{matrix},$$

and

$$\delta := (\varepsilon_1, \ldots, \varepsilon_T) \quad (kp \times T) \quad \text{matrix}.$$

The idea behind using the above notation is that it allows us to write Eq. (10) in a simple form as follows:

$$\Upsilon = BZ + \phi \tag{12}$$

The least squares estimator of B in Eq. (12) as follows:

$$\hat{B} = YZ'(ZZ')^{-1} \tag{13}$$

The $(k \times T)$ matrix of the estimated residuals from the unrestricted regression (12) can be denoted by $\hat{\phi}_U$ and the $(k \times T)$ matrix of residuals from the restricted regression with H_0 imposed can be denoted by $\hat{\phi}_R$. The matrix of the cross products of these residuals can now be defined as $H_U = \hat{\phi}_U{}'\hat{\phi}_U$ and $H_R = \hat{\phi}_R{}'\hat{\phi}_R$, respectively. The multivariate LR-test can now be written as:

$$LR = (T - p) \times \ln\left(\frac{\det H_R}{\det H_U}\right) \tag{14}$$

Here T is the number of observations and p is the order of the lags in the VAR model. ln denotes natural logarithm and det represents the determinant of the respective matrix. The LR-test statistic has a χ^2 distribution with the degrees of freedom commensurate with the number of restrictions to be tested.

3.3.1. Bootstrap simulation technique

The bootstrap testing approach was pioneered by Efron (1979). Several studies (see, *inter alia*, Horowitz, 1994; Mantalos and Sukur, 1998; and Mantalos, 2000) have shown the robustness of the bootstrap critical values. Essentially, in the bootstrap method, the information in the sample is recycled by simulations; the idea being to reduce bias and provide more reliable critical values (see Hatemi, 2002).

From Eq. (12), a direct residual resampling gives:

$$\Upsilon^* = \hat{B}Z^* + \phi^* \tag{15}$$

Here ϕ^* are i.i.d. observations $\phi_1^*, \ldots, \phi_T^*$. This is drawn from the empirical distribution $(\hat{F}_\phi)$ putting mass $1/T$ to the adjusted ordinary least squares residuals $(\hat{\phi}_i - \bar{\phi})$, $i = 1, \ldots, T$. The bootstrap testing procedure works in a straightforward way. It draws a number of bootstrap samples from the model under the null hypothesis and calculates the bootstrap test statistic (T_s^*). Then, one can calculate the bootstrap test statistic (T_s^*) by repeating this step N_b times. Here, following Hatemi (2002) I use $N_b = 1000$ to estimate the P-value. In the aftermath of this process the αth quintile of the bootstrap distribution is taken to obtain the α-level "bootstrap critical values" (c_{ta}^*). What remains now is to estimate the test statistic (T_s) using actual data set. With all the statistics calculated, one rejects the null hypothesis if $T_s \leq c_{i\alpha}^*$.

4. Empirical results

4.1. Unit root tests

While the bounds test for cointegration is applicable irrespective of whether the variables are integrated of order one or order zero, it is important to establish that the variables are not integrated of an order higher than one. As mentioned earlier, our second reason for conducting unit root tests is to gauge whether shocks to saving and investment have a permanent or a transitory effect.

To ascertain the order of integration we begin through applying the Augmented Dickey-Fuller (ADF) and Phillips Perron (PP) unit root tests. The ADF and PP tests suggest that $[I,S]$ are each integrated of order one or $I(1)$. These results are not reported here to conserve space. However, one of our main concerns in this paper is with the implications of structural breaks on unit roots. Given the inability of standard ADF and PP to capture the impact of structural breaks, to circumvent this we report in Tables 1 and 2 the Zivot and Andrews (1992) one-break test results and the Lumsdaine and Papell (1997) two-break test results, respectively.

The ZA test with one structural break finds no additional evidence against the unit root hypothesis relative to the unit root tests without a structural break. In other words,

Table 1
Zivot and Andrews test for unit roots with one structural break

	Investment		Saving	
	Model A	Model C	Model A	Model C
TB	1978	1965	1980	1980
α	−0.4543*** (−3.8096)	−0.41138*** (−3.6097)	−0.9482*** (−4.1576)	−0.9331*** (−3.9670)
θ	−0.0320 (−1.6405)	0.0244 (−0.0279)	−0.0808** (−2.4144)	−0.0753 (0.0009)
γ	–	0.7482 (−1.2050)	–	−2.0519 (−0.4093)
k	4	2	5	5
Ljung-Box Q-statistic	6.0330 (0.9792)	14.1765 (0.5122)	12.2303 (0.6615)	11.6087 (0.7084)
Exact critical values (%)				
1	−6.2711	−6.6322	−6.1936	−6.7499
5	−5.4507	−5.8358	−5.4650	−5.8794
10	−5.0942	−5.4489	−5.0701	−5.4389

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

in models A and C the unit root null hypotheses are not rejected for saving and investment. This result is consistent with the standard ADF and PP test results. The Ljung-Box Q-statistics generally find no evidence of residual serial correlation in the error terms.

While it is true that the Zivot and Andrews (1992) test, by virtue of accounting for one structural break, is an advance over standard ADF and PP tests, it is argued that the Zivot and Andrews test may lose power when confronted with two or more breaks (see Lee and Strazicich, 1999). To address this problem, the procedure devised by Lumsdaine and Papell (1997), which is explained earlier, is a good remedy. Again, we calculate exact critical

Table 2
Lumsdaine and Papell test for unit roots with two breaks

	Saving		Investment	
	Model A	Model C	Model A	Model C
*TB*1	1974	1965	1966	1970
*TB*2	1980	1984	1980	1988
α	−1.4104 (−5.1207)	−0.5054 (−4.4783)	−0.8888 (−4.6585)	−0.9844 (−4.8603)
θ	−0.0365 (−0.1361)	0.0273 (0.0491)	−0.0934 (−0.0656)	0.0522 (0.1129)
ψ	−1.6585*** (−3.7981)	−0.1136 (−0.0009)	−2.9267*** (−2.7066)	−0.0243 (−0.0006)
γ	–	0.7548** (2.2086)	–	1.7968*** (3.3190)
ω	–	−2.7100 (−0.4381)	–	−3.5797 (−0.1670)
k	6	3	5	2
Ljung-Box Q-statistic	7.1598 (0.9531)	17.1851 (0.3079)	8.1023 (0.9196)	8.1047 (0.9195)
Exact critical values (%)				
1	−6.8326	−8.3117	−6.9553	−8.2071
5	−6.1004	−7.3126	−6.1256	−7.3420
10	−5.7457	−6.9085	−5.7535	−6.8973

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

values based on the simulation method as in the one break case. The results are presented in Table 2. Neither model AA nor CC could reject the null hypothesis of unit roots in investment and saving at the 1 percent level for both the sample periods. The Ljung-Box Q-statistics generally show no evidence of residual serial correlation in the error terms.

The unit root hypothesis has far reaching implications with respect to both economic theory and the interpretation of empirical evidence (Perron and Vogelsang, 1992). Under the unit root hypothesis, for instance, random shocks have permanent effect on the system. This is same as saying that fluctuations in a series are not transitory. Our results on investment and saving for Japan in the presence of one break and two breaks clearly reveal that the series are integrated of order one, implying that policies implemented to boost savings and investment are likely to be successful.

4.2. Cointegration test results

The bounds test for cointegration involves the comparison of the F-statistics against the critical values, which are generated using stochastic simulations for $T = 40$ based on 40,000 replications, as explained earlier. Using Eq. (2), each variable in our model—Eq. (1)—is taken as a dependent variable in the calculation of the F-statistics. The calculated F-statistics, both in the presence and absence of a structural break, are reported in Table 3. Sen (2003) argues that model C is statistically more powerful than model A; hence, in testing for cointegration we incorporate the statistically significant structural break from model C.

When investment is the dependent variable for Japan, the calculated F-statistic without the structural break is $F_I(I|S) = 6.137$ and with the structural break is $F_I(I|S) = 6.1936$, which are higher than the upper bound critical value of 4.523 at the 5 percent significance level. However, if saving is the dependent variable over the same period, the calculated F-statistic without the structural break is $F_S(S|I) = 1.241$ and with the structural break is $F_S(S|I) = 1.4964$, which are lower that the lower bound critical value at the 5 percent level.

Table 3
Bounds F-statistics for cointegration relationship

	90% level		95% level		99% level	
	$I(0)$	$I(1)$	$I(0)$	$I(1)$	$I(0)$	$I(1)$
Critical value bounds of the F-statistic						
1	3.210	3.730	3.937	4.523	5.593	6.333
Calculated F-statistics (without a structural break)						
$F_I(I\|S)$			6.1374			
$F_S(S\|I)$			1.2412			
Calculated F-statistics (with a structural break)						
$F_I(I\|S)$			6.1936			
$F_S(S\|I)$			1.4964			

Note: Critical values are calculated using stochastic simulations specific to the sample size based on 40,000 replications. k is the number of regressors.

This suggests that the null hypothesis of no cointegration cannot be accepted for Japan. In other words, we have established that investment and saving are cointegrated for Japan over the 1960–1999 period, when investment is the dependent variable.

4.3. Long-run and short-run elasticities

Since investment and saving are cointegrated the long-run model was estimated using the following ARDL specification:

$$I_t = \alpha_0 + \sum_{i=1}^{m} \alpha_1 I_{t-i} + \sum_{i=0}^{n} \alpha_2 S_{t-i} + \mu_t \tag{16}$$

On the Schwarz Bayesian Criteria (SBC), maximum of 2 lags was used for the model, such that $i_{\max} = 2$. The results are presented in Table 4. The ARDL technique indicates a weak correlation (0.68) between saving and investment for Japan over the 1960–1999 period. This result is contrary to the findings of FH (1980) who found the correlation to be high violating the fact that for open economies with high capital mobility, the correlation should be low. They termed their result as a puzzle. However, our empirical results based on the bounds test and the ARDL technique find no such puzzle in the case of Japan.

The error correction model was also estimated within the ARDL framework. The results show the error correction term—ECM$_{t-1}$—is negative and statistically significant at the 1 percent level, indicating that convergence to long-run equilibrium after a shock to saving is moderate for investment in Japan. For instance, the coefficient of -0.34 suggests that a deviation from the long-run equilibrium level of investment in the current year is corrected by about 34 percent in the next year. Hence it takes around 3 years for external imbalances to stabilise.

Table 4
Long-run and short-run results: dependent variable I_t

	Coefficient		t-statistic
Regressor			
S_t	0.6873***		6.1258
Short-run			
Constant	0.0275***		2.8300
$\Delta \ln S_t$	0.6136***		5.5519
EC_{t-1}	−0.3442***		−2.8149
Diagnostic test			
R^2		0.6246	
$\bar{R}^2$		0.6037	
σ		0.0077	
$\chi^2_{NORM}(2)$		3.7062	
$\chi^2_{SERIAL}(1)$		1.2483	
$\chi^2_{WHITE}(1)$		1.3233	
$\chi^2_{RESET}(1)$		0.0078	

*** Indicates statistical significance at the 1% level.

Table 5
Granger-causality test results

	Null hypothesis	
	Investment does not Granger-cause saving	Saving does not Granger-cause investment
LRE test statistic	15.5161	4.5658
LRE bootstrapped *P*-value	0.0000	0.0780
Critical values (%)		
5	6.1506	5.3164
10	4.6175	4.1647

As for the short-run correlations, it is again low (0.61) consistent with the long-run results (Table 4). However, it should be noted that the short-run correlations reflect business cycle influences—adjustment in demand and supply shocks rather than the degree of capital mobility.

4.4. Granger-causality

The existence of a cointegrating relationship among investment and saving suggests that there must be Granger-causality in at least one direction, but it does not indicate the direction of causality. Table 5 reports the causality results based on the bootstrapped *P*-values.

We find that for the null hypothesis that investment does not Granger-cause saving the LRE test statistic is greater than the critical value at the 5 percent level. This has a *P*-value less than 0.05. This implies that the null hypothesis cannot be accepted and that there is Granger-causality running from investment to saving in Japan. On the other hand for the null hypothesis that saving does not Granger-cause investment we find that the LRE test statistic is greater than the critical value at the 10 percent level with a *P*-value less than 0.10. This implies that the null hypothesis cannot be accepted and that there is Granger-causality running from saving to investment.

Our results on Granger-causality are contrary to the findings of Sinha (2002), who found no evidence of causation based on the *F*-tests.

4.5. Parameter stability

Chow (1960) introduced parameter stability tests into the econometrics literature. Parameter stability tests have been given increasing attention in the recent past in recognition of the fact that systemic instability, which is filtered to the parameters of the estimated model, can lead to misleading inference or forecasting. In this light, it is important to test for parameter stability. To this end, we apply the suite of tests proposed by Hansen (1992). Hansen (1992) recommends three tests—*SupF*, *MeanF*, and L_C—which have the null hypothesis that the parameters are stable. We find that the *F*-statistics do not cross any of the critical values at the 5 percent level of significance, implying stability of the long-run parameters in our model (Fig. 1).

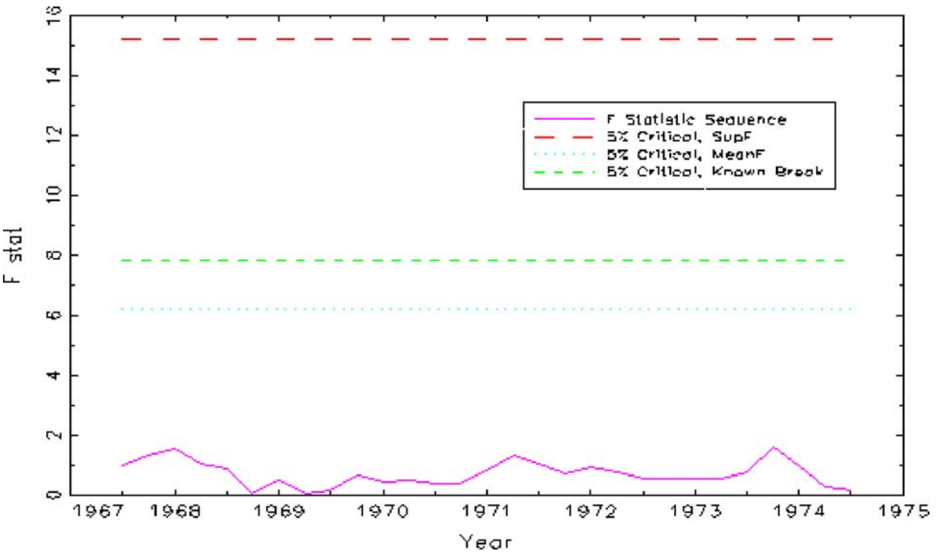

Fig. 1. Hansen instability test.

5. Conclusion

Feldstein and Horioka (1980) built their thesis on the tenet that in a closed economy, because capital mobility is restricted or non-existent, the saving–investment correlation would be high. This idea has puzzled many investigators, for FH (1980) found a high saving–investment correlation for OECD countries despite their more open and integrated markets. Subsequently, two distinct directions of studies have emerged on this subject matter. One group of researchers have attempted to test the relationship with larger sample sizes and more sophisticated econometric techniques while others have tried to offer alternative hypotheses in explaining the puzzle.

We position ourselves among the researchers in the first group and ask whether the saving and investment correlation is really a puzzle for Japan. To answer this question, we examine the relationship using time series data and more advanced econometric techniques. We use the recently developed bounds testing approach to cointegration and find that saving and investment are cointegrated for Japan when investment is the dependent variable. We use the ARDL model, following Pesaran and Pesaran (1997) and Pesaran and Shin (1999), to estimate the long-run elasticity on saving and find that it is around 0.68. On the basis of this result we reject the FH (1980) finding for Japan. In order words it is clear that the relationship between saving and investment is not a puzzle for Japan.

In this paper, we were not only interested in testing the validity of the FH (1980) thesis but we wanted to take the analysis further, and did so in two important ways. First, we conducted a rigorous unit root test in the presence of structural breaks in the data series. The idea here was twofold: to ensure that the series were not integrated of an order higher than one, for the bounds test to cointegration requires variables to be integrated of order one or less; and to establish whether policies implemented to boost saving and investment in

Japan have a permanent effect or a transitory effect on saving and investment. Our results from the unit root tests revealed that both saving and investment were integrated of order one, lending support for policies having a permanent effect on saving and investment.

Second, we were concerned about the direction of causation between saving and investment in Japan. To this end we used the bootstrap approach to Granger-causality and found that saving causes investment and investment causes saving, a result contrary to the findings of Sinha (2002) who used the Granger-causality F-tests.

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